Foreign Exchange Risk and the Flow of International Portfolio Capital: Evidence from Africa's Capital Markets

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Abstract

This dissertation addresses two major issues. First, it investigates whether currency risk commands a significant premium in representative equity markets in Africa. The International Arbitrage Pricing Theory and the Stochastic Discount Factor model respectively provide the analytical frameworks for the unconditional and the conditional asset pricing models used to investigate currency risk pricing. Empirical data analysis uses the Generalized Method of Moments estimation technique. Second, it examines the nexus between real foreign exchange rates and net international portfolio flows in representative capital markets in Africa. Time series and panel data techniques are employed to this end. The study covers seven major African countries: Botswana, Egypt, Ghana, Kenya, Morocco, Nigeria, and South Africa over the period January 1997 through December 2009.

Foreign exchange risk is found to be non-priced unconditionally when returns are measured in the US dollar; weakly priced unconditionally when returns are measured in the euro; and priced with time-varying risk premia in the conditional sense. Africa's equity markets are found to be partially integrated with the rest of the world. Monthly international portfolio flows to Africa are found to be low, non-persistent and relatively volatile. Using monthly data, Granger causality tests and innovation accounting from vector autoregressions (VARs), the study shows that the dynamic relationship between the real exchange rates and net portfolio flows is both country-dependent and time-varying. The findings are robust to alternative VAR specifications. However, annual data exhibit strong causality moving from real exchange rates to net portfolio flows, suggesting that fluctuations in real exchange rates inform the investment decisions of foreign investors in Africa's capital markets.

Among the key policy implications, it is recommended that, in addition to the US dollar and precious metals, Africa's monetary authorities should regard the euro as an important reserve currency; that policies be put in place to expedite the development of private fixed income securities and derivatives markets; that sound monetary policies be instituted to ensure that interest rate changes are market-determined and inflationary pressures are well-managed; and that regional markets integration and financial sector development policies be pursued more meticulously by governments in Africa.

Declaration

This is my original work and has not been presented before for a degree in this or any other University.

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Appreciation

The central theme of this study was conceived back in the late 1990s when, as a budding stock market aficionado, I casually observed that erratic movements in Kenya's stock market barometer, the Nairobi Stock Exchange 20-share index, seemed to correspond with instability in the foreign exchange market. Were these occurrences fortuitous? A few years earlier, the Kenyan government had opened up participation at the Nairobi stock market to foreign investors in a policy move billed by the financial press as a bold attempt to attract the sorely needed foreign capital into the country and to boost the then dwindling foreign exchange reserves. Could the matching fluctuations be related, in any way, to flows of foreign capital into and out of the bourse? These questions essentially commenced the long academic enterprise whose result is this dissertation.

My first appreciation goes to God for granting me good health and divine protection all my life and in the course of writing this dissertation. Dissertation writing is an arduous task, the successful completion of which necessarily involves the input of many more people than just the PhD student. This one has been no exception and has benefitted from the invaluable input of many people. Foremost, I'd like to appreciate the able, incisive, insightful, enthusiastic, and exceptional guidance of my research supervisor, Professor Kalu Ojah. This work has also benefitted immensely from the academically rich comments of reviewers, Professors Nicholas Biekpe, Lloyd Blenmann, and Christopher Malikane (in ascending order of surname).

There are moments in the course of doing this work that called for the greatest fortitude. Such moments arose from many experiences, the most prominent being the data search effort. During such moments, I made appeals to and received assistance from many people. In this connection, Rodwel Mupambirei of *MSCI Barra* deserves special mention for providing me with the MSCI world market equity portfolio dividend yield data series. The tools for analysis of the said data rely virtually on the availability of econometric computer packages. Several such packages were availed to me by Lucas Nkuna of Wits School of Economics and Business Sciences, whose generosity made my analysis possible and to whom I am very grateful. I am also indebted to Beverly Achieng' of the Wits Applied English Language Studies Department for patiently and diligently proof-reading parts of this dissertation.

To my wife, Dorothy and daughters, Valerie and Juliana, whose selflessness, perseverance, prayers and encouragement throughout the several months that I was away from home gave me the inspiration to complete this work, no words can express my deep affection, esteem and gratitude.

Lastly, I sincerely thank Wits University for partially funding my studies through the Postgraduate Merit Award. The usual disclaimer applies.

For my mom, Elizabeth Odongo

CHAPTER ONE

INTRODUCTION

1.1 Background to the Study

International trade and investments necessarily create a situation where one of the parties is dealing in a foreign currency. This creates a need for the existence of a market in which currencies can be exchanged for one another. Foreign exchange markets, which for each country, rely heavily on the *foreign exchange regime* put in place by the country's monetary authorities, exist to serve this purpose. The foreign exchange market is not a physical place; rather, it is an electronically linked network of banks, foreign exchange brokers and dealers whose function is to bring together buyers and sellers of foreign exchange. The market is dispersed throughout leading financial centers of the world. Trading is generally by telephone, fax or the Society for Worldwide Interbank Financial Telecommunications (SWIFT) system. Major participants in the foreign exchange market are the large commercial banks, foreign exchange brokers in the interbank market, commercial customers, and central banks. Foreign exchange regimes, or systems, can be classified according to the degree by which exchange rates are controlled by the government. Two foreign exchange systems, namely, the *fixed* (also known as *hard peg*) and the *floating* foreign exchange rate regimes, stand out at the extreme ends of the continuum.

Calvo and Mishkin (2003) explain that there are two basic ways a government can offer a credible guarantee under a fixed exchange rate system: a *currency board* and *full dollarization*. Under a currency board, foreign exchange rates are either held constant or allowed to fluctuate only within very narrow boundaries against some pre-specified instrument (usually gold or a hard currency) and governments are committed to maintaining the target rates. The central bank consequently guarantees full convertibility because it stands ready to exchange, on demand, domestically issued notes for the instrument and has enough international reserves to do so. The central bank intervenes by actively buying or selling its currency in the foreign exchange market whenever its exchange rate threatens to deviate from its stated par value by more than a predefined percentage. The action of buying results in a sudden surge in the demand for the domestic currency and causes it to rise in value against the instrument against which it is pegged, and by extension, other currencies: this is known as *currency revaluation*. *Currency devaluation* arises from the action of sale of the domestic currency by the central bank. On the other hand, full dollarization involves eliminating the domestic currency altogether and replacing it with a foreign currency like the US dollar, although it could as well involve the use of another currency. In an effort to stabilize its monetary policy, Zimbabwe recently adopted the latter approach, replacing its currency, the Zimbabwean dollar, with the South African rand.

A floating exchange rate system means that the government will allow for some exchange rate flexibility (Calvo and Mishkin, 2003). It rules out a fixed exchange rate regime, but nothing else. A country that allows a floating exchange rate may pursue a number of very different monetary policy strategies: for example, targeting the money supply, targeting the inflation rate (South Africa is an example in Africa) or a discretionary approach in which the nominal anchor is implicit. Under the floating exchange rate regime, the forces of market demand and supply, which depend, to a large extent, on the flow of international capital, are the key drivers of currency values. Since these forces are subject to random fluctuations, it follows that foreign exchange rates also fluctuate randomly.

1.1.1 Foreign Exchange Rate Systems in Africa: A Historical Overview

In the years between 1975 and 1998, the IMF classified member countries' exchange rate regimes based on their officially declared degree of exchange rate flexibility. Thus, countries' foreign exchange rate systems were classified as pegged, limited flexible (i.e. allowed to fluctuate within a narrow band) or more flexible (pure float). The major shortcoming of this system was its failure to capture the difference between what member countries claimed to be officially doing and the exchange rate policy that they were actually pursuing (Bubula and Otker-Robe, 2002). To address this and other weaknesses, the IMF unveiled a new classification system, in 1999, based on the *de facto* policies pursued by each country. The new system ranks exchange rate regimes on the basis of the different degrees of flexibility and clearly distinguishes among the many forms of pegged regimes from the more rigid to the more soft (IMF, 2000).

The 1999 IMF taxonomy, still in use, identified the following exchange rate regimes: (i) exchange arrangements with no separate legal tender, (ii) currency board arrangements, (iii) conventional fixed-peg arrangements, (iv) pegged exchange rates within horizontal bands, (v) crawling pegs, (vi) exchange rates within crawling bands, (vii) managed floating with no pre-determined path for the exchange rate, and (viii) independently floating (see IMF (2000) for a detailed description). While this new scheme has greatly improved on the former classification, it has not been widely used in empirical analysis because a strong historical database is yet to be built based on it. Here, I attempt to trace changes in exchange rate policies over time, spanning across the two IMF nomenclatures, for the sampled African countries.[1]

Africa is represented, in this study, by a sample of seven countries, two in the north (Egypt and Morocco), two in the west (Ghana and Nigeria), one in the east (Kenya) and two in the south (Botswana and South Africa) of the continent. Sampling is based on the relative verve of the foreign investors' segments of the continent's capital markets, with market vibrancy defined as the average annual volume of foreign investors' equity transactions. In addition, the availability of a reasonable long time-series for international portfolio flows, foreign exchange rates and other relevant data is considered in sample selection. On this account, countries with short data series, such as Cote d'Ivoire, Namibia and Tunisia are left out.

Since the introduction of the *pula* in 1976, **Botswana** has adopted a fixed but adjustable peg system. Initially, the pula was pegged only to the US dollar at the same exchange rate at which the South African rand was also pegged to the US dollar then. This implied equality between the pula and the rand. This effect expired when the rand was taken off the US dollar peg and allowed to float. To subdue the effects of exchange rate volatility between the pula and the rand, a trade-weighted basket of currencies comprising the rand and the Special Drawing Right (SDR) was introduced in 1980, to which the pula is still pegged to date. The rand has a greater share in the basket reflecting the need to protect the interests of the majority of Botswana's domestic firms, whose economic decisions have a significant rand-denominated component.

Up to the middle of 2000, the movements of the pula reflected the relatively high weight of the rand in the basket as well as the fact that its fluctuation against the rand was closely managed.

[1] Information on exchange rate systems is sourced from Masalila and Motshidisi (2003) for Botswana; Kamar and Bakardzhieva (2005) for Egypt; Kapur et al. (1991) and Jebuni et al. (1991) for Ghana; Ndung'u (1999) for Kenya; Commission of the European Communities (2004) for Morocco; Ugbebor and Olubusoye (2004), www.emerging-markets-online.com for Nigeria; Dube (2009) and http://www.uneca.org/docs/Major_ECA_Websites/6finmin/experm1.htm for South Africa.

Thus, the pula tended to depreciate and was unstable against the major international currencies as the rand weakened but was relatively stable in nominal terms against the rand. In mid-2000, the band surrounding the rand was removed and focus shifted to maintenance of real exchange rate stability vis-à-vis the basket.

Egypt had, from the 1960s, a "fixed adjustable peg" to the US dollar, combined with foreign exchange controls and multiple exchange rates. With the beginning of the economic reform program in 1991, the Egyptian government suspended foreign exchange controls, unified the exchange rate system and announced the adoption of a "managed floating" regime. Beginning 1997, the Egyptian exchange rate faced numerous external shocks arising from the East Asian crisis in mid-1997, the Luxor (Egypt) terrorist attack in 1997, the fall in world oil prices in 1998, the revival of tensions in the Middle East towards the end of 1990s and the second Palestinian *Intifadah* launched in October 2000. In response to these shocks, the Egyptian government decided, in January 2001, to restore market stability by announcing a new central exchange rate of Egyptian pound 3.85 per US dollar and introducing a "crawling peg" system.

This measure did not however yield much benefit. The September 2001 terrorist attack in New York and the subsequent wars in Afghanistan and Iraq together with the increasing Israeli-Palestinian violence at the Egyptian border further dipped confidence in Egypt as a foreign investment destination and caused a decline in foreign reserves. The Egyptian government responded by re-imposing capital controls, which put the exchange rate under pressure, re-energized the parallel market and effectively left the country with a parallel exchange rate system, with a black market premium of about 10 percent. In January 2003, the crawling peg was abolished as the Egyptian Prime Minister announced a free float of the Egyptian pound.

In November 1958, **Ghana** introduced a new currency, the Ghana pound, which was set at par with the British pound sterling, then exchanging for 2.80 US dollars. Between 1958 and early 1970s, falling cocoa prices and a massive industrialization program initiated by the government occasioned sustained external revenue shortfalls. By 1961, the country's reserve position had become precarious. Pressure from the International Monetary Fund (IMF) and the World Bank for the country to accept an orthodox stabilization policy, including devaluation, was rejected, and Ghana continued to maintain an overvalued fixed exchange rate. In July 1965, the pound was decimalized and the Ghanaian *cedi* was introduced. The Ghana pound was formally demonetized in December 1966.

In 1966, the National Liberation Council (NLC) staged a coup against the government of Kwame Nkrumah. The NLC government accepted the IMF and World Bank economic package and, in July 1967, Ghana had its first official devaluation of the cedi. In June 1978, Ghana introduced a flexible exchange system under which the US dollar value of the cedi was to be adjusted to reflect the underlying economic, financial, and balance of payments situation. The experiment was discontinued in August 1978 when a fixed exchange rate was re-introduced. In April 1983, as a harbinger to the Economic Recovery Program (ERP) to be introduced a few months later, a system of multiple exchange rates based on bonuses and surcharges applied to specified transactions was tried. This system was abandoned in October 1983 when the exchange rate system was unified.

In September 1986 the government introduced a "managed float" in which rates were determined by "market forces" at weekly "retail auctions" in which individuals with import and export licenses would bid exchange rates. The auction system was modified in April 1990 allowing the market exchange rate to be determined in an inter-bank market through a weekly "wholesale" auction. Concurrently with the establishment of the wholesale auction, the restrictions on payments for current international transactions were lifted. This resulted in an essentially full liberalization of Ghana's exchange system, allowing the achievement of significant progress towards the restoration of financial convertibility of the cedi.

Up to 1974, the exchange rate for the **Kenya** shilling was pegged to the US dollar, but the peg was later changed to the SDR. The exchange rate regime was changed to a crawling peg in real terms at the end of 1982. This regime was in place until 1990; a dual exchange rate system was then adopted that lasted until October 1993 when the official exchange rate was abolished after some devaluations. Thus, the official exchange rate was merged with the market rate and the Kenya shilling was allowed to float.

The stated policy objectives in Kenya have been to maintain an exchange rate that would ensure international competitiveness while at the same time keep the domestic rate of inflation at low levels, conduct a strict monetary stance and maintain positive real interest rates. This has been difficult in practice and it has been made even more difficult by a floating exchange rate that at times moves out of line with its fundamentals in the short run. When the exchange rate was put to a float in an environment of excess liquidity, massive depreciation and high and accelerating inflation ensued.

Since the beginning of the 1990s, **Morocco** has pursued an exchange rate policy of pegging the *dirham* to an undisclosed basket of currencies. In April 2001, Moroccan government announced a new basket of currencies comprising of the euro (80%) and the US dollar (20%). This, it was argued, would better reflect Morocco's commercial and financial links with the European Union. The change occasioned a slight depreciation in the Moroccan dirham.

Nurtured by the growing integration of Morocco into the world economy, discussions on a higher degree of flexibility of the exchange rate regime continued. In March 2006, the governor of Bank-Al-Maghrib (Central Bank) announced at an annual press conference that the country would be ready to adopt a floating exchange rate by 2010. Measures to strengthen the banking system and macroeconomic reforms would be introduced to address the potential risk associated with volatile capital inflows. A flexible exchange rate arrangement may be better suited to deal with external shocks, especially in the face of labor market rigidities.

The **Nigerian** pound was introduced in 1959 with its external value fixed at par with the British pound sterling, then valued at 2.80 US dollars. In June 1962, the parity value of the Nigerian pound was fixed at 2.48828 grams of fine gold. The *naira* was introduced as Nigeria's currency in January1973, its par value set at half that of the old Nigerian pound. In February 1978, the system of determining the naira exchange rate against a basket of currencies of Nigeria's main trading partners was adopted. This was the First-tier Foreign Exchange Rate Market (FFEM). In September 1986, following the adoption of the principles of the Structural Adjustment Programs, the Second-tier Foreign Exchange Market (SFEM) was introduced under which the exchange rate was floated. The Second-tier rate was determined by various auction methods.

By March 1992 there was a complete floating of the naira. A change in government in August 1993 ushered in a new fixed exchange rate regime. In 1995, the Autonomous Foreign Exchange Market (AFEM) was introduced under a policy that allowed for Central Bank of Nigeria (CBN) intervention on a predetermined basis instead of arbitrarily. Under AFEM, bureaux de change would buy and sell from privately sourced foreign exchange at the AFEM rate. The fixed exchange rate was reserved for public sector use. In 1993, the Parallel Market and bureaux de change exchange rates were almost double the devalued rate for the naira. Authorities saw this as a signal of a depreciation trend which needed correction. This led to a re-introduction of a fixed exchange rate which pegged the naira at N21.996 to the US dollar in 1994. In January 1997, the naira was formally pegged and a pro-rata system of foreign allocation to end-users was adopted.

In February 2009, CBN issued a set of new controls in the foreign exchange market, reversing much of the liberalization of exchange rates that had been ongoing since 1995. According to the new rules, still in operation, foreign exchange purchased from any authorized dealer cannot be sold on the interbank market, effectively shutting the market down. Foreign exchange from the CBN is to be used only by end-customers, and any foreign exchange purchased by banks from authorized dealers that is not sold to customers for eligible transactions within five days must be sold to the CBN at no more than 1% below the official rate of the previous retail Dutch auction sale. The central bank is committed to keeping the exchange rate in a ±3% band.

As a signatory of the Bretton Woods Agreement of 1945, **South Africa** became party to the system of generalized fixed exchange rate system, with adjustable margins. The par value of one South African pound was established at 4.03 US dollars, equivalent to 3.58143 grams of fine gold. The South African pound was pegged to the pound sterling by buying and selling rates for sterling within a stipulated margin of half percent. In the ensuing years, the currency remained relatively stable and its parity in terms of gold was changed only as a result of the decimalization of the monetary unit. In February 1961, the currency unit of South Africa was changed to the *rand* whose unit gold parity was fixed at 50 percent of that of the old South African pound. In October 1972, South African monetary authorities decided to peg the rand to the US dollar.

Following the collapse of the Bretton Woods system in 1973, South African monetary authorities decided, in June 1974, to adopt a system of independently managed floating. This exchange rate arrangement was in place until 1975 when the authorities adopted a policy of keeping the rand/US dollar rate constant over longer periods and only adjusting it on the basis of major shifts in the country's important economic fundamentals. The link between the rand and the US dollar remained until early 1979 when exchange controls over non-residents were lifted and measures put in place to improve the technical functioning of the spot and forward exchange markets, and to encourage the development of a forward exchange market independent of the central bank.

Political developments after 1984, and the imposition of financial sanctions by the international community on South Africa for the government's apartheid policy, saw monetary authorities re-enter the foreign exchange market. Exchange controls on capital transfers by non-residents were re-introduced in the form of a distinction between the "financial rand" and the "commercial rand." These measures were intensified in 1985. Accordingly, a dual exchange rate system was introduced in South Africa. However, following the multi-party elections of 1992, South Africa once more embarked on reforms of its foreign exchange markets. In March 1995, South Africa abolished the "financial rand" thereby eliminating the dual exchange rate system that had existed since 1984. The rand was then made convertible for all but residents' capital outflows and institutions intent on investing overseas. Further progress is being made to liberalize the country's foreign exchange markets and to move towards market-driven exchange rates.

1.1.2 A Brief Description of Selected Equity Markets in Africa

Some information on capital markets of each of these countries is in order.[2] Located in the capital Gaborone, the **Botswana** share market was established in 1989 and became the Botswana Stock Exchange in 1995. The Exchange has a small, stable listing. It trades in equities, Botswana government bonds and a corporate note. Private investors are estimated to account for fewer than 10% of the total market capitalization, which was approximately \$59.1 billion at the end of 2009.[3] Foreign ownership of the free stock of a local publicly-quoted company trading on the Exchange may not exceed 55% of which no one foreigner may own more than 10%.

[2] The bulk of this information was accessed at http://www.sandpworld.com/markets/exchange (Standard & Poor's Financial Services LLC, 2009) and http://www.mbendi.com/exch/13/p0005.htm#5 (MBendi Information Services, 2010) on Jan 15, 2010.

[3] http://www.bse.co.bw/market_n_statistics/market_statistics.php (accessed on January 15, 2010).

The **Egyptian** Stock Exchange comprises of two exchanges, Cairo (established in 1903) and Alexandria (established in 1883). Both exchanges were very active in the 1940's, and the combined Egyptian Stock Exchange ranked fifth in the world. The two Exchanges remained largely dormant between 1961 and 1992. In the 1990's, the Egyptian government's restructuring and economic reform program revived the Egyptian stock market, and a major change in the organization of the Cairo and Alexandria Stock Exchanges took place in January 1997 with the election of a new board of directors, which brought about significant modernization, culminating in the Exchange winning the award of the most innovative African Exchange in 2008 during the annual Summit organized by Africa Investor. Market capitalization stood at approximately 460 billion Egyptian pounds (approximately $52.9 billion) at the end of 2009.[4] The Exchanges trade in equities, corporate and Egyptian sovereign bonds and mutual fund certificates. There are neither restrictions precluding foreign participation in the market nor any rules against repatriation of profits or capital gain.

The **Ghana** Stock Exchange, located in Accra was incorporated in July 1989 with trading commencing in 1990. In 1993, the Exchange was the sixth best index performing emerging stock market, with a capital appreciation of 116%. In 1994, it was the best index performing stock market in the emerging markets, gaining 124.3% in its index level. All kinds of securities (bonds, notes, equities and certificates) may be listed at the Exchange. Ownership of publicly listed companies by foreign investors (those who are not Ghanaian and who live outside the country) is restricted to a cumulative total of 74%, of which a single foreign investor is allowed to hold up to a maximum of 10%. Market capitalization was 15.9 billion Ghanaian cedi (approximately $10.9 billion) by the end of 2009.

The Nairobi Stock Exchange (NSE) in **Kenya** is small and somewhat speculative. It began in 1954, while Kenya was still a British colony, as an overseas stock exchange with permission of the London Stock Exchange. The Exchange now works in cooperation with the Uganda Securities Exchange and the Dar-es-Salaam Stock Exchange, on various matters including the cross-listing of various equities. The Exchange is sub-Saharan Africa's fourth-largest bourse,

[4] Statistics obtained from the Egyptian Stock Exchange website: http://www.egyptse.com/English/homepage.aspx on January 15, 2010.

with a total market capitalization of approximately 900 billion Kenyan shillings[5] (approximately $11.8 billion). The Exchange was opened to foreign investors for the first time in January 1995, but with a maximum limit of 20% shareholding for institutions and 2.5% for individuals. Later, the ceiling on foreign investment was increased to 75% for foreign investors with no additional restrictions. The exchange trades in equities, Kenyan government and corporate bonds, Treasury bills, and commercial papers. The Exchange recently introduced a central depository system and automated all its operations – a move that is expected to speed up clearing and settlement.

The Casablanca Stock Exchange in **Morocco** is an active stock exchange in Africa. The third oldest stock exchange in Africa, it was established in 1929. The exchange is relatively modern, having experienced reform in 1993. The exchange installed an electronic trading system, and is now organized as two markets: the Central Market and a Block Trade Market, for block trades. It opened a central scrip depository in 1997. There are no restrictions on foreign investment on the Casablanca Stock Exchange, nor on foreign ownership of companies. With a market capitalization of some 510 billion Moroccan dirham (approximately $63 billion) as at end of 2009, the Casablanca stock market is also the sixth-largest in the Arab World.[6]

The **Nigerian** Stock Exchange was established in 1960 as the Lagos Stock Exchange. In December 1977 it became the Nigerian Stock Exchange, with branches established in some of the major commercial cities of the country. At present, the Exchange has six branches, each having a trading floor. The exchange trades in equities, Nigerian federal government bonds, state government and corporate bonds, commercial papers and other money market instruments. The Exchange is an affiliate member of the Federation of International Stock Exchanges (FIBV) and is also a founder member of the African Stock Exchanges Association (ASEA); it operates an Automated Trading System. In order to encourage foreign investment in Nigeria, the government has abolished legislation preventing the flow of foreign capital into the country. This has allowed foreign brokers to enlist as dealers on the Nigerian Stock Exchange and investors of any nationality are free to invest. Nigerian companies are also allowed multiple and cross border listings on foreign markets. The second largest bourse in Africa, the Nigerian Stock Exchange

[5] Obtained from the Nairobi Stock Exchange site: http://www.nse.co.ke/newsite/inner.asp?cat=Stats on January 15, 2010.
[6] See the Wikipedia website: http://en.wikipedia.org/wiki/Casablanca_Stock_Exchange (accessed on January 15, 2010).

had a market capitalization of 7.03 trillion Nigerian naira (approximately $47 billion) by the end of 2009 (Okerere-Onyluke, 2010).

With a market capitalization at the end of 2009 of 5.929 trillion **South African** rand (approximately $790.5 billion), the Johannesburg Securities Exchange (JSE), established in 1887, is the largest stock exchange in Africa. The JSE has evolved from a traditional floor based equities trading market to a modern securities exchange providing fully electronic trading, clearing and settlement in equities, financial and agricultural derivatives and other associated instruments. The JSE is planning to create a pan-African exchange. As a prelude, it has created the Africa Board as a platform where top African companies can be listed and traded. Having been developed to attract foreign capital to the African market, the Africa Board is a long term strategy to promote the growth of capital markets on the African continent. In 1995, substantial amendments were made to the legislation applicable to stock exchanges which resulted in the deregulation of the Exchange through the introduction of limited liability corporate and foreign membership. The bourse has no restrictions on foreign security ownership. However, there is a difference in the treatment between domestic and foreign investment in terms of the local borrowing restrictions imposed by the Exchange Controls authorities.

From the foregoing account, it is clear that all the countries sampled are open to international portfolio investment with differing degrees of openness. In a study of a sample of 134 countries of the world, the Global Enabling Trade Index of the World Economic Forum (2009b: xiii) ranks Kenya as having the most accessible markets with a score of 4.59 out of 10. It is followed by Morocco (4.09), Ghana (3.94), South Africa (3.78), Egypt (3.05), and Nigeria (2.72). Botswana was not covered by the study. However, these markets are characterized by significant disparities in terms of sophistication levels. Having highly sophisticated financial markets, on a par with Belgium and France, and with relatively easy access to capital from various sources, sound banks, and a well-regulated securities market, the Africa Competitiveness Report (World Economic Forum, 2009a: 6) ranks South Africa first in the region and 24th overall. Botswana is ranked in the top third in this pillar ahead of most other countries in the region. Kenya and Nigeria have financial markets that are placed in the top half of the rankings. Further, Kenya and South Africa, with high-quality scientific research institutions and strong investment in research and development, are the top regional performers with respect to innovation. Egypt and Nigeria also feature in the top half of the rankings in this pillar.

Markets for private capital are the product not only of the willingness of investors to place their funds in corporate-issued securities but also, perhaps more importantly, of the existence of private capital demand. In Africa, the desire to utilize technology transfers, the abundant low-cost labor and the availability of natural resources are the key drivers of private capital demand. Demand for private capital provides great opportunities for willing foreign investors to diversify their portfolios, both domestically and internationally. Diversification means spreading one's investable wealth across several assets with a view to reducing their exposure to price variability. International or geographical diversification is a practice whereby investors form portfolios of securities, some of which are issued in foreign markets. Solnik (1974) demonstrates that substantial gains can be realized from international diversification: in terms of variability of return, an internationally well-diversified portfolio is one-tenth as risky as a typical security and half as risky as a well-diversified portfolio of US stocks (with the same number of holdings).

However, international diversification implies that portions of investors' anticipated future returns are denominated in currencies other than their domestic currency. When realized, such future returns are converted to the investor's domestic currency. The return, in domestic currency terms, depends critically on the behavior of the rate at which the foreign currencies in which the returns are denominated can be converted into investor's domestic currency. Since exchange rates fluctuate randomly, one would expect the realized value of such returns to vary randomly from expectations. Solnik (1974) shows that the risk of an international portfolio, that is unprotected against exchange risk, is larger than that of a covered portfolio. The variability in an investment's expected future returns introduced directly by fluctuations in exchange rates is known as *currency risk* or *foreign exchange risk*. Because it affects the value of returns, currency risk is a key factor considered by rational investors interested in international diversification.[7] Consequently, international diversification is considered attractive only to the extent that the potential return from it outweighs the associated currency risk. Now, since international

[7] Currency risk is only one of the sources of risk affecting returns from international portfolio investments. Foreign investors also consider barriers to capital flows created by higher costs of transacting in foreign securities, withholding taxes, political risk, and other factors such as the failure of purchasing power parity, information asymmetries, and regulation.

diversification activities precipitate movement of capital across national borders, there must be a linkage between foreign exchange rate movements and international capital flows.

1.2 Motivation

The floating exchange rate regime has been associated, almost invariably, with instability in the foreign exchange markets. Whether this instability causes foreign exchange risk to be priced[8] in capital markets has remained an elusive empirical question. In advanced capital markets, some studies, especially those that have employed unconditional asset pricing models (Jorion, 1991; Loudon, 1993; Hsin et al., 2007), find no evidence of currency risk pricing, while others, particularly those using conditional models (Dumas and Solnik, 1995; MacDonald, 2000; Cappiello and Panigirtzoglou, 2008), largely report weakly significant, time varying exchange risk premia. In the emerging markets, strong support for the hypothesis that significant unconditional (Carrieri and Majerbi, 2006) as well as conditional (Tai, 1999; Glen 2002; Phylaktis and Ravazzolo, 2004) currency risk premia exist have been documented. Despite the fact that most African countries adopted a flexible foreign exchange rate regime in the early 1990s, there has been no study on this issue devoted exclusively to her capital markets.[9]

Proponents of the flexible exchange rate system argued that its adoption would impact favorably on capital market efficiency and boost the flow of foreign investments into the continent. Notwithstanding its implementation, Africa has not been translated into the destination of choice for international investors. Indeed, IMF/World Bank (2008) not only observes that Africa still lags behind the emerging markets of East Asia and the Pacific in attracting foreign money, but also bemoans the low research effort aimed at addressing the likely causes of the relatively poor performance. The upshot is that the continent's macroeconomic policy makers and potential investors do not have the benefit of sufficient scientific knowledge on the nature of the relationship between currency volatility and portfolio flows. Contradictory findings have, however, been reported by studies that have sought to establish the relationship that exists between the two macroeconomic variables in some emerging markets economies (Agènor, 1998;

[8] A risk factor is said to be priced in the market if it contributes to the overall risk premium of an investment.

[9] The countries covered by Carrieri and Majerbi (2006) were Argentina, Brazil, Chile, Mexico, Greece, India, Korea, Thailand and Zimbabwe. Only Zimbabwe happens to be in Africa.

Kim and Singal, 2000; Ahmed et al., 2005) as well as in some advanced market economies (Kohlhagen, 1977; Brennan and Cao, 1997; Hau and Rey, 2004, 2006).

1.3 Purpose Statement

The purpose of this study is to investigate whether currency risk commands a significant premium in equity issues in African capital markets and to establish whether a significant relationship exists between foreign exchange rates and the flow of international portfolio capital into and out of the continent's markets. Specifically, these issues are explored in a sample of seven countries: Botswana, Egypt, Ghana, Kenya, Morocco, Nigeria, and South Africa. Sampling is based on the relative verve of the foreign investors' segments of capital markets in the continent. The findings of the study are used to draw implications for macroeconomic policy.

1.4 Research Questions

Put succinctly, the research questions of the study are:

(i) Is currency risk priced in Africa's capital markets?

(ii) Is there any linkage between volatility in foreign exchange rates and the flow of portfolio investment funds in Africa?

Both questions speak to the relative efficiency of capital markets in African economies. Consequently, the study's findings point to market efficiency enhancing, financial markets development and institutional infrastructure provision policies that, if implemented, present the potential to encourage international investors to increase the inflow (and stability) of sorely needed investment funds into Africa's capital markets.

1.5 Significance/Contributions of this Study to the Literature

The 1990s witnessed a surge in international capital flows into emerging market economies, particularly in Latin America and East Asia. This development coincided with, and was perhaps underpinned by, the removal of capital controls and the broader liberalization of financial markets in developing countries. To date, the two regions continue to enjoy substantial amounts of inflows of foreign portfolio investments. Gross Foreign Capital Flows (GFCF) as a proportion of Gross Domestic Product (GDP) for Latin America and the Caribbean for the year 2007 stood at 19.6%. For East Asia and Pacific, the GFCF as a proportion of GDP was 24%.[10] However, international capital flows to Africa compare unfavorably with flow to other developing regions of the world. The GFCF as a proportion of Gross Domestic Product, in the year 2007 was 9.5% for Botswana, 13.7% for Egypt, 17.5% for Ghana, 15.2% for Kenya, 17% for Nigeria (author's estimate), 25.3% for Morocco and 15.7% for South Africa. The average GFCF as a proportion of GDP based on this sample is 16.3%.

The flow of portfolio capital (a component of GFCF and the focus of this study) to African countries is not very impressive either. As a percentage of GDP, net annual portfolio flows for the period 1997 through 2009 averaged 0.23 for the six sampled countries including Botswana, Egypt, Kenya, Morocco, Nigeria and South Africa (see Table 33 in Chapter 5). Data is unavailable for Ghana, the other country in this study. Although a number of reasons (such as the presence of some level of capital controls, unstable interest rates, and weak national institutions) might be cited to explain the low flow of investment funds, it is important to observe that Africa is characterized by volatile exchange rates; most of the countries' currencies are soft and vulnerable to changes in macroeconomic variables and external shocks.

The question that arises is: Is there a linkage between the behavior of foreign exchange rates and the flow of foreign portfolio funds into and out of Africa's capital markets? To answer this question, I sought to establish, first, whether foreign exchange risk is priced in the equity markets of Africa. In this respect, this dissertation is a pioneering study in Africa and it has contributed to the literature in several ways: (i) it finds that foreign exchange risk is not unconditionally priced in Africa's stock markets; (ii) however, it finds that foreign exchange risk is conditionally priced in the same equity markets; and (iii) as a natural extension to the main question, I also tested for segmentation/integration of Africa's equity markets with the rest of the world's equity markets. Results suggest that Africa's equity markets are partially segmented.

Secondly, I investigated the nature of the relationship between foreign exchange rates and the flow of foreign portfolio capital. In this regard, my study contributes to the literature in three

[10] The data available as at March 2009 from the World Bank website: http://web.worldbank.org/WBSITE/EXTERNAL/DATASTATISTICS/0,,contentMDK:20394897~menuPK:64133163~pagePK:64133150~piPK:64133175~theSitePK:239419~isCURL:Y~isCURL:Y~isCURL:Y~isCURL:Y,00.html

major ways: (i) using monthly data (available only for four countries, namely, Egypt, Morocco, Nigeria and South Africa), the study finds that the dynamic relation between foreign exchange rates and net portfolio flows is time-varying and country-dependent; (ii) portfolio flows are found to be low, volatile and non-persistent and hence largely constitutes "hot money," the term used to describe flows with these characteristics in the literature; and (iii) the analysis of annual data (six countries including the foregoing four plus Botswana and Kenya) finds strong panel causality from foreign exchange rates to net portfolio flows. Several policy implications (presented in Chapter Six) are drawn from these findings.

1.6 Organization of the Study

The rest of this work is organized as follows. Chapter Two surveys the existing theoretical and empirical literature on international capital flows, the pricing of currency risk, and the nature of the relationship between foreign exchange rates and international capital flows. Chapter Three reviews the existing econometric models for the analysis of the pricing of currency risk and those that relate currency exchange rates to international capital flows. The chapter also presents the models used in this study and describes the data. Chapter Four describes the basic statistical characteristics of the data, presents and discusses empirical estimation results on the pricing of foreign exchange risk in Africa's equity markets. Chapter Five presents and discusses empirical findings on the bivariate relationship between (real) foreign exchange rates and the flow of international portfolio capital. Chapter Six recapitulates salient findings of the study and highlights the evident policy guides of these results, in addition to pointing out areas of future research interventions on the subjects addressed by the study.

CHAPTER TWO

FOREIGN EXCHANGE RATES, FOREIGN EXCHANGE RISK AND THE FLOW OF INTERNATIONAL CAPITAL: EXTANT EVIDENCE

2.1 Introduction

The purpose of this chapter is to articulate the conceptual foundations of the research. The chapter is organized as follows. Section 2.2 surveys existing theoretical and empirical literature on international capital flows. Section 2.3 discusses the theories and empirical evidence relating to the pricing of foreign exchange risk. Section 2.4 explores the nature of the relationship between the currency fluctuations and the flow of international portfolio capital. Section 2.5 discusses the sub-Saharan African case and section 2.6 concludes.

2.2 International Capital Flows

In this dissertation, the term 'capital flows' refers to net inflows - that is, gross inflows minus repatriation. Capital inflows are characterized as foreign direct investment (FDI) if the investor acquires a lasting management interest (10 percent or more of the voting stock) in the foreign enterprise (Verdier, 2008). Portfolio investment flows include portfolio debt flows (that is, purchase or sale of domestic bonds and/or short term debt securities) and non-debt-creating portfolio equity flows. This study focuses on portfolio flows.

2.2.1 Benefits and Costs of International Capital Flows

Although controversy exists in the literature regarding the nature of benefits of foreign portfolio investments, scholars and policy-makers have more or less agreed that *some* benefits may accrue to the destination country. Errunza (1986) has classified the potential benefits emanating from foreign portfolio investments into three major categories, namely; the developmental effect, the resource effect and the welfare effect.

On the *developmental effect*, Errunza suggests three types of major benefits. First, he argues that foreign investors would demand more efficient market regulation, protection of minority interest, fair trading and brokerage practices, and adequate listing and disclosure requirements. Implementation of these demands would result in major benefits to the destination market.

Secondly, foreign participation would necessitate development of new institutions and investments on one hand and result in transfer of knowledge and training of local executives on the other hand. Finally, sustained foreign investment would help dampen speculative movements and reduce imperfections that afflict many Less Developed Country (LDC) markets. The improvement in regulatory, institutional and market environment in conjunction with private foreign capital would increase domestic investor confidence and participation, thereby providing further stimulus and support for these national markets.

On the *resource effect*, Errunza (1986) posits that foreign capital may substantially reduce the loss of domestic savings through capital flight by increasing domestic investor confidence as well as by making available a wider selection of securities initially designed to satisfy foreigners. Reduction in capital flight together with capital inflows in the form of portfolio investments would augment total domestic savings.

On the *welfare effect*, Errunza argues that the removal of barriers, as regimes liberalize their markets to allow foreign portfolio investments, may enable the achievement of complete integration of world markets. Complete integration would mean that the prices of LDC securities will in general go up with a consequent decline in the cost of equity capital. Since market integration will expand the investment opportunity set, the ability to invest in the world market portfolio will result in a positive diversification effect. Finally, if LDC investors hold optimal portfolios, their welfare would increase following integration. The net effect on LDC security prices and investor welfare would therefore be positive.

Economic theory has also identified a number of channels through which openness to international financial flows could raise productivity growth. Kose et. al (2009) provide a comprehensive analysis of the relationship between financial openness and total factor productivity (TFP) growth using an extensive dataset that includes various measures of productivity and financial openness for a large sample of countries. They find that *de-jure* capital account openness has a robust positive effect on TFP growth. The effect of *de-facto* financial integration on TFP growth is less clear from their study, but this masks an important and novel result. They find strong evidence that foreign direct investments and portfolio equity liabilities boost TFP growth while external debt is actually negatively correlated with TFP growth. The negative relationship between external debt liabilities and TFP growth is attenuated in economies with higher levels of financial development and better institutions.

Singling out one major component of international financial flows, Choi and Baek (2006) show that the volatility of portfolio flows significantly raises the level of reserve holdings. Further, volatility of portfolio balance (net flows) is more sensitive than other types of portfolio flows volatility in determining reserve holdings. These results imply that monetary authorities have accumulated more precautionary reserve balances against increased volatility of capital flows as capital account liberalization progresses and more frequent international crises occur.

In a broader and more recent investigation, Ferreira and Laux (2009) examine the importance of portfolio investment flow levels and volatilities as determinants of subsequent economic growth in cross-country data. They use a design that includes several methodological innovations to allow for the interpretation of results as evidence of the relation between financial market openness and growth. They find that openness to portfolio flows is statistically conducive to growth, in that a country's GDP grows after both positive flows of funds and also, strikingly, after some types of large negative flows of funds. This evidence is present for the full sample, and also for subsamples of countries defined by their development status – though the specific pattern differs across the subsamples. The evidence is strongest for the less developed countries in the sample. Overall, results of this study indicate that openness to flows in both directions is associated with growth, and that the portfolio flow volatility that might come with openness is not harmful for any set of countries.

Notwithstanding these positive effects, theory is also replete with conjectures about negative implications of foreign investment flows to the recipient country. Among other researchers, Dornbusch and Park (1995) contend that the trades of foreign investors are affected by past returns, so that they buy when prices have increased and sell when they have fallen. Such a practice is called *positive feedback trading*, and it leads to what is known in investment parlance as *herding*. In a study of the potential destabilizing effects of positive feedback trading, DeLong et al. (1990) claim that the sign of arbitrage positions taken by rational speculators can be the opposite of what one needs to move asset prices toward fundamentals. They conclude that positive feedback trading has the potential of moving asset prices further away from the fundamentals and thus exerting a destabilizing influence on the stock market.

However, Choe et al. (1999) dispute these findings. Using order and trade data for the period between end of November 1996 and end of 1997, they find strong evidence of positive feedback trading and herding by foreign investors before the period of Korea's economic crisis. During the crisis period, herding falls, and positive feedback trading by foreign investors mostly disappears. Their study does not, however, find any evidence that trades by foreign investors had a destabilizing effect on Korea's stock market over the sample period. Rather, they find that the market adjusted quickly and efficiently to large sales by foreign investors, and that these sales were not followed by negative abnormal returns.

Other harmful effects on the recipient country's economy associated with international capital flows have been identified by Kim (2000), who observes that a surge in capital inflow tends to cause inflationary pressure and increase current account deficits. The real exchange rate tends to appreciate in the capital-receiving country while the traded goods sector of the economy loses competitiveness in international trade. The increase in the current account deficit and the appreciation of the real exchange rate also make the economy more vulnerable to foreign shocks. When the inflow of foreign capital is interrupted, the economy has to go through reverse adjustments in the current account and real exchange rate. The process of adjustment to adverse shocks in capital movement has been highlighted by the widespread costly debt crisis of the 1980s, the Mexican crisis of 1994–95, and the Asian crisis of 1997–2000.

2.2.2 Determinants of International Capital Flows

Among the early researchers to investigate the causes of international investment flows was Stulz (1983). Using an intertemporal model in which it is assumed that investors maximize their expected utility of lifetime consumption, the study demonstrates that the *expected return* and *variance* of the return on investment in foreign technology play significant but opposing roles in determining the level of foreign investment in the domestic country. The study also shows that an increment in *foreign wealth* accompanied by an equivalent decline in *domestic wealth* increases net foreign investment in the domestic country. Finally, it shows that an increase in the *risk tolerance* of the domestic investor has, at first approximation, no effect on net foreign investment in the domestic country. Although Stulz's model does not include exchange rate changes, he remarks that exchange rate dynamics affect the demands for risky assets, and therefore, affect net foreign investment.

Uppal (1992) investigates the impact of deviations from purchasing power parity on international portfolio choice and capital flows in a model where the world interest rate is determined endogenously. He shows that in a two-country world, a deterioration in the terms of trade leads to a current account deficit, if the elasticity of substitution is less than one. He demonstrates that the current account deficit is accompanied by an increase in lending and a decrease in investment in risky equity. His conclusion is that the volume of capital flows is determined by the absolute deviation from purchasing power parity.

The surge in private capital inflows into emerging economies, which began around 1989, elicited further research and debate. One important lesson that was drawn from the increased investment activity of that time was that private (bond and equity) flows, as opposed to official flows, had become a crucial source of financing large current account imbalances in developing countries. Bruno (1993) observes that close to half of all aggregate external financing of developing economies comes from private sources and goes to private destinations. These trends raise important issues concerning the factors that motivate capital flows and their effect on performance, especially of developing countries.

In the debate that ensued, economists and policy experts were split on whether destination-country (domestic or "pull") factors outweighed the source-country (foreign or "push") factors in encouraging portfolio investment inflows from the developed world into developing economies. Those in favor of the "pull" factor dominance argued that the sustainability of these flows were, to a large extent, a function of the "success" of domestic policies. This school of thought argued that favorable domestic factors improved the creditworthiness of the destination nations and hence attracted voluntary private capital flows. Chuhan et al. (1993), using a model in which country creditworthiness features prominently as a domestic "pull" variable, concludes that domestic factors are at least as important as external factors in explaining portfolio capital inflows to Latin America and Asia.

The polar school of thought to the "favorable-domestic-policies" reasoning is that "push" factors are at play. Adherents to this school point out that private capital flows are volatile because they respond more to changes in fortunes in the source-country markets, factors beyond the control of recipient countries' policy-makers. Among the proponents of the "push-factors" argument are Calvo et. al (1993), who attribute the surge in private capital inflows to Latin American countries to a fall in interest rates, recession and balance of payments developments in advanced countries, particularly the USA. The authors point out that an improvement in the economic situation of developed (or source) countries could lead to a future capital outflow from Latin America. However, the findings supporting these arguments have been challenged on the basis of the weaknesses in the methodology used. In the study, Calvo et al. (1993) use international reserves as a proxy for capital inflows with their inferences based only on the statistical analysis of common factors.

Kim (2000) develops a method that overcomes some of the limitations of the Calvo et al. (1993) study. He uses a structural vector autoregressive model with data broadly similar to that of Chuhan et al. (1993). His findings differ sharply from those of Chuhan et al. (1993) but lend credence to the findings of Calvo et al. (1993). He provides evidence that capital movements to developing countries in Latin America and East Asia are explained largely by external reasons, such as decreases in the world interest rate or output in industrial countries. Domestic factors, including country-specific productivity shocks and demand shocks, are relatively less important. Another interesting finding is that the fundamental causes of capital flows appear to differ little across developing countries. These results suggest that developing countries need to pay attention to the global financial arrangements associated with capital flows, to exchange rate policy and international macroeconomic fundamentals to avoid financial crises in a world of increased capital mobility.

Earlier, Fernandez-Arias (1996) had advanced a powerful argument in support of the contention that external influences are more responsible for inflows of international portfolio capital to developing economies than internal factors. Highlights of his findings are summarized thus: (i) improvements in country creditworthiness appear to be a significant variable in explaining investors' behavior and the surge of capital inflows in middle-income developing countries. The exception is sub-Saharan African countries, where improvements in country creditworthiness have not been large enough to lead to levels exceeding the credit rationing threshold; (ii) to a large extent, developing country creditworthiness has been in turn driven by external factors, especially international interest rates. In this regard, improvements in country creditworthiness is best seen as a channel through which the underlying (external) factors induced capital inflows; (iii) in terms of causal underlying factors, the surge of capital inflows in most countries appears to have been largely pushed by low returns in developed countries, both directly or through the country creditworthiness channel, as opposed to pulled by domestic factors; (iv) consequently, most developing countries are vulnerable to adverse exogenous developments that would render capital inflows unsustainable. Capital flows into the recipient country are largely dependent on unfavorable international interest rates and, *ceteris paribus*, cannot be sustained if they improve.

Taylor and Sarno (1997) use cointegration techniques to investigate the determinants of the large portfolio flows from the United States to Latin American and Asian countries during the period 1988–92. Their study reveals that both domestic and global factors explain bond and equity flows to developing countries and represent significant long-run determinants of portfolio flows. Their study also investigates the dynamics of portfolio flows by estimating seemingly unrelated error-correction models. Their findings indicate that global and country-specific factors are equally important in determining the long-run movements in equity flows for both Asian and Latin American countries, while global factors, particularly USA interest rates are much more important than domestic factors in explaining the dynamics of bond flows.

The impact of institutional factors on capital flows has also been investigated. Taylor (1999) for instance, points out that that policy responses in the 1930s, and subsequent decades of relative economic retardation, can be better understood as the cause and effect of the creation of long-run barriers in international capital markets. To support this notion, he discusses the quantitative extent of these barriers and their effects on economic growth in Latin America. He argues that the political economy of institutional changes in the 1930s in developing countries might be understood in similar terms to those economic historians have used to discuss the macroeconomic crisis in the developed world. He concludes that a political-economy model might thus have universal (rather than core-specific) use: it might predict the "reactive" and "passive" responses by developing countries to external shocks, and the persistence of such shocks in the postwar (Second World War) period.

Similarly, Montiel and Reihart (1999) argue that capital flows respond to the countercyclical policies adopted by countries faced with surges in capital inflows. Their study focuses on two such policies, sterilized intervention and capital controls which target short-term or portfolio flows. On sterilized intervention, they find that while portfolio flows and foreign direct investments do not appear to be responsive to the intensity of sterilization, sterilized intervention significantly alters the composition of capital flows, reducing the share of foreign direct investments in total flows and increasing the share of short-term and portfolio flows. On the other hand, capital controls do appear to alter the composition of capital flows, reducing the share of short-term and portfolio flows while increasing that of foreign direct investments. On the push factors-versus-pull factors debate, they find that foreign interest rates appear to have a significant effect on both the volume and composition of flows: foreign interest rates appear to significantly alter the composition of capital flows, with a rise in US interest rates tending to reduce the share of short-term and portfolio flows. Finally, they find that portfolio flows appear to be responsive to the depth of the equity market, measured by the number of listed companies in the stock exchange; suggesting that bond and equity flows gravitate to those countries that have the more developed financial markets.

More empirical evidence in support of the "push-factors" hypothesis is provided by Antzoulatos (2000). His study, which focuses on bond flows to Latin American countries, suggests that, in the assessment of the prospects of private capital flows to LDCs, policy-makers and market-participants should pay attention not only to interest rates in the industrial world, but also to the global supply of funds. The latter condition, he notes, has been fueled by other factors including financial deregulation and the reduction in budget deficits in the industrial world, the trend towards greater international portfolio diversification, financial innovation, and technological advances in communication and computing. He postulates that these factors will likely continue to affect the global supply of funds. And since they are independent from developments in the developing world, the global supply of funds should not be affected significantly by adverse developments in developing countries. He concludes that a rising global supply of funds may help bond flows to these developing countries withstand the shocks of cyclical interest rate rises and country-specific difficulties.

Empirical work also provides strong evidence that there is a very important linkage between geographical distance and international asset flows. The effect of a specific geographical pattern of international asset transactions was uncovered in a series of papers by Portes and Rey (2000, 2005) and Portes et al. (2001). In the 2005 study, they analyze a panel data set on bilateral gross cross-border equity flows between fourteen countries, 1989–1996. Deriving an estimated equation from a simple micro-founded model of asset trade, their results show that a gravity model[11] explains transactions in financial assets at least as well as trade in goods. To investigate further their hypothesis that distance enters in the equation as a proxy for information asymmetries, they use other variables which plausibly represent international information flows (telephone traffic, number of bank branches, and index of insider trading) and show that these variables are also significant. The study finds weak evidence of a diversification motive in asset trade at yearly frequency. The results of this study help to prove further that international capital markets are segmented by informational asymmetries or familiarity effects; and that countries have different information sets, which heavily influence their international transactions.

Hernández et al. (2001) analyze the determinants of private capital flows toward the developing countries in the 1970s and 1990s, and tests for the possibility of contagion based on trade linkages and country macroeconomic similarities. Their results show that private capital flows are determined mainly by a country's own characteristics and that external or 'push' factors are not significant in explaining the inflows. They also find strong evidence of contagion based on trade linkages for both foreign direct investment and portfolio flows, and some evidence of

[11] This is a static general equilibrium model in which bilateral trade is determined by the wealth and size of countries, the distance between them, and other factors that distort trade (Tamirisa, 1999). In the static model, the volume of trade between two countries increases with the product of their GDPs and decreases with their geographical distance. The idea is that countries with a larger economy tend to trade more in absolute terms, while distance represents a proxy for transportation costs and it should depress bilateral trade (Dell'ariccia, 1999). The microeconomic foundations of the model can be directly linked to the theory of trade under imperfect competition and, more specifically, to intra-industry trade theory.

contagion in private capital flows based on country macroeconomic similarities, with the latter depending on the flow type. They also find strong evidence of contagion in foreign direct investment and portfolio flows for countries in the same geographical region. The contagion effect has also been reported by Lozovyi and Kudina (2007) whose study covers the Commonwealth of Independent States (CIS) region. They find that a growth in portfolio flows to Central and Eastern European countries (which are now members of the European Union) seems to have also enhanced capital flows to the CIS. Though the coefficient measuring this influence is small, it is robust to the inclusion of other explanatory variables. They also find that political stability is an important determinant of portfolio investment into the CIS.

An empirical framework that allows disentangling the relative weight of country-specific and global factors in determining capital flows is presented by Fiess (2003). Through multivariate cointegration methods, the author uses pure country risk and global risk as explanatory variables accounting for the observed pattern of capital flows to Argentina, Brazil, Mexico and Venezuela during 1990 and 2001, to find strong evidence that idiosyncratic ('pull') factors play a significant role in the observed capital inflows. The author finds conclusive evidence that global ('push') factors play a major role in capital flows to Brazil, Mexico and Venezuela. This contrasts with results for Argentina where no such evidence exists. He further finds that the contribution of push factors and pull factors has not been stable during the 1990s.

In a study with less mundane focus, Gande and Parsley (2004) use data from 85 countries for 1996-2002 to investigate the effect of sovereign ratings and corruption on portfolio flows. They find that the effects of sovereign ratings are asymmetric: sovereign downgrades are strongly associated contemporaneously with outflows of capital from the country being downgraded while improvements in a country's sovereign rating are not associated with discernable changes in equity flows. Low levels of corruption however, are associated with a statistically significant reduction in the responsiveness of equity flows to downgrades. The results hold even after adjusting for country size, legal traditions, market liquidity, or crisis versus non-crisis periods, and are robust to different assumptions regarding the within-month distribution of equity flows, monthly predicted benchmark flows, or persistence of equity flows. These results get support from Kim and Wu (2008) who find strong evidence that their sovereign credit rating measures affect financial intermediary sector developments and capital flows. In particular, they find long-term foreign currency sovereign credit ratings important for encouraging financial intermediary development and for attracting capital flows in emerging markets.

Liljeblom and Löflund (2005) investigate the determinants of foreign portfolio investment flows into the Finnish stock market where restrictions for foreign investments were removed in 1993. After the removal of the restrictions, the relative share of the Finnish stock market owned by foreign investors grew rapidly and was, in December 1998, 53% of the total market value of the listed shares. Using company-specific data on the degree of foreign ownership, they report that foreign investment flows are significantly related to variables linked to (i) investment barriers, as proxied by variables such as dividend yield, liquidity, and firm size, and (ii) profitability or risk related variables. Additional analysis of subsequent portfolio performance do not provide robust evidence of apparent informational differences, which would result in either group (foreign or domestic investors) systematically outperforming the other.

In an ambitious study, De Santis (2006) investigates a number of factors influencing equity and bond portfolios in the euro area over the turbulent 1998-2001 period and documents findings suggesting that (i) a decline in home bias generates portfolio outflows; (ii) the higher the initial non-linear degree of misallocation (which might be due to higher fixed transaction costs and information asymmetries) the greater the incentive to reduce them and, consequently, the larger the subsequent bond flows; (iii) asset allocators engage in trend chasing activities in both equity and bond markets in the long term; (iv) after controlling for diversification benefits eliminating exchange rate risk, cross-border portfolio flows among euro area countries have increased due to the catalyst effect of European Monetary Union (i.e., reduction of legal barriers, sharing of common platforms, and simplification of cross-border regulations).

Additional evidence on the "push factors" versus "pull factors" debate is provided by De Vita and Kyaw (2008). Using data of thirty-two developing countries for the period 1990-2004, they employ a variety of panel estimators that control for individual and time effects, potential heterogeneity across individual members of the panel, endogeneity and serial correlation. Their results suggest that, for foreign direct investment flows, domestic productivity growth increases considerably a developing country's attractiveness for investment while foreign output growth exerts a significantly negative influence. For portfolio flows, monetary factors appear to become more important, with domestic money growth emerging as the dominant 'pull' factor.

The quest to understand the determinants of international capital flows has recently moved to international equity funds. Zhao (2008) investigates a data set of retail international equity funds from 1992 to 2001 and documents several findings: (i) diversification benefits appear to be a major reason why investors choose international equity funds. Funds less correlated with the US markets and funds that invest in a diversified portfolio of securities from different regions in the world tend to be preferred by US investors; (ii) risk-adjusted return exerts a greater effect on flows into international equity funds than raw return; (iii) international equity funds from fund families offering a greater number of investment objectives receive higher flows, suggesting that investment in these funds might be affected by investors' general asset allocation strategies; (iv) international equity fund investors do not appear to be sensitive to expenses or exchange rates.

Additional interesting findings on factors influencing portfolio flows have been reported in more recent studies. De Santis and Lührmann (2009), using a panel covering a large number of countries from 1970 to 2003, show that population ageing, institutions, money and deviations from the Uncovered Interest Parity (UIP) influence developments in net capital flows. They present evidence suggesting that population ageing is associated with net equity inflows, net outflows in debt instruments and current account deficits. This finding can be interpreted in two ways: either it corroborates the hypothesis that investors prefer to hold and purchase safer assets when they get older, re-allocating part of their investments towards fixed income portfolios, or it endorses the hypothesis that foreign investors may reduce their investment in bonds issued by ageing countries, with potential negative consequences on future domestic bond prices. It is not, however, clear from this study which of the two interpretations is dominant.

Further findings of the De Santis and Lührmann (2009) study are as follows. First, better institutions favor net capital inflows. Second, higher money to GDP ratio – associated with lower interest rates – enhances international investments in domestic stocks to the detriment of the less attractive domestic bonds. Third, a rise in the short-term domestic interest rate above its trend brings about an equilibrating portfolio shift away from domestic debt instruments. Clearly, this study emphasizes the "pull" factors in attracting foreign capital flows.

Egly et al. (2010) examine the relationship of net foreign portfolio investment inflows (that is, stocks and corporate bonds) to two pull factors – investor risk aversion and the US stock market. Using a vector autoregressive model, they find that positive shocks to the stock market elicit an insignificant response to the net corporate bond inflow and a significant short term positive response to the net corporate stock inflow. The net corporate stock inflow does not respond to risk aversion, while bond inflows do exhibit a significant midterm response to an increase in risk aversion. They also report results showing that internal country-specific factors may influence foreign portfolio inflows.

On the theoretical front, Tille and van Wincoop (2010) develop a simple two-country Dynamic Stochastic General Equilibrium (DSGE) model with portfolio choice in order to shed light on the implications of portfolio choice for international capital flows. They show that capital flows are driven by three factors: portfolio growth, portfolio reallocation associated with time-varying expected returns, and portfolio reallocation associated with time-varying second moments; the last two playing a sizable role. They also find that several important factors that determine equilibrium expected return differences have no effect on capital flows. The model stresses the relevance of endogenous variations in second moments, even though the standard deviation of model innovations is constant. These changing second moments affect capital flows only to the extent that they affect domestic and foreign investors differently, and lead to positive co-movements between capital inflows and outflows.

2.2.2 Barriers to International Capital Flows

Investment theory suggests that the ability to diversify risk – by investing in internationally diversified portfolios of stocks – can influence investment decisions. Levine and Zervos (1996) explain, however, that the ability to diversify risk internationally may be impeded by barriers to international capital flows such as taxes, regulatory restrictions, information asymmetries, and sovereign (country) risk. These barriers will reduce capital market integration and keep arbitrageurs from equalizing risk internationally.

To an individual investor, the benefits of an investment in any security ultimately depend on a trade-off between the expected rate of return and its associated risk. According to Claessens (1995), this trade-off can be assessed by considering the underlying factors driving the rate of return and its variability; the efficiency of domestic stock market; the regulatory, accounting, and enforcement standards in the host country; the relative ease of investing in the country; the different forms of transfer risk (for example, the imposition of capital controls which affects the ability to repatriate capital out of the country); and taxes and other transaction costs. Where these factors are unfavorable in a country, they may act as impediments to foreign capital inflows. Concurring with this view are Eun and Janakiramanan (1986), who contend that barriers to international investment may take many forms such as exchange and capital controls by governments which restrict access of foreigners to local capital markets, reduce their freedom to repatriate capital and dividends, and limit the fraction of local firms' equity that foreigners own.

However, in a study of emerging equity markets, Bekaert (1995) argues that formal barriers in the form of ownership restrictions matter little, suggesting that they are either not binding or are circumvented. His study finds that the emerging markets exhibit differing degrees of market integration with the U.S. market, and the differences are not necessarily associated with direct barriers to investment. Secondly, the most important *de-facto* barriers to global equity-market integration are poor credit ratings, high and variable inflation, exchange rate controls, the lack of a high quality regulatory and accounting framework, the lack of sufficient country funds or cross-listed securities, and the limited size of some stock markets.

In a study conducted at the Stock Exchange of Thailand, Bailey and Jagtiani (1994) identify foreign ownership restrictions, liquidity, information availability, and foreign investor 'familiarity' as some of the factors that generate significant price premiums to Alien Board.[12] With the Alien Board price premium computed as the logarithm of the ratio of the Alien Board price to the Main Board price, they find significant differences in the risk exposures and expected risk premiums faced by Thai and non-Thai investors. The Alien Board monthly price

[12] The Alien Board in Thailand is a distinct market where investors of non-Thai nationalities trade stocks which have reached foreign ownership limits. Thus, an investor's nationality determines whether or not he/she trades in stocks listed in this Board.

premium varies from an average of zero at inception of the Board to 10 percent or more some months later. On the average, the premia are significantly different from zero over the study period. The Thai evidence illustrates how substantial differences in expected returns and prices can arise even if segmentation from international capital markets is only partial: values for risky investments vary depending on the nationality of a prospective investor.

Bailey and Jagtiani (1994) conclude that the financial markets generally signal fundamental information about the underlying economy: price discounts on shares restricted to locals are consistent with a booming local economy, high demand for capital, and high required returns on investments. Their study has specific implications for investors, regulators, and policy-makers involved with equity markets in developing countries: price differentials will exist whenever foreign ownership restrictions are significant and effective. Furthermore, the evidence on downward-sloping demand, liquidity, and information availability suggests that high trading activity, good information flow, and privatization of large, well known companies will attract foreign investors. Securities markets in developing countries should include primary and secondary market institutions designed to accommodate and benefit from such interest.

Of the many barriers to international trade and investments mentioned above, the one that has received much attention in the international finance literature is *exchange and capital controls*. Theoretically, exchange controls act as a tax on the foreign currency required for purchasing foreign goods and services and, by raising the domestic price of imports, they tend to reduce trade. Besides this basic effect, exchange and capital controls can influence trade through other channels, for example, transaction costs, exchange rates, foreign exchange risk hedging, and trade financing. Capital controls, in particular, can affect trade in goods by reducing intertemporal trade and portfolio diversification.

Applying the gravity model to 1996 world trade data, Tamirisa (1999) finds that exchange and capital controls represent a significant barrier to trade, which, however, depends on the level of development in each country and the type of exchange and capital controls in place. Defining "controls on current payments and transfers" as exchange controls over current international transactions, and "capital controls" as encompassing controls pertaining to capital account transactions, the study finds that capital controls reduce bilateral trade for developing and transition economies, but not for industrial countries while controls on current payments and transfers are a negligible impediment to trade.

Cardoso and Goldfajn (1998) estimate a vector autoregression with capital flows, controls and interest differentials and show that controls had been temporarily effective in altering levels and composition of capital flows but have no sustained effects in the long run. Almost a decade later, Neumann (2006) makes more or less similar observations. Her study focuses on the impact of capital controls on the composition and volume of capital flows. In an effort to improve the understanding of international capital flows, she incorporates the consequences of asymmetric information into a small open economy model where external funds are required to support future domestic output. She develops a model of international capital flows that ensures a unique debt-equity combination in equilibrium. Using this model, and imposing barriers to capital flows as taxes on first-period transfers or second-period payments of interest or capital gains, she demonstrates that capital controls are effective at changing the composition and volume of international capital flows. As a result, capital controls may be useful in changing the composition of capital flows towards longer-term equity flows. Her findings on the effect of capital controls on composition of capital flows may be particularly important for developing economies subject to information asymmetries that heighten the risk of sudden capital outflows.

Another deterrent to portfolio diversification that has received attention in international investment literature is 'transaction costs'. Rowland (1999) employs an intertemporal portfolio-choice model that incorporates proportional transaction costs to examine two features of portfolio allocation: the domestic bias of equity holdings and the relationship between domestic and international turnover rates. The model demonstrates that the rate of portfolio diversification decreases as the magnitude of the transaction costs increases. As costs increase, active portfolio reallocation decreases and is replaced by passive portfolio reallocation, which is costless and is accomplished through the realization of capital gains. The study also showed that the international turnover rate is greater than the domestic turnover rate because the average holdings of the international asset are small, not because the volume of trading is large.

2.3 Currency Risk: Theory and Evidence

2.3.1 The Behavior of Foreign Exchange Rates

The analysis of investment opportunities needs to take into account several factors that drive the return-generating process. These factors include: the pay-off of the investment, the investor's value judgment and investment horizon, composition of the portfolio, consumption preferences, and the time value of money. In addition, for investors considering putting their money in portfolios of foreign securities, political risk considerations and the behavior of foreign exchange rates must be given adequate attention. Consequently, international finance researchers have, over the years, spent a lot of effort trying to understand the behavior of foreign exchange rates and/or to ascertain whether a distinct probability distribution can adequately describe it.

In theory, financial economists believe that foreign exchange rate volatility, like a number of random economic variables, can be described by a normal probability distribution. However, empirical evidence on this ideal theoretical assumption has been elusive. Giddy and Dufey (1975), using flexible exchange rate data from several Western European countries and North America, find evidence of significant non-normality in the distribution of exchange rates. They conclude that exchange rate forecasting based on the sequence of observed past rates is futile.

Westerfield (1977) examines the underlying probability models that can best be used to describe the observed variability of foreign exchange rates. He uses a data file of weekly foreign exchange rates for Canada, the UK, West Germany, Switzerland and Netherlands covering the period between January 1962 and July 1975 – thus, both fixed and floating exchange rate regimes are examined. After several distributional tests, he rejects the hypothesis that a normal probability model is an adequate description of the sample exchange rates data. The alternative hypothesis, that the sample data are drawn from a member of the non-normal stable Paretian family of distributions, provides a more adequate description of the real world.[13] The major implication of a non-normal stable model is that means and variances do not adequately describe the probabilistic properties of foreign exchange rates; exchange rate movements and speculative risks in the various currencies must be assessed using other measures of variability than variance.

[13] The stable Paretian family, of which the normal distribution is a special case, is characterized by 'fat tails' (a greater probability of the distribution occurring in the tails of the distribution).

Rogalski and Vinso (1978), however, disagree with Westerfield's (1977) conclusion that a symmetric stable model describes adequately both fixed and floating rate changes. Their investigation finds that the distribution of the underlying stochastic process for foreign exchange rate changes was stable Paretian during fixed rate periods, whereas a Student model provides a better description of floating rates. In addition, the impact of the shift from fixed to flexible regimes appears to reduce the 'peakedness' of the distribution of exchange rate changes. Coupled with the reduced 'peakedness' is substantially greater variability during the floating period. This larger dispersion, however, was accompanied by greater average exchange rate changes, which is consistent with an efficient international market.

Friedman and Vandersteel (1982) examine six years of daily foreign exchange spot rate movements for nine major currencies. Their analysis shows that spot rates of exchange are not normally distributed as very large fluctuations (and also very small fluctuations) are more common than one might expect. The data appear to support the hypothesis that there is an underlying normal process which generates the fluctuations, but that the process changes over time. An interpretation is that both the trend and volatility of exchange rate movements are affected by changing economic and institutional factors. These findings cast real doubt on the validity of any data analysis employing statistical time series techniques which presume the normality and/or stationarity of the exchange rate series. Beyond this, the results also suggest that an autonomous two-parameter model cannot adequately characterize foreign exchange risk.

An analysis of fourteen major currencies, also during the flexible exchange rate period, by Calderon-Rossell and Ben-Horim (1982) suggests two major generalizations: (i) there is no unique distribution that represents the behavior of foreign exchange rates of major currencies; and (ii) the behavior of foreign exchange rates is strongly determined by both the foreign exchange rate management policies pursued by the monetary authorities of the respective countries and the underlying economic forces determining foreign exchange rates. From the findings, they infer that different policies for managing the foreign exchange rate bounded by the underlying economic forces result in different probability distributions of foreign exchange rates.

Using a different analytical framework with roughly the same data-set, So (1986) rejects the findings of Calderon-Rossell and Ben-Horim (1982). His analysis does not find strong support for the arguments concerning the effects of foreign exchange management policy on exchange rates. Instead, he argues that economic forces may be more important than foreign exchange management policies in determining the behavior of exchange rates. He also finds that exchange rates can be described by a non-normal, non-stationary distribution. Consequently, the use of ordinary least squares to test foreign exchange market efficiency should result in inefficient parameter estimates. Similarly, the usual mean-variance types of currency risk management practices may not be fully appropriate.

In general, therefore, there seems to be consensus among researchers that the distribution of exchange rates is both non-normal and non-stationary and that parametric investigations employing mean and variance and least squares regression methods may yield inefficient results.

2.3.2 The Pricing of Currency Risk in Capital Markets

Foreign exchange rate volatility and the risk that it generates, foreign exchange risk, are some of the most investigated macroeconomic issues.[14] Modern capital market theory defines foreign exchange (or currency) risk as "the systematic risk associated with a foreign currency denominated return (or cost) stream and measured by the covariance between the rate of change of the exchange rate and the domestic market return" (Jacque, 1981).

An asset, liability, profit or expected future cash flow stream – whether certain or not – is said to be exposed to foreign exchange risk when a currency movement would change, for better or for worse, its parent or home currency value (Buckley, 1990). The term exposure, used in the context of foreign exchange, means that a firm has assets, liabilities, profits or expected future cash flow streams such that the domestic currency value of these items changes as exchange rates change. For purposes of risk measurement and management, Hekman (1985) has defined foreign exchange exposure as the sensitivity of a specific investment's value in reference currency to changes in exchange rate *forecasts*.

[14] An exchange rate is simply the price of one country's currency in terms of another currency. Volatility refers to the random fluctuations in the rates of exchange.

Risk arises because currency movements may alter domestic currency values. In this sense, assets, liabilities and expected future cash flow streams denominated in foreign currencies are clearly exposed to currency risk. But a share of cash flows denominated in the domestic (or reference) currency which is affected by future exchange rates can also generate sensitivity. For instance, a domestic firm selling in its home market may be competing with foreign-based firms. In such circumstances, exchange rate changes may affect the present value of the domestic company's expected cash flows by strengthening or weakening its competitive position against its foreign-based rivals. This study, however, focuses on currency risk exposure associated with foreign currency-denominated financial flows. Accordingly, the study uses Buckley's (1990) definition of foreign exchange risk.

Considerable research effort has been devoted to discerning whether foreign exchange risk is priced in the capital markets. In an examination of the US stock market, Jorion (1991), using two-factor and seven-factor arbitrage pricing models, presents evidence that the relation between stock returns and the value of the U.S. dollar differs systematically across industries. He, however, finds little evidence to suggest that foreign exchange risk is priced in the stock market. He posits that currency risk appears to be diversifiable and that reasons other than pricing must explain why firms in the US actively manage foreign exchange risk. Loudon (1993), in a replication study, but using only the two-factor model, corroborates Jorion's findings. His findings suggest that that a large proportion (30 percent) of Australian industries exhibit significant positive foreign exchange rate exposure, a fact that provides a *prima facie* case for currency hedging. However, this case is somewhat nullified by the study's failure, like Jorion's, to detect any premium for foreign exchange rate risk in Australian equity returns. The latter finding implies that investors are not willing to reward companies for hedging this source of risk.

Both Jorion (1991) and Loudon (1993) use models that rely on the assumption that the currency risk premium is constant over time. Perhaps the use of a model which relaxes this assumption could determine whether the failure to find a significant foreign exchange premium is attributable to time variation. This argument is underpinned by the fact that many researchers, among them Mun and Morgan (2003), have presented evidence to indicate that there exists a risk premium on foreign exchange and the risk premium is time-varying rather than constant.

In recognition of this major setback of the Jorion (1991) model, Dumas and Solnik (1995) use a conditional approach that allows for time variation in the rewards for exchange rate risk to investigate whether foreign exchange rate risks are priced in international asset markets. Using a parsimonious econometric specification, their results for equities and currencies of the world's four largest equity markets, Japan, USA, the UK and Germany, show that foreign exchange risks premia are a significant component of securities rates of return in the international financial markets. The duo concludes that stochastic changes in foreign exchange rates are associated with changes in equity prices and constitute additional sources of risk in asset pricing models. The model used by Dumas and Solnik (1995), although robust, has been criticized (see, for example, De Santis and Gerard, 1998) on the following grounds. First, because it does not specify the dynamics of the conditional second moments, it cannot evaluate the economic magnitude of the exchange risk premiums relative to the market premium. Second, without second moments, it cannot measure several quantities of interest to the investor, such as correlations, betas, and hedge ratios. Lastly, their test should be interpreted as a test of some of the unconditional implications of the conditional model rather than as a direct test of the conditional model.

Choi and Rajan (1997) perform a joint test of market segmentation and currency risk pricing based on individual stock data from seven major countries, outside of the USA, for the period January 1981 to December 1989. They use a multifactor model with domestic and world market factors and an exchange risk factor. Employing the maximum likelihood method to estimate risk premia and factor analysis to provide further evidence on the pricing of risk factors, their results indicate that the factor structure of asset returns is internationally heterogeneous and that many national capital markets can be described as partially segmented, rather than the polar cases of complete segmentation or integration. Importantly, they find that currency risk is a significant factor affecting asset returns in addition to the domestic and world market risk factors.

Similar results are obtained by Choi et al. (1998) using monthly Japanese data for the period January 1974 through December 1995. With both unconditional and conditional multi-factor asset pricing models, they provide results indicating that the foreign exchange risk is generally priced in Japan. More specifically, they provide evidence, in the unconditional model, that the exchange risk is priced in both weak and strong yen periods, when the bilateral yen/US dollar exchange rate measure is used. However, the foreign exchange risk pricing results from the model are sensitive to the choice of sub-periods, suggesting a time-varying nature to the price of the exchange risk. No evidence of pricing is found with this model when the multilateral trade weighted rate is used. For the conditional model, the exchange risk is priced regardless of whether the bilateral or the multilateral trade-weighted exchange rate measure is used.

De Santis and Gerard (1998) propose the use of a parametric approach to test the conditional version of the international CAPM and assess whether currency risk premiums significantly affect international returns. Applying their approach to the markets for equity and one-month Eurocurrency deposits of Germany, Japan, UK, and USA, they find strong support for a specification of the international CAPM that includes both worldwide market risk and foreign exchange risk. They also show that the relevance of market and currency risk as pricing factors in the conditional model is detected only when their prices are allowed to change over time. In addition, they find that the components of the risk premiums vary significantly over time and across markets. However, the study also finds evidence to suggest that the average premium for currency risk is only a small fraction of the average total premium, measured as the sum of market and currency premiums. This may explain why studies that use the unconditional version of the international CAPM are likely to conclude that foreign exchange rate risk is not priced.

Doukas et al. (1999) test whether foreign currency exposure is priced in the capital market of Japan using an intertemporal multifactor asset pricing model that relies on the assumption that the currency risk premium changes through time in response to changes in business conditions and investors' perception of risk. Their results show that currency risk exposure commands a significant risk premium for MNCs and high-exporting Japanese firms although it is less influential in explaining the behavior of average returns for low-exporting and domestic firms. More importantly, they find that Japanese stock returns are associated with significant currency risk premia. That is, it exhibits a large return volatility that is likely to be perceived by investors, who wish to control portfolio risk, as an important underlying source of risk. The study therefore identifies currency risk as one of the factors of special hedging concern to investors in Japan.

Using weekly data observed over the period January 2, 1985 through to December, 12 1991, MacDonald (2000) empirically models risk premia for four foreign exchange markets of Germany, Japan, the UK and USA. From the findings, he reports that survey-based risk premia are time-varying, volatile, and stationary, but exhibit considerable persistence. Secondly, he uses ARCH- and GARCH-based models to test a version of the general equilibrium asset pricing model of the risk premium and finds that foreign exchange risk premia are significantly related to the conditional variance of forecast errors. Finally, he uses the survey-based measures of the risk premium to test a variant of the portfolio balance model and again reports statistically significant relationships between risk premia and the conditional variance of stock market volatility. He concludes, in contrast to much of the extant literature based on the assumption of rational expectations, that "risk premium is alive and well in the foreign exchange market."

Dominguez and Tesar (2001) adopt a data-driven approach to measuring exposure and study a relatively broad sample of countries (Chile, France, Germany, Italy, Japan, Netherlands, Thailand and the UK) over a 19-year period. Their results are consistent with high degrees of foreign exchange rate exposure at both the firm and industry level across eight countries studied. They posit that the absence of evidence (or weak evidence) on the relationship between international stock prices and foreign exchange rates in previous studies may be due to restrictions imposed on empirical specifications used in those studies. Contrarily, Robotti (2001) finds that inflation and foreign exchange risks do not seem to be priced either unconditionally or conditionally in international equity markets. His international intertemporal capital asset pricing model, tested in the presence of deviations from purchasing power parity, finds evidence in favor of at least mild segmentation of international equity markets in which only global market risk commands a significant and highly positive unconditional risk premium as indicated by its relative Sharpe ratio. Foreign exchange risk and global market risk both exhibit time-variation.

From the Australian equity market, Iorio and Faff (2002) generate results that are somewhat mixed and inconclusive. They implement several variations of a two-factor asset pricing model, using a systems GMM approach on monthly data for the period 1988 to 1998. First, regardless of the model specified, the GMM test results are largely statistically insignificant. Foreign exchange risk appears to be priced for the full sample period. However, when they partition the sample into four major sub-periods, they observe that foreign exchange risk is only priced in two of sub-periods both of which, coincidentally, marked times of relative weakness and uncertainty in the Australian economy, and a secularly weak Australian dollar.

Hsin et al. (2007) examine why the burgeoning literature finds no prevailing evidence of significant exposures to exchange rate risk for US stock returns. Their evidence reveals the crucial role played by the lagged exposure: the inclusion of the lagged effect into the exposure measurement alters the significance of the currency risk for individual stocks. Nevertheless, despite altering the exposure significance for individual stocks, the inclusion of the lagged effect in the exposure measurement still fail to raise the significance of exchange rate risk with regard to the pricing for the overall sample of stocks. In about 50 percent of the return series, the reactions to currency changes are revised in the opposite direction during the next period.

More recent evidence from the U.S. stock market seems to identify a higher level of significance for exchange rate risk. Kolari et al. (2008) show that foreign exchange risk is priced in the cross-section of USA stock returns during the period from 1973 to 2002. They initially demonstrate that stocks with extreme absolute sensitivity to foreign exchange have lower required rates of return than other stocks. Next, departing from previous studies, they test whether exchange rate risk is priced in the cross-section of US stock returns by forming a zero-investment factor based on the absolute foreign exchange-sensitivity, and show that this factor can significantly reduce mean pricing errors for foreign exchange-sensitive portfolios. Also, various two-dimensional sorts of asset-pricing factors indicate that estimated coefficients of the zero-investment factor are generally significant. Finally, their ex-ante tests show that risk premia associated with foreign exchange sensitivity are significant and negative during the sample period.

Cappiello and Panigirtzoglou (2008) use a general no-arbitrage pricing kernel/stochastic discount factor model that permits estimation of market prices of risk. In an international framework, the model allows the computation of foreign exchange risk premia. They find that market prices of risk are time-varying and increase during periods of financial turmoil. They conjecture that investors become more risk-averse during turbulent financial markets. Importantly, their findings suggest that foreign exchange risk premia are also time-varying and exhibit most variation from the early 1970s onwards, when the Bretton Woods exchange rate system collapsed.

Emerging Markets' Evidence

Literature on the pricing of exchange risk in the emerging markets is not quite extensive. Claessens et al. (1998), in their examination of the cross-sectional pattern of returns in the emerging markets, were among the early researchers to provide a role for foreign exchange risk. Their work, covering eighteen developing country markets, suggest that, in addition to beta, two factors, namely, size and trading volume, have significant explanatory power in a number of these markets; dividend yield and earnings/price ratios were also important, but in slightly fewer markets. For a number of the markets studied, however, the relationship between all four of these variables and stock returns contradicts the relationships documented in the USA and Japanese markets. Importantly, their findings also suggest that exchange rate risk is a significant factor in explaining stock returns in several emerging markets countries.

Tai (1999), in a study covering five Asia-Pacific countries and the USA finds support for the idea that the predictable component in deviations from uncovered interest parity (UIP) is due to a time-varying foreign exchange risk premium, and not to irrationality among market participants. His evidence of significant foreign exchange risk pricing supports the idea that foreign exchange risk is not diversifiable and hence investors should be compensated for bearing this risk. It also supports the role of deviations from purchasing power parity (PPP) in pricing foreign exchange rates and equity. Furthermore, his empirical results suggest that a multi-factor asset pricing model, especially in its conditional form, outperforms a single-factor asset pricing model.

Glen (2002) investigates stock market performance over a sample of 24 devaluation events covering eighteen emerging market countries over the period 1980–1999. The analysis compares stock market performance before and after the devaluation event with the general distribution of returns for these markets. The findings are interesting: on average, stock returns are reduced in the period leading up to a devaluation event, but the period following these events is characterized by normal return behavior. There is considerable variation across events, however, and much of this variation can be explained by economic growth, the size of the devaluation, and the industry and country in which the event occurs.

Phylaktis and Ravazzolo (2004) use a parsimonious multivariate GARCH-in-Mean process to estimate the conditional dynamics of a system of equations, which also allow for the examination of the effects of capital market liberalization and the Asian Financial crisis of mid-1997 on the volatilities of stock and currency returns. They find strong support for the specification of an International CAPM that includes both market and currency risk. Currency risk is priced in both pre- and post-liberalization periods: thus, omitting foreign currency risk in pricing international assets might give rise to model misspecification. They present evidence that reveals significant variation, over time and across markets, in the components of the risk premiums. Currency risk premium is substantial and forms a big part of the total risk premium, dominating it at times. It is also bigger and more variable when markets are segmented. In general, these results show that currency risk is an important component in international capital asset pricing models even during periods when markets are not officially open to international investors.

Carrieri and Majerbi (2006) also provide empirical evidence on the pricing of exchange risk in emerging stock markets. They use an unconditional framework to investigate whether exchange risk represents, on average over the long run, an important component of expected equity returns. They conduct tests at the market-, portfolio- and firm-level and use real exchange rate specifications to fully account for the effects of PPP deviations. Their results support the hypothesis that exchange risk is globally priced and commands a significant unconditional risk premium in emerging stock markets. The estimated exchange risk pricing coefficients are generally higher than those estimated in similar frameworks for developed markets and, with cross-sectional data at the firm level, there is indication that the size and sign of exchange risk premia vary across countries and regions.

More recent studies continue to show a close relationship between exchange rates and stock prices. In an examination of the dynamic linkages between the foreign exchange and stock markets, Pan et al. (2007) show a significant causal relation from exchange rates to stock prices before the 1997 Asian financial crisis for four of the seven East Asian countries studied. They also find a causal relation from the equity market to the foreign exchange market for three countries. The findings also indicate that the linkages could vary across economies with respect to exchange rate regimes, the trade size, the degree of capital control, and the size of equity market. These results are corroborated by Dube (2008) using South African data. In his study, he finds evidence supporting the existence of a long-run relationship of plausible magnitudes between the Rand/US dollar exchange rate and fundamental variables, including stock prices.

Overall, empirical evidence suggests that, unlike in the developed markets where weak significance is largely reported, currency risk appears to be significantly priced in the emerging capital markets. Since currency risk, like market risk, is systematic, the flow of foreign portfolio investments into these markets most likely reflects investors' willingness to trade off this risk against the potential geographical diversification gains. The next section explores literature on the linkage between foreign portfolio flows and foreign exchange rates.

2.4 Foreign Exchange Rates and the Flow of International Portfolio Investments

If interest rate parity relationship is violated, as is often the case in international financial markets, then the overall rate of return earned from the holding of foreign securities is comprised of the investment return (dividends and capital gains) on the equity securities in question plus the gain or loss which results from a change in the exchange rate during the holding period.[15] In such situations, fluctuations in foreign exchange rates will be a source of potential gain or loss. The introduction of floating exchange rate regimes in the 1970s in developed countries and late 1980s and early 1990s in the emerging markets have worsened foreign exchange rate fluctuations and significantly increased the uncertainty associated with foreign investments. These developments have increased interest among international finance researchers on the linkage between foreign exchange rate volatility and international capital flows. At the general level, Landon and Smith (2009) remark that although inflation may eventually erode the impact of a nominal depreciation on the real exchange rate, the observed long periods of currency over- and under-valuation suggest that this may take considerable time and that, in the interim, foreign exchange rate changes could have a significant impact on investment activity.

[15] Other than diversification benefits, violation of the IRP theory is often seen as the main motivation behind international portfolio investments. If IRP holds, returns from foreign investments, adjusted for exchange rate movements, would equal returns on domestic investments and there would be no incentive for external capital flows.

A theoretical relationship between international capital flows and foreign exchange rate changes has been proposed by Bailey et al. (2001). Their model traces the linkage between the two variables to productivity shocks in the economy. In the simplest version of the model, there is only one good and purchasing power parity holds at all times. In this case, a productivity shock will generate capital inflows without affecting the real exchange rate. In the more developed version of the model, each economy is endowed with two goods, one that is tradable with the other country and one that is not tradable. In this case, a shock that raises the productivity of both sectors in one country will always lead to long-run real exchange rate depreciation. If, on the other hand, the shock affects only the tradable sector, the real exchange rate will appreciate in both the short and the long run.[16] Provided the productivity shock has some effect on the tradable sector, capital flows towards the country experiencing the shock will still be observed.[17]

2.4.1 Evidence from Equity Markets of Advanced Economies

The seminal work of Solnik (1974) advances a strong case for international diversification, arguing that it presents more risk reduction benefits than domestic diversification. Using data drawn from eight countries in Europe and North America, he generates portfolios containing an increasing number of stocks, including several of the same size, and obtains an average measure of risk for each portfolio. Plotting these results on a plane of risk against number of securities in

[16] Within the uncovered interest parity (UIP) framework, a shock to productivity that raises the future level of productivity, would lead to an increase in the domestic real interest rate relative to the world real interest rate; that, in turn, would prompt a jump appreciation of the real exchange rate. As productivity growth returns to trend, bringing the domestic real interest rate back into line with the world real interest rate, the real exchange rate would depreciate back to its equilibrium value. The effect on the equilibrium real exchange rate would depend importantly on whether the productivity shock is concentrated in the tradable or the non-tradable sector. A productivity shock that affects both sectors equally is likely to lead to a depreciation of the equilibrium real exchange rate. This happens because such a shock implies an increase in the relative supply of domestic goods and services; given this, their relative price must fall. On the other hand, a productivity shock concentrated in the tradable sector is likely to lead to an appreciation of the equilibrium real exchange rate. This occurs because product market arbitrage between countries equilibrates prices for the tradable goods and services at the same time as labor market arbitrage within economies means that wages are equalized at the margin between the tradable and non-tradable sectors. If one country has an increase in the productivity of its tradable sector, other things equal, real wages will increase in both the tradable and non-tradable sectors. Because there has been no productivity change in the non-tradable sector this leads to a rise in the price of non-tradable goods and services relative to tradables in the home economy, and an appreciation of the real exchange rate (Bailey et al. 2001).

[17] Assuming that consumers are sufficiently forward-looking to wish to smooth their consumption over the present and future time periods, a productivity shock that raises expected future output in the home country will tend to lead to capital inflows. This is because the expected increase in future productivity would raise expected future profits, which would, in turn, lead to an increase in equity prices. This would encourage investment. Residents of the home country would want to take advantage of current investment opportunities that enhance future output but without forgoing current consumption. So the increase in investment demand that is not financed by current domestic savings would be financed by inflows of capital. And inflows of foreign direct investment and foreign equity investment are particularly likely to increase as overseas investors also take advantage of the higher rates of return to capital in the home country (Bailey et al. 2001).

a portfolio, he demonstrates that risk reduction benefits would be greater for internationally diversified portfolios. However, he cautions that such risk reduction benefits might be reduced by many institutional, political and psychological factors such as foreign exchange controls, capital restrictions on portfolio holdings, and the existence of foreign exchange risk. He suggests that investors could remove exchange rate risk from foreign portfolio holdings through hedging techniques, such as forward contracts. In the absence of such hedging, he claims, the investor would be speculating on the foreign currency itself. Although he does not conduct an investigation to verify this latter claim, the essence of his argument is that currency risk exists in international portfolio holdings and therefore has the potential of influencing portfolio flows.

Levy and Sarnat (1975) examine the implications of different countries' viewpoints on the composition of the efficient internationally diversified portfolio. In a study of portfolio composition for American and Israeli investors they show that, given a series of devaluations of the Israeli currency, investment in Israeli equities would not be part of the efficient set for US investors. Essentially, the study finds that exchange risk is priced and foreign investors would require commensurate compensation to lure them into Israeli stocks. They suggest several incentive schemes that could be offered to American investors to remove foreign exchange risk and encourage diversification into the Israeli stock market.

Kohlhagen (1977) specifies exchange rate expectations as a function of real economic variables that are endogenous to a fully specified model of capital flows. He estimates both the capital flow equation and an expectations function for the fixed and floating Canadian dollar and for selected short-run periods. The expectations function enables the analysis of the implications of exchange rate expectations without resorting to highly mechanistic, *ad hoc*, and/or exogenous representations of speculation. The model demonstrates that the parameter values of this expectations function determine the responsiveness of capital flows to domestic monetary policies and external disturbances. Empirical results imply that international capital flows do not have a well-defined, stable relationship with the exogenous variables of such a model and at least suggest that part of the reason lies in the variability of foreign exchange rate expectations.

Biger (1979) evaluates the systematic risk of foreign exchange by deriving efficient sets of international portfolios from six national viewpoints. The composition of these portfolios is examined and the effect of different exchange rate risks is discussed theoretically and tested empirically. The study shows that in the context of international portfolios, currency risk matters much less than would be expected. Biger's results contradict those of Levy and Sarnat (1975): his conclusion gives no role to foreign exchange rate volatility in international investment flows.

Utilizing international stock market and exchange rate data accumulated over a decade of the flexible foreign exchange rates regime, Eun and Resnick (1985) examines international portfolio diversification in the context of flexible exchange rates for fifteen numeraire currencies. Country funds of Sweden, the Netherlands, and Japan dominated the optimal ex-post portfolio of each national investor, comprising 65 to 100 percent of investment. Exchange rate changes affect the desirability of each national stock market, leading to substantial compositional variations in the optimal international portfolio across investors' countries. The potential gains from international diversification are likely to be diminished by exchange rate variations, but the gains still appear to be substantial. In yet another study on the same subject, Eun and Resnick (1988) demonstrate that exchange rate risk is non-diversifiable to a large extent due to the high correlations among the changes in the exchange rates and, as a result, substantially contributes to the overall risk of the international portfolio. Their analysis shows that fluctuating exchange rates make foreign investments more risky and, at the same time, aggravate estimation risk; thereby diminishing the gains from international diversification. They suggest that the US investor can substantially increase the gains from international diversification by using a hedging strategy. This suggestion arises from the finding that all of the hedging strategies, designed to control both estimation and foreign exchange rate risks, substantially outperform any of the 'unhedged' strategies.

Brennan and Cao (1997) find little evidence that US capital flows are associated with foreign exchange rate changes except insofar as these are impounded in the US dollar returns. The results are considerably different for investment by residents of the developed countries in the USA. Their likelihood ratio test rejects the null hypothesis that these flows do not depend on exchange rate changes in addition to the US market return measured in US dollars. When they test whether the results are affected by including lagged values of the investment flows as an independent variable, they find no association between current and lagged US flows to either developed or emerging markets. However, they find lag effects for investment in the US from the developed countries. They attribute these differences to information asymmetry; that is, they find some evidence that residents of the USA are at an informational disadvantage in the equity markets of other countries, while no similar evidence suggests that residents of those other countries are at an information disadvantage in the US equity markets exists.

Using an augmented Vector Autoregression (VAR) to estimate dynamic models that includes a measure of net capital flows, the nominal exchange rate, equity return differentials and interest rate differentials for five OECD countries, Siourounis (2003) presents evidence that, for four major countries (Germany, Switzerland, the UK, and USA), a good portion of their exchange rate movements can be explained by net cross-border equity flows. For Japan, however, the evidence is inconsistent with theory since net purchases of US assets from Japanese residents are associated with a strong yen. In terms of the predictive content of the empirical model in short to medium-horizon forecasts (1 month to 2 years), Siourounis shows that dynamic forecasts from an equity augmented-VAR provide support for exchange rate predictability and outperform a random walk and a standard VAR that includes only exchange rates and interest rate differentials. Overall, the study shows that as net purchases of cross-border equities increase their share in total flows (including foreign direct investment, bank flows etc.), their effect on nominal exchange rates become increasingly important.

In a study linking equity returns, equity flows and foreign exchange rates across the world's largest stock markets (France, Japan, Germany, the UK and USA), Hau and Rey (2004), use a variance-covariance decomposition approach, to show that global investors repatriate foreign equity wealth either because of foreign-equity excess returns or after an unexpected appreciation of the foreign currency. Moreover, these equity flows move the exchange rate in line with a price-inelastic supply of foreign exchange balances. Portfolio flow shocks appreciate the foreign-exchange rate and create foreign equity-market excess returns.

In a bid to shed some light on a significant puzzle in international currency markets in which the US dollar appreciated while the euro and the Japanese yen weakened significantly despite a record US current account deficit, Brooks et al. (2004) ascribe the puzzle to sharp increase in large capital flows among the three currency areas. In particular, their study finds that observed large inflows into the US equity markets and direct investment flows financed the current account deficit and allowed the dollar to remain strong. Conversely, large and initially unanticipated outflows from the euro area appear to account for a substantial part of its fall and persistent weakness. Their analysis of the relationship between foreign exchange rates and portfolio flows suggests that in the second half of the 1990s, in the case of the euro area, there was a strong relationship between exchange rate movements and equity flows, with an increase in equity flows to the USA associated with a clear depreciation of the euro (or synthetic euro prior to 1999) vis-à-vis the US dollar. However, such a clear relationship is not seen in the case of Japan, the bulk of whose portfolio flows were in the form of US sovereign bond purchases.

An apparent contradiction of the conventional belief that strong equity markets are accompanied by currency appreciation comes from Hau and Rey (2006). In a study of a pooled data of 17 OECD countries they derive a negative correlation between foreign equity excess returns (in local currency) and the corresponding foreign exchange rate returns. The negative relationship is induced by the rebalancing of the portfolio of global investors who decrease the exposure of their investments to exchange rate risk. Such a negative correlation decreases the risk of foreign investment in home currency terms as negative foreign equity returns tend to be compensated by positive exchange rate returns. This automatic hedge reduces the home bias and facilitates international equity risk sharing. The cross-sectional evidence also points to the role of financial market development. Countries with a higher equity market capitalization relative to Gross Domestic Product (GDP) tend to have a more negative return correlation. They also explore the correlation between exchange rate returns and net equity flows. Their model predicts a positive correlation and the authors explain that net equity flows are tied to foreign exchange order flows.

Heimonen (2009) estimates the equity flow equation and the exchange-rate equation simultaneously using the Full Information Maximum Likelihood (FIML) in a system in which both the exchange rate and equity flows are treated as endogenous. The results indicate that an increase in euro area equity returns with respect to USA equity returns caused an equity capital outflow from the euro area to the USA. This equity flow generates an order flow in the foreign exchange markets, which leads to appreciation of the dollar. The author explains that the equity flows between the US and the euro area are deviations from the minimum variance portfolio.

Thus, an increase in the return on foreign-country equities with respect to domestic returns increases the relative share of foreign equities in agents' total wealth and implies a deviation from the minimum variance portfolio of foreign equities. As a result, the investor decreases his holdings of foreign country equities. There is an equity outflow from the country with the excess equity returns. This equity outflow generates an outflow in the foreign exchange market, which finally causes depreciation of the currency with excess returns.

Consistent with the theory, both foreign exchange rate returns and order flows into the overseas market have explanatory power for the domestic stock market returns. Dunne et al. (2010) recently developed a model that can account for observable asymmetries in the correlation structure between equity returns and foreign exchange rates. They derive a closed-form solution for equity returns in domestic and foreign equity markets, which relates equity returns to the foreign exchange rate and to order flows in both the local and the overseas market. The model can potentially explain asymmetry across countries in the correlations between domestic equity returns and the exchange rate return conditional on order flows. They test the model with 5 years of daily US (domestic) and French (foreign) equity data. The respective daily order flows for the S&P100 and the CAC40 index are constructed based on the aggregation of approximately 800 million individual equity transactions. They find that an extraordinarily high percentage of aggregate equity return variation is explained jointly by macroeconomic order flows and foreign exchange rate returns.

2.4.2 Evidence from Emerging Countries' Equity Markets

Unlike the voluminous literature on the determinants of foreign direct investment (FDI) inflows, research on the determinants of portfolio inflows to emerging markets is limited. Most of the empirical analyses focus on industrial countries, owing, in part, to data availability and the fact that portfolio flows, especially equity inflows, to emerging markets and developing countries, began only recently, mostly in the late 1980s. It has been argued that large inflows create difficulties in containing monetary and credit expansion, and may have adverse effects on inflation, the external current account, and the real exchange rate. Indeed, several of the main recipient countries in Asia and Latin America have experienced an increase in domestic inflation and a significant real exchange rate appreciation. In low-income Africa, where successful stabilization has been associated with large and persistent increases in aid and private capital flows, Buffie et al. (2004) have observed that foreign-denominated assets constitute an important share of private financial wealth. The incipient capital inflow places acute short-run pressure on the foreign exchange market, dramatically undermining the case for a floating exchange rate. Such portfolio pressures produce a nominal appreciation that is an order of magnitude larger than the required real appreciation, and unless the prices of non-traded goods are perfectly flexible, the real exchange rate overshoots and substitution effects produce a potentially deep recession.

Using country mutual funds, Frankel and Schmukler (1996) investigate the chain of events surrounding the Mexican crisis of 1994. Their results indicate that the Mexican devaluation of 1994 may have been different from other exchange rate changes and can only partially explain changes in country fund discounts. In other words, the fall in the discount in December 1994 was greater than would be expected from the magnitude of the devaluation and the usual pattern associated with exchange rate changes. They interpret this as a loss in confidence by Mexican investors (relative to US investors). This supports the hypothesis that the change in discounts was partly due to less optimistic Mexican investors, and not simply to the devaluation itself.

The influential work of Edwards (1998), conducted with quarterly data from eight Latin American countries, documents a negative relationship between capital inflows and the real exchange rate, i.e. increases in capital inflows are associated with real exchange rate appreciation while declines in inflows are associated with real exchange rate depreciation. Results of the Granger Causality tests show in seven out of the eight countries, that it is not possible to reject the hypothesis that capital flows cause real exchange rates; and in three of the eight countries, that it is not possible to reject the two-way causality hypothesis. More strikingly, the data present no evidence that the real exchange rate causes capital inflows. Edwards interprets these results to lend support, at a preliminary level, to the view that a surge in capital inflows is responsible for generating loss in real international competitiveness.

Kim and Singal (2000) investigate the effects of opening up of stock markets to foreign investors on changes in the level and volatility of stock prices, inflation rates and exchange rates. They use both non-parametric and parametric tests on a sample of twenty emerging market economies over a twenty year period, 1976–1996. The non-parametric tests suggest that there was a significant decrease in volatility for three years after market opening when compared with the corresponding pre-opening periods. Results of the parametric tests are also consistent with the non-parametric tests, except that the decreases are significant only for the second and third years after opening. However, none of the tests imply an increase in the volatility of changes in nominal exchange rates. The reduction in currency risk implies that foreign investors exert a calming influence on volatility. They conclude by highlighting two ways in which the lower volatility of changes in exchange rates is useful: first, the volume of trade is likely to increase as a result of less risk related to trade. Second, the lower currency risk will encourage foreign investors to invest more at a lower required rate of return.

Froot at al. (2001) use daily data of portfolio flows for forty four countries, from both developed and emerging markets, to examine the covariance of equity returns with cross-border flows. They find a statistically positive contemporaneous covariance between net inflows and both dollar equity and currency returns. The data also reveal strong evidence of correlation between net inflows and lagged equity and currency returns, with the sign generally positive. This pattern suggests that international investors engage in positive feedback trading, or 'trend chasing.' The flows are also correlated with future equity and currency returns in emerging markets.

From the developing world, Ndung'u and Ngugi (1999) estimate a VAR for Kenya with real effective exchange rate, domestic inflation, real interest rates differential, money supply and volatility of capital flows. The results showed that the real exchange rate movements *and* real interest rate differential absorbs over half of the forecast error variance of the volatility in the private capital flows at the end of the forecast horizon. For the impulse response functions, a unit shock in the volatility of capital flows leads to an initial decline in the real exchange rate followed by a continuous rise, with no signs of the effects dying out. Volatility in capital flows only accounts for about 7 per cent of the innovations from the real exchange rate implying that there is weak feedback from the real exchange rate movements to the volatility in capital flows.

Following the observation of an imbalance between foreign direct investments and portfolio inflows, into South Africa, in favor of the latter, Ahmed et al. (2005) conduct an investigation whose results suggest that a number of policy variables contribute to the lower share of foreign direct investment and higher share of portfolio flows. The policy variables unearthed by the study include failure to fully liberalize trade, poor growth and infrastructure and weak observance of law and order. The results also suggest that lower currency volatility would contribute to an increase in the share of foreign direct investment. Put differently, a stable currency would encourage foreign direct investment as it deemphasizes portfolio inflows!

In their analysis of real exchange rate volatility as a key determinant of international portfolio allocation and home bias, Fidora et al. (2007) take a global perspective (40 investor countries, covering all major industrialized and emerging markets economies, and up to 120 destination countries). They analyze the importance of real exchange rate volatility in explaining cross-country differences in home bias, and in particular as an explanation for differences in home bias across financial asset classes (that is, between equities and bonds). They use a Markowitz-type international capital asset pricing model (CAPM) which incorporates real exchange rate volatility as stochastic deviations from PPP. Given a mean-variance optimization which implies risk aversion of investors, real exchange rate volatility is found to induce a bias towards domestic financial assets because it puts additional risk on holding foreign securities from a domestic (currency) investors' perspective, unless foreign/local currency real returns and the real exchange rate are sufficiently negatively correlated.

Through dynamic panel data techniques and a panel of 85 developing and developed economies for the sample period 1997–2006, Saborowski (2009) provides strong evidence for the hypothesis that the exchange rate appreciation effect of capital inflows is lower in countries with a higher level of financial development. Using a Behavioral model of the exchange rate that includes different types of capital inflows and interaction terms between the inflow variables and indicators of financial sector development, the author shows that the real appreciation effect of foreign direct investment on the exchange rate is significantly attenuated if an economy disposes of a deep financial sector as well as large and active stock markets. However, the study does not find similar evidence for other types of capital inflows.

Recent dynamic panel data evidence is provided by Jongwanich (2010) for nine emerging Asian countries, including the People's Republic of China; India; Indonesia; the Republic of Korea; Malaysia; Philippines; Singapore; Taipei-China; and Thailand during 2000–2009. The estimation results show that compositions of capital flows matter in determining impacts on real exchange rates. Other forms of capital flows, both portfolio investment and other investment (including bank loans), bring in a faster speed of real exchange rate appreciation than foreign direct investment inflows. The nature of foreign direct investment flows, which are relatively stable and concentrated mostly in tradable and export-oriented sectors, leads to the slower adjustment of non-tradable prices and the real exchange rate. However, the magnitude of appreciation among capital flows tends to be close to each other. The estimation results also show that during the estimation period, capital outflows bring about a greater degree of exchange rate adjustment than capital inflows. The latter evidence is found for all types of capital flows.

The nature of the relationship between capital flows and exchange rate fluctuations has also received some attention in the Sub-Saharan Africa region. Kasekende et al. (1996) observe that private capital inflows, though predominantly short-term, have led to short-term exchange rate appreciations in almost all countries in his sample – Kenya, South Africa, Tanzania, Uganda, Zambia and Zimbabwe. They observe, for instance, that Kenya and Uganda experienced very sharp appreciations in their exchange rates in 1993/94 and 1994/5 at a time when private transfers and access to short-term credits markedly increased.

2.4.3 Lessons from Studies on Bond Returns

Among the early researchers to investigate the pricing of economic variables in bond markets were Ibbotson et al. (1982). Using return data from bonds and stocks of eighteen countries drawn from North America, Europe and Asia between 1960 and 1980, their study indicates that deviations from the international parity theorems occur often, especially over short periods of time. Consequently, risk is generally rewarded in both stock and bond markets. The study's major findings are that (i) inflation hurts both the stock and long-term bond markets in most countries, while a country's short-term securities tend to track its inflation rate; and (ii) non-U.S. bonds benefitted from appreciations against the dollar in the 1970s, making them superior investments from a U.S. dollar investor's perspective. These findings suggest that the economic relationships often posited between international stock and bond expected returns, inflation and foreign exchange rates hold only imperfectly.

Using a beta asset pricing model with regression analysis, Adler and Simon (1986) investigate exchange risk exposure in international portfolios. They estimated exposures of long-term foreign bonds from monthly bond indexes for the period 1973-1980 and a series for foreign exchange rates and local currency market indexes for the period January 1976 to December 1982 for nine countries. Their analysis shows that some foreign stock market indexes were more exposed to foreign exchange risk than foreign bonds during 1976–1979 period and that their exposures have generally risen since October 1979 compared to the earlier period.

According to Dym (1992), foreign exchange rates affect the risk of a foreign bond in two ways. First, the coupon and face value of the bond are paid in units of the foreign currency. Looking at the issue from a US investor's perspective, he points out that the US investor will be directly affected by changes in the foreign currency's exchange rate with respect to the US dollar. In addition, changes in the foreign exchange rate may be associated with movements in the bond's yield. Dym suggests a risk measure that is adapted to reflect the foreign currency-denomination of a bond's cash flows. The measure accommodates both direct and indirect effects of foreign exchange rate movements.

An analysis by Sturges (2000) explores the relationship between foreign bond and currency returns using return data from Canada, Germany, Japan, the UK, and USA for the period January 1980 through October 1997. Employing a sticky-price model, he estimates two-country variance autoregressive systems to analyze the fundamental components and covariances of bond and currency returns. His findings suggest that a relationship exists between innovations in bond return fundamentals and currency premia. Thus, for all countries analyzed, news that increases expectations of future inflation and interest rates (bond return fundamentals) also increases the currency risk premium. Further, an expectation of increases in future foreign interest and inflation rates lowers the current period's excess bond return, while the higher risk premium lowers the current period's excess currency return.

Min et al. (2003) investigate the determinants of bond spreads for eleven emerging markets economies during the period 1990 to 1999. Based on panel data regression with a simple beta-pricing model, they find that emerging economies' liquidity-related variables (especially the international reserves-to-GDP ratio) play an important role in determining bond spreads. Several macroeconomic fundamentals like domestic inflation rate, net foreign assets (as measured by the cumulative current account), terms of trade, and real exchange rate are also found to be significant in determining bond yield spreads. However, external shocks, when measured by the real oil price, are found to be insignificant while international interest rate is significant in determining yield spreads. Ferrucci (2003), using estimation procedures and data roughly similar to that of Min et al. (2003), makes similar conclusions in an investigation of the determinants of emerging markets sovereign bond spreads. Ferrucci's results suggest that a debtor country's fundamentals and external liquidity conditions are important determinants of market spreads. However, the diagnostic statistics also indicate that the market assessment of a country's creditworthiness is more broad based than that provided by the set of fundamentals included in the model. The study also find that the generalized fall in sovereign spreads seen between 1995 and 1997 cannot be entirely explained in terms of improved fundamentals.

Utilizing a GARCH(3,3) specification that allows for adjustments for the time-varying volatility structure of return series, Batten et al. (2006) investigate the factors affecting yields on bonds in nine emerging markets countries from East Asia and the Pacific region. They use a sample of daily yields for the period beginning December 30, 1999 and ending November 28, 2002. Consistent with theory, they find the interest rate factor to be statistically and economically significant with a negative coefficient in all the nine cases. However, they find the asset factor in only three of the nine cases, with a negative coefficient in two cases and positive in the remaining one. A positive relationship suggests that a rise in the stock market is associated with an increase in the spread. Although this result is inconsistent with theory, they attribute it to the possibility of portfolio rebalancing between bonds and stocks held by international portfolio managers. Finally, their results find the foreign exchange rate variable insignificant in eight of the nine cases, implying that currency risk is not priced in the bond markets.

Young et al. (2007) examine return volatility among the Swiss, German, UK, and US bond markets by comparing both in-country and US-based returns and risk measures. Among other important results, they find that low correlations between US bond market returns and European bond market returns offer potential diversification benefits to bond-portfolio investors and that, although adding to return volatility, currency returns have generally enhanced overall returns to US-based investors in European bonds. Their results also show that global bond market investors can achieve the greatest efficiency in terms of risk per unit of return by hedging currency risk. Unlike Batten et al. (2006) therefore, they find currency risk to be priced in bond markets.

Noting that emerging markets bond issues have recently become popular with developed countries' institutional investors (particularly mutual funds), Xiao (2007) demonstrates that such investors show strong preferences towards bonds with certain country economic and bond financial characteristics. Specifically, mutual funds prefer to invest in bonds issued by countries with sound fundamentals and more openness to trade. Sound fundamentals such as fiscal discipline, high reserves, and favorable current account position reduce the countries' balance sheet risk, strengthen their repayment capacity, and lower their leverage. More trade openness boosts the countries' visibility and increases foreign investors' familiarity with their bonds. Mutual funds also favor bonds with high past returns and yields while shunning bonds with high transaction costs and idiosyncratic risks. Although Xiao does not give a specific role to exchange rate risk, it should be observed that foreign currency reserves and current account balances have an influence on the value of a country's currency and can therefore proxy for currency risk.

Ebner (2009) investigates possible determinants of the yield difference between ten-year Euro-denominated Central and Eastern European (CEE) bonds and the ten-year sovereign bond issued by the German government as the "risk-free" benchmark. He finds that international risk, captured by the market volatility, is the single most important explanatory factor. This common factor drives bond spreads in CEE. He concludes that important political and economic events contribute to a better understanding of sovereign spreads. However, the study does not provide a distinct role to currency risk as a determinant of sovereign bond spreads.

2.5 The State of Affairs in Sub-Saharan Africa[18]

The IMF/World Bank (2008) report on sub-Saharan Africa presents some interesting insights into the relationship between foreign exchange rates and portfolio investment flows into the region. The report notes that response to rising foreign trade and capital inflows has been a pressing challenge for many countries in the sub-Saharan African region. While the inflows have helped raise investment and growth in some countries, they have also put pressure on prices and the real exchange rate. In the CFA franc zone[19] where currencies are pegged to the euro and monetary policy is determined at the currency union level, rising inflows and the strengthening of the euro against the US dollar in 2007 have led to a modest appreciation of the real effective exchange rates of both the Central African Economic and Monetary Community (CEMAC) and the West African Economic and Monetary Union (WAEMU). Among all countries with a flexible exchange rate regime, the real effective exchange rate has been appreciating only for oil exporters and some low-income countries. In many countries, exchange rate adjustments have only partly reflected their current account positions. In South Africa, under the inflation-targeting regime, continued inflation pressures led the South African Reserve Bank to resume its monetary tightening in 2007 to fight inflation pressures. The Reserve Bank has continued to strengthen its international reserves without an explicit exchange rate objective. In Nigeria, foreign reserve accumulation has helped stabilize the official exchange rate against the dollar since 2004.

Private capital flows to sub-Saharan African countries have increased almost five-fold over the past seven years, from US$11 billion in 2000 to US$53 billion in 2007. The increase in portfolio flows to US$23 billion in 2006 was particularly rapid, reaching about 14 times the 2003 level. Private debt flows have also increased rapidly since 2004. However, these flows remain small compared with total global capital inflows of about US$6.4 trillion in 2006. In 2006, private capital flows to sub-Saharan Africa overtook official aid for the first time. The bulk of these flows went to South Africa and Nigeria, but portfolio flows are also trending up in a small group of other countries notably, Ghana, Kenya, Tanzania, Uganda, and Zambia, in response to

[18] Unless otherwise specified, the literature reviewed in this section is based on a report by IMF/World Bank (2008).

[19] CFA stands for *Communauté Financière Africaine* (African Financial Community). The CFA Franc was created on December 26, 1945. The reason for its creation was the weakness of the French Franc immediately after World War II. When France ratified the Bretton Woods Agreement in December 1945, the French Franc was devalued in order to set a fixed exchange rate with the US dollar. New currencies were created in the French colonies to spare them the devaluation (history obtained from http://en.wikipedia.org/wiki/CFA_franc#History on April 16, 2009).

improved risk ratings and attractive yields. The acceleration of private capital flows presents a challenge to policy-makers, because significant inflows could lead to increased macroeconomic volatility and the buildup of balance sheet vulnerabilities, and over time, to real exchange rate appreciation and loss of external competitiveness.

The foregoing observations receive the support of Sayeh (2008). He remarks that sub-Saharan Africa remains less integrated in global financial markets and the direct impact of global financial turmoil is likely to be less severe than in the advanced and emerging economies. But Africa is not immune from global events. Sayeh opines that in the short term, many countries in Sub-Saharan Africa are more vulnerable to a number of shocks including the tightening of global credit conditions, which is likely to lower foreign direct investment flows and reduce or reverse portfolio inflows, as investors flee into more liquid or safer assets.

In a bid to explain the poor performance in attracting international private capital, Brink and Viviers (2003) identify the following as the main obstacles to foreign portfolio inflows in the Southern African region: the underdevelopment of domestic financial markets, macroeconomic instability, interest rate structures, exchange rate risk, country risk, exchange control, tax structures, inadequate availability of information, and an underdeveloped telecommunications infrastructure. These factors either reduce the expected rates of return or increase the perceived risk of investments. They recommend policy measures to deal with each of these barriers.

2.6 Concluding Remarks

To conclude, the literature survey indicates that some relationship exists between exchange rate volatility/risk and the flow of international portfolio capital. It is also clear from the review that many of the studies that have been conducted in the emerging market economies, often citing data constraints, have had a bias towards the relatively more prosperous Latin American and East Asian countries. However, as a segment of the emerging markets, Africa's financial markets appear not to have received adequate attention by researchers on these issues. The dearth of literature to establish the exact nature of the relationship between foreign exchange rates and international capital flows in Africa is so worrying that even the IMF/World Bank (2008) summary for sub-Saharan Africa paints a rather grim picture on the situation. This study was a pioneer attempt at filling this significant research gap.

CHAPTER THREE

EMPIRICAL MODELS AND ESTIMATION PROCEDURES

3.1 Introduction

This chapter presents the set of methodologies used to investigate the subjects addressed by this study. The chapter highlights, and explains the rationale of an appropriate paradigm on which the methods of inquiry rests; hypothesizes relationships among variables under investigation; and proposes appropriate scientific models to test for the existence of the hypothesized relationships.

3.2 Research Paradigm and Design

In an attempt to discover the truth, as a condition of knowledge, a researcher is faced with two philosophical viewpoints of science: a viewpoint that considers the world as socially constructed and subjective and in which science is driven by the viewpoint of the researcher, and a strong empiricists' viewpoint which considers that it is possible for a researcher to have no preconception when gathering data and deriving theories from the data. To a large extent, the latter viewpoint informs the epistemological foundations for the conduct of this study.

This viewpoint emphasizes the development of scientific knowledge through empirical observations and measurements, carefully examined and tested using a set of statistical techniques to provide evidence that conjectures made cannot be refuted. Philosophers, led by Comte (1865), have referred to this approach to science as *positivism*. As a research paradigm, positivism emphasizes careful empirical observations, discovery of cause-effect laws, and value-free research. Also known as *logical empiricism*, positivist social science is an organized method for combining deductive logic with precise empirical observations of study subjects in order to discover and confirm a set of probabilistic causal laws that can be used to predict general patterns of activity of the subjects (Neuman, 2006: 82). Logical empiricists seek rigorous exact measures and objective research and test hypotheses by carefully analyzing numbers from those measures, their ultimate goal being to discover and document "universal causal laws" (Turner, 1985).

Easterby-Smith et al. (2002) point out that the key idea of positivism is that the world exists externally, and that its properties should be measured through objective methods. A positivist assumption is different from the *interpretivist epistemology* which assumes that experience of the world is subjective and best understood in terms of individuals' subjective meanings rather than the researcher's objective definitions. By adopting a positivist approach, the researcher assumes that the contextual factors and knowledge strategies are objective phenomena with known properties or dimensions. By choosing the assumption of objectivity, it means the research concepts can be defined and measured with standard instruments. The researcher will be objective in collecting, analyzing and making interpretations about the data in a value-free manner.

Popper (1972) explains that the determination of what to observe/gather must be guided by the formulation of theory and hypotheses; thus scientific methods entail testing of theories in ways in which outcomes can potentially justify theory. This reasoning is characterized by the belief that theory precedes research and statistical justification of conclusions derived from empirically testable hypotheses form the core tenets of scientific knowledge development. In accord with this thinking and consistent with studies conducted earlier on this subject, the design of this study takes the form of a model-based statistical inference approach. Econometric tools are used to estimate and test variants of existing financial models in order to provide answers to research questions. Central to both the theoretical foundation and empirical implementation of these models is the concept of uncertainty. As Campbell et al. (1997: 3) put it, "the very existence of financial economics as a discipline is predicated on uncertainty." It is an accepted fact in financial literature that investors operate in environments in which outcomes depend on variables that are subject to continuous random fluctuations which cannot be predicted with precision. Foreign exchange rates and international portfolio flows are such variables.

Logical empiricism is not without shortcomings. Critics argue that its obsession with complex formulae and models abstract from real life and might, to that extent, be irrelevant. They point out that however sophisticated and dynamic, mathematical and statistical models that have become the tools of choice in scientific inquiry, are incapable of incorporating all relevant information and cannot explain, nay predict, naturally occurring phenomena accurately and without bias. These setbacks obviously constitute the generalized limitations of this and similar scientific studies.

3.3 Modeling Foreign Exchange Risk Pricing

3.3.1 The Multi-Beta Asset Pricing Model

The Arbitrage Pricing Theory, initially proposed by Ross (1976) and later extended to the international framework by Solnik (1983), provides the theoretical foundation on which the *unconditional* empirical model used in this study is built. According to the model, the return-generating process for an asset i, in terms of a given numeraire (or reference) currency, is a linear function of k international common factors:

$$\tilde{R}_i = \bar{R}_i + \beta_{i1}F_1 + \beta_{i2}F_2 + \cdots + \beta_{ik}F_k + \varepsilon_i \qquad i = 1, 2, \dots, n;\;\; k < n \tag{1}$$

where $\tilde{R}_i$ is the random rate of return on asset i; $\bar{R}_i$ is the expected return on asset i, viewed as the normal return on the asset given the set of information available to investors at the beginning of the investment period; β_{ij} (j = 1, 2, … k) are the sensitivities of returns of asset i to the international zero-mean common risk factors F_j: in the language of factor models, the β_{ij} are also known as factor loadings; ε_i are the zero-mean residual terms of the assets, assumed to be uncorrelated with F_j and with each other;[20] n is the total number of assets under consideration; k is the total number of factors. The model in equation (1) is typically expressed in matrix form:

$$\tilde{\boldsymbol{R}} = \bar{\boldsymbol{R}} + \boldsymbol{\beta F} + \boldsymbol{\varepsilon} \tag{1$'$}$$

where $\tilde{\boldsymbol{R}}$, $\bar{\boldsymbol{R}}$, and $\boldsymbol{\varepsilon}$ are n-vectors of random returns, excess returns and errors respectively, $\boldsymbol{\beta}$ is an $(n \times k)$ matrix of factor loadings, and $\boldsymbol{F}$ is a k-vector of factors. Now, suppose that the number of assets, n, is sufficiently large that investors can form well-diversified portfolios. Further, assume that the number of risk factors, k, is much smaller than n. Ross (1976) demonstrates that the pricing relation for an arbitrage portfolio[21] with weights, X_i, invested in asset i can be derived as follows. First, the requirement that an arbitrage portfolio requires no additional funds from the investor can be modeled thus:

[20] Mathematically, $E(\varepsilon_i, \varepsilon_j) = 0; and\ E(\varepsilon_i, F_j) = 0\ for\ all\ i\ and\ j.$

[21] An arbitrage portfolio is one that allows an investor to increase her expected return without increasing its risk. Thus, the net investment on such a portfolio must equal zero (i.e. the portfolio does not require additional funds from the investor), and the portfolio must have zero sensitivity to each of the k factors.

$$X_1 + X_2 + \ldots + X_n = \sum_{i=1}^{n} X_i = \boldsymbol{X'\iota} = 0 \qquad \text{(i)}$$

where $\boldsymbol{X}$ is a *n*-vector of asset weights, $\boldsymbol{\iota}$ is a *n*-vector of ones. Second, the arbitrage portfolio has no sensitivity to any factor. Since the sensitivity of a portfolio to a factor is the weighted average of the sensitivities of the securities in the portfolio to that factor, this requirement is expressed as

$$X_1\beta_{1j} + X_2\beta_{2j} + \ldots + X_n\beta_{nj} = \sum_{i=1}^{n} X_i\beta_{ij} = 0 \qquad \text{for each } j \qquad \text{(ii)}$$

Requirement (ii) is commonly expressed as $\boldsymbol{X'\beta} = \boldsymbol{0}$. Because the portfolio is constructed with a sufficiently large *n*, it is well-diversified so that it's residual (or idiosyncratic) risk is negligible:

$$X_1\varepsilon_1 + X_2\varepsilon_2 + \ldots + X_n\varepsilon_n = \sum_{i=1}^{n} X_i\varepsilon_i = \boldsymbol{X'\varepsilon} \sim 0 \qquad \text{(iii)}$$

Now, the return on a portfolio is the weighted average of the returns on securities comprising it. Thus, equation (1) can be transformed into a portfolio return-generating process:

$$\sum_{i=1}^{n} X_i\tilde{R}_i = \sum_{i=1}^{n} X_i\bar{R}_i + \sum_{i=1}^{n} X_i\beta_{i1}F_1 + \cdots + \sum_{i=1}^{n} X_i\beta_{ik}F_k + \sum_{i=1}^{n} X_i\varepsilon_i \qquad (2)$$

The compact form of equation (2) is $\boldsymbol{X'\tilde{R}} = \boldsymbol{X'\bar{R}} + (\boldsymbol{X'\beta})\boldsymbol{F} + \boldsymbol{X'\varepsilon}$. Given the results in (ii) and (iii), the relationship in equation (2) can be expressed as

$$\boldsymbol{X'\tilde{R}} = \boldsymbol{X'\bar{R}} + (\boldsymbol{X'\beta})\boldsymbol{F} + \boldsymbol{X'\varepsilon} \sim \boldsymbol{X'\bar{R}} \qquad (3)$$

By relationships (i), (ii) and (iii) this portfolio is almost riskless. Investors' arbitrage activities will drive prices of the securities in the portfolio to equilibrium, eventually eliminating arbitrage opportunities. At that point, the expected return on the arbitrage portfolio will equal zero:

$$\sum_{i=1}^{n} X_i\bar{R}_i = \boldsymbol{X'\bar{R}} = 0 \qquad \text{(iv)}$$

The foregoing conditions are really statements in linear algebra. Copeland et al. (2005: 178) observe that any vector that is orthogonal to the constant vector (equation (i)), and orthogonal to each of the coefficient vectors (equation (ii)), must also be orthogonal to the vector of expected returns (equation (iv)). Ross (1976) shows that the algebraic consequence of this statement is that the expected return vector must be a linear combination of the constant vector and the coefficient vectors. That is, there exist a set of $(k+1)$ coefficients $\lambda_0,\ \lambda_1, \dots, \lambda_k$ such that:

$$\bar{R}_i = \lambda_0 + \lambda_1\beta_{i1} + \lambda_2\beta_{i2} + \dots + \lambda_k\beta_{ik} \tag{4}$$

$$\bar{\boldsymbol{R}} = \lambda_0\boldsymbol{\iota} + \boldsymbol{\lambda}'\boldsymbol{\beta} \tag{4$'$}$$

where $\boldsymbol{\lambda}$ is a k-vector of coefficients. Equations (4) and (4′) characterize the Arbitrage Pricing Theory (APT). Since the model does not assume an ability to lend and borrow freely at the risk-free rate of return, the λ's can be interpreted as follows(Ross et al., 2002): λ_0 is the expected return on an asset with zero systematic risk, commonly proxied by R_f, the risk-free rate of return; $\lambda_j (= R_j - R_f)$ represent the risk premia related to each of the common factors j, such that $R_j$is the expected return on a security whose beta with respect to factor j is unity, and whose beta with respect to all other factors is naught. The form of the Arbitrage Pricing Theory (APT) in equation (4) suggests that the expected return is a summation of the risk-free rate and the compensation for each type of risk that the security bears. If a risk-free asset with return R_f exists, then $R_f = \lambda_0$ so that equation (4) can be expressed in the following form:

$$\bar{R}_i - R_f = \lambda_1\beta_{i1} + \lambda_2\beta_{i2} + \dots + \lambda_k\beta_{ik} \tag{5}$$

The left hand side of equation (5) is the expected excess return on security *i*. Now, defining $r_i = \bar{R}_i - R_f$, then the resulting multi-beta asset pricing model can be expressed as

$$r_i = \lambda_1\beta_{i1} + \lambda_2\beta_{i2} + \dots + \lambda_k\beta_{ik} \tag{6}$$

It is believed in the asset pricing literature that the model in equations (4) and (6) hold only as an approximation, particularly in a finite economy. In a large economy with infinitely many assets, Dybvig and Ross (1985) explain that the model holds as an exact equality under certain conditions. The magnitude of mispricing due to the approximation should be mitigated in the international context by the fact that there are more assets in the world economy than in any particular national economy (Cho et al., 1986). Solnik (1983) explains that the multi-beta asset pricing structures in equations (4) and (6) can be applied to the international setting in much the same way as it relates to nominal returns in the domestic setting. Solnik demonstrates that the structure is invariant to the currency chosen and applies to a set of international assets just as it applies to a set of domestic assets. He also posits that the model can fit in a structure consisting of a few international factors common to all assets, or where the sets of common factors strictly differs across national markets; it can also be applied to a situation with a combination of international factors common to all or specific types of assets plus national factors affecting only domestic markets. For these reasons, the model has become popular in the empirical testing of international asset pricing relationships.

Finally, equation (6) is restated in a manner that allows for time variation in excess returns. In this form, the model allows foreign exchange and other risk premia to vary with time in response to changes in fundamental variables in the economy. The restated model follows.

$$r_{it} = \lambda_1 \beta_{i1} + \lambda_2 \beta_{i2} + \ldots + \lambda_k \beta_{ik} \tag{7}$$

where r_{it} is the expected excess return on asset *i* at time *t*; $\lambda_j = R_{jt} - R_{ft},\ j = 1, 2, \ldots, k$.

Empirical Specification of the Unconditional Asset Pricing Model

Plugging equation (7) into the return-generating process in equation (1) gives the multi-beta asset pricing model in a form that is amenable to empirical estimation:

$$r_{it} = \lambda_1 \beta_{i1} + \lambda_2 \beta_{i2} + \ldots + \lambda_k \beta_{ik} + \beta_{i1} F_{1t} + \beta_{i2} F_{2t} + \cdots + \beta_{ik} F_{kt} + \varepsilon_i \tag{8}$$

where r_{it} is the one-period US dollar-equity market return for country i, in excess of the corresponding US dollar return on US Treasury bills at time *t*; $i = 1, 2, \ldots, n;\ \ k < n;\ t = 1, 2, \ldots, T$. An important observation about the foregoing multi-beta asset pricing equations is that they do not specify the particular risk sources that affect asset returns. Consequently, empirical studies have the latitude to select the relevant or appropriate factors. Two decision approaches have gained prominence among researchers using these models. The first approach uses statistical techniques, such as the asymptotic principal components method of factor analysis, to identify factors from a hypothesized set, performs statistical significance tests on the identified factors, then runs a regression on those which turn out to be statistically significant (Loudon, 1993). The traditional approach, however, has been to hypothesize the number and identity of factors and then test whether they are priced. Because Purchasing Power Parity may fail to hold exactly, international investors may be faced with a priced foreign exchange risk. Thus, exchange risk is typically included among the hypothesized factors. This study employs the latter approach.

Ferson and Harvey (1994) present a framework that allows for the study of the *unconditional* version of the multi-beta pricing model. In their framework, one starts by "cleaning" the factor returns, F, by removing the mean factor return, $\bar{F}$, from each return in the series. The resulting *demeaned* factor returns are plugged into equation (8) to allow the model to be tested as a restricted seemingly unrelated regression model (SURM). The regression is restricted by imposing the requirement that the intercept term be zero. The following model results:

$$\begin{aligned} r_{it} &= \lambda_1\beta_{i1} + \lambda_2\beta_{i2} + \dots + \lambda_k\beta_{ik} + \beta_{i1}f_{1t} + \beta_{i2}f_{2t} + \dots + \beta_{ik}f_{kt} + \varepsilon_{it} \\ &= \sum_{j=1}^{k}\beta_{ij}(\lambda_j + f_{jt}) + \varepsilon_{it} \qquad (i = 1, 2, \dots, n) \end{aligned} \tag{9}$$

where $f_{jt} = F_{jt} - \bar{F}_j$ are the demeaned returns on risk factors; $\bar{F}_j$ is the sample mean return on factor j; $E(f_{it}) = 0$. Because of the latter condition, Ferson and Harvey (1994) explain that the assumption that the means of the factors F_j are not related in any way to the expected risk premia λ_j must hold. Further, from the theoretical model in equation (8), the assumption that $E(\varepsilon_{it}) = 0$ must be true for equation (9) as well. These two assumptions, together with the assumption that $(\varepsilon_{it}, F_{jt}) = 0$, spell out the orthogonality conditions for the model. Accordingly, economic variables, such as exchange rates, can be used in the model as if they were extracted factors.

Following Carrieri and Majerbi (2006), equation (9) is first tested as a two-factor return-generating process where the ith country/portfolio excess return (r_i) is a linear function of the demeaned excess return on the world market portfolio (r_w) and the percentage change in the real exchange rate component (r_s) orthogonal to the world market portfolio returns:[22]

$$r_{it} = \beta_{iw} r_{wt} + \beta_{is} r_{st} + \lambda_w \beta_{iw} + \lambda_s \beta_{is} + \varepsilon_{it} \tag{10}$$

By construction, the currency risk factor, (r_s), has zero mean. For each of the two risk factors, the parameters to be estimated are the unconditional betas, (β_{iw}, β_{is}) and the risk premia (λ_w, λ_s). Foreign exchange rate exposure is said to be priced if the coefficient λ_s is non-zero.[23]

The model in equation (10) implicitly assumes that world equity markets are fully integrated so that there are no barriers to cross-border investments, no taxes or transaction costs, and no delays or costs associated with information flow. African markets may still be reasonably segmented as many of the region's countries are still in the middle of the liberalization process and some do not allow unfettered access to their financial markets by foreign investors. The stringent market integration assumptions are therefore commonly violated and the model may fail to account for idiosyncratic risks associated with each market. Thus, there is need for a model that can reasonably capture segmentation as well as currency risk in the continent's markets.

Various authors have argued that the flow of international money is, to a large extent, restricted by various pecuniary and non-pecuniary barriers. Black (1974) proposes a market segmentation model in which barriers to international investment correspond to a tax on the net value of an investor's holdings of foreign risky assets. The tax represents various kinds of barriers including, the possibility of expropriation of foreign holdings, direct controls on the import or export of capital, reserve requirements on bank deposits and other assets held by foreigners, restrictions on

[22] Orthogonalization is achieved by regressing real exchange rate changes on excess world market returns, then using the resulting residuals in the model. The procedure is described in equations (13) and (14).

[23] In the standard multi-beta asset pricing model, of which the APT is the most prominent, factor returns, F, represent news or innovations in the security market and are therefore observed *ex-post*. In this regard, their role in asset pricing is largely that of price revision at the period's end to reflect the actual state of affairs in the market at that time. Investors, however, make their purchase and sale decisions *ex-ante* and, naturally, would make their forecasts at that time. Forecasts of the expected returns on systematic factors and the extent to which they differ from the risk-free rate (this difference is known as a factor's risk premium or price) therefore form a fundamental source of information for computing a security's expected return. Asset pricing models are typically concerned with the estimation of the risk premia (λ) attached to each factor of interest.

the fraction of business that can be foreign owned, and barriers created by the unfamiliarity that residents of one country have with other countries. Stulz (1981) presents a partial market segmentation model in which barriers to international investment are represented as taxes on the absolute value of an investor's holdings of risky foreign assets.

Errunza and Losq (1985) develop a model of international capital asset pricing whose primary distinguishing feature is the unequal access assumption. Their empirical test on the model shows that the required return on a security, the holding of which an international investor is ineligible, is higher than what the standard CAPM would suggest whereas that on an "eligible" security is consistent with the CAPM. The results lend support to the mild segmentation hypothesis. Eun and Janakiramanan (1986) argue that the existence of barriers to international investment constrains investors' portfolio choice and distorts market equilibrium. Assuming a two-country world – one domestic and one foreign – in which foreign investors are constrained to own only a fraction of the outstanding shares of domestic firms while the domestic investors do not face such restrictions, they demonstrate the existence of a two-tier pricing relationship in which foreign investors pay a premium over and above the equilibrium price for securities with no such constraint. In all these cases, the authors demonstrate that higher equilibrium asset returns, arising from investment barriers, cause partial segmentation in international capital markets.

These models or their variants have been investigated empirically by a number of researchers. Bailey and Jagtiani, (1994) provide evidence that substantial differences in expected returns exist between local and foreign investors at the Thai stock market. The differences are correlated with proxies for the severity of foreign ownership limits, liquidity, and information availability; suggesting that the Thai stock market is partially segmented. Using a conditional asset pricing methodology with GARCH parameterization, Gerard et al. (2003) find little support for the hypothesis that exposure to residual country risk is rewarded across major East Asian equity markets. Since they find that exposure to world market risk carries a significant premium, they reject the market segmentation hypothesis in East Asia over the period 1985–1998. Antell and Vaihekoski (2007) find the local market risk to be relevant for the pricing of Finnish stocks and suggest that one should consider partially segmented asset pricing models for smaller stock markets. Similar results have been documented by Saleem and Vaihekoski (2008) who show that the local market risk factor is priced in the Russian stock market and conclude that the market is at least partly segmented. These studies show that equity markets are mostly partially segmented.

In this study, I take the view that fully integrated or fully segmented markets do not exist anywhere in the world today. Accordingly, on the basis of a partial segmentation framework, I utilize two-factor and three-factor models akin to those in Choi and Rajan (1997), which assume that idiosyncratic market risk is priced in each country. The two models follow.

$$r_{it} = \lambda_w \beta_{iw} + \lambda_m \beta_{im} + \beta_{iw} r_{wt} + \beta_{im} r_{mt} + \varepsilon_{it} \tag{11}$$

$$r_{it} = \lambda_w \beta_{iw} + \lambda_m \beta_{im} + \lambda_s \beta_{is} + \beta_{iw} r_{wt} + \beta_{im} r_{mt} + \beta_{is} r_{st} + \varepsilon_{it} \tag{12}$$

Possible contemporaneous correlations among factors are eliminated by separately running the following regressions, then using the resulting residuals as risk factors (also see Jorion, 1991):

$$r_{Mt} = \gamma_0 + \gamma_1 r_{wt} + \epsilon_{Mt} \tag{13}$$

$$s_t = \gamma_0 + \gamma_1 r_{wt} + \gamma_2 r_{Mt} + \epsilon_{st} \tag{14}$$

where r_{Mt} is the excess return on the local equity-market index; s_t is the change in real exchange rates; ϵ_{Mt} the residual from the first regression, is known as the pure local market factor (hereafter labeled, r_{mt}); ϵ_{st} is the pure currency risk factor (hereafter labeled, r_{st}). By construction, the residual factors have zero mean. Thus, all the factors in the two models are orthogonal to each other and all returns are stated in excess of the return on the risk-free asset. In equations (11) and (12), β_{im} is the sensitivity of the idiosyncratic local market risk factor; λ_m is the corresponding risk premium parameter. The rest of the terms are as defined earlier. The hypotheses are $\lambda_w \neq 0$ and $\lambda_m = \lambda_s = 0$ i.e., equity markets are fully integrated and foreign exchange rate risk is not priced. The result $\lambda_w = \lambda_m = 0$ points to model misspecification since markets cannot be non-integrated and non-segmented at the same time.

In each of equations (10), (11) and (12), coefficients are jointly estimated using the iterated Generalized Method of Moments (GMM), presented in section 3.3.3. The iterated GMM is an improvement on the GMM of Hansen (1982) and produces estimates with better small sample properties: Ferson and Foerster (1994) demonstrate that the iterated GMM produces approximately unbiased coefficient estimates with goodness-of-fit statistics that conform well to the asymptotic distribution even with time series observations as few as sixty. Following Ferson and Harvey (1994), this study uses a constant and the contemporaneous values of factors F_{jt} as instruments in the GMM regression. The study also assumes that the data vector $\{r_{ij}, f_{jt}: i = 1, 2, \ldots, n; j = 1, 2, \ldots, k\}$ is generated by a strictly stationary and ergodic stochastic process. Orthogonality conditions for the GMM are specified by $E(\varepsilon_{it}) = E\left(\varepsilon_{it}, F_{jt}\right) = 0$.

3.3.2 The Stochastic Discount Factor (SDF) Model

The unconditional models used in the preceding section are not without drawbacks. The models' underlying assumption that foreign exchange risk is unconditionally priced contrasts with bourgeoning empirical evidence that currency risk commands nonzero time-varying and conditional risk premia (Dumas and Solnik, 1995; De Santis and Gerard, 1998; Choi et al. 1998; Doukas et al., 1999; Vassalou, 2000). Indeed, Dumas and Solnik (1995) aver that "it is natural to test any asset pricing model in its conditional form," and point out that, since conditioning information is obviously available to investors, it cannot be ignored by researchers and must be incorporated in empirical tests, in the form of instrumental variables. Thus, the compelling evidence against unconditional asset pricing models, together with the need for triangulation of methods, necessitate that an alternative empirical strategy be employed on the same data in order to establish the exact nature of foreign exchange risk pricing in Africa's equity markets.

The stochastic discount factor (SDF) model is rapidly gaining ground in the literature as a convenient and general asset pricing approach. By specifying the discount factor suitably, this approach encompasses most of the theories currently in use, including the Capital Asset Pricing Model (CAPM), the Consumption-based CAPM and the Intertemporal CAPM (Smith and Wickens, 2002). The SDF can be used as a single-factor model or as a multiple-factor model, with either latent or observed factors.

To motivate the model, suppose that a consumer-investor has wealth, W, to be split between today's consumption, C_t, and investment, $W - C_t$. The amount invested today provides the consumer-investor with consumption at time $t + 1$ equal to C_{t+1}. The consumer-investor would like his/her investment to consist of Y shares of a common stock i, each currently trading at the price P_{it} in the market place. It follows that

$$W = C_t + YP_{it} \qquad \Rightarrow \qquad C_t = W - YP_{it} \tag{15}$$

Let $X_{i,t+1}$ denote the payoff of a share of stock at time $t+1$. Now, $X_{i,t+1}$ is composed of the price, $P_{i,t+1}$ and dividend, $D_{i,t+1}$, anticipated from the stock at time $t+1$. Both the future price and dividend on the stock are state-dependent so that $X_{i,t+1}$ is a random variable. Since the investor holds an amount Y of the stock, the investor's payoff (or consumption) at time $t+1$ is given as $C_{t+1} = YX_{i,t+1}$. With the assumption that the investor prefers a consumption stream that is steady over time and across states of nature, Cochrane (2000: 14) models the investor's utility function over current and future values of consumption thus:

$$U(C_t, C_{t+1}) = u(C_t) + \beta E_t[u(C_{t+1})] \tag{16}$$

where $E_t(\cdot)$ is an expectation operator conditional on market information set, Ω_t, available at time t; the parameter β is known as the investor's *subjective discount factor*, a measure of risk aversion and intertemporal substitution. Substituting $C_t = W - YP_{it}$ and $C_{t+1} = YX_{i,t+1}$ into equation (16), the investor's utility function can be expressed as

$$U(Y) = u(W - YP_{it}) + \beta E_t\left[u\left(YX_{i,t+1}\right)\right] \tag{17}$$

The first order condition of the consumer-investor is obtained from the model in equation (17) by taking the partial derivative of $U(Y)$ with respect to Y, and equating it to zero:

$$\frac{\partial U}{\partial Y} = u'[W - YP_{it}](-P_{it}) + \beta E_t\left[u'\left(YX_{i,t+1}\right)\left(X_{i,t+1}\right)\right] = 0$$

$$= -P_{it}u'(C_t) + \beta E_t\left[X_{i,t+1}u'(C_{t+1})\right] = 0 \tag{18}$$

where $u'(\cdot)$ denotes partial derivative. Solving for P_{it} and rearranging the result yields

$$P_{it} = E_t\left[\beta \cdot \frac{u'(C_{t+1})}{u'(C_t)} \cdot X_{i,t+1}\right] \tag{19}$$

This result is the first order condition for an investor's optimal consumption and portfolio choice. At this point, it is not uncommon for some researchers to specify a functional form of utility (the time-separable power utility function, $u(C) = C^{1-\gamma}/(1-\gamma)$; $u'(C) = C^{-\gamma}$ is commonly used).[24] When this is plugged into equation (19), the resulting equation is the well-known consumption-based asset pricing model: $P_{it} = E_t\left[\beta(C_{t+1}/C_t)^{-\gamma}X_{i,t+1}\right]$, which does not however, perform well in empirical tests. Hansen and Singleton (1983) find that the model cannot simultaneously explain the time variation of interest rates and cross-sectional pattern of average returns of bonds and stocks in the USA. Their study rejects the model. Mehra and Prescott (1985) find that the Consumption-CAPM is incapable of explaining the equity premium puzzle. They attribute this result partly to errors associated with measuring the inflation rate and partly to variations in income taxes. Inflation measurement errors, they argue, bias the estimates of consumption growth rates and the real risk-free rate and hence bias tests results. On income taxes, they argue that the theory implicitly considers after-tax returns but fails to capture the fact that income taxes vary over income classes and change through time.

Wheatley (1988) tests the joint hypothesis that equity markets are integrated and that the model holds internationally using a discrete time version of the consumption-based asset pricing model. With monthly equity data for the period January 1960 through December 1985, he rejects the hypothesis. Using artificial stock data, Campbell and Cochrane (1999, 2000) find that the static CAPM, and its various multi-beta extensions, is a much better approximate asset pricing model than the canonical power utility consumption-based model. Campbell and Cochrane (2000) conjecture that these results would hold even if the true investors' utility function were known and demonstrate that the CAPM performs better because stock returns are more closely unconditionally correlated with the marginal rate of substitution than is consumption growth.

[24] This power utility function is known as the Constant Relative Risk Aversion function. In the function, γ, a constant, is known as the coefficient of relative risk aversion. Notice that the marginal utility of consumption, $u'(C) = C^{-\gamma}$, has a constant elasticity equal to $-\gamma$. Thus, γ is a measure of how rapidly marginal utility declines when consumption is increased (Bailey, 2005: 256). The larger is γ, the more sensitive is the investor's utility to fluctuations in consumption. The power utility function holds only when $\gamma \neq 1$; when $\gamma = 1$, the function follows the logarithmic form: $u(C) = \ln C$, the limit of the function as $\gamma \to 1$.

In light of the highlighted empirical weaknesses of the consumption-based CAPM, this study avoids specifying a utility function and takes the alternative route, which is even more commonly used in the asset pricing literature. In the spirit of Lucas (1978), I define the *stochastic discount factor* as follows:

$$M_{t+1} = \beta \cdot \frac{u'(C_{t+1})}{u'(C_t)} \tag{20}$$

The stochastic discount factor is a random variable whose realizations are always positive (Campbell, 2000). It generalizes the familiar notion of a discount factor to a world of uncertainty: if there is no uncertainty, or if investors are risk-neutral, M_{t+1} is just a constant that converts tomorrow's expected payoffs to today's values. In this respect, the stochastic discount factor is also called the *intertemporal marginal rate of substitution*. The intertemporal marginal rate of substitution is the rate at which the investor is willing to substitute consumption at time *t* +1 for consumption at time *t*. In the asset pricing literature, it has also been described as the *pricing kernel* (Dumas and Solnik, 1995; Tai, 1999; and Cochrane, 2000). Using the definition in equation (20), equation (19) can be expressed in a more parsimonious form:

$$P_{it} = E_t[M_{t+1}X_{i,t+1}] \tag{21}$$

Dividing through equation (21) by P_{it} yields $E_t\left[M_{t+1} \cdot \frac{X_{i,t+1}}{P_{it}}\right] = 1$, where $\left(\frac{X_{i,t+1}}{P_{it}} = R_{i,t+1}\right)$ is the *gross* random return on the investor-consumer's asset at time $t+1$. Thus, equation (21) becomes

$$E_t\left[M_{t+1}R_{i,t+1}\right] = 1 \tag{22}$$

Equation (22) simply expresses the price of an asset as the expected discount value of its payoffs, discounting being done by the investor-consumer's intertemporal marginal rate of substitution. A relationship similar to equation (22) can be developed for the risk-free asset with return, R_{ft}:

$$E_t\left[M_{t+1}R_{ft}\right] = 1 \tag{23}$$

Subtracting equation (23) from equation (22) gives

$$E_t\left(M_{t+1}r_{i,t+1}\right) = 0 \tag{24}$$

where $r_{i,t+1} = R_{i,t+1} - R_{ft}$ is the expected excess return on stock *i* at time *t*. Now, the random rate of return, $\tilde{r}$, on an asset is defined as $\tilde{r}_{t+1} = \frac{X_{t+1} - P_t}{P_t} = \frac{X_{t+1}}{P_t} - 1 = R_{t+1} - 1$. Utilizing this result and denoting the typical asset i $(i = 1, 2, \dots, N)$, equation (23) can also be expressed as

$$E_t\left[M_{t+1}(1 + r_{ft})\right] = 1 \quad \Longrightarrow \quad (1 + r_{ft})E_t(M_{t+1}) = 1 \tag{25}$$

From equation (25), it is clear that $E_t(M_{t+1}) = 1/\left(1 + r_{ft}\right) = 1/R_{ft}$.[25] Ferson (1995) argues that without more structure, the stochastic discount factor model has little empirical content because it is easy to find some random variable M_{t+1} for which the foregoing equations hold. It is the specific form of M_{t+1} implied by a model (referred to as the *benchmark pricing variable*) that gives the equations further empirical content. Consequently, empirical tests of asset pricing models often work directly with equations (22) through (25) plus the relevant definition of M_{t+1}.

A Digression: Expected Excess Return

Define the conditional covariance at time *t* as follows:[26]

$$Cov_t\left(M_{t+1}, r_{i,t+1}\right) = E_t\left(M_{t+1}r_{i,t+1}\right) - E_t(M_{t+1})E_t(r_{i,t+1}) \tag{26}$$

Thus, equation (24) can be expressed as

$$Cov_t\left(r_{i,t+1}, M_{t+1}\right) + E_t\left(r_{i,t+1}\right)E_t(M_{t+1}) = 0 \tag{27}$$

[25] The implication of this result is that the discount factor is a random variable of the form: $M_{t+1} = 1/1 + r_{ft} + \xi_{t+1}$, where the random variable ξ_{t+1} has conditional mean $E_t(\xi_{t+1}) = 0$ (Smith and Wickens, 2002).

[26] For any two random variables *X* and *Y*, the Covariance can be mathematically expressed as:

$$\begin{aligned} Cov(X,Y) &= E\left[\left(X - E(X)\right)\left(Y - E(Y)\right)\right] = E[(XY - XE(Y) - YE(X) + E(X)E(Y)] \\ &= E(XY) - 2E(X)E(Y) + E(X)E(Y) \\ &= E(XY) - E(X)E(Y) \end{aligned}$$

Subtracting $Cov_t(r_{i,t+1}, M_{t+1})$ from both sides and dividing the result by $E_t(M_{t+1})$ gives

$$E_t(r_{i,t+1}) = \frac{-Cov_t(r_{i,t+1}, M_t)}{E_t(M_{t+1})} = \frac{1}{E_t(M_{t+1})}\left[-Cov_t(r_{i,t+1}, M_{t+1})\right] \tag{28}$$

Since $E_t(M_{t+1}) = 1/(1 + r_{ft})$, equation (28) can also be expressed as

$$E_t(r_{i,t+1}) = (1 + r_{ft})\left[-Cov_t(r_{i,t+1}, M_{t+1})\right] \tag{29}$$

Equation (29) is the expected excess return on asset *i* at time *t* + 1, conditional on the information set available to investors at time *t*. The equation specifies the no-arbitrage condition that all correctly priced assets must satisfy (Smith and Wickens, 2002). The term on the right hand side of equation (29) represents the risk premium, that is, the extra return over the risk-free rate that is required to compensate investors for holding the risky asset.

If M_{t+1} represents the aggregate intertemporal marginal rate of substitution as was earlier explained, then equation (29) says that a security will earn a positive risk premium if its return is negatively correlated with the marginal rate of substitution. Negative correlation means that the asset is likely to return more than expected when the marginal utility in the future period, *t* + 1, relative to the current period, *t*, is lower than expected (Ferson, 1995: 148). The more negative is the covariance with M_{t+1}, the less desirable is the distribution of the random return, and the larger must be the expected compensation for holding the asset. Intuitively, an asset whose payoff covaries negatively with consumption helps to smooth consumption and is therefore valuable to an investor-consumer (Cochrane, 2000: 23).[27]

[27] This argument can be seen more clearly by restating the discount pricing model in the cash flow form (equation 21): $P_t = E[M_{t+1}X_{t+1}]$. Using the covariance decomposition presented in footnote 26, this relationship can be expressed in the form: $P_t = E(M_{t+1})E(X_{t+1}) + Cov_t(M_{t+1}, X_{t+1}) = \frac{X_{t+1}}{R_{ft}} + Cov_t(M_{t+1}, X_{t+1})$. Noting from equation (20) that $M_{t+1} = \beta u'(C_{t+1})/u'(C_t)$, the pricing relationship can be represented as: $P_t = \frac{X_{t+1}}{R_{ft}} + Cov_t([\beta u'(C_{t+1})/u'(C_t)], X_{t+1})$. Because of the concavity of the utility function, $u'(C)$ declines as C (consumption) rises. Thus, an asset's expected return increases (declines) if its payoff covaries negatively (positively) with consumption. As Cochrane (2000: 23) explains, investors do not like uncertainty about consumption: if an investor buys an asset whose payoff increases when the investor is already feeling rich and declines when the investor is feeling poor, the asset will make the investor's consumption stream more volatile. The investor will require a lower price to be induced to buy the asset. An asset whose payoff varies negatively with consumption therefore helps to smooth intertemporal consumption and is more valuable.

3.3.2.1 The Stochastic Discount Factor Model as a Linear Factor Model

Using the foregoing covariance definition, equation (22) can be expressed as follows:[28]

$$E_t(R_{i,t+1}) = \frac{1}{E_t(M_{t+1})} - \frac{Cov_t(R_{i,t+1}, M_{t+1})}{E_t(M_{t+1})} \tag{30}$$

Multiplying and dividing through equation (30) by $Var(M_t)$ yields:

$$E_t(R_{i,t+1}) = \frac{1}{E_t(M_{t+1})} + \left[\frac{Cov_t(R_{i,t+1}, M_{t+1})}{Var_t(M_{t+1})}\right]\left[-\frac{Var_t(M_{t+1})}{E_t(M_{t+1})}\right] = \lambda_{t,0} + \beta_{t,iM}\lambda_{t,M} \tag{31}$$

where $\lambda_{t,0} = 1/E_t(M_{t+1})$ is the price of the zero-beta asset (or the risk-free asset if one exists): it is assumed that $\lambda_{t,0}$ is uncorrelated with M_{t+1};[29] $\lambda_{t,M}$ is the risk per unit of the expected return on some economy-wide benchmark pricing variable, M_{t+1}; $\beta_{t,iM}$ is the regression coefficient of $R_{i,t+1}$ on M_{t+1} and can be loosely interpreted as a general measure of *systematic risk* of asset *i*. Cochrane (2000: 26) contends that $\beta_{t,iM}$ can be interpreted as the *quantity of risk* in each asset while $\lambda_{t,M}$ is the *price of risk*.[30] Specifically, Ferson (2003: 749) argues that the covariance with M_{t+1} (a component of the beta term in the equation) can be viewed as a very general measure of systematic risk because it measures the component of the return that contributes to fluctuations in the marginal utility of wealth: if we regressed the asset return on the SDF, the residual in the regression would capture the "unsystematic" risk and would not be "priced" (i.e., command a risk premium). If the conditional covariance is zero for a particular asset, the expected excess return on that asset should be zero. As expected, this must be the case for a zero-beta portfolio or the risk-free asset if one exists. Equation (31) gives the expected gross return on asset *i* at time *t*, conditional on information available to the investor at time *t*.

[28] Also note that $Cov_t(R_{i,t+1}, -M_{t+1})/E_t(M_{t+1}) = -Cov_t(R_{i,t+1}, -M_{t+1})/E_t(M_{t+1})$

[29] Being the risk-free rate of return, $\lambda_{t,0}$ is known with certainty at time *t*. However, M_{t+1} represents some widely-accepted pricing variable whose value (to be realized at time *t+1*) is state-dependent and therefore unknown at time *t*. In Investments literature, the returns on a risk-free asset are uncorrelated with returns on a risky asset/portfolio. Consequently, the two values ($\lambda_{t,0}$ and M_{t+1}) must be uncorrelated with each other.

[30] The coefficient λ_M is the same for all assets *i*, while $\beta_{i,M}$ varies from one asset to another. The model says that expected returns on asset *i* ($i = 1,2, \dots, N$) should be proportional to their betas in a regression of returns ($R_{i,t+1}$) on the discount factor (M_{t+1}). Equation (31) can be easily extended to the multi-factor case: $E_t(R_{i,t+1}) = \lambda_0 + \beta_{i,1}\lambda_1 + \beta_{i,2}\lambda_2 \dots + +\beta_{i,k}\lambda_k$.

Equation (31) shows that the SDF model can be expressed as a factor pricing model. Cochrane (2000: 104) demonstrates that such factor pricing models as the one in equation (31) are equivalent to linear stochastic discount factor models. Therefore, the pricing kernel can be expressed as a linear combination of factors in the following form:

$$M_{t+1} = a_t + b_{1t}F_{1,t+1} + b_{2t}F_{2,t+1} + \ldots + b_{kt}F_{k,t+1} = a_t + \boldsymbol{b}_t'\boldsymbol{F}_{t+1} \quad (32)$$

where $\boldsymbol{b}_t$ is a $(k \times 1)$ vector of factor loadings and, $\boldsymbol{F}_{t+1}$ is a $(k \times 1)$ vector of factors. The next discussion strives to obtain a connection to the beta representation based on covariances.

Incorporating Conditioning Information

As already explained, the stochastic discount factor model assumes that returns are conditional on the full set of information, Ω_t, available to the investor at time *t*.[31] Since the market-wide information set, Ω_t, at the disposal of investors at time *t* is not observable, moment conditions in this specification are not directly testable. In practice, therefore, econometric tests of the stochastic discount factor model are performed by proxying Ω_t by a set of a few carefully selected and judiciously transformed instrumental variables, $Z_t(\in \Omega_t)$, assumed to contain time *t* information. The set of instruments comprises of observable economic variables or portfolios of assets whose returns mimic the returns on the economic variables.

The conditioning information set, Z_t, can be incorporated in econometric modeling in a number of ways. One, they can be used to scale the asset returns: the scaled returns can be interpreted as the returns on managed portfolios in which the manager invests more or less according to the signal Z_t (Cochrane, 1996). Two, the instrument set can be used to scale the factors in a situation in which the factors are expected only to conditionally price assets. In the latter case, the model's parameters are first modeled as linear functions of Z_t: $a_t = \boldsymbol{a}'\boldsymbol{Z}_t$ and $b_t = \boldsymbol{b}'\boldsymbol{Z}_t$ (Cochrane, 1996; Lettau and Ludvigson, 2002; Schrimpf et al., 2007; Drobetz et al., 2008), where $\boldsymbol{Z}_t$ is the vector of s instruments (the choice of which is discussed in section 3.6). Thus, the scaled factor representation of the model in equation (32) factor is given by

[31] Thus, the conditional asset pricing model in equation (23) can also be presented in the form $E(M_{t+1}, R_{t+1}|\Omega_t) = 1$.

$$M_{t+1} = \boldsymbol{a}'\boldsymbol{Z}_t + (\boldsymbol{b}'\boldsymbol{Z}_t)\boldsymbol{F}_{t+1} \tag{33}$$

Now, let $\boldsymbol{F}_{t+1} \equiv \boldsymbol{f}_{t+1}$ where $\boldsymbol{f}$ represents scaled and unscaled factors (see, Iqbal et al., 2010) and let $\boldsymbol{a} = (a_0, a_1)'$; $\boldsymbol{b}' = [\boldsymbol{b}_0', \boldsymbol{b}_1']$; further, define $\boldsymbol{Z}_t = (1, z_t)'$. In this setup, z_t is, effectively, a scalar quantity, implying that in actual application, one conditioning variable is used at a time (also see Hodrick and Zhang, 2007). It follows that equation (33) can be restated thus:

$$M_{t+1} = (a_0 \quad a_1)\begin{pmatrix}1\\ z_t\end{pmatrix} + \left([\boldsymbol{b}_0' \quad \boldsymbol{b}_1']\begin{bmatrix}1\\ z_t\end{bmatrix}\right)\boldsymbol{f}_{t+1} = a_0 + a_1 z_t + (\boldsymbol{b}_0' + \boldsymbol{b}_1' z_t) \otimes \boldsymbol{f}_{t+1} \tag{34}$$

where $\otimes$ is the Kronecker product (i.e., multiply each term in the bracket by every factor). If there is only one factor, equation (34) reduces to:

$$M_{t+1} = a_0 + a_1 z_t + (b_0 + b_1 z_t) f_{t+1} = a_0 + a_1 z_t + b_0 f_{t+1} + b_1 (f_{t+1} \times z_t) \tag{35}$$

The results in equation (35) demonstrate that scaling the prices of factors is equivalent to scaling of factors (Cochrane, 2000: 135). Notice that, instead of a single factor model with time-varying factor weights, scaling results in a four-factor model with constant (or fixed) weights. However, parameters are allowed to vary with time through their interaction with the instrumental variables, z_t. By extension, a scaled two-factor model representation gives the following six factor model with constant weights:[32]

$$M_{t+1} = a_0 + a_1 z_t + b_{01} f_{1,t+1} + b_{02} f_{2,t+1} + b_{11}\left(f_{1,t+1}\, z_t\right) + b_{12}\left(f_{2,t+1}\, z_t\right) \tag{36}$$

A Multivariate Beta Representation

Plugging equation (32) into equation (30) gives

$$E_t\left(R_{i,t+1}\right) = \frac{1}{E_t(M_{t+1})} - \frac{Cov_t\left(M_{t+1}, R_{i,t+1}\right)}{E_t(M_{t+1})} = \frac{1}{E\left(a_t + f_{t+1}^{'}\boldsymbol{b}_t\right)} - \frac{Cov_t\left([a_t + \boldsymbol{b}_t'\boldsymbol{f}_{t+1}], R_{i,t+1}\right)}{E\left(a_t + f_{t+1}^{'}\boldsymbol{b}_t\right)} \tag{37}$$

[32] In practice, Cochrane (1996) suggests that the terms without factors can be dropped from the analysis so that the one-factor model generates the following three-factor constant-weights model: $M_{t+1} = a_0 + b_0 f_{t+1} + b_1(f_{t+1} \times z_t)$. In the same way, the model in equation (36) can be reduced to a five-factor representation.

Following Cochrane (2000), I assume that all factors are demeaned so that $E_t(M_{t+1}) = E\left(a_t + \boldsymbol{f}'_{t+1}\boldsymbol{b}_t\right) = a_t$. Thus, equation (37) can be written as:

$$E_t\left(R_{i,t+1}\right) = \frac{1}{a_t} - \frac{Cov_t\left(\boldsymbol{f}'_{t+1}, R_{i,t+1}\right)\boldsymbol{b}_t}{a_t} \tag{38}$$

Now, let $\boldsymbol{\beta}_i$ be the vector of regression coefficients of $R_{i,t+1}$ on $\boldsymbol{f}_{t+1}$ and a constant. Then, from factor models, $\boldsymbol{\beta}'_i = Cov_t\left(\boldsymbol{f}'_{t+1}, R_{i,t+1}\right)Var_t(\boldsymbol{f}'_{t+1})^{-1}$ are the factor sensitivities (or betas). Thus, the covariance vectors in equation (38) can be replaced with the vectors of appropriate betas:

$$E_t\left(R_{i,t+1}\right) = \frac{1}{a_t} - \frac{Cov_t(\boldsymbol{f}'_{t+1}, R_{i,t+1})Var_t(\boldsymbol{f}'_{t+1})^{-1}Var_t(\boldsymbol{f}'_{t+1})\boldsymbol{b}_t}{a_t} = \frac{1}{a_t} - \frac{\boldsymbol{\beta}_t Var_t(\boldsymbol{f}'_{t+1})\boldsymbol{b}_t}{a_t} \tag{39}$$

By equation (31), $\frac{1}{E_t(M_{t+1})} = \frac{1}{a_t} = \lambda_0$. Thus, equation (39) can be expressed as:

$$E_t\left(R_{i,t+1}\right) = \lambda_0 + \left[-\lambda_0\boldsymbol{\beta}'_i Var_t(\boldsymbol{f}'_{t+1})\boldsymbol{b}_t\right] = \lambda_0 + \boldsymbol{\beta}'_i\boldsymbol{\lambda} \tag{40}$$

where $\boldsymbol{\lambda} = -\lambda_0 Var_t(\boldsymbol{f}'_{t+1})\boldsymbol{b}_t$ are the factor risk premia. Cochrane (2000: 107) notes that the factors as defined here need not be returns (though they may be); they need not be orthogonal, and they need not be serially uncorrelated or conditionally or unconditionally mean zero. Such properties may occur as natural special cases, or as part of the economic derivation of specific factor models, but they are not required for the existence of a factor pricing representation. Using the law of iterated expectations, equation (40) suggests that the unconditional expected return on asset *i* can be written as follows:

$$E\left(R_{i,t+1}\right) = \lambda_0 + \lambda_1\frac{Cov\left(f_{1,t+1}, R_{i,t+1}\right)}{Var\left(f_{1,t+1}\right)} + \lambda_2\frac{Cov\left(f_{2,t+1}, R_{i,t+1}\right)}{Var\left(f_{2,t+1}\right)} + \ldots + \lambda_k\frac{Cov\left(f_{k,t+1}, R_{i,t+1}\right)}{Var\left(f_{k,t+1}\right)} \tag{41}$$

The GMM approach is used to estimate the pricing coefficients in equation (41).

3.3.2.2 Empirical Strategy

Estimation of Factor Sensitivities

The procedure in equation (35) can be easily extended to a multifactor, multi-instrument framework to give the following compact form of the scaled factor model representation:

$$M_{t+1} = a_0 + (\boldsymbol{b}'\boldsymbol{z}_t) \otimes \boldsymbol{f}_{t+1} \tag{42}$$

where, as before, $\boldsymbol{b}$ is a vector of coefficients and $\otimes$ is the Kronecker product. Since the weights are fixed, the scaled factor model can be tested unconditionally by applying the Law of Iterated Expectations.[33] Thus, combining equations (23) and (42), the following unconditional moment restrictions, to be evaluated using the GMM method (discussed in section 3.3.3), are obtained:

$$E\{[a_0 + (\boldsymbol{b}'\boldsymbol{z}_t) \otimes \boldsymbol{f}_{t+1}]\boldsymbol{R}_{i,t+1}\} = 1 \qquad i = 1,2, \dots, n \tag{43}$$

The system to be estimated, obtained by rearranging equation (43) and stating it in the form of sample moment restrictions, is as follows:

$$g_T(\varphi) = \frac{1}{T}\sum_{t=1}^{T}\{[a_0 + (\boldsymbol{b}'z_t) \otimes \boldsymbol{f}_{t+1}]\boldsymbol{R}_{i,t+1} - 1\}) = \boldsymbol{0} \qquad i = 1,2, \dots, n \tag{44}$$

where φ is the set of all parameters (a, $\boldsymbol{b}$) to be estimated and $\boldsymbol{0}$ is a n-vector of zeros. Equation (44) shows that by scaling factors, one can investigate the unconditional implications of the conditional stochastic discount factor model. Non-zero elements of $\boldsymbol{b}$ indicate the importance of a factor as a determinant of the pricing kernel. I model equation (44) with only two factors: the world market risk factor and a foreign exchange risk factor. Inclusion of the foreign exchange risk factor is informed by Zhang (2006), who finds that exchange risk premiums contribute significantly to the excess returns on international assets, and that the conditional International Capital Asset Pricing Model (ICAPM) with exchange risk performs better than all international asset pricing models that she investigates.

[33] The law states that taking an expected value using less information of an expected value that is formed on more information, gives the expected value using less information (Cochrane, 2000: 129). For instance, $E[E_t(X)] = E(X)$.

Estimation of Factor Risk Premia

Using the parameterization in equation (41) and in line with previous studies (Harvey and Kirby, 1995; Shanken and Zhou, 2007), the sample moment conditions to be evaluated follow:

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T}\begin{bmatrix} \boldsymbol{R}_{t+1} - \widehat{\boldsymbol{\mu}}_r \\ \boldsymbol{f}_{t+1} - \widehat{\boldsymbol{\mu}}_f \\ \left(\boldsymbol{f}_{t+1} - \widehat{\boldsymbol{\mu}}_f\right)^2 - \widehat{\boldsymbol{\sigma}}_f^2 \\ \boldsymbol{R}_{t+1} - \boldsymbol{\iota}\lambda_0 - \sum_{j=1}^{k}\lambda_j \frac{(\boldsymbol{R}_{t+1}-\widehat{\boldsymbol{\mu}}_r)(\boldsymbol{f}_{j,t+1}-\widehat{\boldsymbol{\mu}}_{j,f})}{\widehat{\boldsymbol{\sigma}}_{j,f}^2} \end{bmatrix} = \boldsymbol{0} \tag{45}$$

where a circumflex denotes that the parameter is an estimate; θ is the vector of all parameters; $\boldsymbol{\iota}$ is an $(n \times 1)$ vector of ones; $\boldsymbol{R}_{t+1}$ is an $(n \times 1)$ vector of asset returns with mean vector $\widehat{\boldsymbol{\mu}}_r$, $\boldsymbol{f}_{t+1}$ is a $(k \times 1)$ vector of scaled and unscaled factors with mean vector $\widehat{\boldsymbol{\mu}}_f$, and $\widehat{\boldsymbol{\sigma}}_{j,f}^2 = Var\left(\boldsymbol{f}_{j,t+1}\right)$. There are a total of $2(n+k)$ moment conditions in system (45). Parameters in the system are estimated using the sequential GMM of Ogaki (1993). Following Harvey and Kirby (1995) and Shanken and Zhou (2007), the moment conditions $g_T(\theta)$ are partitioned into two sub-vectors:

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T}\begin{pmatrix} h_{1,t+1}(\theta_1) \\ h_{2,t+1}(\theta_1, \theta_2) \end{pmatrix} \tag{46}$$

The first sub-vector $h_{1,t+1}(\theta_1)$ contains $n + 2k$ moment conditions and yields $n + 2k$ parameters θ_1 consisting of means of returns, and means and variances of factors. The first sub-system is therefore exactly identified so that its GMM estimator $\hat{\theta}_1 = \left(\overline{\boldsymbol{R}}', \bar{\boldsymbol{f}}', \widehat{\boldsymbol{\sigma}}_f^{2\prime}\right)'$ is independent of the weighting matrix. The second sub-vector is defined as $h_{2,t+1}(\theta_1, \theta_2)$, where $\theta_2 = \boldsymbol{\lambda}$ consists of the risk free rate of return and factor risk premia. By plugging $\hat{\theta}_1$ into the last n moment conditions and setting $1/T\sum_{t=1}^{T}\left[h_{2,t+1}\left(\hat{\theta}_1, \theta_2\right)\right] = 0$, the estimator $\hat{\theta}_2 = \boldsymbol{\lambda}$ is obtained. Since $\boldsymbol{\lambda}$ is a $k + 1(< n)$-vector, the second sub-system is over-identified so that the weighting matrix used in the GMM estimation matters. Shanken and Zhou (2007) suggest use of the identity matrix or the inverse of estimates of the variance-covariance matrix of the n moment conditions. This procedure yields estimates of the risk premia per unit of each factor instead of $n \times 1$ vector of asset betas. Harvey and Kirby (1995) demonstrate that the GMM estimator $\hat{\theta}_2$ is fully efficient. Iqbal et al. (2010) explain that the estimator is not subject to errors-in-variables problem because generated regressors are not employed in the estimation of the risk premia; rather, only the means and variances of returns and factors from the first stage of the sequential GMM process are used.

Stability Tests in Conditional Asset Pricing Models

Unconditional asset pricing models have proved ineffective in capturing time-varying risk premia. This is largely because those models impose strong assumptions on the underlying probability distributions and investors' attitudes toward risk to obtain the time-invariant linear factor structures. Consequently, conditional models have become attractive in empirical asset pricing investigations. However, conditional asset pricing models only work well if they are correctly specified in the sense that the instrumental variables used can correctly capture the dynamics of risk premia. Ghysels (1998) points out the dangers of committing serious pricing errors if the factor risks are inherently misspecified in conditional asset pricing models. He shows that conditional models may have larger pricing errors than unconditional models and attributes this result to structural shifts, the existence of which causes parameter instability.

Garcia and Ghysels (1998) further document the importance of testing for structural changes in the context of emerging markets, especially given the strong political and economic idiosyncrasies that have disrupted these markets in comparison with world events. The duo proposes the use of the sup-LM test of Andrews (1993) as an effective way of checking for structural instability in the stochastic discount factor parameters. The null hypothesis for the sup-LM test is that there are no structural shifts so that parameters are stable. This hypothesis is tested against the alternative that there is a single structural break at some unknown point in time. The sup-LM statistic is the largest of the LM statistics computed, in this study, at 5% increments between 15% and 85% of the sample. The calculated sup-LM is evaluated against critical values in Table 1 of Andrews (2003).

3.3.3 A Brief Description of the Generalized Method of Moments (GMM)

As already mentioned, parameter estimation in this study makes use of the Generalized Method of Moments (GMM). The method is particularly suitable because of its robustness to heteroskedasticity (Smith and Wickens, 2002). Further, Hall (1993: 404) explains that under the GMM approach, one need not make an explicit specification of the data generating process. This can be contrasted with the traditional techniques of estimation, such as the maximum likelihood procedure, where a data generating process, typically informed by an explicitly defined probability density function (PDF), necessarily forms a part of the estimation process. If the model is misspecified, that is, if the specified distribution function turns out not to accurately reflect the true data generating process, techniques such as the maximum likelihood procedure are likely to give biased parameter estimates: clearly, the GMM does not suffer this shortcoming. Hall (1993: 403) also points out that estimation of the required parameters using the GMM procedure is computationally less demanding and more efficient than the computational requirements for such techniques as the Maximum Likelihood estimation.

In the GMM terminology, $g_T(\theta)$ represents the *sample* average pricing errors. The subscript T (the sample size) indicates the dependence of this statistic on the sample; θ represents all the parameters to be estimated. The Generalized Method of Moments (GMM) selects consistent and asymptotically normal and efficient estimators $\hat{\theta}$ of θ so as to minimize an objective function of the following quadratic form (Cochrane, 2000: 179):

$$Q_T(\theta) = g_T(\theta)' W_T \, g_T(\theta) \tag{47}$$

where $g_T(\theta)$ denotes the $(m \times 1)$ vector of pricing errors with m being the number of sample moments; W_T is a positive definite symmetric weighting matrix, which converges in probability to a positive symmetric and non-singular matrix W (Hall, 1993: 395). The optimal weighting matrix, W_T^*, is the one that minimizes $g_T(\theta)$. As Tauchen (1986) notes, the dimension of the right side of equation (47) is mr, hereinafter referred to as the number of moment conditions, such that r is the number of components of the instrument vector.

An order condition for identification is that $mr \geq p$, where p is the number of parameters to be estimated. If mr is equal to the number of parameters to be estimated, then, in general, it is possible to choose $\hat{\theta}$ so that $g_T(\hat{\theta})$ equals zero exactly and the value of the objective function (47) will be zero at the estimate. In this case, Hall (1993: 396) explains that θ, is "just identified" because there is just enough information (that is, moment conditions) to estimate θ. On the other hand, if mr exceeds p, then $g_T(\theta)$ cannot in general equal zero for any θ, and only asymptotically will the minimized value of the objective function vanish (Tauchen, 1986). Hall (1993: 396) explains that this situation results in θ being "over-identified" because there is more information than is necessary to obtain an estimate of θ. In this case, Hansen (1982) demonstrates that equation (47) is minimized by $\hat{\theta}$, the consistent estimator of θ, whose asymptotic variance-covariance matrix, $\sum_W$, is informed by the limiting weighting matrix, W. Tauchen (1986) explains that, among all possible W's, the one that minimizes (in the matrix sense) the limiting variance-covariance matrix is $W_T^* = (E[g_T(\theta)g_T(\theta)'])^{-1}$, which is simply the inverse of the variance-covariance matrix of the random variable $g_T(\theta)$.

To test for the overall goodness-of-fit of the model (that is, the null hypothesis that the estimated coefficients are all zero), the J_T-statistic (known as the test for over-identifying restrictions) is used. When the asymptotically optimal weighting matrix is used, the J_T-statistic equals the sample size, T, times the minimized value of the objective function in equation (47), $Q_T(\hat{\theta})$, that is, $J_T = T.Q_T(\hat{\theta})$. When an arbitrary weighting matrix other than the asymptotically optimal one is used, the statistic is given as $J_T = T.g_T(\hat{\theta})'[(I - d(ad)^{-1}a)S(I - d(ad)^{-1}a)]^{-1}g_T(\hat{\theta})$, where $a \equiv p\, lim a_T$ with a_T being the matrix that defines the linear combination of $g_T(\theta)$ that will be set to zero and d is the population moment (Vassalou, 2003).[34] S is known as the spectral density matrix at frequency zero of u_t, the residuals, at time t (Cochrane, 2000: 182); it takes on the value, $S = \sum_{j=-\infty}^{\infty} u_t u'_{t-j}$. In either case, the J_T-statistic takes the form of a chi-square distribution with degrees of freedom equal to the number of moment conditions minus the number of parameters. A high J_T-statistic implies that the disturbances are correlated with the instrumental variables and signifies model misspecification.

Several studies have investigated the finite sample behavior of the GMM estimator. Tauchen (1986) uses sample sizes of 50 and 75 in the consumption-based asset pricing model with a vector of ones and lagged endogenous variables ($L = 1, 2, 3, \dots, 4$) as the instruments. The study

[34] Vassalou (2003) notes that the resulting variance-covariance matrix is singular because the terms $I - d(ad)^{-1}$ set some linear combinations of $g_T(\theta)$ to zero in order to estimate the parameters. Because the variance-covariance matrix is singular, it requires pseudo-inversion.

finds that choice of L has an impact on the estimator, which exhibits enhanced efficiency but increased bias as L increases. He advises that L should be chosen to be small. The study also finds that the J_T-statistic is well approximated by its asymptotic distribution. Kocherlakota (1990) using a similar experimental design as Tauchen (1986) with L held constant at zero, confirms many of Tauchen's (1986) findings but reports that the J_T-statistic tends to reject too frequently. Kocherlakota's (1990) key finding is that the performance of GMM estimator worsens with large instrument sets; in particular, the author reports biasness for some estimators and too narrow confidence intervals based on the asymptotic distribution.

With real financial data, Ferson and Foerster (1994) perform a simulation experiment using the seemingly unrelated regression model. They consider both a two-step procedure and the iterated procedure to derive GMM estimators. In the iterative procedure, one obtains further estimates from the original until convergence occurs between parameter estimates and the minimized value of the objective function. They find that GMM estimators exhibit relatively small biases even with sample sizes as small as sixty. However, the sample sizes based on asymptotic theory understate the variability of the estimator; the bias increasing with decreasing sample size and increasing dimension of the system. Their results also show that the test of over-identifying restrictions rejects too often when the two-step procedure is used; with the iterative procedure, the test though undersized, is much closer to the nominal size. To enhance the finite sample properties of the estimates, the authors advise that the iterated GMM should be used in practice. Following this advice, I test the two asset pricing models in this study using the iterated GMM.

3.4 Testing the Linkage between Foreign Exchange Rates and Foreign Portfolio Flows

The uncovered interest parity relationship, presented in equations (A1) and (A4) of Appendix I, forms the bridge between the exchange rate and the asset market equilibrium. For asset markets to remain in equilibrium, ceteris paribus, a decrease in domestic output must be accompanied by a currency depreciation (Granger et al., 2000).[35] Clearly, if this relationship holds, the expected returns on assets would be influenced by anticipated changes in exchange rates and vice versa. In turn, changes in return expectations have an impact on the willingness of investors to undertake investments, and by extension, the flow of capital across borders.

[35] See Appendix I for a detailed exposition.

The nature of causality between foreign exchange rates and asset prices has not been resolved in the literature. On the one hand are scholars such as Aggarwal (1981) who argue that exchange rate movements affect firms' values and that such effects are manifested in the prices of assets, such as common stocks, issued by those firms. From this viewpoint, exchange rate changes are expected to give rise to asset price changes. Such a causal relation is known as the *traditional approach* (Granger et al., 2000). On the other hand, proponents of the *portfolio balance approach* believe that asset price fluctuations influence exchange rate movements. The latter approach views exchange rates like the price of any traded commodity determined by the market mechanism: a flourishing stock market or an improvement in a country's investment climate attracts capital flows from foreign investors and hence causes an increase in the demand for a country's currency (Pan et al., 2007). Consequently, rising (declining) asset prices are related to appreciating (depreciating) exchange rates. Under the assumptions of this approach therefore, asset prices are expected to lead exchange rate with a negative correlation (Granger et al., 2000).

Because these two competing schools of thought conflict in their explanation of the direction of causality between foreign exchange rate fluctuations and asset prices (and, by extension, international capital flows), it is of interest to know the exact nature of the linkage between these variables in the African context. The definition of causality provided by Granger (1969) has traditionally formed the basis for determining the relationship between two stationary time series. The Granger causality measure is predicated on the two series being stationary. However, economic variables such as foreign exchange rates and asset (especially common stock) prices typically exhibit a random walk[36] and are therefore non-stationary (Gujarati, 2004: 800). If asset prices are non-stationary, it is unreasonable to assume that international capital flows, which depend largely on the changes on those prices, would be stationary. A random walk is also known loosely as a *unit root process* in time series literature.

[36] A stochastic variable Y is said to follow a random walk without drift if its value at time t can be mathematically expressed as the sum of its value at time t - 1 and a random shock, or white noise, ϵ (with zero mean and constant variance, σ^2): $Y_t = \rho Y_{t-1} + \varepsilon_t$, where ρ is a constant. If $\rho = 1$, the random walk model gives rise to a unit root process (Gujarati, 2004: 799).

3.4.1 Unit Root Test

The Dickey and Fuller (1979) and the Augmented Dickey and Fuller (ADF) methodologies are popular methods of testing for the presence of a unit root (that is, absence of stationarity). To see the logic behind these two tests, consider the following first order autoregressive process, $AR(1)$:

$$Y_t = \rho Y_{t-1} + \varepsilon_t \qquad -1 \leq \rho \leq 1 \tag{48}$$

Subtracting the term Y_{t-1} from both sides of equation (48) gives the first difference form of the random walk model:

$$Y_t - Y_{t-1} = \rho Y_{t-1} + \varepsilon_t - Y_{t-1}$$

$$\Delta Y_t = (\rho - 1)Y_{t-1} + \varepsilon_t = \alpha Y_{t-1} + \varepsilon_t \tag{49a}$$

where $\Delta Y_t = Y_t - Y_{t-1}$ is the first difference of the random variable Y at time t; $\alpha = \rho - 1$ and ε_t is the white noise term at time t. Equation (49a) is restricted in the sense that it ignores the possible presence of a constant term that may cause the series Y_t to drift away from the origin. Introducing a constant term gives random walk model with a drift:

$$\Delta Y_t = \beta_1 + \alpha Y_{t-1} + \varepsilon_t \tag{49b}$$

Finally, the model can be presented in a manner that allows for a drift around a trend as follows:

$$\Delta Y_t = \beta_1 + \beta_2 t + \alpha Y_{t-1} + \varepsilon_t \tag{49c}$$

For each of the three equations (49a, 49b and 49c), the standard Dickey-Fuller procedure tests the null hypothesis that $\alpha = 0$, that is, $\rho = 1$ against the alternative that $\alpha < 0$, that is, $\rho < 1$.[37] Rejection of the null hypothesis implies that the series is stationary. If the null hypothesis is not rejected, one concludes that the series has a unit root, meaning that it is non-stationary. The τ

[37] According to Patterson (2000: 229), the null hypothesis "is straightforward since it corresponds to the existence of a single unit root." The author explains that since an alternative hypothesis should be chosen so as to maximize the power of the test in the likely direction of departure from the null, a two-sided alternative, $\alpha \neq 0$, implying $\alpha < 0$ *and* $\alpha > 0$, is not generally chosen because $\alpha > 0$ corresponds to $\rho > 1$, whose implication is that the data generating process for Y_t is unstable. Consequently, the left-tailed alternative, $\alpha < 0$, which is the more likely deviation from the null, is typically chosen.

(tau) statistic, whose critical values have been developed by Dickey and Fuller (1979), is used to test the null hypothesis. The standard Dickey-Fuller test assumes that the white noise terms ε_t are serially uncorrelated (Gujarati, 2004: 817). If this assumption is violated, the augmented version of the Dickey-Fuller test (ADF) is more appropriate. The ADF test constructs the *Y*-series to follow an $AR(p)$ process by including p lagged difference terms of the dependent variable *Y* on the right hand side of the model as follows:

$$\Delta Y_t = \beta_1 + \beta_2 t + \alpha Y_{t-1} + \theta_1 \Delta Y_{t-1} + \theta_2 \Delta Y_{t-2} + \ldots + \theta_p \Delta Y_{t-p} + \varepsilon_t$$

$$= \beta_1 + \beta_2 t + \alpha Y_{t-1} + \sum_{j=1}^{p} \theta_j \Delta Y_{t-j} + \varepsilon_t \tag{50}$$

where $\Delta Y_{t-j} = Y_{t-j} - Y_{t-(j+1)}$ are the lagged differences. The number of lagged differences is determined in such a way as to ensure that the error term in the resulting model is serially uncorrelated. In principle, the choice of an appropriate lag length, or order size, can be done by a visual "analysis" of a graphical plot of the autocorrelation (ACF) and partial autocorrelation (PACF) functions. If the data generating process of a variable exhibits a theoretical, stationary autoregressive process, the ACF plot should peter out geometrically as lag length increases. In such situations, the appropriate lag length is obtained as the point at which the ACF becomes insignificant at some predetermined level of significance (say 5%).

Unfortunately, real financial data may be too "messy" to exhibit an ACF pattern sufficiently smooth that it can be visually identified with a particular autoregressive and moving average (ARMA) process. In such cases, it becomes necessary to employ certain scientific procedures, known as *information criteria*, to make an informed choice of an appropriate lag length. Several procedures, differing only on the stiffness of the penalty for increasing the number of lags, are available for choosing the appropriate lag length. As Koreisha and Pukkila (1995) explain, each of the criteria works on the basis of the minimization of a function of the following form:

$$\delta(k) = n \ln \hat{\sigma}_k^2 + k.g(n) \tag{51}$$

where $n = (T - k); k = 1, 2, \dots, k^*$ is the number of degrees of freedom. k^* is an *a priori* determined upper limit. The information criteria are therefore minimized subject to the upper bound, specified as the number of moving average (q^*) and/or the number of autoregressive (p^*) terms to be considered (Brooks, 2005: 257). Since T is the number of data points, n is regarded as the effective number of observations.

$k = p + q + 1$ is the total number of parameters estimated. In an ARMA process, p is the number of lags of the autoregressive (AR) component while q is the number of lags of the moving average (MA) component. If $q = 0$, the resulting process is a pure autoregressive process of order p, so that $k = p + 1$.

$\hat{\sigma}_k^2$ is the residual variance (equal to the residual sum of squares divided by the effective number of observations or number of degrees of freedom). Thus, $\hat{\sigma}_k^2 = \sum \hat{\varepsilon}^2 / n$.

$k.g(n)$ is known as the penalty term and generally increases as more lags are included.

The term $g(n)$ is defined separately by different information criteria: when $g(n) = 2/n$, the resulting model is the Akaike's (1974) Information Criterion (AIC); when $g(n) = \ln(n)/n$, the model that results is the Schwarz (1987) Bayesian Information Criterion (SBIC); finally, the Hannan and Quinn (1979) Information Criterion (HQIC) is obtained when $g(n) = \frac{2}{n}\ln(\ln(n))$.

From these relationships, it is to be noted that, since the penalty term, $\ln(n)/n$, is the greatest of the three penalty terms (for any 'usable' sample size), the marginal cost of adding lagged differences is greater under SBIC than under both the AIC and the HQIC. Consequently, of the three criteria, the SBIC tends to choose the most parsimonious model (Enders, 2004: 70). However, Brooks (2005: 258) observes that although strongly consistent asymptotically, the SBIC is less efficient. The consistency property implies that, on the average, SBIC tends to deliver the correct model order in large samples. On the other hand, weak efficiency means that model orders selected by the SBIC from different samples within a given population tend to exhibit larger average variation. The author also observes that the AIC is not consistent but is generally more efficient. Dickey and Fuller (1979) show that the asymptotic distribution of the τ (tau) statistic is independent of the number of lagged difference terms used in the ADF regression. Hence, the same critical values are used for both the DF and the various forms of the ADF tests. This study uses the SBIC method in the choice of lag lengths.

3.4.2 Cointegration Test

A time series is said to be integrated of order d, $I(d)$, if stochastic trends or unit roots can be removed by differencing the series d times and a stochastic trend still remains after differencing only $d-1$ times (Lütkepohl, 2007: 279). Accordingly, a variable without a stochastic trend or unit root is also said to be integrated of order zero, $I(0)$. A set of variables of the same integration order d (typically 1), are said to be cointegrated if a linear combination of the variables exists which is $I(0)$. In econometric parlance, two (or more) economic variables are said to be cointegrated if a long-run, or equilibrium, relationship exists between (or among) them.

Two of the most commonly used methods for testing for cointegration are the Engle and Granger (1987) and the Johansen (1988, 1995) procedures. The Engle-Granger procedure is predicated on the notion that two or more economic variables are cointegrated if the residuals from the regression of the two variables exhibit stationarity, i.e. if the residuals are integrated of order zero, $I(0)$. Therefore, the Engle-Granger cointegration tests try to establish whether the noise term ε_t is $I(0)$. The Engle-Granger and the Augmented Engle-Granger forms, respectively, of the cointegration test proceed by running regressions of the following forms on the residuals:

$$\Delta\hat{\varepsilon}_t = \gamma\hat{\varepsilon}_{t-1} + \vartheta_t \tag{52a}$$

$$\Delta\hat{\varepsilon}_t = \gamma\hat{\varepsilon}_{t-1} + \sum_{k=1}^{q}\delta_k\Delta\hat{\varepsilon}_{t-k} + \vartheta_t \tag{52b}$$

where, as before, Δ is the first difference operator whereas a circumflex denotes an estimated or fitted value. Regression coefficients are tested for significance using the τ (tau) statistic for γ. The null hypothesis is that the variables are not cointegrated, i.e. the residuals from their regression are not $I(0)$. The null hypothesis that the residuals ε_t are not $I(0)$ is rejected if the computed τ statistic is more negative than the critical τ statistic.

The Engle-Granger procedure, although simple and easy to implement, is not short of defects. First, as Enders (2004: 347) observes, estimation of the long-run equilibrium relationship requires that the researcher places one variable on the left-hand side and use the other(s) as regressor(s). In practical situations, where only small samples are available to the researcher, it is possible to find one regression indicating the existence of cointegration, whereas reversing the order indicates no cointegration! Because the test for cointegration should be invariant to the choice of the variable selected for normalization, this is a very undesirable feature of the procedure. Where there are more than two variables, there is always the possibility that more than one cointegrating vectors exist; the Engle-Granger approach has no systematic procedure for separate estimation of the multiple cointegrating vectors. The second defect arises from the procedure's reliance on the two-step estimation: generating the residual series first and then using them to estimate the regressions in equations (42a) and (42b). The problem here is that any errors introduced in the first estimation are carried into the second step.

Johansen (1988, 1995) has developed a maximum likelihood estimation procedure based on the reduced rank regression method that addresses some of these drawbacks. First, it relaxes the assumption that the cointegrating vector is unique and secondly, it takes into account the short-run dynamics of the system when estimating the cointegrating vectors (Dolado et al., 2001). The approach is based on estimating the p^{th} order Vector Autoregressive (VAR) model in *n* variables:

$$y_t = \pi_1 y_{t-1} + \pi_2 y_{t-2} + \ \dots + \pi_p y_{t-p} + \varepsilon_t \ = \textstyle\sum_{i=1}^{p} \pi_i y_{t-i} + \varepsilon_t \qquad (53)$$

where $t = 1, 2, \dots, T$, $\varepsilon_t \sim IN(0, \sum_\varepsilon)$[38], such that $\sum_\varepsilon$ is the *n*-dimensional variance-covariance matrix. Since levels of time series y_t might be non-stationary, it is better to transform equation (53) into a dynamic vector error correction model (VECM) of the form (Kozhan, 2009: 110):

$$\Delta y_t = \pi y_{t-1} + \Gamma_1 \Delta y_{t-1} + \Gamma_2 \Delta y_{t-2} + \ \dots + \Gamma_{p-1} \Delta y_{t-(p-1)} + \varepsilon_t$$

$$= \pi y_{t-1} + \textstyle\sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + \varepsilon_t \qquad (54)$$

where $\pi = -(I - \pi_1 - \pi_2 - \ \dots - \pi_p) = \ -(I - \sum_{i=1}^{p} \pi_i)$ and $\Gamma_i = -\sum_{i=1}^{p-1} \pi_i$. Now suppose that y_t contains some non-stationary $I(1)$ time series components. In order to get the stationary

[38] This notation means that the error term has an Independent (I) and Normal (N) distribution. The mean and variance of the distribution are shown in parentheses.

error term, ε_t, the term πy_{t-1} should also be stationary. This requires that πy_{t-1} must contain $r < n$ cointegrating relationships. Now, if the VAR(p) process in equation (54) has unit roots, then the matrix π has reduced rank given by $rank(\pi) = r < k$. Accordingly, Johansen (1995) argues that testing for cointegration of the system is equivalent to checking out the rank of the matrix π.[39] The following relationships can be drawn (Enders, 2004: 352): if $rank(\pi) = 0$, the matrix is null and equation (44) is the usual VAR in first differences; if $rank(\pi) = n$ the vector process is stationary; if $rank(\pi) = 1$, there is a single cointegrating vector and the expression πy_{t-1} is the error correction term; for other cases in which $1 < rank(\pi) < n$, multiple cointegrating vectors exist.

Kozhan (2009: 111) explains that if $1 < rank(\pi) = r < n$, the implication is that y_t is $I(1)$ with r linearly independent cointegrating vectors and $k - r$ non-stationary vectors. Since π has a rank r, it can be written as the product, $\pi = \alpha\beta'$, where α and β are $n \times r$ matrices such that $rank(\alpha) = rank(\beta) = r$. β is a matrix of long-run coefficients and α represents the speed of adjustment to disequilibrium. Thus, the VECM becomes:

$$\Delta y_t = \alpha\beta' y_{t-1} + \Gamma_1 \Delta y_{t-1} + \Gamma_2 \Delta y_{t-2} + \ \dots + \Gamma_{p-1} \Delta y_{t-(p-1)} + \varepsilon_t$$

$$= \alpha\beta' y_{t-1} + \textstyle\sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + \varepsilon_t \qquad (55)$$

with $\beta' y_{t-1} \sim I(0)$. The maximum likelihood estimates of β equals the matrix of eigenvectors corresponding to r largest eigenvalues (or characteristic roots) of a $n \times n$ residual matrix. In practice, only estimates of π and its eigenvalues are obtained. Let the n estimated eigenvalues of the matrix π be ordered such that $\hat{\lambda}_1 > \hat{\lambda}_2 > \ \dots > \hat{\lambda}_n$. The number of distinct cointegrating vectors can be obtained by checking the significance of the resulting eigenvalues. To test for the significance of the eigenvalues, two likelihood ratio statistics are used: The first is known as the *trace statistic* and is given as:

$$\lambda_{trace}(r_0) = \ -T \textstyle\sum_{i=r+1}^{n} \ln\left(1 - \hat{\lambda}_i\right) \qquad (56)$$

[39] The rank of a matrix is is the number of linearly independent rows, or columns, of the matrix. Technically, it is equal to the number of characteristic roots (eigenvalues) that differ from zero.

where $T = n - 2$ is number of usable observations; r_0 is the anticipated number of cointegrating vectors and $\hat{\lambda}_i$ are the estimated eigenvalues. This statistic tests the null hypothesis, $H_0: r \leq r_0$ against the alternative, $H_1: r > r_0$, such that *r* is the number of distinct cointegrating vectors. The trace statistic will equal zero if all λ_i are zero. The further the estimated eigenvalues are from zero, the more negative is $\left(1 - \hat{\lambda}_i\right)$ and the larger is the trace statistic (Enders, 2004: 353). The second statistic is known as the *maximum eigenvalue statistic*. This statistic tests the null hypothesis $r_0 = r$ against the alternative $r_0 = r_0 + 1$. If $\hat{\lambda}_{r_0+1}$ is close to zero, the *maximum eigenvalue statistic* will be small (Dolado et al., 2001). This statistic is computed thus:

$$\lambda_{max}(r_0, r_0 + 1) = -T \ln\left(1 - \hat{\lambda}_{r_0+1}\right) \tag{57}$$

The Engle-Granger and the Johansen approaches yield consistent results if there is a single (unique) cointegrating vector. However, as already pointed out, the former cannot be used in the presence of more than one cointegrating vectors. Similarly, some element of ambiguity is present in the Engle-Granger procedure as error terms from *any* equilibrium relationship among the variables can be tested – i.e. it has no mechanism for identifying the particular relationship of interest. The Johansen procedure is therefore preferred to test for cointegration in this study.

3.4.3 Granger Causality Test

Granger (1969) suggested a method for determining how much of the current value of a variable, *Y*, can be explained by past values of *Y* and whether adding lagged values of another variable, *X*, can improve the explanation. Then, *Y* is said to be "Granger-caused" by *X* if *X* helps to predict *Y*. Similarly, *X* is said to be "Granger-caused" by *Y* if *Y* helps to predict *X*. If cointegration does not exist between *Y* and *X*, then the Granger Causality test is performed by running a bivariate vector autoregression (VAR) of the following form:

$$\begin{aligned} Y_t &= \alpha_{Y0} + \sum_{i=1}^{L} \beta_i Y_{t-i} + \sum_{i=1}^{L} \gamma_i X_{t-i} + \varepsilon_{Yt} \\ X_t &= \alpha_{X0} + \sum_{i=1}^{L} \delta_i X_{t-i} + \sum_{i=1}^{L} \theta_i Y_{t-i} + \varepsilon_{Xt} \end{aligned} \tag{58}$$

where *Y* and *X* represent the variables of interest, in this case international portfolio flows foreign and exchange rates, respectively. The null hypotheses are $H_{01}: \gamma_1 = \gamma_2 = \cdots = \gamma_L = 0$ and

$H_{02}: \theta_1 = \theta_2 = \cdots = \theta_L = 0$. Failure to reject H_{01} implies that foreign exchange rates (*X*) do not Granger-cause international portfolio flows (*Y*); whereas failure to reject H_{02} implies that international portfolio flows (*Y*) do not Granger-cause foreign exchange rates (*X*).

If *Y* and *X* are cointegrated, the VAR in system (58) is misspecified since it excludes the long-run equilibrium relationships, among the variables, that are contained in $\pi_X X_{t-1}$ and $\pi_Y Y_{t-1}$ (Enders, 2004: 358). Because of the misspecification, the coefficient estimates from equation (58) and the associated tests are not representative of the true data generating process. To deal with this problem, an error correction term is incorporated in the Granger causality test, resulting in the following form of the bivariate regression (see also Granger et al., 2000):

$$\begin{aligned} Y_t &= \alpha_{Y0} + \pi_Y(Y_{t-1} - \omega X_{t-1}) + \textstyle\sum_{i=1}^{L} \beta_i Y_{t-i} + \sum_{i=1}^{L} \gamma_i X_{t-i} + \varepsilon_{Yt} \\ X_t &= \alpha_{X0} + \pi_X(X_{t-1} - \omega Y_{t-1}) + \textstyle\sum_{i=1}^{L} \delta_i X_{t-i} + \sum_{i=1}^{L} \theta_i Y_{t-i} + \varepsilon_{Xt} \end{aligned} \tag{59}$$

In this formulation, the terms π_Y and π_X denote the speeds of adjustment and are incorporated in the null hypotheses. Thus, the revised null hypotheses are $H_{01}: \gamma_1 = \gamma_2 = \cdots = \gamma_L = 0$ *and* $\pi_Y = 0$; and $H_{02}: \theta_1 = \theta_2 = \cdots = \theta_L = 0$ *and* $\pi_X = 0$. Thus, *Y* does not Granger-cause *X* if lagged values Y_{t-i} do not enter X_t equation *and* if X_t does not respond to the long-run equilibrium. One can, accordingly, draw the implication that X_t is weakly exogenous. In short, if $\pi_X = 0$ in equation (59), then *X* is weakly exogenous and is not Granger-caused by *Y*. The resulting *F*- statistics are interpreted in a manner similar to those of equations (58).

3.5 Data Description and Measurement

In this study, the USA serves as the domestic country while the African countries under investigation serve as the foreign investment destinations. Except where explicitly stated otherwise, all categories of data and all returns are measured in US dollars and cover the period January 1997 through December 2009. This is the period during which foreign investors' participation has been allowed in the financial markets of most countries in Africa. Coincidentally, this is also the period over which financial time series data is consistently available for the sampled countries. Observations are sampled at monthly intervals. Excess returns are in respect of one-month yields on three-month US Treasury bills.

The end-of-month rates of exchange between the US dollar and the selected African currencies and annual flow of foreign portfolio capital for various African countries are obtained from the International Financial Statistics (IFS) database provided by the International Monetary Fund (IMF) and availed online via the Wits University Library. Monthly international portfolio flows of some of the sampled African countries are obtained from the USA Treasury department database. As pointed out by Kasekende et al. (1996), data on African countries' capital flows are prone to recording errors and sometimes outright manipulation as a result of which the portfolio capital flows may be unreliable. Whereas data issues in other regions distinguish private from official inflows; gross and net inflows; short- from long-term; bank from non-bank; and debt from non-debt, African data are mostly muddled up, with recording of many capital inflows in the current account or under errors/omissions being a common feature. Consequently, it has been common practice to rely fairly extensively on IMF (IFS), Organization for Economic Cooperation and Development, World Bank or USA Treasury department.

The exchange rate is defined as the foreign currency price of the US dollar so that a positive change means an appreciation of the US dollar. Following Carrieri and Majerbi (2006), we account for inflation rate differentials, in the GMM estimation, by converting nominal exchange rates to real exchange rates using consumer price indices obtained from IFS. The real exchange rate of a foreign currency is calculated as the product of the nominal exchange rate with respect to the US dollar and the CPI of the USA relative to the CPI of the foreign country: for instance, the real exchange rate of the Kenya shilling (KES) with respect to the US dollar ($) is computed as $RER_{KES/\$} = NER_{KES/\$} \times (CPI_{USA}/CPI_{Kenya})$, where RER is the real exchange rate, and NER is the nominal exchange rate.[40] The reason for using real exchange rates is based on the belief that there is a strong relationship between nominal exchange rate movements and movements in

[40] Real exchange rates can be defined in many ways. The definition used here conforms to Kreinin (2002: 267), with foreign and local prices proxied by the respective CPI. Other authors define the real exchange rate in logarithmic form [e.g., Chinn (2006), equation (1)]: in that case, $RER_{KES/\$} = \ln(NER_{KES/\$}) + \ln(CPI_{USA}) - \ln(CPI_{Kenya})$. In either case, the change in the real exchange rate is mathematically defined, in continuous time, as:

$$\Delta RER_{KES/\$} = \left[\ln\left(NER_t^{KES/\$}\right) - \ln\left(NER_{t-1}^{KES/\$}\right)\right] + [\ln(CPI_t^{USA}) - \ln(CPI_{t-1}^{USA})] - \left[\ln\left(CPI_t^{Kenya}\right) - \ln\left(CPI_{t-1}^{Kenya}\right)\right].$$

This definition is used to compute the changes in real exchange rates used in Tables 4 and 8 of this dissertation.

inflation rates in Africa, so that failure to adjust for purchasing power changes would have a significant effect on the quality of the series. Another way of dealing with inflation is to include it as a variable in the model but multicollinearity in the resulting model may adversely affect the efficiency of estimates. The relative change in the value of foreign currency, s_t, is calculated as

$$s_t = \ln S_t - \ln S_{t-1} \tag{60}$$

where S_t is the real rate of exchange at time *t*. National stock market indices for Botswana, Ghana, Kenya and Nigeria are obtained from the *Datastream*. Market indices for the other countries are obtained from the *MSCI Barra*.[41] World equity market portfolio and emerging markets composite indices are, similarly, procured from the MSCI Barra. MSCI Barra has a uniform base period of 1969 for all its indices; it provides index values in US dollars. Both MSCI and *Datastream* indexes are value-weighted and are calculated with dividends reinvested. Harvey (1991) observes that the MSCI indexes, although weighted towards larger capitalization stocks, are similar to widely quoted country index returns. This means that the country indexes reported by *Datastream* and the MSCI indexes are largely comparable. Indeed, correlation analysis reports a coefficient of 97.33% between MSCI Kenyan returns and the value-weighted Nairobi Stock Exchange 20-share Index returns (reported by *Datastream*) for the period May 2002 through December 2008. For South Africa, there is a correlation of 83% between MSCI index returns and the Johannesburg All-share Index returns reported by *Datastream* for the period between May 2005 and December 2008. For Nigeria, the correlation coefficient is 99.5% between MSCI index returns and the Nigerian Stock Exchange Index reported by *Datastream* for the period May 2002 through December 2008.

The gross return on an index is computed by adding one to the monthly random rate of return. The monthly random rate of return is, in turn, calculated thus:

$$\tilde{r}_{i,t} = \ln I_{i,t} - \ln I_{i,t-1} \tag{61}$$

where $I_{i,t}$ is the value of index *i* at time *t*. The excess return on an index is defined as the random rate of return on that index minus the dollar one-month nominal risk-free rate. Since the study

[41] Harvey (1991) and Ferson and Harvey (1994) have discussed some of the major characteristics of MSCI indices.

looks at the US investor as the domestic investor, the risk free rate is proxied by one-month yield on the three-month US Treasury Bill, obtained from US Federal Reserve Bank database. Because the data are provided as annualized percentages, the monthly yields are calculated using the continuous compounding formula:[42]

$$r_{f,monthly} = \exp(r_{f,annualized}/12) - 1 \tag{62}$$

Next, the issue of the choice of instrumental variables is addressed. A major weakness of the Stochastic Discount Factor (SDF) and other conditional asset pricing models is that they do not specify the instrumental variables (conditioning information) to include in empirical analysis. However, Dumas and Solnik (1995) explain that instrumental variables can be proxied by endogenous variables, such as financial market variables, that are observed frequently. The general criteria for inclusion of such economic and financial variables are that they must be predictors of return or leading business cycle indicators (Drobetz et al., 2002). Importantly, the instrumental variables should approximate the information set used by investors in setting prices. In principle therefore, instrumental variables can be chosen by performing an analysis of predictability between hypothesized variables and variation in first and second conditional moments (Robotti, 2001).

Several studies have investigated the predictability of foreign exchange rate, bond and equity returns in the industrial and emerging markets. Fama and French (1988) find that *dividend yields* have systematic forecast power across different time periods and return horizons in the USA. Additionally, Fama and French (1989) indicate that dividend yields also forecast bond returns. Predictable variation in stock returns is, in turn, tracked by variables commonly used to measure *default* and *term* (or maturity) *premiums* in bond returns.[43] They conclude that dividend yield and the default spread capture similar variation in expected bond and stock returns. However,

[42] Vaihekoski (2009) points out that this method may produce small errors in the risk-free return series. For US Treasury bills, he proposes the formula: $ln[(100dpy - R_t^P\langle dtm - d\rangle)/(100dpy - R_t^P dtm)]$. Assuming 365 days in a year (*dpy*), 91 days to maturity (*dtm*) for the three-month bill, 30 days in a month (*d*), and annualized three-month rate (R_t^P) of 4.5%, this formula gives a monthly yield of approximately 0.37504% against 0.37570% given by my formula. The resulting computation error (as a proportion of 0.37504%) of 0.176% is small enough not to significantly affect the key results of the tests conducted here.

[43] The default-premium variable (the default spread) is defined in the Fama-French study as the difference between the yield on a market portfolio of corporate bonds and the yield on Treasury bonds (or Aaa-rated corporate bonds) of the same maturity. The term- or maturity-premium variable (the term spread) is the difference between the Aaa yield (or treasury notes), usually of ten-year maturity, and the one-month Treasury bill rate.

Campbell and Yogo (2006) find that the predictive power of the dividend yield is considerably weak but the predictive power of the *short rate* of interest is robust. Ang and Bekaert (2007) find that dividend yields predict excess returns only at short horizons and do not have any long-horizon predictive power. They also report that the short rate strongly negatively predicts returns at short horizons. Their results are robust in international data as well.

A study by Keim and Stambaugh (1986) provides evidence suggesting that a *bond yield spread*[44] contains ex ante information about expected risk premiums on many assets, including stocks and bonds. They also find that there is at best a weak positive *January seasonal* in the market beta of the difference in returns between small and large firms, based on the conditional estimates. Gultekin and Gultekin (1983) also find evidence of seasonal patterns of stock returns in many industrial countries. The seasonality is manifested in significantly large mean returns in the turn of the tax year, usually January. Systematically higher returns in the month of January, christened *the January effect* in the literature, have also been reported in some emerging market economies (Claessens et al., 1993). Giovannini and Jorion (1987) find that increases in *interest rates* are associated with predictable increases in the volatility of returns in the foreign exchange market and in the US stock market, and that expected returns both in the stock market and in the foreign exchange market are negatively correlated with nominal interest rates. Finally, Geske and Roll (1983) find, contrary to economic theory, that stock returns are negatively related to both *expected* and *unexpected inflation*. They argue that this puzzling finding does not indicate causality; instead, stock returns are negatively related to contemporaneous changes in expected inflation because they signal a chain of events which results in a higher rate of monetary expansion. They attribute the apparent negative effect of inflation on subsequent stock returns to an empirical illusion driven by spurious causality.

In practice, the selection of instrumental variables can also be guided by previous studies of a similar design. In addition to worldwide factors (such as the dividend yield on world market portfolio and the Eurodollar deposit rate) assumed to apply uniformly across countries, Harvey (1991) considers the following local instruments: lagged own-country equity market index

[44] This is measured, in the Keim and Stambaugh (1986) study, as the difference between yields on Baa-rated bonds and short-term (approximately one month) Treasury bills.

returns, country-specific dividend yield, country-specific short-term interest rates, foreign exchange rate changes, and local maturity spreads. Dumas and Solnik (1995) find that the inclusion of non-dollar interest rates reduces the finite-sample properties of estimates. They conclude that instrument choice is important in determining the success of tests of asset pricing models. In contrast, Buckberg (1995) reports substantial improvements in return prediction following an inclusion of lagged local market instruments. Accordingly, Buckberg's instruments set includes lagged local dividend yield and the lagged return on the dollar-local currency exchange rate. In their study, Cappiello and Panigirtzoglou (2008) also use lagged values of local instruments variables: dividend yields, dollar-local currency exchange rate, local market risk-free interest rate, and local market index return. Many studies also include a January dummy, a variable that is not suitable for this study because, unlike most industrial countries, the fiscal year-end for most countries sampled for the current study is not December 31.[45]

In preliminary investigations, I have considered both common (worldwide, emerging markets and US market) and country-specific instrumental variables. Ordinary least squares regression results presented in Table 1, show that common factors, in isolation, have relatively low predictive power for equity market returns in the representative African equity markets (first row in each case). This probably indicates that capital markets of the sample countries are not fully integrated with those of the rest of the world. Dropping the worst performers (measured by the relative significance of their regression coefficients) and adding lagged values of local market index return improves the model's fit (notice the improvement in adjusted R-squared in the second and third rows) and yields better coefficient estimates.[46] The hypothesis that inflation could be an important factor in return predictability is not supported by the data. The negative association between equity returns and changes in inflation, reported by Geske and Roll (1983), is also largely not supported by the data.

[45] The fiscal year-end is June 30 for Kenya and Egypt, and March 31 for Botswana, Nigeria and South Africa. It is not convenient to include separate dummy variables for different countries. However, the January dummy is used, together with other variables, in this study as a check on the robustness of conditional asset pricing tests results.

[46] It is important state that obtaining "good quality" local return predictors proved difficult. Dividend yields data are only available for South Africa; short-term interest rates (treasury bills, bank-on-bank and principal rates) are available from the IFS, but in most cases, remained constant for long periods of time and could not therefore generate meaningful series. Therefore, this study considers IFS monthly CPI indexes and lagged market index returns, as the only country-specific instrumental variables.

Table 1
OLS regression results for excess equity market index returns on possible instrumental variables

	Cons	WLDLMR	EMPLMR	USALMR	WLDLDY	USATPL	EURDRL	(*i*)LCPI	(*i*)LMR	Adj R^2
BOTSWANA	0.023	0.675*	0.0539	-0.450	-15.079	4.220	3.649			0.0729
	(0.51)	(1.69)	(0.50)	(-1.35)	(-1.29)	(0.43)	(0.51)			
	-0.012	0.329***				5.010	6.077		0.036*	0.0772
	(-0.37)	(3.50)				(0.52)	(0.96)		(1.77)	
	-0.021		0.204***			6.895	8.282		0.041**	0.0763
	(-0.67)		(3.47)			(0.71)	(1.27)		(2.04)	
	-0.009	0.343***				3.395	4.934	0.445		0.0604
	(-0.29)	(3.61)				(0.35)	(0.77)	(0.60)		
EGYPT	0.148**	0.432	-0.001	-0.112	-12.622	-26.854*	-27.666**			0.0664
	(2.15)	(0.70)	(-0.01)	(-0.22)	(-0.69)	(-1.74)	(-2.49)			
	0.094**	0.128				-21.290	-20.039**		0.331***	0.1690
	(2.03)	(0.86)				(-1.50)	(-2.13)		(4.12)	
	0.099**		-0.006			-23.133	-21.213**		0.358***	0.1649
	(2.12)		(-0.06)			(-1.61)	(-2.22)		(4.25)	
	0.125**	0.350**				-26.697*	-26.133***	-0.557		0.0772
	(2.49)	(2.40)				(-1.78)	(-2.61)	(-0.55)		
GHANA	0.145***	0.264	-0.161	-0.202	-28.466**	-17.734*	-26.263***			0.1003
	(3.19)	(0.65)	(-1.44)	(-0.59)	(-2.37)	(-1.75)	(-3.60)			
	0.048	-0.070				-10.450	-13.049**		0.328***	0.1654
	(1.54)	(-0.75)				(-1.09)	(-2.05)		(4.29)	
	0.054*		-0.068			-11.770	-14.326**		0.322***	0.1698
	(1.70)		(-1.18)			(-1.22)	(-2.20)		(4.21)	
	0.065**	-0.086				-13.290	-18.123***	-0.064		0.0641
	(1.99)	(-0.87)				(-1.32)	(-2.73)	(-0.22)		
KENYA	0.073	0.791	0.059	-0.698	5.386	-21.907	-18.906*			0.0509
	(1.19)	(1.42)	(0.39)	(-1.50)	(0.33)	(-1.59)	(-1.90)			
	0.108**	0.206				-27.764**	-24.852***		-0.059	0.0486
	(2.42)	(1.49)				(-2.02)	(-2.72)		(-0.69)	
	0.100**		0.138			-26.094*	-23.059**		-0.052	0.0514
	(2.21)		(1.64)			(-1.89)	(-2.49)		(-0.62)	
	0.110**	0.166				-27.668**	-24.530***	-0.388		0.0522
	(2.49)	(1.26)				(-2.04)	(-2.75)	(-1.03)		
MOROCCO	0.052	0.520	0.063	-0.482	5.846	-21.289**	-10.424			0.0336
	(1.11)	(1.23)	(0.55)	(-1.36)	(0.47)	(-2.03)	(-1.38)			
	0.077**	0.119				-24.262**	-14.073**		0.015	0.0289
	(2.28)	(1.18)				(-2.31)	(-2.05)		(0.18)	
	0.071**		0.092			-23.018**	-12.801*		0.014	0.0338
	(2.08)		(1.47)			(-2.18)	(-1.84)		(0.17)	
	0.076**	0.143				-23.572**	-13.573**	-1.107		0.0438
	(2.33)	(1.43)				(-2.31)	(-2.02)	(-1.55)		
NIGERIA	0.063	0.191	0.367	-0.670	-0.274	-19.352	-13.670			-0.0046
	(0.59)	(0.20)	(1.39)	(-0.83)	(-0.01)	(-0.81)	(-0.79)			
	0.093	0.021				-26.270	-20.347		0.060	-0.0095
	(1.23)	(0.09)				(-1.12)	(-1.32)		(0.74)	
	0.076		0.130			-21.870	-16.719		0.051	-0.0041
	(0.99)		(0.91)			(-0.92)	(-1.06)		(0.63)	
	0.093	0.021				-26.899	-20.736	0.222		-0.0123
	(1.22)	(0.09)				(-1.14)	(-1.34)	(0.37)		

Table 1 (*Continued*)

S. AFRICA	0.083	-0.379	0.086	0.411	3.300	-20.348	-19.609*			0.0098
	(1.18)	(-0.60)	(0.50)	(0.78)	(0.18)	(-1.30)	(-1.73)			
	0.093*	0.187				-20.739	-20.712**		-0.047	0.0190
	(1.87)	(0.92)				(-1.36)	(-2.04)		(-0.42)	
	0.086*		0.187			-19.307	-19.223*		-0.110	0.0223
	(1.72)		(1.17)			(-1.26)	(-1.88)		(-0.79)	
	0.092*	0.119				-20.062	-19.828*	-0.679		0.0194
	(1.87)	(0.79)				(-1.32)	(-1.97)	(-0.49)		

This table uses monthly returns data the period from 1997:1 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). *, **, and *** respectively indicates that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. The dependent variables are the respective excess returns on the local country equity market index, measured in US dollars (Botswana – *BOTEMR*; Egypt – *EGTEMR*; Ghana – *GHAEMR*; Kenya – *KENEMR*; Morocco – *MOREMR*; Nigeria – *NIGEMR*; and South Africa – *ZAREMR*). The explanatory variables are the lagged values of returns on the world market portfolio (*WLDLMR*), lagged values of returns on the emerging market index (*EMPLMR*), lagged values of excess returns on the MSCI index of the USA stock market (*USALMR*), the lagged values of changes in the monthly dividend yield on the world market portfolio (*WLDLDY*), lagged values of the monthly USA term premium (*USATPL*), lagged values of the one-month Eurodollar deposit Rates (*EURDRL*), the lagged values of local equity market index returns (Botswana – *BOTLMR*; Egypt – *EGTLMR*; Ghana – *GHALMR*; Kenya – *KENLMR*; Morocco – *MORLMR*; Nigeria – *NIGLMR*; and South Africa – ZARLMR), and the lagged values of changes in the monthly Consumer Price Index (CPI) (Botswana – *BOTLCPI*, Egypt – *EGTLCPI*, Ghana – *GHALCPI*, Kenya – *KENLCPI* Nigeria – *NIGLCPI* and South Africa – *ZARLCPI*). "Cons" is short form for the regression constant. "Adj R^2" is adjusted R-squared.

Regression results presented in Table 1 are predictive in nature because the regression procedure uses lagged values of the instrumental variables as regressors. Specifically, current values of the dependent variables are regressed against past (lagged) values of independent variables. If the relationship between the two is statistically significant, one can reasonably infer that known values of the regressors inform future (and unknown) values of the dependent variables. Accordingly, a high R-squared can be interpreted to imply high return (dependent variable) predictability; conversely, a low R-squared would imply poor return predictability. Thus, stock market returns appear to be more predictable for Ghana and Egypt (with R-square values higher than 10% in many instances) than for any of the other countries. This result may have implications for informational efficiency in the two markets. Results for Nigeria depart markedly from the general tendency in the African markets: all her models appear to be poorly fitted with all R-squared values being negative. This may be attributed largely to the irregular behavior of the Nigerian naira-US dollar exchange rate over the study period (discussed in section 4.2.3).

Based on these results, I isolate five potential common instrumental variables, namely, the lagged values of world equity market portfolio returns (*WLDLMR*), lagged values of world equity market portfolio dividend yields (WLDDYL), the lagged values of emerging markets composite portfolio returns (*EMPLMR*), lagged values of the one-month Eurodollar rates (*EURDRL*), and lagged values of the USA term premium (*USALTP*); and "one" country-specific instrumental variable, namely, lagged values of local equity market index returns. However, because of the rapid proliferation of orthogonality restrictions in GMM estimation, not all of these instruments can be used.[47] From their coefficient estimates, lagged values of the one-month Eurodollar rates (*EURDRL*), and lagged values of the USA term premium (*USALTP*) are clear favorites. Next, the seven lagged values of local market returns have to be dropped because their inclusion in the GMM system of equations is not practically plausible. Lastly, choice between the lagged values of world market portfolio returns (*WLDLMR*) and the lagged values of emerging market portfolio returns (*EMPLMR*) is enabled by a Pearson's Product Moments Correlation analysis. The results follow:

Table 2
Correlation matrix for excess returns on world market portfolio, and lagged values of returns on the world market portfolio and emerging equity markets index

	WLDEMR	WLDLMR
WLDLMR	0.2141**	
EMPLMR	0.1521**	0.8353**

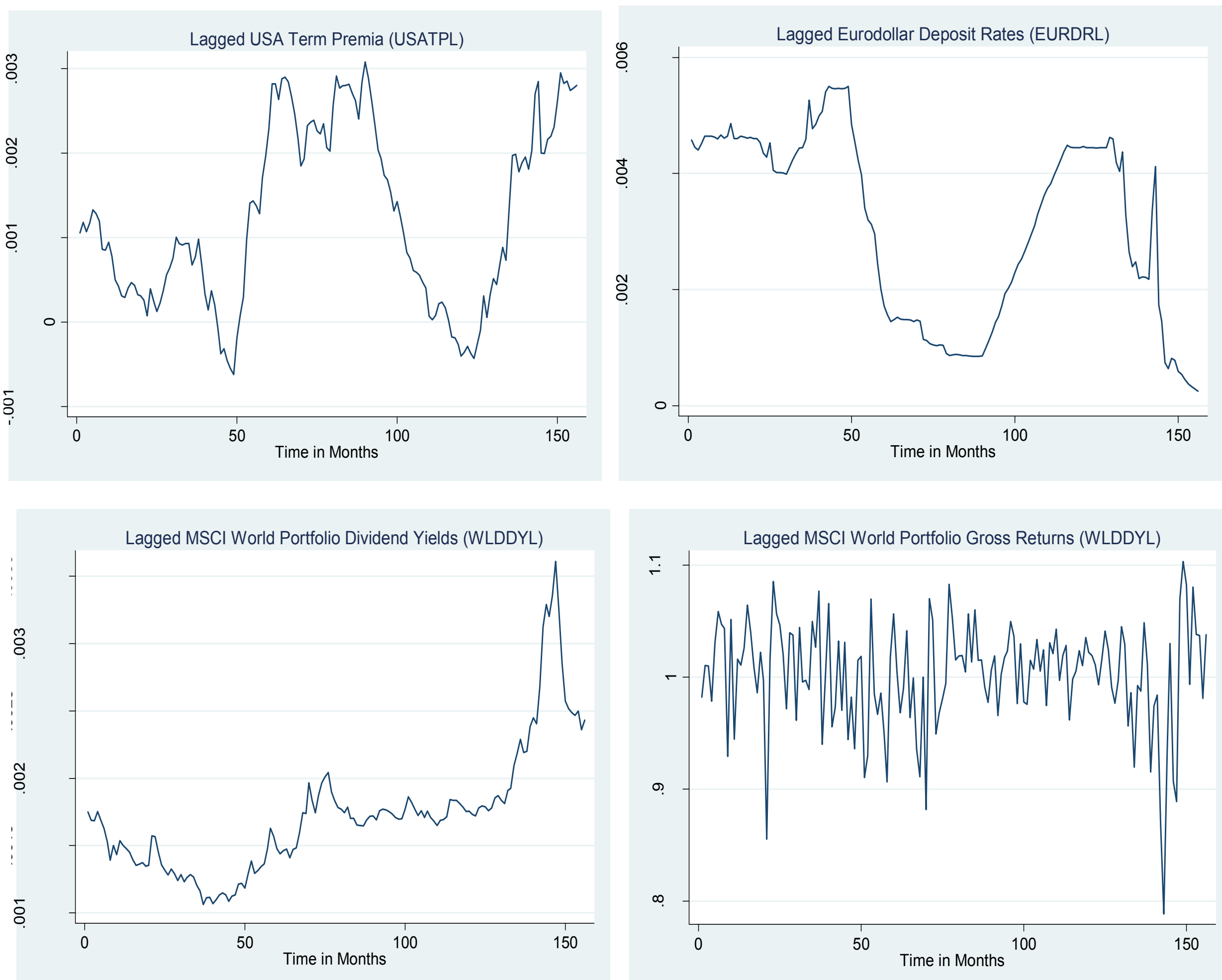
** indicates that the reported coefficient is statistically significant at the 5% level. All data run from 1997:1 through 2009:12. The variables correlated are the excess returns on world market portfolio (*WLDEMR*), the lagged values of world market portfolio excess returns (*WLDLMR*) and the lagged values of emerging market portfolio excess returns (*EMPLMR*).

The correlation results indicate that lagged values of returns on the world market portfolio and lagged values of emerging equity markets index are collinear (correlation coefficient of 0.8353). It is not advisable therefore to use them jointly as conditioning variables in GMM regression since multicollinearity has a biasing effect on the values of the computed standard errors. Similarly, it is also superfluous to use them separately as their information content is likely to be largely similar. The results also show that the lagged values of world market portfolio returns are more closely related to the excess return on the world market portfolio, one of the explanatory variables in the GMM regression. Consequently, "lagged values of emerging market portfolio

[47] In addition to the proliferation of orthogonality conditions in GMM estimation, it will be recalled that many studies, cited earlier (for instance, Tauchen, 1986 and Kocherlakota, 1990), have suggested that usage of few instrumental variables in GMM estimation improve model fit and reduces the tendency of the *J*-statistic to reject the estimates.

returns" is eliminated. Thus, for the stochastic discount model, the study proceeds with four common instrumental variables: *WLDLMR*, *WLDDYL*, *USATPL* and *EURDRL*.

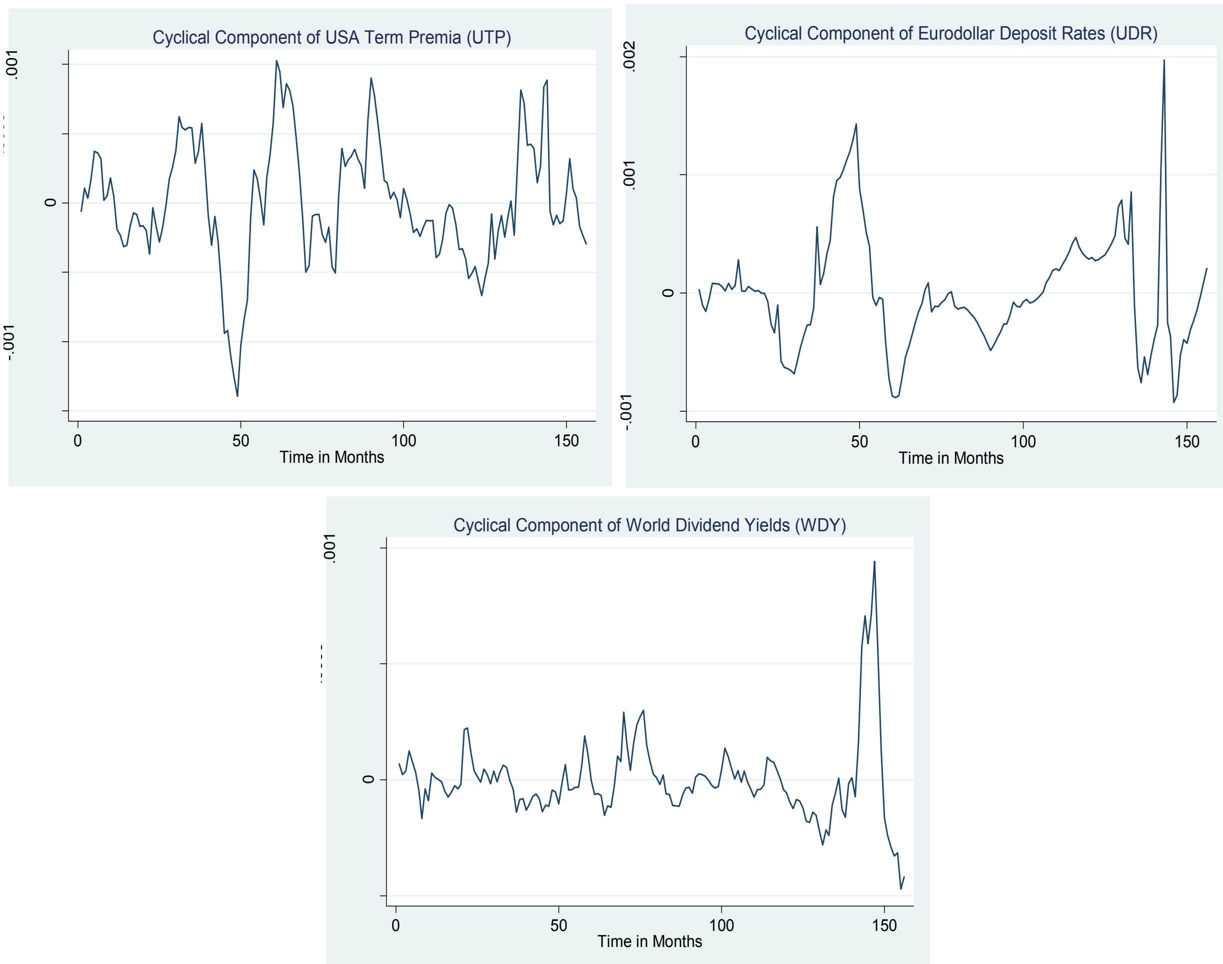
Figure 1
Time series properties of instrumental variables



USATPL, EURDRL, WLDDYL and WLDLMR represent, respectively, the lagged values of the USA term premia, Eurodollar deposit rates and MSCI world equity portfolio dividend yields and MSCI world equity portfolio's gross returns. Data runs from January 1997 through December 2009.

GMM estimation rests on the assumption that the variables used are stationary. However, there is a possibility that some of the chosen instrumental variables may not be stationary. Figure 1 displays, graphically, the time series properties of the common instrumental variables. The figure shows that the lagged values of the USA term premia, the Eurodollar deposit rates and world market dividend yields do not appear to be stationary while the lagged values of world equity portfolio returns appear stationary.

Figure 2
Cyclical components of the USA term premia, Eurodollar deposit rates and the MSCI world equity portfolio dividend yields



UTP, UDR and WDY respectively, represent the cyclical components of the USA Term premia, Eurodollar deposit rates and the MSCI world equity portfolio dividend yields. Cyclical components are extracted using the Hodrick and Prescott (1997) filter. Data runs from January 1997 through December 2009.

Augmented Dickey-Fuller (ADF) unit root tests confirm these observations: the calculated Dickey-Fuller test statistics are –0.933 for the USA term premia, –0.670 for the Eurodollar deposit rates, –0.830 for MSCI world equity portfolio dividend yield and –9.928 for lagged gross returns on the MSCI world market equity portfolio. The critical values are –3.492, –2.886 and –2.576 respectively at 1, 5, and 10 percent levels of significance. Clearly, one cannot conclude that the first three variables (*USATPL, EURDRL* and *WLDDYL*) contain no unit roots. Contrarily, the world market equity portfolio's lagged values do not exhibit the unit root problem.

To deal with the problem of non-stationarity and to ensure that the conditioning variables have the ability to capture the time series properties of the risk factors, empirical implementation of the lagged values of the USA term premia (*UTP*), the Eurodollar deposit rates (*UDR*) and world dividend yield (*WDY*) makes use of their cyclical components. The cyclical components of the three conditioning variables are extracted using the Hodrick and Prescott (1997) filter. Visual investigation (Figure 2) shows that the variables, so constructed, are stationary. The Dickey-Fuller statistics are –3.115, –3.806 and –2.919 respectively for USA term premia, Eurodollar deposit rates and MSCI world market portfolio dividend yields. The hypothesis that unit roots exists in the series is rejected at the 5% level for *UTP* and *WDY* and at the 1% level for *UDR*.

Table 3
Regression of national equity markets gross returns on conditioning variables

	Intercept	WLR	WDY	UTP	UDR	Chi Sq.	R^2
Botswana	0.73***	0.27**	-28.93	-5.80	-3.88	14.61	0.0857
	(6.94)	(2.74)	(-1.06)	(-0.44)	(-0.31)	[0.006]	
Egypt	0.66***	0.35**	3.30	-18.52	-33.92*	11.74	0.0700
	(4.03)	(2.15)	(0.08)	(-0.90)	(-1.74)	[0.019]	
Ghana	1.23***	-0.24**	-91.91***	-16.50	-35.26***	14.58	0.0855
	(11.39)	(-2.20)	(-3.27)	(-1.22)	(-2.73)	[0.006]	
Kenya	0.69***	0.31**	66.44*	-20.24	-4.07	8.28	0.0504
	(4.67)	(2.14)	(1.74)	(-1.10)	(-0.23)	[0.082]	
Morocco	0.87***	0.14	2.44	-28.39**	-13.52	7.34	0.0449
	(7.82)	(1.24)	(0.08)	(-2.04)	(-1.02)	[0.119]	
Nigeria	0.92***	0.08	2.17	-25.18	-0.07	1.19	0.0076
	(3.62)	(0.31)	(0.03)	(-0.79)	(-0.01)	[0.880]	
S. Africa	0.78***	0.22	64.47	-17.41	-22.99*	6.70	0.0412
	(4.76)	(1.36)	(1.51)	(-0.84)	(-1.77)	[0.153]	

This table uses monthly returns data the period from 1997:1 to 2009:12. The values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). *, **, and *** respectively indicates that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. The dependent variables are the respective gross returns on the local country equity market index, measured in US dollars. The explanatory variables are lagged values of returns on the world market portfolio (*WLR*), and the Hodrick-Prescott-filtered lagged values of changes in the monthly dividend yield on the world market portfolio (*WDY*), lagged values of the monthly USA term premium (*UTP*), and lagged values of the one-month Eurodollar deposit Rates (*UDR*). "Chi sq." is the Newey-West statistic from the test that all regression coefficients are zero; its accompanying *p*-value is in square brackets. R^2 is the coefficient of determination.

To test for the ability of the resulting conditioning variables to predict gross returns (used in conditional asset pricing tests), I run an ordinary least squares regression of equity indexes in the sampled countries on the set of conditioning variables. Results are displayed on Table 3. The coefficients of adjustment indicate that the four conditioning variables explain up to 8.6 percent of gross returns for Botswana and Ghana and as low as 0.8% for Nigeria. The recorded return predictability is generally consistent with those reported in advanced equity markets (see, for example, Dumas and Solnik, 1995; Harvey, 1991). The Newey-West (chi-square) test that all coefficients equal zero rejects the null in four countries: Botswana, Egypt, Ghana and Kenya; it fails marginally to reject the null for Morocco and South Africa. The hypothesis that all coefficients equal zero is cannot be rejected in the case of Nigeria whose predictability of gross returns is especially low. It is important to note that each of the conditioning variables is significant in at least one of the regression equations.

The 'Afro' Real Exchange Rate Index

Table 4 displays pairwise correlation coefficients between percentage changes in real exchange rates for various African currencies. Each country's exchange rate is bilateral vis-à-vis the US dollar. Because foreign exchange rates and inflation rates of African countries tend to move together over time, it is observed that many of the values reported are statistically significant at the 5% level. An implication of the high correlations is that it is unwise to include the various country-specific exchange rates in one regression model because of attendant multicollinearity problems. Second, Vassalou (2000) explains that inclusion of a large number of exchange rate variables compromises the efficiency of the resulting exchange rate premia estimates.

Vassalou (2000) proposes a method to deal with the problem of multicollinearity and obviate the need to estimate a large number of currency risk premia. The procedure involves decomposing the bilateral exchange rates into their common (or systematic) and residual components, then using these components to form two equally-weighted indexes to be used jointly as exchange rate factors in the regression. Other researchers (Ferson and Harvey, 1994; Choi et al., 1998; Carrieri and Majerbi, 2006) use broad, widely quoted, exchange rate indexes to circumvent similar problems. In many of these studies, the indexes include more currencies than the ones investigated. For want of a widely quoted index that can represent the various currencies investigated in this study, I develop a broad trade-weighted real exchange rate index for African currencies vis-à-vis the US dollar. I develop a similar index for the euro and used it to investigate the robustness of the study's findings to alternative return measurements.

Currencies are included in the index on the basis of the issuing country's competitiveness ranking, obtained from the Africa Competitiveness Report (World Economic Forum, 2009b). The report is drawn from a study that covered a sample of 134 countries from around the globe. To have as broad a sample as possible for the construction of the index, countries are considered for inclusion if they ranked 110 or less. Thus, a total of fifteen countries, namely, Tunisia (rank, 36), South Africa (45), Botswana (56), Mauritius (57), Morocco (73), Namibia (80), Egypt (81), Gambia (87), Libya (91), Kenya (93), Nigeria (94), Senegal (96), Algeria (99), Ghana (102) and Cote D'Ivoire (110) qualify for inclusion. However, three countries, Namibia, Libya, and Senegal later drop out because of lack of essential data. Thus, the twelve remaining countries' currencies constitute what this study refers to as the *Afro currency index*.

Table 4
Correlation matrix for bilateral real exchange rates

	Botswana	Egypt	Ghana	Kenya	Morocco	Nigeria
Egypt	0.0758					
Ghana	0.0087	-0.0111				
Kenya	0.1957*	0.1362	0.1114			
Morocco	0.4472*	0.0129	0.1380	0.2200*		
Nigeria	0.0207	-0.0173	0.0808	0.0125	0.0501	
South Africa	0.8732*	0.1457	0.0135	0.2133*	0.3771*	0.0397

Percentage changes in real exchange rates are for the period 1997:1 to 2009:12. * denotes statistical significant at the 5% level.

The dollar index covers the period between January 1997 and December 2009 with a base value of 100 at end of December 1996. The euro index covers the period between January 1999 and December 2009, the base month being January 1999. To begin, real exchange rates are calculated for each country using consumer price indexes and the bilateral nominal foreign exchange rates provided by the International Financial Statistics. Bilateral nominal rates of exchange between the euro and African currencies are constructed using cross rates with the US dollar as the intervening currency. The real exchange rates are then fed into the following formula to obtain real exchange rate indexes.

$$I_t = I_{t-1} \prod_{j=1}^{n} \left(RER_{j,t} / RER_{j,t-1} \right)^{W_j} \tag{63}$$

where I_t is the real exchange rate index at time *t*; $RER_{j,t}$ is the real exchange rate for country *j* at time *t*, and W_j is the weight of country *j*. The time-invariant weights used in the formula are computed based on the 2008 (the latest available) bilateral import and export trade data obtained from the World Trade Organization website. The weight for country *j* is computed as

$$W_j = \frac{X_j}{\sum X_j}; \quad X_j = \varphi_j + \phi_j, \qquad \sum_{j=1}^{n} W_j = 1 \tag{64}$$

where φ_j is country *j*'s import market share commanded by the USA, or the European Union, as a proportion of the total share commanded by country *j*'s biggest five trading partners; ϕ_j is the share of country *j*'s exports to the USA, or the European Union, as a proportion of all the share of exports to country *j*'s five major trading partners.[48] That is,

$$\varphi_j = \frac{Proportion\ of\ country\ j's\ total\ imports\ emanating\ from\ USA\ (or\ European\ Union)}{Proportion\ of\ country\ j's\ total\ imports\ from\ five\ major\ trading\ partners}$$

$$\phi_j = \frac{Proportion\ of\ country\ j's\ total\ exports\ to\ USA\ (or\ European\ Union)}{Proportion\ of\ country\ j's\ total\ exports\ to\ five\ major\ trading\ partners}$$

In a way therefore, these weights measure competitiveness of the USA and the European Union as trading partners with the African countries in the index. The *Afro* index compares favorably with other potential index measures. In particular, notwithstanding the fact that more countries than the ones whose currencies are the subject of this study are included in its construction, the first differences in logs of the *Afro* index has correlations of 0.923 and 0.492, with residual and common indexes respectively, constructed following Vassalou (2000). Further, the *Afro* index is similar in principle to the US Federal Reserve Bank's Other Important Trading Partners (OITP) index. The OITP index is a trade weighted exchange rate of the US dollar against nineteen emerging markets currencies and is deemed as the indicative real exchange rate of emerging markets currencies against the US dollar. The OITP index was recently used successfully by Carrieri and Majerbi (2006) to estimate the pricing of currency risk in the emerging markets.

[48] I considered portfolio flows-based index as an alternative to the trade-based one used here; however data difficulty hampered this attempt.

CHAPTER FOUR

THE PRICING OF CURRENCY RISK IN AFRICA'S EQUITY MARKETS

4.1 Introduction

This chapter presents the summary statistics, the results from empirical investigations and the accompanying interpretations of the statistics emanating therefrom. Inferences are drawn and comparisons are made with findings from studies of a similar nature done elsewhere.

4.2 Preliminary Analysis

4.2.1 Stock Market Returns in Africa

Table 5 presents a summary of stock market index returns and their autocorrelations. Using the MSCI world market portfolio as the benchmark, the Sharpe ratios indicate the presence of abnormal risk-adjusted equity returns for many of the African countries over the sample period. The presence of abnormal returns is also observed when the emerging markets composite index is used as the benchmark portfolio. Use of the Sharpe measure is predicated on the modern portfolio theory, which assumes that returns are normally distributed and can therefore be described adequately by their mean and standard deviation. This assumption appears to be violated by the distributions of returns for all the national market indices (except Kenya) as well as the emerging markets composite and the world market portfolio indexes.

This observation conforms to emerging empirical evidence which suggests that the assumption of normality is frequently violated in asset price returns. For instance, Harvey and Siddique (2000) argue that the presence of large positive skewness in the distribution of asset returns may induce investors to hold a portfolio even if it has a negative expected return. The duo develop an asset pricing framework that gives a role to skewness, test it empirically and conclude that such a model has greater success, in explaining the negativity of the expected risk premium for the world market portfolio, than models that do not incorporate skewness. Brooks et al. (2005) remark that leptokurtosis is almost universally observed in financial asset returns irrespective of the frequency of observation. They develop and empirically test a model that provides strong evidence for the dependence of asset returns on conditional kurtosis. Leon et al. (2005) propose a GARCH-type model allowing for time-varying volatility, skewness and kurtosis. They estimate the model using two different data sets: daily returns on four foreign exchange rates series (Sterling Pound/USD, Yen/USD, German Mark/USD and Swiss Franc/USD) and five stock indices [S&P500 and NASDAQ100 (USA), DAX30 (Germany), IBEX35 (Spain), and the MEXBOL emerging market index (Mexico)]. Their results indicate significant presence of conditional skewness and kurtosis. They also find that specifications allowing for time-varying skewness and kurtosis outperform specifications that assume them constant. The presence of non-normal return behavior in African capital markets implies that conditional negative skewness in these markets should attract a premium *ex ante*, causing an increment in the desired rates of return beyond that dictated by the standard deviation of returns.

Table 5
Descriptive statistics for excess returns on market indexes

Panel A: Summary Statistics

	Mean	Std. Dev.	Sharpe Ratio	Skew	Kurt	Jarque-Bera
Botswana	0.0130	0.0575	22.57	-0.0854	0.6966	34.68***
Egypt	0.0074	0.0894	8.25	0.0010	-0.3108	71.25***
Ghana	-0.0082	0.0599	-13.73	-0.4947	3.3934	7.37**
Kenya	-0.0032	0.0795	-4.04	-0.1417	2.9421	0.54
Morocco	0.0035	0.0597	5.82	0.0494	1.1566	22.15***
Nigeria	-0.0027	0.1343	-2.02	-7.1628	76.2449	36205.26***
South Africa	0.0026	0.0885	2.95	-1.0875	2.3152	33.80***
EM Index	0.0019	0.0791	2.48	-1.2504	3.2798	41.16***
World Portfolio	-0.0005	0.0482	-0.95	-1.0385	2.2368	31.82***

Panel B: Autocorrelations

	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_{12}	ρ_{24}
Botswana	0.350‡	0.156‡	-0.020	0.000	-0.051	-0.005	-0.059	-0.036
Egypt	0.386‡	0.208‡	0.180‡	0.164‡	0.153‡	0.047	0.068	0.007
Ghana	0.380‡	0.282‡	0.172‡	0.222‡	0.127	0.181‡	-0.034	-0.042
Kenya	0.047	0.065	0.099	0.140	0.030	0.033	0.173‡	0.080
Morocco	0.077	0.122	0.089	0.046	0.056	0.107	0.142‡	0.052
Nigeria	0.068	0.023	0.078	0.071	-0.105	0.065	-0.024	-0.054
South Africa	0.069	0.034	-0.003	-0.038	-0.125	0.024	0.031	-0.130
EM Index	0.194‡	0.129	0.082	0.033	-0.003	-0.126	-0.022	0.015
World Portfolio	0.215‡	0.004	0.113	0.136	0.022	-0.058	0.069	0.008

‡ indicates that the autocorrelation is significantly different from zero. ** and *** indicates that the reported value is statistically significant at the 5% and 1% levels respectively. "Skew" and "Kurt," respectively, stand for skewness and kurtosis. The Sharpe ratio (reward to variability) is reported in percentages. US Treasury bills rates are obtained from US Federal reserve bank website. The World market equity portfolio indices, the Emerging Markets (EM) composite portfolio indices and national market indexes for Egypt, Morocco and South Africa returns are obtained from MSCI. The national market indexes for Botswana, Ghana, Kenya and Nigeria are obtained from the *Datastream* database. Monthly observations for the period 1996:12 through 2009:12 are used. All returns are in US dollars and are expressed in excess of one-month yields on US Treasury bills with maturity closest to one month; they cover the period 1997:1 to 2009:12.

For a few of the African countries, return autocorrelations are high, implying high return predictability. High positive first order autocorrelation (above 10%) is observed for Botswana, Egypt, Ghana, the world market portfolio and the emerging markets composite index. For the rest of the countries, however, first order autocorrelations are not statistically significant. Ghana exhibits the highest persistent return predictability; its autocorrelation function remains significant up to the ninth lag (0.152, not reported). Positive autocorrelation of security/portfolio returns implies that below-average returns tend to be followed by other below-average returns. If negative returns can be predicted, we would expect to see them in the months that follow large negative returns. Similar implications can be drawn in respect of above-average returns.

The high autocorrelation of returns may be indicative of low levels of informational efficiency in the sample markets.[49] Many studies have established a linkage between autocorrelations in stock market returns and market efficiency. Mech (1993) presents evidence that portfolio return autocorrelation is not caused by time-varying expected returns, non-trading, stale limit orders, or market-maker trading strategies. He attributes portfolio return autocorrelation to transaction costs, which cause delays in price adjustment. He concludes that markets are inefficient, in that prices do not always fully reflect all available information, but not irrational because there are costs that prevent the mispricing from being exploited. Froot and Perold (1995) find that autocorrelation in short-run index returns appear to be the result of inefficient processing of market-wide information and that technological and institutional improvements in information processing removes much of the autocorrelation. On their part, Campbell et al. (1993) attribute high first order autocorrelation in returns to trading volumes. Using daily return data, they find that the first daily autocorrelation of stock returns is lower on high-volume days than on low-volume days and tends to decline with volume.

Table 6 displays the correlation matrix for excess returns on equity market indexes. The correlations among African stock markets are generally low, suggestive of possible low levels of

[49] A word of caution is necessary here. Although high autocorrelation may be associated with market inefficiency, the autocorrelation function is not a formal test for market efficiency and any "hunches" emanating from observed high autocorrelation need be formally investigated and tested to justify any statistical inferences made.

contagion. Further, very low correlations are observed between the USA and African markets, presenting rich opportunities for portfolio and geographical diversification. A notable exception is the South African market, which exhibits a high correlation of returns with the USA market (59.8%) and with the world market portfolio (69.1%). This can be given two interpretations: firstly, that of the African stock markets, the South African market is probably the most accessible and known to foreign investors, and secondly, that it is, accordingly, the most integrated with markets in the rest of the world. The second conjecture is tested unconditionally, in this study, by the three-factor asset pricing model. Results are presented in section 4.3.3.

Table 6
Correlation matrix for excess returns on equity market indexes

	Botswana	Egypt	Ghana	Kenya	Morocco	Nigeria	S. Africa	USA	EM
Egypt	0.0708								
Ghana	0.0232	0.0507							
Kenya	0.1152	0.2261**	-0.0397						
Morocco	0.0513	0.2553**	0.1898**	0.0654					
Nigeria	0.1498	0.0410	-0.1258	0.0110	0.0897				
S. Africa	0.1170	0.3978**	-0.0547	0.2318**	0.2584**	0.0407			
USA	0.1718**	0.3347**	-0.0048	0.3062**	0.1221	-0.0063	0.5983**		
EM	0.0821	0.4869**	-0.0204	0.2959**	0.1757**	0.1339	0.8249**	0.7608**	
World	0.2060**	0.3832**	-0.0065	0.3403**	0.1892**	0.0477	0.6907**	0.9579**	0.8359**

** Indicates that correlation is significant at the 5% level. "EM" stands for the MSCI Emerging markets composite; "World" represents the MSCI world market equity portfolio. Treasury bill rates are obtained from US Federal reserve bank website. The World market equity portfolio indices, the Emerging Markets (EM) composite portfolio indices and national market indexes for Egypt, Morocco and South Africa returns are obtained from MSCI. The national market indexes for Botswana, Ghana, Kenya and Nigeria are obtained from the *Datastream* database. Monthly observations for the period 1996:12 through 2009:12 are used. All returns are in US dollars and are expressed in excess of one-month yields on US Treasury bills with maturity closest to one month; they cover the period 1997:1 through 2009:12.

4.2.2 Portfolio-level Stock Returns – South Africa

Table 7 displays the summary statistics, at firm level, from the Johannesburg Stock Exchange. South Africa was the only country for which firm-level data was long enough, and spanned over a number of firms. Data was availed for the period between December 1994 and November 2008 only those firms with data spanning over the entire period available were selected. Return characteristics are presented for four equal-weighted portfolios formed and ranked on the basis of market capitalization as at the end of December 2008. All data are from the *Datastream*.

Table 7
Descriptive statistics for firm-level excess stock returns

Panel A: Summary Statistics

	Obs	Mean	Std. Dev	Sharpe Ratio	Skew	Kurt	Jarque-Bera
Portfolio 1	167	-0.0106	0.0786	-13.49	-0.574	4.795	31.59***
Portfolio 2	167	-0.0012	0.0852	-1.41	-0.863	4.953	47.27***
Portfolio 3	167	-0.0046	0.0799	-5.76	-1.122	6.680	129.27***
Portfolio 4	167	-0.0022	0.0836	-2.63	-0.847	5.786	73.98***
Average		-0.0021	0.0819	-5.68			

Panel B: Autocorrelations

	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_{12}	ρ_{24}
Portfolio 1	0.257‡	0.042	0.011	-0.034	0.007	0.107	0.075	-0.125
Portfolio 2	0.277‡	0.120	0.034	-0.088	-0.021	-0.033	-0.003	0.010
Portfolio 3	0.221‡	0.023	-0.048	-0.011	0.003	0.010	0.002	-0.074
Portfolio 4	0.115	-0.037	-0.063	-0.144‡	0.005	0.000	-0.016	-0.061

‡ indicates that the autocorrelation is significantly different from zero. ** and *** indicates that the reported value is statistically significant at the 5% and 1% levels respectively. "Skew" and "Kurt," respectively, stand for skewness and kurtosis. The Sharpe ratio (reward to variability) is reported in percentages. Treasury bill rates are obtained from US Federal reserve bank website. All firm-level data are obtained from the *Datastream* database. Monthly observations are for the period 1994:12 through 2008:12. All returns are in US dollars and are expressed in excess of one-month returns on US Treasury bills with maturity closest to one month; they cover the period 1995:1 through 2008:11.

The Sharpe ratios are, on the average, higher than similar market-level ratios for the USA and the world market equity portfolio, again indicating the potential for diversification benefits for foreign investors. The autocorrelation functions show significant serial return dependency in the first month for most of the portfolios. The return dependency largely dies out by the third month. The data does not also bring out any seasonal influence on return behavior. This is not surprising given that any such influences are idiosyncratic risk sources, which are expected to be eliminated, or substantially reduced, by blending assets together in portfolios. Like in the case of market-level data, the Jarque-Bera statistic rejects, for all the portfolios, the hypothesis that returns are described by the normal distribution.

4.2.3 The Behavior of Foreign Exchange Rates

Table 8 depicts summary statistics for the change in bilateral real exchange rates of the sampled African currencies with respect to the US dollar. Exchange rates are defined as the number of units of the foreign currency per unit of the US dollar. With the exception of Morocco, the mean values of change in exchange rates are all positive, indicating that the US dollar was appreciating against the African currencies during the study period. The variability in exchange rate changes is fairly low with the standard deviation ranging from 2.19% (Egypt) to 11.26% (Nigeria).

Table 8
Descriptive statistics for change in real exchange rates

Panel A: Summary Statistics

	Obs	Mean	Std Dev	Skew	Kurt	Jarque-Bera
Botswana	156	0.0086	0.0364	1.4108	6.470	130.01***
Egypt	156	0.0067	0.0219	3.5202	28.000	4384.56***
Ghana	156	0.0249	0.0309	1.5945	3.044	66.12***
Kenya	156	0.0085	0.0307	1.8307	9.395	352.97***
Morocco	156	-0.0011	0.0245	0.4922	2.545	7.65**
Nigeria	156	0.0196	0.1126	11.5213	139.726	124962.77***
South Africa	156	0.0059	0.0491	0.7603	1.973	21.89***
Afro Index	156	0.0095	0.0297	8.0536	84.528	44891.11***

Panel B: Autocorrelations

	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_{12}	ρ_{24}
Botswana	0.082	0.104	-0.046	0.019	0.005	-0.137	-0.085	0.052
Egypt	0.005	0.117	0.135	0.047	-0.008	0.031	0.011	-0.086
Ghana	0.674‡	0.539‡	0.439‡	0.417‡	0.278‡	0.162‡	0.116	0.052
Kenya	0.172‡	-0.103	0.102	0.099	-0.042	-0.071	0.101	0.100
Morocco	0.129	0.006	0.051	-0.148‡	-0.034	0.069	-0.105	-0.012
Nigeria	0.001	0.026	0.006	-0.030	-0.019	-0.011	-0.018	-0.002
South Africa	0.134	0.055	0.013	0.008	-0.039	-0.128	-0.054	0.012
Afro Index	0.137	0.095	-0.021	-0.042	0.002	-0.037	0.008	0.063

‡ Indicates that the autocorrelation is significantly different from zero. ** and *** indicate that the reported figure is statistically significant at the 10%, 5% and 1% levels respectively. "Obs" is short form for number of observations; "Skew" and "Kurt" respectively stand for skewness and kurtosis. End-of-month real foreign exchange rates for all countries are with respect to the US dollar and are computed from nominal exchange rates and consumer price index data obtained from the International Financial Statistics. Changes in foreign exchange rates are computed as the first differences in the natural logarithms of the ending and the beginning rates. Monthly observations for the period 1997:1 through 2009:12 are used.

The estimated autocorrelation coefficients are generally small in absolute terms. For the first lag, the coefficients are statistically significant, at the 5% level, only for Ghana and Kenya, and all of them are positive. Ghana exhibits significant serial correlation in exchange rate returns up to the ninth lag (0.173). Overall, exchange rate changes appear not to be predictable in the sampled African markets. Nevertheless, one cannot conclude on the basis of these results that the foreign exchange markets are efficient since the autocorrelation function "may be an inappropriate indicator of market efficiency" (Cornell and Dietrich, 1978).

The Jarque-Bera statistics suggest that the data generating process for foreign exchange rates in these markets cannot be described by the normal distribution. This is consistent with the numerous studies that have found the distribution of foreign exchange returns to be non-normal. For example, McFarland et al. (1982), document results suggesting that the distribution of foreign exchange rate changes is too complex to be summarized neatly by either the normal or the non-normal stable distributions. The "abnormal" behavior of the observed Nigerian naira-US dollar and the Egyptian pound-US dollar nominal exchange rate returns may emanate from the fact that "official," rather than "market determined" rates of exchange are reported for both countries. In the case of Egypt, nominal exchange rates remain constant for long periods of time before adjustments (largely devaluations) are made. This has the implication that changes in the real rates of exchange remain very low for long periods, hence the low standard deviation of real exchange rate returns.

In the case of Nigeria, some devaluation events are observed over the study period. The first devaluation causes the rate of exchange to jump from N21.886 at the end of December 1998 to N85.570 at the end of January 1999, a change of 136 percent. In the second instance, the exchange rate hovers around N117 to the US dollar for one year between December 2007 and November 2008 then suddenly moves to N132.56 to the US dollar. The combined influence of these irregularities is witnessed in the huge standard deviation of exchange rate returns (for Nigeria) and the conspicuously large Jarque-Bera statistic for both Egypt and Nigeria. These characteristics also find their way into the *Afro* index.

4.2.4 Instrumental Variables and Risk Factors

In this section, I conduct further tests to establish the suitability and appropriateness of the chosen conditioning variables. Also included is a preliminary analysis of the relationship between the conditioning variables and the risk factors that they scale in the GMM estimation. Table 9 displays the summary statistics for the conditioning variables – lagged values of gross returns on the world equity portfolio (WLR); and Hodrick-Prescott (1997) filtered lagged values of the MSCI world equity portfolio dividend yields (WDY), USA term premia (UTP), and Eurodollar deposit rates (UDR). From the table, it is seen that, save for WLR, serial dependence of the instrumental variables is high and persistent. This can be explained by the filtering process, which removes the trend component of the series; the remaining cyclical component are time-dependent, a fact that is well captured by their persistent autocorrelation. This feature is desirable as it enhances the ability of the instrumental variables to capture the time varying properties of the risk-factors that they condition. It is also observed that all of the variables, except UTP cannot be described by the normal distribution; however, this is not expected to affect the outcome of tests performed on the data.

Table 9
Descriptive statistics for instrumental variables

Panel A: Basic Characteristics

	Obs	Mean	Std Dev	Skew	Kurt	Jarque-Bera
WLR	156	1.0020	0.0482	-1.049	5.191	59.81***
WDY	156	0.0000	0.0002	1.967	10.464	462.72***
UTP	156	0.0000	0.0004	-0.185	3.615	3.35
UDR	156	0.0000	0.0005	0.9278	4.754	42.38***

Panel B: Autocorrelations

	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_{12}	ρ_{24}
WLR	0.214[a]	0.000	0.116	0.136	0.019	-0.057	0.069	0.004
WDY	0.849[a]	0.626[a]	0.439[a]	0.247[a]	0.041	-0.135	-0.250[a]	0.012
UTP	0.877[a]	0.706[a]	0.587[a]	0.469[a]	0.340[a]	0.240[a]	-0.460[a]	0.179[a]
UDR	0.825[a]	0.650[a]	0.524[a]	0.402[a]	0.309[a]	0.230[a]	-0.089	-0.145[a]

USATPL, EURDRL and WLDLMR represent, respectively, the lagged values of the USA term premia, Eurodollar deposit rates and world market excess equity returns. Data runs from January 1997 through December 2009. *** and ** represents statistical significance at 1 and 5 percent levels respectively. [a] means the autocorrelation is significantly different from zero.

Table 10 reports the correlation matrix for the risk factors filtered and instrumental variables. The correlation coefficient between instrumental variables, *UTP* and *UDR* is quite high (-0.6092); the rest of coefficients are moderate in magnitude but statistically significant in some cases. This may introduce multicollinearity in the estimation process; to avoid the multicollinearity problem, implementation of the model is done by conditioning the risk factors with one instrumental variable at a time. The correlation coefficient between WLR and WLD is 22% and statistically significant; this is largely expected as the former variable is the lagged version of the latter. In the remaining cases, the coefficients are low; this establishes the orthogonality conditions required for GMM estimation.

Table 10
Correlation structure for instrumental variables and risk factors

	WLR	WDY	UTP	UDR	WLD
WDY	-0.3656**				
UTP	-0.0447**	0.0710			
UDR	-0.1660**	-0.1840**	-0.6092**		
WLD	0.2159**	0.0667	-0.0298	-0.1696**	
NXR	0.0009	-0.0225	-0.0010	0.0724	-0.3002**

The instrumental variables *WLR*, *WDY*, *UTP* and *UDR* represent, respectively, the lagged values of the gross returns on the MSCI world equity portfolio, the Hodrick-Prescott-filtered lagged values of the MSCI world equity portfolio dividend yield, USA term premia, and Eurodollar deposit rates. The risk factors are the gross returns on the MSCI world equity portfolio (*WLD*) and the equal-weighted nominal foreign exchange rates (*NXR*) and Data runs from January 1997 through December 2009 representing 156 data points for all variables. ** denotes that the reported correlation coefficient is significant at the 5 percent level.

4.3 Empirical Results for the Unconditional Multi-beta Asset Pricing Model

The multi-factor asset pricing models presented in section 3.3.1 are estimated as seemingly unrelated regression models using the Generalized Method of Moments procedure implemented through the SAS 9.2 package. All available information is used to estimate both the factor loadings and the associated risk premia in one step. As demonstrated by Burmeister and McElroy (1988), joint estimation of both the factor sensitivities and the associated risk premia generates more efficient estimates than those obtained through the factor analysis techniques or the traditional two-step procedure of Fama and MacBeth (1973). Results are presented separately for the two-factor and the three-factor models.

4.3.1 Tests Results for the Two-factor Models

1. Two-factor Foreign Exchange Risk Model

Table 11 summarizes tests results for the model in equation (10). This is the first model used to test for the pricing of foreign exchange risk in African capital markets. Excess returns on the MSCI world market equity portfolio and returns on the dollar real exchange rate index for Africa are assumed to be the only variables explaining the cross section of returns on the aggregate markets, as proxied by the appropriate equity market indexes, of the sampled African countries. Inclusion of the exchange rate term in the multi-beta asset pricing structure is motivated by a number of international asset pricing models (Solnik, 1974; Stulz, 1981; Adler and Dumas, 1983; Ikeda, 1991) which have suggested that fluctuations in foreign exchange rates constitute an important risk source for international investors.

Table 11
GMM regression results for the two-factor currency risk model

$$r_{it} = \beta_{iw} r_{wt} + \beta_{is} r_{st} + \lambda_w \beta_{iw} + \lambda_s \beta_{is} + \varepsilon_{it}$$

Country	β_W	β_S	RMSE	Adj-R^2	APE
Botswana	0.2686*** (3.45)	-0.3395** (-2.55)	0.0566	0.0321	0.0122
Egypt	0.6882*** (5.24)	-0.0694 (-0.26)	0.0830	0.1375	0.0048
Ghana	0.0988 (0.92)	-0.0454 (-0.31)	0.0610	-0.0370	-0.0086
Kenya	0.5664*** (3.32)	-0.1622 (-0.95)	0.0751	0.1072	-0.0053
Morocco	0.2765** (2.27)	-0.1563 (-0.81)	0.0587	0.0336	0.0025
Nigeria	0.0022 (0.01)	-3.6392*** (-6.36)	0.0839	0.6097	-0.0000
S. Africa	1.2833*** (10.78)	-0.2944 (-1.04)	0.0636	0.4835	0.0022

λ_W	λ_S	J-Statistic
0.394 (1.08)	0.075 (0.41)	8.1906 [0.1460]

The table uses data covering the period 1997:1 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses),which The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; $r_{wt}$is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index. β_{iW} and β_{iS} are, respectively, the sensitivities of African market equity returns to the world portfolio and foreign exchange rate factors and λ_W and λ_S are the respective risk premia, in percentages, for the two factors. RMSE is the root mean squared error and Adj-R^2 is the adjusted coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollar and expressed in excess of one-month yields on the US Treasury bills closest to one-month maturity.

The two macroeconomic factors proved important in explaining the excess equity market returns in most of the countries studied. The betas relating to the excess return on the world equity market index range from 0.0022 for Nigeria to 1.2833 for South Africa. All the world equity market betas are positive and, with the exception of Ghana and Nigeria, statistically significant. Thus, equity market indexes in these countries have a tendency of moving in the same direction as the world market equity portfolio. The betas in respect of the foreign exchange rates factor are all negative and range from -3.6392 for Nigeria to -0.0454 for Ghana. On the average, therefore, there is an inverse relationship between foreign exchange rate movements and movements in the equity markets of all the countries investigated. Because real exchange rates are stated such that they rise with an appreciating US dollar, these results imply that declining excess returns, in US dollar terms, are expected from the equity markets of these countries when the US dollar appreciates. In absolute terms, exchange rate betas are generally smaller than those of the world equity market portfolio and only two of them (Botswana and Nigeria) are statistically different from zero. This suggests that the exchange rate factor only exerts a weak influence on stock market returns in Africa's equity markets.

The estimated average unconditional monthly risk premia are, respectively, 0.394 percent and 0.075 percent, for the world equity market portfolio and foreign exchange rates. Neither is statistically different from zero. Therefore, the results do not support that either one of the risk factors is priced in Africa's stock markets. This finding contrasts recent evidence provided by Carrieri and Majerbi (2006) that unconditional foreign exchange risk pricing is alive and well in emerging stock markets. These findings can be explained in several ways.

First, because of home bias among international investors, fear of political uncertainties, institutional inadequacies, information asymmetries, and poor regulatory enforcement and weak accounting framework, foreign portfolio investors have apparently ignored the African markets despite the many incentives offered and the many reform efforts put in place by many a government in the continent. Consequently, it would appear that African capital markets in general and equity markets in particular, are largely driven by domestic investors whose returns are denominated in local currencies and therefore shielded from the influence of changes in foreign exchange rates and developments in the world markets. Second, perhaps the two-factor model is simply not suitable to describe the cross-section of equity returns in the continent. To test the last conjecture, we perform a number of diagnostic checks on the two-factor model. First, the *J*-statistic of Hansen (1982), a test of over-identifying restrictions, fails to reject the model's goodness-of-fit at any conventional level of significance, yielding a right-tailed *P*-value of 0.1460. However, it is clear that the model is sufficiently weak that it only marginally escapes rejection, a first indication that the two-factor model could do with some improvement.

Next, the average pricing error (APE) is computed for each country. The APE is defined as the difference between the average returns and the expected returns provided by the model, evaluated at the sample estimates (Ferson and Harvey, 1994). A negative APE implies that the model tends to overprice securities. For this model, the APE is positive for four countries and negative for the remaining three. The APE is very large for Botswana, showing an absolute error of 1.22 percent in unexplained monthly average excess returns. Following Botswana is Ghana, with an average absolute monthly pricing error of 1.63 percent. Nigeria has the lowest average pricing error of 0.002 percent (rounded to zero in the table) per month. Sawyer et al. (2010) explain that the pricing error in the multi-beta asset pricing structure is neither a sampling error nor a misspecification error; it is a theoretical pricing error that relates to expected returns and which is present even when the factor structure is properly specified. However, it must be pointed out that the pricing error tends to be high for incorrectly specified models. The two-factor model tends to underprice African equity returns and the mispricing error appears substantial in many cases. This motivates the search for an alternative specification.

Third, the Root Mean Square Error (RMSE) is also computed for each country. The RMSE is useful in determining the change in variation of returns, as measured by the standard deviation, resulting from the application of the model. Table 4 reported the standard deviations of index returns before the model's application. From that table and the reported RMSE, one observes that the model results in an increment in the standard deviation of returns of 3.81 percent per annum $\left[(0.0599 - 0.0610) \times \sqrt{12}\right]$ for Ghana's market index. In the rest of the cases, there is a reduction in annual return variability as follows. Leading the pack is Nigeria with an annual standard deviation reduction of 17.47 percent. Nigeria is followed by South Africa, 8.62 percent; Botswana, 3.28 percent; Egypt, 2.20 percent; Kenya, 1.51 percent and Morocco, 0.36 percent. An important implication that can be drawn from the RMSE is that the two-factor model does not fully account for the variation in excess returns in Africa's equity markets.

Finally, the coefficients of determination indicate that, with the exception of Nigeria (60.97%) and South Africa (48.35%), the two factors do not perform very well in explaining changes in the excess returns on equity market indices across Africa. Still, the adjusted R-square for Nigeria may be explained by its commanding position in the African real exchange return index: Nigeria leads the African continent in bilateral trade with the USA. Low coefficients of determination are a common feature of tests of this nature (see, for example Choi and Rajan, 1997; Carrieri and Majerbi, 2006) as well as in conditional asset pricing tests (Dumas and Solnik, 1995). The pitfalls discussed above motivate the use of the three-factor model specification, whose results are presented in section 4.3.2.

2. Two-factor Market Segmentation Model

This model assumes that excess returns in Africa's equity markets are explained by the world market and the local market risk factors. Partial market segmentation models, such as the ones employed here, are inspired by various authors who have argued that the flow of international money is, to a large extent, restricted by various pecuniary and non-pecuniary barriers. The partial segmentation hypothesis is tested here through two-factor and three-factor models akin to Choi and Rajan (1997). Table 12 reports empirical results for the model in equation (11).

Table 12
GMM regression results for the two-factor market segmentation model

$$r_{it} = \beta_{iw} r_{wt} + \beta_{im} r_{mt} + \lambda_w \beta_{iw} + \lambda_m \beta_{im} + \varepsilon_{it}$$

Country	β_W	β_m	RMSE	Adj-R^2	APE
Botswana	0.2557*** (3.02)	-0.1619 (-1.65)	0.0563	0.0433	0.0073
Egypt	0.6731*** (5.28)	0.5491*** (4.50)	0.0787	0.2248	0.0073
Ghana	0.0822 (0.78)	0.0090 (0.08)	0.0611	-0.0377	-0.0091
Kenya	0.4414** (2.55)	0.0230 (0.95)	0.0757	0.0918	-0.0085
Morocco	0.2499** (2.08)	0.0106 (0.11)	0.0589	0.0280	0.0004
Nigeria	0.0594 (0.31)	0.4719** (2.00)	0.1329	0.0218	0.0034
S. Africa	1.2751*** (16.90)	0.9626*** (8.06)	0.0503	0.6774	0.0006
	λ_W	λ_m	J-Statistic		
	1.276* (1.80)	-1.476* (-1.78)	7.4608 [0.1886]		

The table uses data covering the period 1997:1 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; $r_{wt}$is the demeaned excess return on the world market equity portfolio; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW} and β_{im} are, respectively, the sensitivities of African market equity returns to the world portfolio and pure local market risk factors; λ_W and λ_m are the respective risk premia, in percentages, for the two factors. RMSE is the root mean squared error and Adj-R^2 is the adjusted coefficient of determination; APE is the average pricing error. All index returns are measured in US dollars and expressed in excess of one-month yields on the US Treasury bills closest to one-month maturity.

As seen in the table, the unconditional two-factor model is not rejected by the monthly excess index returns data: the *J*-statistic is 7.4608 with a *p*-value of 0.1886. The monthly rewards for world market and pure local market risk factors are 1.276 percent and -1.476 percent respectively; both are statistically significant at the 10 percent level. The local market risk factor is assumed to capture all risk sources idiosyncratic to each African country market. Chen et al. (1986) explain that the market factor is basically a summary measure of the various fundamental economic factors. The factor is constructed as the component of excess MSCI emerging markets equity portfolio index returns orthogonal to the excess returns on the MSCI world market equity portfolio index. These results suggest that Africa's equity markets can be described as partially segmented/integrated during the study period. The negative sign of the local market risk premium deserves further attention. The equity risk premium is the extra return that equity holders expect to achieve, on average, over risk-free assets, such as Treasury bills. The market efficiency assumption, which informs asset pricing models such as the one tested here, implicitly rules out situations where risky assets (common stocks in this case) have lower expected returns than risk-free assets; this is the situation implied by the negative risk premium documented here.

However, negative risk premia have been documented in many other empirical investigations of asset pricing models, both in advanced and emerging equity markets (Fama and Schwert, 1977; Jorion, 1991; Choi and Rajan, 1997; Choi et al, 1998; Carrieri and Majerbi, 2006). Additionally, tests of the positivity restriction in asset pricing models have provided evidence of negative equity risk premia in various markets. Tests of the positivity of the US market risk premium using 188 years of annual data by Boudoukh et al., (1993) find evidence that the expected return on the US market is less than the risk free rate in some periods. Using the same data but a different methodology, Boudoukh et al. (1997) examine the characteristics of the US *ex ante* risk premium conditioned on the slope of the US term structure and find evidence of negative US *ex ante* risk premiums. Ostdiek (1998) directly assesses the non-negativity restriction in international asset pricing models. The evidence indicates that the *ex ante* world market (proxied by the MSCI dollar-denominated world portfolio) risk premium can be negative. The results are robust to market proxies that are hedged and unhedged with respect to currency risk. The author also uses a local currency-denominated portfolio as a proxy to allow a test of the risk premium of the underlying market portfolio of risky assets. The evidence again indicates that the *ex ante* risk premium on the market portfolio of risky assets is not always positive. Theoretically, negative equity risk premia can be explained by high Treasury bills rates especially during times in which the term structure is downward-sloping (Boudoukh et al., 1993).

Coefficient estimates for the world market factor are all positive and many of them are statistically significant. The absolute magnitudes range from 0.0594 for Ghana to 1.2751 for South Africa. The positive values indicate that equity returns in all the sample countries vary directly with returns on the world market equity portfolio. Four (and one) of the seven countries have statistically non-significant (and negative) local market factor coefficient estimates. In all the seven countries, but Nigeria, coefficient estimates for the local market factor are smaller in absolute magnitudes than those of the world market factor: thus, the residual local market factor appears to play a less important role than the world market factor in influencing equity market index returns in the representative African bourses.

The ability of the two factors to explain return variation in each of the countries appears weak. The degrees-of-freedom-adjusted coefficients of determination (R-square) are less than 10 percent for all the countries except Egypt and South Africa. Similarly, the root mean square errors (RMSE) are large and range from 0.0503 (South Africa) to 0.1329 (Nigeria). Also still large in size are the average pricing errors, implying that a lot of variations in excess returns still remain explained: they are positive for five countries and negative for the remaining two.

4.3.2 Tests Results for the Three-factor Model

The three-factor model tests for the pricing of foreign exchange rate risk as well as integration of African equity markets with the world equity markets. Thus, in addition to the world market and the pure local market risk factors, this model includes residuals of the projection of the changes in the *Afro* real exchange rate index on excess returns on the world market equity portfolio index, the emerging markets composite portfolio index and a constant: this is the currency risk factor.

The test results for the model in equation (12) are presented on Table 13. Having all the three factors in one model appears to have greatly improved the model specification. The *J*-statistic now produces a large right-tail *p*-value of 0.5359. Thus, the hypothesis that the three factor model does not adequately describe the return-generating process in African equity markets is rejected by the data. Inclusion of the two factors appears to diminish the strength of the world market factor in explaining excess market index returns. Consequently, world market factor sensitivities have reduced in size and now range from -0.2659 for Nigeria to 1.25 for South Africa. As is clear, some have changed signs from positive to negative further indicating the reduced influence of the variable on excess returns. With the exception of Ghana, all exchange rate coefficients are still negative, implying that even with the influence of the local market factor, excess dollar returns decline with US dollar appreciations. The exchange rate coefficients range from -3.4196 for Nigeria to 0.0150 for Ghana and are significant, at the 1 percent level, only for Botswana and Nigeria. The rest are not statistically different from zero.

Table 13
GMM regression results for the three-factor model

$$r_{it} = \beta_{iw}r_{wt} + \beta_{im}r_{mt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_m\beta_{im} + \lambda_s\beta_{is} + \varepsilon_{it}$$

Country	β_W	β_S	β_m	RMSE	Adj-R^2	APE
Botswana	0.1673* (1.92)	-0.4361*** (-2.84)	-0.1673* (-1.80)	0.0545	0.1029	-0.0028
Egypt	0.7146*** (6.14)	-0.0239 (-0.08)	-0.5370*** (-5.25)	0.0788	0.2216	-0.0049
Ghana	-0.0421 (-0.39)	0.0150 (0.12)	0.0237 (0.23)	0.0607	-0.0263	-0.0054
Kenya	0.1930 (1.34)	-0.0509 (-0.35)	0.0859 (0.67)	0.0777	0.0450	-0.0068
Morocco	0.1515 (1.43)	-0.1816 (-0.88)	0.0196 (0.22)	0.0590	0.0240	-0.0020
Nigeria	-0.2659* (1.80)	-3.4196*** (-4.64)	0.4813*** (3.03)	0.0839	0.6095	0.0006
S. Africa	1.2500*** (16.50)	-0.2251 (-0.66)	0.9733*** (8.14)	0.0504	0.6758	-0.0005
	λ_W	λ_S	λ_m	J-Statistic		
	3.95** (2.52)	-0.268 (-0.71)	-4.813** (-2.46)	6.0334 [0.5359]		

The table uses data covering the period 1997:1 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; $r_{wt}$is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index and the pure local market factor; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW}, β_{iS} and β_{im} are, respectively, the sensitivities of African market equity index returns to the excess return on the world market equity portfolio index, foreign exchange rate changes and pure local market factor. λ_W, λ_S and $\lambda_m$are the respective risk premia, in percentages, for the three factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollars and are expressed in excess of one-month yields on US Treasury bills closest to one-month maturity.

The local market factor sensitivities are statistically significant for four of the seven countries. The magnitudes vary from very large negative and significant (-0.5370 for Egypt) to very large positive and significant (0.9733 for South Africa). Others are very small and statistically insignificant (0.0196 for Morocco). This huge variation may be attributed to the construction of this factor. Recall that, due to the absence of a widely quoted African equity market return index, the local market factor was proxied by the residuals from the projection of the emerging markets equity portfolio index on a constant and the world market equity portfolio. This measure may not be the appropriate local market factor for some of the countries, such as Kenya, Ghana and Nigeria, which are yet to be classified under the emerging markets category by MSCI Barra. To establish the appropriateness of this factor, the MSCI Africa Frontier Markets Index is later used in sensitivity analysis. The results are reported in section 4.3.4.

A brief explanation of the negative beta signs for the world market portfolio and the local market risk factors is in order. The beta coefficient is interpreted in the asset pricing literature as the sensitivity of a security's (or, in our case here, specific country equity portfolio's) returns to changes in the factors named. Theoretically, if the equity markets in the individual countries examined are positively correlated (or high cointegrated) with the MSCI world market equity portfolio as well as the MSCI emerging markets equity portfolio, used in this study to generate the idiosyncratic local market factor series, one would expect *a priori*, that the beta coefficients to be all positive. In that case, the negative beta coefficients may be viewed as having wrong signs. However, given the composition of the two MSCI portfolios,[50] which do not incorporate many of the countries in Africa, it is possible that the assumption of cointegration may not be held for some of the countries in my sample. Thus, the negative coefficients need not be interpreted as incorrect signs. Further, it is important to note that similar studies in the literature (e.g., Carrieri and Majerbi, 2006) have also reported negative world market portfolio beta coefficients for countries, such as India.

[50] The *MSCI Emerging Markets Index* consists of the following 21 emerging market country indices: Brazil, Chile, China, Colombia, Czech Republic, Egypt, Hungary, India, Indonesia, Korea, Malaysia, Mexico, Morocco, Peru, Philippines, Poland, Russia, South Africa, Taiwan, Thailand, and Turkey. The *MSCI World Index* consists of the following 24 developed market country indices: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Hong Kong, Ireland, Israel, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, the United Kingdom, and the United States. (Source: http://www.msci.com/products/indices/tools/index.html#EM).

The estimated unconditional monthly risk premia are 3.95 percent for the world market factor, -0.268 percent for the foreign exchange rate factor and -4.813 percent for the local market factor. The local and world market factor premia are both statistically significant at the 5 percent level. Thus, the evidence suggests that Africa's equity markets are partially segmented, partially integrated. Although bigger in magnitude and with a higher *t*-statistic than under the two-factor model case, foreign exchange risk premium is still not statistically different from zero. Thus, the three factor model fails to reject the hypothesis of a non-priced foreign exchange risk factor in Africa's equity markets. This finding is consistent with Choi et al. (1998) who find, using a three-factor model that includes a local market factor and an interest rate factor, that multilateral trade-weighted exchange risk does not command a risk premium in the Japanese stock market.

Similar to the two-factor model, some diagnostics are performed on the three-factor specification as well. First, in comparison with results from the two factor models, the average pricing error declines for Botswana, Ghana and South Africa, remains steady for Egypt and Nigeria and increases for Kenya and Morocco. The increment for Kenya is particularly substantial, perhaps a further indication that the emerging markets index does not accurately capture the country's idiosyncratic risks. The Root Mean Square Error however increases for Kenya and Morocco, remains steady for Nigeria and declines for the rest of the countries, further testimony that the three-factor model prices assets in these markets better than the two-factor model. Finally, intuition from the adjusted coefficient of determination shows that the addition of a third factor does not introduce a lot of new information to the return-generating process for Nigeria's stock market index. However, the proportion of excess stock market index returns explained by the three-factor model relative to the two-factor model is enhanced for all the other countries, except Morocco, which records a decline. This is a further indication that the three-factor model performs relatively better than the two-factor models.

4.3.3 The South African Case

The above models were also tested at the portfolio level using data from South Africa – the only country that had usable firm-level data available from the *Datastream* database. Data are obtained for the period between June 1995 and December 2008, yielding a total of 161 excess return observations for each of the 78 firms with data spanning over the sample period. Using market capitalization statistics, the firms are grouped into four size-based portfolios, the lowest market capitalization firms falling into the first portfolio, and so on. All returns are expressed in US dollars. Table 14 displays GMM regression results for the two- and three-factor models.

Table 14
GMM regression results for South Africa's portfolios

Panel A: *Two-factor Model* $r_{it} = \beta_{iw}r_{wt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_s\beta_{is} + \varepsilon_{it}$

Portfolio	β_W	β_S	RMSE	Adj-R^2	APE
Portfolio 1	0.5652***	-1.0491**	0.0619	0.4012	-0.0015
	(5.81)	(-7.68)			
Portfolio 2	0.8351***	-1.0768***	0.0632	0.4673	-0.0018
	(7.74)	(-8.29)			
Portfolio 3	0.7810***	-1.1626***	0.0532	0.5683	0.0001
	(7.14)	(-10.74)			
Portfolio 4	0.8545***	-1.1971***	0.0558	0.5711	-0.0001
	(7.10)	(-9.76)			
	λ_W	λ_S	J-Statistic		
	4.092*	3.095*	0.5525		
	(1.77)	(1.72)	[0.7586]		

Panel B: *Three-factor Model* $r_{it} = \beta_{iw}r_{wt} + \beta_{im}r_{mt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_m\beta_{im} + \lambda_s\beta_{is} + \varepsilon_{it}$

Portfolio	β_W	β_S	β_m	RMSE	Adj-R^2	APE
Portfolio 1	0.5710***	-0.7312***	0.7092***	0.0595	0.4466	-0.0010
	(5.99)	(-6.33)	(6.85)			
Portfolio 2	0.8584***	-0.8189***	0.6716***	0.0617	0.4935	-0.0012
	(7.11)	(-5.61)	(6.01)			
Portfolio 3	0.7907***	-0.7787***	0.8159***	0.0485	0.6410	0.0003
	(7.49)	(-7.58)	(10.35)			
Portfolio 4	0.8553***	-0.6962***	0.9453***	0.0472	0.6930	-0.0004
	(9.38)	(-5.55)	(14.39)			
	λ_W	λ_S	λ_m	J-Statistic		
	4.093*	2.732	-1.875	0.3561		
	(1.74)	(1.20)	(-1.07)	[0.5507]		

The table uses data covering the period 1995:7 to 2008:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th portfolio; r_{wt} is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the South African rand-US dollar real exchange rate, orthogonal to the world portfolio index and the pure local market factor; r_{mt} is the pure local market factor, constructed as the excess return on the South African equity market index orthogonal to the excess return on the world portfolio index. β_{iW}, β_{iS} and β_{im} are, respectively, the sensitivities of portfolio returns to the world market equity portfolio index, foreign exchange rate changes and pure local market factor. λ_W, λ_S and λ_m are the respective risk premia, in percentages, for the three factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollars and expressed in excess of one-month yields on US Treasury bills closest to one-month maturity.

Results for the two-factor model (Panel A) show that all the beta coefficients for the world equity market factor are positive, significant at the 1 percent level, and have a tendency to increase with market capitalization. They range from 0.5652 for the low market capitalization portfolio to 0.8545 for the large market capitalization portfolio. Thus, larger firms appear to be more sensitive to changes in the world equity market than do smaller firms. On the average, therefore, an increase in return on the world market equity portfolio is associated with an increase in portfolio returns in South Africa's equity market.

The existence of an inverse relationship between foreign exchange rate fluctuations and excess dollar returns on the portfolios is captured by the negative foreign exchange risk factor loadings. Thus, excess dollar returns on South African portfolios have a tendency to decline with dollar appreciations. Factor loadings are generally higher for the exchange rate factor than for the world equity market factor, implying that portfolio returns are more sensitive to changes in foreign exchange rates. Like their world market risk factor counterparts, the absolute magnitude of foreign exchange risk factor betas increase with market capitalization and are all statistically different from zero.

When the local market factor is incorporated into the model, the resulting three-factor model appears to boost the importance of the world equity market factor in explaining portfolio returns in South Africa. The world market factor coefficient estimates increase for all the four portfolios, the greatest increment being observed for portfolio 2, the lowest for portfolio 4. The importance of the foreign exchange risk in return prediction diminishes but the inverse relationship is maintained. The absolute magnitudes of the foreign exchange risk factor betas decline for all the portfolios but the decline is more pronounced for portfolio 4. All the betas in respect of the local market factor are positive and statistically significant at the 1 percent level. They range from 0.6716 for portfolio 2 to 0.9453 for portfolio 4. Clearly, large firms appear to be more exposed to idiosyncratic risk factors within the local market than their smaller counterparts.

Overall, both the two-factor and the three-factor models account for more of the return variation associated with the large-firm portfolio than the smaller-firm portfolios. The coefficients of determination, adjusted for degrees of freedom, for portfolio 4 are 69.30% and 57.11%

respectively under the three-factor and two-factor models; the associated root mean square errors are 0.0472 and 0.0558. For portfolio 1, the adjusted coefficient of determination is 44.66% with a root mean square error of 0.0595 under the three-factor structure and 40.12% and 0.0619 respectively under the two-factor structure. The average pricing errors are mostly negative suggesting that the two models therefore have a tendency to overprice portfolios in South Africa's equity market. In absolute terms, portfolio 3 attracts the lowest average pricing error.

With a large *p*-value of 0.7586, the *J*-statistic fails to reject the overall goodness-of-fit for the two-factor model specification. Foreign exchange risk factor attracts a lower average monthly premium, at 3.095 percent, than the world equity market factor, which attracts an average monthly premium of 4.092 percent. Both risk premia are statistically significant at the 10 percent level. The consistency of the data with restrictions imposed by the model, as illustrated by the *J*-statistic, is still upheld for the three-factor structure. Under this specification, the foreign exchange rate factor yields a large monthly risk premium of 2.732 percent, which is not statistically significant at any conventional level. It is, therefore, apparent that the pricing of this factor under the two-factor model specification may have been fortuitous. This contention, as well as other earlier inferences, is the subject of further investigation in the proceeding section.

Finally, the global risk factor has a monthly premium of 4.093 percent which is significantly different from zero at the 10 percent level. The local market factor yields a risk premium of -1.875 percent, which is however not significantly different from zero. Thus, the South African equity market is fully integrated under the assumptions of the three-factor model – a change from the two-factor model, which gives a verdict of partial segmentation, partial integration. These results are consistent with Kabundi and Mouchili (2009) who use a multivariate approach based on Dynamic Factor Model of Forni et al. (2005) to find moderate synchronization of the South African stock market with the world common equity market between 1997 and 2006. In their study, the world return explains 55 percent of variance of South African stock returns.

4.3.4 Further Explorations

To further delve into the role played by foreign exchange risk in stock return prediction in Africa, I conduct the following additional experiments.

1. Robustness to Local Market Factor and Included Countries

To begin, I perform three checks for robustness of the unconditional multi-beta model structures as follows. Firstly, market index data, in US dollars, for six countries are obtained from the *MSCI Barra* for the period between May 2002 and December 2009. This sub-period has witnessed relative tranquility in Africa's foreign exchange markets and conveniently allows us to check whether time-variation in foreign exchange risk premium can be captured by the unconditional model specifications. Coincidentally, this also happens to be the period when national stock market index data for all countries except Botswana and Ghana are available from one source – MSCI Barra. In addition to the five countries, I now incorporate Mauritius in the investigations. The change in country composition allows us to check whether tests results are sensitive to included countries. Finally, since the MSCI African Frontier Markets index series is available for the entire sub-period, it is now used in place of the emerging markets composite index to proxy for the African local market. Consequently, the local market factor is constructed as the residuals from the projection of excess returns on the African Frontier Markets Index on a constant and the excess returns on the world market equity portfolio.

Results from these tests are displayed on Tables 15 and 16. Table 15 shows that the pure local market risk factor attracts a monthly reward of 2.275 percent, which is statistically significant at the 5 percent level. However, the world market risk factor commands a zero premium in the equity markets during the period from June 2002 through December 2009. Accordingly, the markets exhibit full segmentation during the period. With a p-value of 0.6561, the J-statistic does not reject the model. The other diagnostic statistics also show that the model fits well at the individual country level albeit with substantial, largely positive, pricing errors.

Panel A of Table 16 shows results for the two-factor foreign exchange risk model. Unlike the results on Table 15, but consistent with previous findings from the same model, the average pricing errors are very large and mostly positive; the root mean square errors are also large; the adjusted R-squares small. Importantly, however, the J-statistic, with a p-value of 0.4822, does not reject the model. Consistent with earlier findings, foreign exchange risk has a small, statistically insignificant, monthly premium of only 0.270 percent. The world market risk factor is priced, with a monthly premium of 1.308 percent, statistically significant at 5 percent.

Table 15
GMM regression results for robustness checks: two-factor market segmentation model

$$r_{it} = \beta_{iw}r_{wt} + \beta_{im}r_{mt} + \lambda_w\beta_{iw} + \lambda_m\beta_{im} + \varepsilon_{it}$$

Country	β_W	β_m	RMSE	Adj-R^2	APE
Egypt	0.8543*** (6.29)	0.8247*** (6.88)	0.0679	0.4341	0.0040
Kenya	0.7818*** (3.68)	0.5017*** (3.75)	0.0850	0.2181	0.0040
Mauritius	0.7940*** (4.51)	0.6276*** (3.87)	0.0661	0.3836	0.0030
Morocco	0.4526*** (3.20)	0.2202** (2.25)	0.0589	0.1087	0.0055
Nigeria	0.6433*** (2.95)	0.6233** (2.44)	0.0994	0.2504	-0.0098
S. Africa	1.3251*** (11.77)	0.1316 (1.31)	0.0517	0.6120	0.0043
	λ_W	λ_m	J-Statistic		
	0.171 (0.40)	2.275** (2.47)	3.2855 [0.6561]		

The table uses data covering the period 2002:6 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; r_{wt} is the demeaned excess return on the world market equity portfolio; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW} and β_{im} are, respectively, the sensitivities of African market equity returns to the world portfolio and pure local market factors and λ_W and λ_m are the respective risk premia, in percentages, for the two factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollars and are expressed in excess of one-month yields on US Treasury bills closest to one-month maturity.

Results for the three-factor model are shown in Panel B of Table 16. *P*-value of the *J*-statistic now stands at 0.99. The additional diagnostic statistics also show impressive improvements: the monthly average pricing error for Egypt, for instance, drops from 1.51 percent in return deviation to a meager 0.07 percent; the root mean square error for Nigeria drops from 0.1035 to 0.0957. The monthly premium on the foreign exchange risk factor rises to 0.901 percent but, statistically, it is still not different from zero at any conventional level of significance.

The results also show that the pure local market factor has diminished in importance, recording an insignificant premium in the three-factor model. Only the world market factor is priced at the

Table 16
GMM regression results for robustness checks: two- and three-factor structures

Panel C: Two-factor Model: $r_{it} = \beta_{iw} r_{wt} + \beta_{is} r_{st} + \lambda_w \beta_{iw} + \lambda_s \beta_{is} + \varepsilon_{it}$

Country	β_W	β_S	RMSE	Adj-R^2	APE
Egypt	0.9004*** (5.50)	-1.1330** (-2.05)	0.0811	0.1930	0.0155
Kenya	0.9553*** (4.93)	-0.7898 (-1.18)	0.0862	0.1955	0.0064
Mauritius	1.0615*** (6.30)	-1.3429*** (-2.91)	0.0710	0.2879	0.0084
Morocco	0.4810*** (3.69)	-1.0490*** (-2.92)	0.0591	0.1017	0.0079
Nigeria	0.7779*** (4.03)	-3.5214*** (-3.61)	0.1035	0.1883	0.0048
S. Africa	1.2516*** (13.64)	-1.2874*** (-3.79)	0.0476	0.6701	-0.0033
	λ_W	λ_S	J-Statistic		
	1.308** (2.00)	0.270 (0.70)	4.4826 [0.4822]		

Panel B: Three-factor Model: $r_{it} = \beta_{iw} r_{wt} + \beta_{im} r_{mt} + \beta_{is} r_{st} + \lambda_w \beta_{iw} + \lambda_m \beta_{im} + \lambda_s \beta_{is} + \varepsilon_{it}$

Country	β_W	β_S	β_m	RMSE	Adj-R^2	APE
Egypt	0.8285*** (6.75)	0.5547 (1.15)	0.8292*** (6.27)	0.0681	0.4311	0.0007
Kenya	0.8292*** (6.27)	-0.3277 (-0.53)	0.4033*** (2.76)	0.0849	0.2192	0.0012
Mauritius	0.9100*** (5.80)	-0.3988 (-0.87)	0.5344*** (3.58)	0.0659	0.3858	0.0028
Morocco	0.4557*** (3.61)	-0.3589 (-0.90)	0.2221** (1.99)	0.0588	0.1111	0.0049
Nigeria	0.7502*** (4.41)	-1.5282** (-2.03)	0.9200*** (3.27)	0.0957	0.3058	0.0020
S. Africa	1.2887*** (13.85)	-1.7520*** (-5.26)	0.0281*** (0.25)	0.0472	0.6759	-0.0003
	λ_W	λ_S	λ_m	J-Statistic		
	1.986* (1.85)	0.901 (1.50)	0.252 (0.24)	1.2326 [0.9902]		

The table uses data covering the period 2002:6 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the J-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; $r_{wt}$is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index and the pure local market factor; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW}, β_{is} and β_{im} are, respectively, the sensitivities of African market equity index returns to the excess return on the world market equity portfolio index, foreign exchange rate changes and pure local market factor. λ_W, λ_S and $\lambda_m$are the respective risk premia, in percentages, for the three factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollars and are expressed in excess of one-month yields on US Treasury bills closest to one-month maturity.

10 percent level with a monthly premium of 1.986 percent. This suggests that African markets are fully integrated for the period under review. Since the two-factor segmentation model

suggested full segmentation, these results are mixed and inconclusive. This conflicting finding and relatively poor performance of the local market factor may be traced to the African frontier market index to which the unconditional multi-factor models appear to be non-robust. It is apparent that this index does not explain equity returns in the continent as well as does the Emerging markets composite portfolio index. With the exception of Kenya, the rest of the countries, including Botswana and Ghana which were left out of this sensitivity analysis, now seem to possess market characteristics that are closer to those of other emerging markets.

2. Robustness to Changes in Reference Currency

This section is a check on the sensitivity of the model to various definitions of the foreign exchange rate. If a real exchange rate index similar to the one used in the preceding tests were developed for a currency different from the US dollar and if it gives different results from those obtained using the US dollar index, one can infer that the pricing of foreign exchange risk is dependent on the currency against which it is tested. The euro is used to conduct these checks. Due to historical and geographical distance reasons, the European Union accounts for a large proportion of external bilateral trade in African countries. World Trade Organization data reveal the European Union is the single largest trading partner of many African countries. For instance, in the year 2008, 60.4 percent of Botswana's and 44 percent of Ghana's exports went to the European Union, while 51.9 percent of Morocco's imports originated from the European Union.

Table 17 displays results for the two-factor foreign exchange risk model with returns measured in the euro. It is clear, from the table, that the hypothesis of a zero foreign exchange rate risk premium in Africa's equity markets is rejected by the data. The monthly foreign exchange rate risk premium is -1.926 percent, which is significant at the 10 percent level. Thus, with returns measured in the euro, foreign exchange risk appears weakly priced unconditionally in Africa's equity markets. The world market equity portfolio factor is also priced, commanding a monthly premium of 1.932 percent that is significant at the 1 percent level. The *J*-statistic yields a large *p*-value of 0.6676, confirming that the model's goodness-of-fit is not in dispute. Recall that the *p*-value of the *J*-statistic for the equivalent model with returns measured in US dollars was only 0.1460. This demonstrates that the unconditional two-factor foreign exchange risk model performs remarkably well with the euro as the reference currency.

Table 17
GMM regression results for the euro: two-factor currency risk model

$$r_{it} = \beta_{iw}r_{wt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_s\beta_{is} + \varepsilon_{it}$$

Country	β_W	β_S	RMSE	Adj-R^2	APE
Botswana	0.2880*** (4.08)	-0.0537 (-0.37)	0.0501	0.2572	0.0016
Egypt	0.0724 (0.99)	-0.1794 (-0.57)	0.0897	-0.0023	0.0043
Ghana	-0.5418*** (-5.79)	0.3626 (1.60)	0.0652	0.4319	-0.0008
Kenya	-0.0780 (-0.85)	0.5793* (1.71)	0.1059	0.0097	-0.0012
Morocco	0.6822*** (5.27)	-0.0031 (-0.02)	0.0593	0.5913	0.0032
Nigeria	0.2371* (1.84)	-0.3182 (-0.73)	0.1397	0.0199	0.0100
S. Africa	-0.9540*** (-30.10)	-0.1883 (-1.06)	0.0486	0.8189	0.0024
	λ_W	λ_S	J-Statistic		
	1.932*** (4.65)	-1.926* (-1.70)	3.2101 [0.6676]		

The table uses data covering the period 1999:2 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; $r_{wt}$is the demeaned excess return on the world market equity portfolio; r_{mt} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index. β_{iW} and β_{im} are, respectively, the sensitivities of African market equity returns to the world portfolio and foreign exchange rate factors and λ_W and λ_m are the respective risk premia, in percentages, for the two factors. RMSE is the root mean squared error and Adj-R^2 is the adjusted coefficient of determination; APE is the average pricing error. All index returns are measured in the euro and are expressed in excess of the one-month Eurodollar rate.

Coefficient estimates for many of the countries reveal that foreign exchange rates play a less important role in determining equity returns than the world equity portfolio. Only one of the foreign exchange risk factor loadings (Kenya) is statistically significant, and only at the 10 percent level. In the case of the world market portfolio, five of the seven coefficient estimates are non-zero. The other performance metrics also yield better results with the euro than the US dollar. The absolute values of the average pricing error now range from 0.08 percent per month (Ghana) to 0.10 percent (Nigeria). The highest APE with the US dollar as the reference currency was 1.22 percent per month. However, mixed results are observed with the RMSE and the coefficient of determination, with Egypt, Kenya and Nigeria, appearing to be more poorly fitted by the euro model than the dollar model whereas Botswana, Ghana, Morocco and South Africa record tremendous improvements when the reference currency changes from dollar to euro.

Table 18
GMM regression results for the euro: market segmentation model and three-factor model

Panel A: Two-factor Model $r_{it} = \beta_{iw}r_{wt} + \beta_{im}r_{mt} + \lambda_w\beta_{iw} + \lambda_m\beta_{im} + \varepsilon_{it}$

Country	β_W	β_m	RMSE	Adj-R^2	APE
Botswana	0.3026***	0.0007	0.0502	0.2537	0.0084
	(4.44)	(0.02)			
Egypt	0.0703	-0.2063***	0.0849	0.1020	0.0041
	(1.02)	(-3.46)			
Ghana	-0.5420***	-0.0506	0.0658	0.4218	-0.0005
	(-5.96)	(-1.27)			
Kenya	- 0.1146	0.5670***	0.0718	0.5447	0.0044
	(-1.64)	(13.05)			
Morocco	0.6769***	0.0360	0.0596	0.5869	-0.0065
	(5.37)	(0.97)			
Nigeria	0.2246***	-0.8101***	0.0721	0.7391	-0.0146
	(3.62)	(-18.81)			
S. Africa	-0.9611***	0.1384***	0.0464	0.8351	0.0187
	(27.53)	(4.73)			

λ_W	λ_m	J-Statistic
1.523***	-1.043**	4.5591
(5.20)	(-2.21)	[0.4720]

Panel B: Three-factor Model $r_{it} = \beta_{iw}r_{wt} + \beta_{im}r_{mt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_m\beta_{im} + \lambda_s\beta_{is} + \varepsilon_{it}$

Country	β_W	β_S	β_m	RMSE	Adj-R^2	APE
Botswana	0.2854***	-0.0890	-0.0099	0.0504	0.2491	0.0020
	(4.09)	(-0.50)	(-0.29)			
Egypt	0.0681	-0.1106	-0.2067***	0.0851	0.0974	0.0042
	(0.98)	(-0.35)	(-3.39)			
Ghana	-0.5350***	0.4402	0.0187	0.0655	0.4261	-0.0028
	(-5.71)	(1.64)	(-0.43)			
Kenya	-0.0911	0.4513*	0.5666***	0.0713	0.5517	0.0004
	(-1.31)	(1.70)	(12.66)			
Morocco	0.6633***	-0.0372	0.0222	0.0596	0.5865	0.0049
	(5.16)	(-0.25)	(0.56)			
Nigeria	0.2308***	-0.0827	-0.8133***	0.0722	0.7384	0.0070
	(3.71)	(-0.27)	(-17.25)			
S. Africa	-0.9606***	-0.1412	0.1316***	0.0465	0.8344	0.0031
	(-27.06)	(-0.71)	(4.15)			

λ_W	λ_S	λ_m	J-Statistic
1.696***	-1.491	-1.058*	2.6995
(3.97)	(-1.43)	(-1.89)	[0.9113]

The table uses data covering the period 1999:2 to 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; $r_{wt}$is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index and the pure local market factor; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW}, β_{iS} and β_{im} are, respectively, the sensitivities of African market equity index returns to the excess return on the world market equity portfolio index, foreign exchange rate changes and pure local market factor. λ_W, λ_S and $\lambda_m$are the respective risk premia, in percentages, for the three factors. RMSE is the root mean squared error and Adj-R^2 is the adjusted coefficient of determination; APE is the average pricing error. All index returns are measured in the euro and are expressed in excess of the one-month Eurodollar rate.

Tests results for the two- and three-factor market segmentation models are presented in Table 18. The two-factor model (Panel A) betas show that the world market equity portfolio still plays a more important role in influencing equity returns in Africa than does the pure local market factor. The monthly reward on the world market risk factor is 1.523, statistically significant at the 1 percent level. The monthly reward on the pure local market risk factor is -1.043, statistically significant at the 10 percent level. Thus, these results are consistent with partially segmented, partially integrated equity markets in Africa.

Results of the three-factor model in Panel B show that foreign exchange rate risk is not priced in Africa's equity markets when tested jointly with the pure market risk factor. The monthly foreign exchange risk premium falls to -1.491 percent (from -1.926 percent) and is not significant. Second, both the world and local market risk sources command significant risk premia, implying, once again, that African equity markets are partially segmented from, or partially integrated with, the rest of the world's equity markets. The *J*-statistic generates a large *p*-value of 0.9113 and fails to reject overall goodness-of-fit of the three-factor model. The apparently priced exchange risk factor under the two factor model, like in the case of South Africa, seems to be fortuitous, and most likely arises from the fact that the factor proxies for other factors not represented in the model. Once the idiosyncratic market factor is introduced, it better represents those other factors and the currency risk premium becomes insignificant, once again.

The better performance of the unconditional asset pricing models with returns measured in the euro cannot be ascribed to chance. The significance of the European Union to trade in Africa is such that the euro may be regarded as the dominant invoice currency. Empirical evidence shows that trade activities between developing and industrialized countries are predominantly invoiced in the industrialized countries' currencies (Grassman 1973). This finding, which is part of what is widely known in the literature as the Grassman's law, is predicated on the intuition that a firm with more bargaining power will choose its own currency to avoid foreign exchange rate risk.

Tavlas (1997) provides further support for the dominance of stable currencies in bilateral trade. He points out that because of their better storage of value, currencies with low inflation and inflation variability are preferred for invoicing. Silva (2004) also finds evidence that the strength of a currency, the depth of the financial market and the absence of high inflationary tendencies enhance the use of a country's currency in trade. Finally, Kamps (2006) provides evidence not only of the increasing importance of the euro as a world currency but of a slightly diminishing role of the US dollar. Further, he finds that the introduction of a common currency in the euro area increased the invoicing in euro at the expense of the US dollar.

These documented currency invoicing findings must be understood within the context of my earlier statement that the European Union dominates the rest of the world in bilateral trade with many African countries. Since cash flows arising from trade activities are likely to have an impact on foreign exchange rate movements, such fluctuations are likely to be witnessed more on the rate of exchange between African currencies and the dominant trading partner currency – the euro. It is no surprise, therefore, that the models investigated here perform better for Africa's equity markets with euro returns than with US dollar returns.

3. Robustness to Changes in Time

Finally, I run a check on the robustness of the models to different sub-periods of the study period. For this purpose, the study period is split into two equal parts, the first being the period during which a lot of noise was observed on foreign exchange rates and dollar-denominated market index returns for most countries. The second period witnessed relative tranquility for both variables. This check is conducted to establish whether the results obtained earlier can be replicated for the two time periods separately. Failure to obtain similar results for the two sub-periods may indicate that the risk premia are time-variant: this will be the case especially if changes in signs of the risk premia are observed between the two time periods or if risk premia signs for the entire period are different from the signs for either or both sub-periods.

Estimation results for the two-factor foreign exchange risk model are reported in Table 19. Panel A shows results for the first sub-period. The world market factor betas are all positive except for Ghana. Foreign exchange factor betas are negative except for Egypt and Ghana. The world market factor premium is negative and statistically insignificant. It was positive and insignificant when the entire period was used in the analysis. Foreign exchange risk remains non-priced, but the sign has also changed to negative. The *J*-statistic has a low *p*-value of 0.1644, which does

Table 19
GMM regression results for the sub-periods: the two-factor currency risk model

$$r_{it} = \beta_{iw}r_{wt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_s\beta_{is} + \varepsilon_{it}$$

Panel A: First Sub-period (1997:1 to 2003:6)

Country	β_W	β_S	RMSE	Adj-R^2	APE
Botswana	0.1174 (1.02)	-0.2035*** (-3.94)	0.0612	-0.0776	0.0171
Egypt	0.4547*** (2.68)	0.2796*** (4.72)	0.0788	0.0426	-0.0104
Ghana	-0.0540 (-0.49)	0.1138*** (2.68)	0.0622	-0.0871	-0.0160
Kenya	0.2269 (1.61)	-0.0093 (-0.15)	0.0645	-0.0274	-0.0131
Morocco	0.1385 (1.01)	0.0430 (0.75)	0.0524	-0.0303	-0.0035
Nigeria	0.4651*** (4.01)	-4.143*** (-14.98)	0.0763	0.7910	-0.0211
S. Africa	1.1491*** (5.27)	-0.0976 (-0.50)	0.0760	0.3147	-0.0057
	λ_W	λ_S	J-Statistic		
	-0.088 (-0.12)	-0.118 (-0.50)	7.8547 [0.1644]		

Panel B: Second Sub-period (2003:7 to 2009:12)

Country	β_W	β_S	RMSE	Adj-R^2	APE
Botswana	0.3034** (2.59)	-0.9823** (-2.14)	0.0504	0.1928	0.0059
Egypt	0.8390*** (5.93)	-1.3799*** (-3.24)	0.0822	0.2405	0.0201
Ghana	0.0939 (0.52)	-0.4641 (-1.60)	0.0601	-0.0187	-0.0014
Kenya	0.9341*** (3.49)	0.2644 (0.37)	0.0821	0.2032	0.0032
Morocco	0.2979* (1.75)	-1.2436*** (-3.03)	0.0620	0.1276	0.0076
Nigeria	-0.0171 (-0.05)	-1.9109*** (-3.38)	0.0859	0.0851	0.0085
S. Africa	1.3376*** (14.28)	-0.8360*** (-2.84)	0.0461	0.7033	0.0047
	λ_W	λ_S	J-Statistic		
	0.442 (0.91)	-0.145 (-0.39)	1.8277 [0.8724]		

The table uses data covering the period 1997:1 to 2009:12, split equally into 1997:1 to 2003:6 (first sub-period) and 2003:7 to 2009:12 (second sub-period). The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; r_{wt} is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index. β_{iW} and β_{iS} are, respectively, the sensitivities of African market equity returns to the world portfolio and foreign exchange rate factors and λ_W and λ_S are the respective risk premia, in percentages, for the two factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All returns are measured in the US dollar and expressed in excess of one-month returns on the US Treasury bills closest to one-month maturity.

not, however, reject the model. Other diagnostic statistics do not give a clean bill of health to individual equations in the system. Four equations (Botswana, Ghana, Kenya and Morocco) record negative *R*-squares, meaning that the two factors hardly explain the returns on those countries' aggregate market equity indices during the sub-period. Lackluster performance of the two factors during the sub-period is further confirmed by the very high average pricing errors, implying that the bulk of the variations in returns are not explained by the model. Generally, the model fit for this sub-period appears weak.

Panel B reports results for the second sub-period. Correct signs are reported for all factor betas except for Nigeria's world market risk factor. The foreign exchange factor risk premium is -0.145 percent and is statistically insignificant. Thus, for the two sub-periods, the two-factor model confirms that currency risk does not command a risk premium in Africa's equity markets. At 0.442 percent per month, word market factor premium is also statistically insignificant. Notice, however, that the world market factor premium is now positive, demonstrating its time-varying properties. With the exception of Egypt, the average pricing errors are relatively smaller during this sub-period than they were in the first sub-period. Similarly, the coefficients of determination are bigger in size not only than those of the first sub-period, but also those in the entire period. The RMSE are also smaller. The *J*-statistic's *p*-value realizes a large change, to 0.8724 from 0.1644 in the first sub-period. It is also better than that of the entire period, 0.1460. The model fit is much better for this sub-period. Overall, there is insufficient evidence to suggest that two-factor currency risk model is not robust to changes in time period. The observed changes in sign for factor premia are indicative of their time-varying properties.

Empirical results for the two-factor market segmentation tests are reported in Table 20. Panel A shows that that during the period between January 1997 and June 2003, the monthly world market factor commands a premium of 2.463 percent, significant at the 5 percent level. The monthly reward for the local market factor is -1.573 percent that is significant at the 1 percent level during the same period. The *J*-statistic manages a right-tailed *p*-value of 0.3306 and hence fails to reject the model's goodness-of-fit. However, very large root mean square errors and very low coefficients of determination are reported signifying the weak ability of the two factors to explain aggregate equity market returns in the African continent's bourses during this period.

Table 20
GMM regression results for the sub-periods: two-factor market segmentation model

$$r_{it} = \beta_{iw} r_{wt} + \beta_{im} r_{mt} + \lambda_w \beta_{iw} + \lambda_m \beta_{im} + \varepsilon_{it}$$

Panel A: First Sub-period (1997:1 to 2003:6)

Country	β_W	β_m	RMSE	Adj-R^2	APE
Botswana	0.1429 (1.45)	-0.1164 (-1.10)	0.0598	-0.0294	0.0119
Egypt	0.4051** (2.35)	0.5768*** (5.46)	0.0733	0.1710	-0.0120
Ghana	-0.0585 (-0.56)	0.1172 (0.95)	0.0611	-0.0494	-0.0128
Kenya	0.1496 (1.05)	0.0066 (0.06)	0.0655	-0.0598	-0.0169
Morocco	0.0183 (0.15)	-0.0939 (-1.22)	0.0522	-0.0220	-0.0056
Nigeria	-0.1171 (-0.76)	0.5081** (2.01)	0.1664	0.0071	-0.0057
S. Africa	1.1344*** (8.64)	1.0208*** (6.61)	0.0591	0.5863	-0.0185
	λ_W	λ_m	J-Statistic		
	2.463** (2.52)	-1.573*** (-2.87)	5.7571 [0.3306]		

Panel B: Second Sub-period (2003:7 to 2009:12)

Country	β_W	β_m	RMSE	Adj-R^2	APE
Botswana	0.4393*** (4.80)	-0.3538* (-1.87)	0.0513	0.1632	0.0030
Egypt	1.0033*** (7.73)	0.4796* (1.80)	0.0820	0.2450	0.0208
Ghana	0.2178 (1.25)	-0.4322** (-2.13)	0.0590	0.0168	-0.0048
Kenya	0.8795*** (3.84)	0.0448 (0.18)	0.0820	0.2039	0.0002
Morocco	0.4837*** (3.14)	0.3561* (1.81)	0.0633	0.0912	0.0090
Nigeria	0.2639 (0.86)	0.4465 (1.19)	0.0887	0.0238	0.0118
S. Africa	1.4021*** (19.96)	0.8411*** (6.55)	0.0398	0.7787	0.0059
	λ_W	λ_m	J-Statistic		
	0.794* (1.92)	-0.619* (-0.87)	3.1342 [0.6793]		

The table uses data covering the period 1997:1 to 2009:12, split equally into 1997:1 to 2003:6 (first sub-period) and 2003:7 to 2009:12 (second sub-period). The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; r_{wt} is the demeaned excess return on the world market equity portfolio; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW} and β_{im} are, respectively, the sensitivities of African market equity returns to the world portfolio and pure local market factors and λ_W and λ_m are the respective risk premia, in percentages, for the two factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollar and expressed in excess of one-month yields on the US Treasury bills closest to one-month maturity.

Similarly, the average pricing errors are large and negative indicating the tendency of the model not only to overprice securities but also its weak ability to explain returns during the sub-period.

Panel B displays two-factor market segmentation tests results for the period July 2003 through December 2009. All the betas for the world market factor are positive; residual local market factor betas are negative for Botswana and Ghana and positive for the remaining countries; they are generally lower than those of the world market factor, an indication of the superior role played by the world market factor in determining equity returns in the sampled countries.

For all countries, the coefficients of determination are manifestly better than the first sub-period values. The average pricing errors are mainly positive and smaller although they still leave a lot of returns unexplained in general and compared to the reported values for the entire period. The right-tailed *p*-value of the *J*-statistic is 0.6793, indicating that the model cannot be rejected. The world market factor attracts a low monthly reward of 0.794 percent, which is statistically significantly different from zero at the 10 percent level. The residual local market factor similarly commands a statistically significant (at 10 percent) reward of -0.619 percent per month.

These results confirm again that Africa's equity markets exhibit partial segmentation during the two sub-periods, but the level of segmentation varies with time. Clearly, we have little evidence to conclude for Africa's equity markets that the two-factor market segmentation model is not robust to changes in time periods.

Table 21 reports regression results for the three-factor model. From Panel A, we see that the model does not fare very well in explaining equity returns during the first sub-period. Two of the world market factor and two foreign exchange risk factor sensitivities have incorrect signs. Three of the countries have negative adjusted *R*-squares, the RMSE are also high for many countries. Similarly, large pricing errors are observed, especially for Nigeria and South Africa: they are all negative. The premium commanded by foreign exchange risk is statistically not different from zero. The world market and residual local market factors command very large risk premia (7.326 percent and 7.796 percent respectively), both of which are statistically significant at 5 percent. These large rewards can be explained by the high volatility in both exchange rates and dollar-

Table 21

GMM regression results for the sub-periods: the three-factor model

$$r_{it} = \beta_{iw}r_{wt} + \beta_{im}r_{mt} + \beta_{is}r_{st} + \lambda_w\beta_{iw} + \lambda_m\beta_{im} + \lambda_s\beta_{is} + \varepsilon_{it}$$

Panel A: First Sub-period (1997:1 to 2003:6)

Country	β_W	β_S	β_m	RMSE	Adj-R^2	APE
Botswana	0.0676	-0.3958***	-0.1448	0.0576	0.0469	-0.0020
	(0.55)	(-3.03)	(-1.27)			
Egypt	0.4826***	0.2554*	0.5486***	0.0729	0.1803	-0.0017
	(2.76)	(1.83)	(5.47)			
Ghana	-0.1338	0.1028	0.0630	0.0603	-0.0205	-0.0006
	(-1.24)	(1.09)	(0.53)			
Kenya	0.0691	-0.0772	0.1507	0.0654	-0.0560	-0.0072
	(0.52)	(-0.58)	(1.50)			
Morocco	-0.0598	-0.0386	-0.0585	0.0526	-0.0387	-0.0042
	(-0.59)	(-0.31)	(-0.77)			
Nigeria	0.4315***	-3.6470***	0.6392***	0.0761	0.7923	-0.0263
	(3.05)	(-5.29)	(3.68)			
S. Africa	1.1452***	-0.1838	1.0255	0.0598	0.5758	-0.0120
	(7.89)	(-0.49)	(6.64)			
	λ_W	λ_S	λ_m	J-Statistic		
	7.326**	-0.765	-7.796**	4.2971		
	(2.22)	(-1.39)	(-2.30)	[0.7450]		

Panel B: Second Sub-period (2003:7 to 2009:12)

Country	β_W	β_S	β_m	RMSE	Adj-R^2	APE
Botswana	0.3001**	-1.1552**	-0.3089*	0.0490	0.2372	-0.0088
	(2.53)	(-2.48)	(-1.72)			
Egypt	0.8467***	-1.1955**	0.5125*	0.0815	0.2543	-0.0126
	(5.74)	(-2.15)	(1.96)			
Ghana	0.0942	-0.6688*	-0.4068**	0.0590	0.0188	-0.0080
	(0.51)	(-1.90)	(-2.06)			
Kenya	0.9315***	0.2697	0.0128	0.0827	0.1913	-0.0349
	(3.35)	(0.35)	(0.05)			
Morocco	0.3048*	-1.1320**	0.3550*	0.0622	0.1244	-0.0037
	(1.77)	(-2.49)	(1.74)			
Nigeria	-0.0139	-1.7779***	0.4341	0.0862	0.0788	0.0103
	(-0.04)	(-2.72)	(1.21)			
S. Africa	1.3487***	-0.4891	0.8364***	0.0391	0.7867	-0.0464
	(17.49)	(-1.65)	(6.44)			
	λ_W	λ_S	λ_m	J-Statistic		
	4.551	-0.203	-0.497	1.7932		
	(0.88)	(-0.49)	(-0.69)	[0.9704]		

The table uses data covering the period 1997:1 to 2009:12, split equally into 1997:1 to 2003:6 (first sub-period) and 2003:7 to 2009:12 (second sub-period). The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. Prob-values of the *J*-statistic are in square braces. *, **, and *** respectively indicate that the reported coefficients are statistically significant at 10%, 5%, and 1% levels. r_{it} is the excess return on the *i*th country equity market index; r_{wt} is the demeaned excess return on the world market equity portfolio; r_{st} is the change in the African real exchange rate index for the dollar, orthogonal to the excess return on the world market equity portfolio index and the pure local market factor; r_{mt} is the pure local market factor, constructed as the excess return on the emerging markets equity portfolio orthogonal to the world market equity portfolio index. β_{iW}, β_{iS} and β_{im} are, respectively, the sensitivities of African market equity index returns to the excess return on the world market equity portfolio index, foreign exchange rate changes and pure local market factor. λ_W, λ_S and λ_m are the respective risk premia, in percentages, for the three factors. RMSE is the root mean squared error and Adj-R^2 is the coefficient of determination; APE is the average pricing error. All index returns are measured in the US dollars and are expressed in excess of one-month returns on US Treasury bills closest to one-month maturity.

denominated returns during the sub-period: the monthly standard deviation of returns in excess of the risk free rate ranges from 5.17% (Morocco) to 16.69% (Nigeria) during the sub-period. The *J*-statistic fails to reject the goodness of fit.

Panel B shows some improvement in the diagnostic statistics. None of the coefficients of determination are negative: they show that the three factors jointly explain between 1.88 percent (Ghana) and 78.67 percent (South Africa) of equity returns. The root mean square errors have generally fallen. However, the average pricing errors are generally larger during the second period than during the first period. The data fail to reject the model's goodness-of-fit since the *p*-value of the *J*-statistic is very large at 0.9704. The world market factor attracts a very large risk premium of 4.551 percent during the second sub-period, which is, however, not statistically significant. The local market and foreign exchange rates attract very low risk premia of -0.497 and -0.203 percent respectively, none of which is statistically significant. For this period, therefore, the model appears misspecified in respect of market segmentation.

Robustness is therefore established with the three-factor model in respect of the foreign exchange rate factor. On the other hand, there is strong evidence in favor of partial equity market segmentation in the first sub-period, but little evidence in support of any form of segmentation or integration during the second sub-period. This result is hardly surprising given that the unconditional model setting has been used in the analysis. In a similar study of the US equity market, Jorion (1991) finds the market risk factor priced in only one of the four sub-periods studied. Similarly, Choi et al. (1998) and Iorio and Faff (2002) find unconditional estimates of currency risk premia to be sensitive to choice of sub-periods. In emerging markets, Carrieri and Majerbi (2006) find unconditional risk premia to be sensitive to model specification. In each case, instability in risk premia estimates is attributed to the unconditional model's inability to capture time-varying risk. This study also documents sign changes in risk premia for all factors.

4.3.5 Concluding Remarks

Using unconditional models, I have, in this study, investigated whether fluctuations in foreign exchange rates introduce a risk that is perceived as significant by foreign investors to African equity markets. The study further tested whether those markets are regarded as segmented from, or integrated with the rest of the world's equity markets. Results from market segmentation studies generally conclude that equity markets in Africa are partially segmented/partially integrated. In broad terms, the evidence presented suggests that foreign exchange risk is not priced in African equity markets. This is the case for all the models investigated when returns are measured in the US dollar. When returns are measured in the euro, however, the two-factor model finds foreign exchange risk to be weakly priced while the three-factor model does not reject the hypothesis that currency risk commands a zero premium in African equity markets. One implication of these findings is that African monetary authorities should consider the euro as an important reserve currency. This is supported by the fact that the euro appears to be to be the currency in which the bulk of cash flows into and out of the continent are denominated. It is also strengthened by findings of Kamps (2006) which suggest that the euro's star, as an important invoicing currency, seems to be rising worldwide at the expense of the US dollar. At the portfolio level, studies with data from South Africa's equity market conclude that foreign exchange risk is only priced when the two-factor model is used.

Clearly, in this unconditional setting, the pricing of foreign exchange risk depends on the measurement currency as well as the model specification. Instability in the pricing of foreign exchange risk is an indictment on the suitability of unconditional asset pricing models in explaining variability in equity returns. Still, it is possible that use of aggregate market data masks important information on the response of stock returns to foreign exchange rate fluctuations that would be better captured through firm-level data. This will be explored in future as usable firm-level data for each of the countries becomes available. In lieu of this, conditional asset pricing models, to which I turn in the next section, may prove more successful in explaining the role of currency risk in African stock markets. This is because, as Choi et al. (1998) observe, conditional asset pricing models can incorporate the effects of intertemporal changes in market environments and are free from biases caused by small sample sizes and orthogonalization processes inherent in unconditional models.

4.4 Empirical Results for the Conditional SDF Asset Pricing Model

This section presents empirical results for the estimates of the stochastic discount factor models discussed in Section 3.3.2 (Chapter 3). In order to guard against over-fitting, each conditioning variable is incorporated separately into the stochastic discount factor (SDF) model. Results are presented separately for each of the different specifications. Inferences are based on the *J*-statistic of Hansen's (1982) optimal iterative GMM.[51] Additionally, this section provides results for some robustness checks.

4.4.1 Parameter Estimates

Table 22 contains estimation outputs for parameters ($\boldsymbol{b}$) of the stochastic discount factor as defined in equation (44). These parameters provide information on the importance of each factor in determining the pricing kernel. The table also contains estimates of factor risk premia ($\boldsymbol{\lambda}$) as presented in system (45). The latter parameters impart information on the relative importance of each factor in influencing expected returns on equity securities in Africa's capital markets. The model includes two risk factors, namely, the world market equity portfolio (*WLD*) returns and nominal foreign exchange rate changes (*NXR*). This specification is informed by Zhang (2006), who finds it to be the "most successful" model in an international equity pricing framework. Consistent with Cochrane (1996), model specifications in which the conditioning variables are not included in the pricing kernel are used here. Many authors, including Lettau and Ludvigson (2001), Drobetz et al. (2002) and Iqbal et al. (2010) have employed this approach successfully in conditional asset pricing studies in advanced and emerging equity markets. The implications of this model specification are investigated further in section 4.4.2.

Table 22 shows that only when interacted with the world dividend yield (Panel B) does the foreign exchange factor become statistically significant, at 1 percent, in explaining the pricing kernel of equity securities. In all the remaining model specifications, none of the risk factors significantly explains the pricing kernel for equity securities in Africa. These results are not surprising given that the model does not incorporate any local market factors.[52] In segmented or partially segmented equity markets, local factors are expected to exert more influence on the equity pricing kernel, than international/foreign factors (see, for example, Drobetz et al., 2002).

[51] Were the performance of different models in asset return prediction the study's objective, the distance metric of Hansen and Jagannathan (1997) would be more appropriate than the optimal GMM's *J*-statistic. However, the study's purpose is only to establish whether or not foreign exchange risk is conditionally priced in the equity markets, with the conditioning variable taking on different definitions. Further, Hodrick and Zhang (2001) explain that inferences on the validity of a model based on the test of the HJ-distance equal zero are always similar to inferences based on the *J*-test for the optimal GMM.

[52] As was explained earlier, available local factors were not of good quality and could not be used to conduct these tests. None of them showed any tendency to explain local equity returns using ordinary least squares regression.

Table 22
Parameter estimates for the SDF model

Panel A: Factors Scaled by returns on *MSCI* world equity portfolio (WLR)					
SDF Parameters	Constant	b_{WLD}	b_{NXR}	$b_{WLD \cdot WLR}$	$b_{NXR \cdot WLR}$
Coefficient	-1.156	1.903	0.910	0.289	-1.006
(*t*-value)	(-0.24)	(0.37)	(1.19)	(0.43)	(-1.40)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot WLR}$	$\lambda_{NXR \cdot WLR}$
Coefficient	1.011***	0.018*	5.147	-0.041***	-4.633
(*t*-value)	(235.47)	(1.82)	(0.70)	(-3.47)	(-0.63)
Model Tests		J-Statistic	Wald1	Wald2	Sup-LM
Chi-square		12.8248	12.48	5.25	5.084
(*p*-value)		[0.1707]	[0.0141]	[0.0725]	
Panel B: Factors Scaled by *MSCI* world equity portfolio Dividend Yield (WDY)					
SDF Parameters	Constant	b_{WLD}	b_{NXR}	$b_{WLD \cdot WDY}$	$b_{NXR \cdot WDY}$
Coefficient	2.047	-1.030	-0.042	0.146	0.812***
(*t*-value)	(0.71)	(-0.35)	(-1.40)	(0.40)	(2.80)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot WDY}$	$\lambda_{NXR \cdot WDY}$
Coefficient	1.006***	0.001	0.258	0.003*	-0.111*
(*t*-value)	(249.22)	(0.16)	(0.95)	(1.91)	(-1.78)
Model Tests		J-Statistic	Wald1	Wald2	Sup-LM
Chi-square		12.3883	18.83	3.28	5.260
(*p*-value)		[0.1923]	[0.0008]	[0.1935]	
Panel C: Factors Scaled by Eurodollar Deposit rate (UDR)					
SDF Parameters	Constant	b_{WLD}	b_{NXR}	$b_{WLD \cdot UDR}$	$b_{NXR \cdot UDR}$
Coefficient	1.891	-0.883	-0.038	0.003	-0.592
(*t*-value)	(0.72)	(-0.34)	(-1.55)	(0.33)	(-0.60)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot UDR}$	$\lambda_{NXR \cdot UDR}$
Coefficient	1.007***	0.021**	2.333***	0.785**	-122.199***
(*t*-value)	(632.32)	(2.89)	(3.46)	(2.20)	(-5.91)
Model Tests		J-Statistic	Wald1	Wald2	Sup-LM
Chi-square		23.9844	58.09	42.83	21.333*
(*p*-value)		[0.0043]	[<0.0001]	[<0.0001]	
Panel D: Factors Scaled by USA Term Premium (UTP)					
SDF Parameters	Constant	b_{WLD}	b_{NXR}	$b_{WLD \cdot UTP}$	$b_{NXR \cdot UTP}$
Coefficient	-1.678	2.662	0.009	-0.021	0.133
(*t*-value)	(-0.55)	(0.88)	(0.21)	(-0.33)	(0.80)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot UTP}$	$\lambda_{NXR \cdot UTP}$
Coefficient	1.011***	0.018	1.621***	0.264***	0.518**
(*t*-value)	(251.76)	(1.57)	(2.65)	(3.27)	(2.46)
Model Tests		J-Statistic	Wald1	Wald2	Sup-LM
Chi-square		10.4182	13.81	7.41	9.307
(*p*-value)		[0.3177]	[0.0079]	[0.0246]	

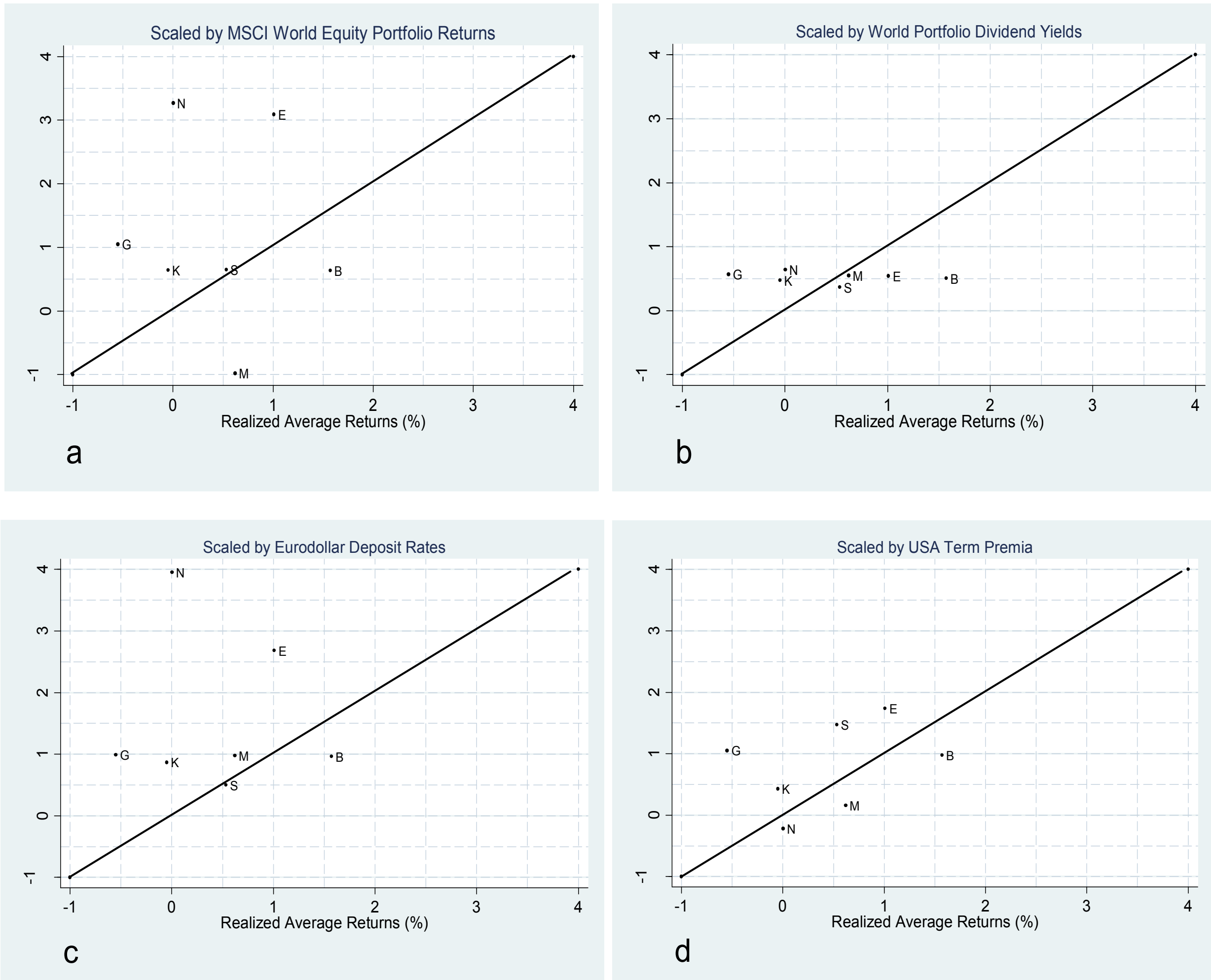
The table uses monthly nominal gross returns for the period 1997:1 to 2009:12. Returns are denominated in US dollars. The table reports GMM estimates of parameters (*b*) of the stochastic discount factor model and factor risk premia (λ). The reported *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. *J*-statistic is Hansen's (1982) test for overidentifying restrictions; Wald1 is the joint Wald test that all factor pricing parameters equal zero; Wald2 is the joint Wald test that parameters λ_{NXR} and $\lambda_{NXR.CON}$ equal zero (where *CON* is a specific conditioning variable). *WLD* and *NXR* stand for gross monthly returns on MSCI world equity portfolio and monthly nominal foreign exchange rates respectively. The foreign exchange rate factor is the equal-weighted average of binomial exchange rates between the US dollar and currencies of each of the seven African countries. All conditioning variables are lagged one period. Numbers in square brackets are *prob*-values of the test statistics. Sup-LM is Andrews (1993) test of structural stability of the parameters of the stochastic discount factor model: critical values are obtained from Table 1 of Andrews (2003). *, ** and *** indicate statistical significance at 10%, 5% and 1% respectively.

Parameter estimates for factor risk premia yield interesting results. The zero-beta asset return averages between 0.6% and 1.1% per month, which is reasonable for Africa's money markets. Proxying the zero-beta asset by Treasury bills, the explanation for the moderate to high zero-beta rate draws from the heavy demand for domestic debt to finance short-term budget deficits in most of the economies studied. The average annualized Treasury bills rate for these countries over the study period was 11.895% (approximately 1.0% per month).[53] Thus, the estimated zero-beta rate is very close to the observed risk-free rates. This finding satisfies one of the criteria prescribed by Lewellen et al. (2010), that conditional asset pricing models must be evaluated against their ability to estimate risk premia that are close to observed values. I use Wald statistics to examine the joint significance of the factor risk estimates: for all model specifications, the test strongly rejects the hypothesis that all factor risk coefficients are zero (Wald1).

The time-varying component of the foreign exchange risk factor is significantly priced in all the specifications excluding the one in which factors are scaled by lagged returns on the world market equity portfolio. Further, except for the case where scaling is done with dividend yields on the world equity market portfolio, Wald tests of joint parameter significance of the foreign exchange risk premia (that is, both the time-invariant and time-varying foreign exchange risk premia coefficients) rejects the hypothesis that they are both zero (Wald2). These results therefore suggest that the foreign exchange risk factor commands a significant premium in the equity markets studied. The world market risk factor also appears to be conditionally priced in all the specifications.

[53] The average annualized percentage Treasury bills rates for each country were: Egypt: 8.89; Ghana: 25.01; Kenya: 10.51; Morocco: 4.12; Nigeria: 12.39 and South Africa: 10.45 (source: International Financial Statistics). No Treasury bills rates were available for Botswana for the period.

Figure 3
Pricing errors for the Stochastic Discount Factor model



The figure shows pricing error (monthly percentages) plots for the stochastic discount factor model of equity returns with MSCI world equity portfolio and a foreign exchange variable as risk factors. The foreign exchange rate factor is measured as a weighted average of binomial exchange rates between the US dollar and currencies of each of the seven African countries. The risk factors are scaled by lagged values of MSCI world equity portfolio (Panel a), MSCI world equity portfolio dividend yield (Panel b), Eurodollar deposit rates (Panel c), and USA term premium (Panel d). Cyclical components of all conditioning variables, except the MSCI world equity portfolio returns, are extracted using the Hodrick and Prescott (1997) filter. Mean realized returns (horizontal axis) are plotted against mean returns implied by the respective model specification (vertical axis). The test assets are equity market aggregate returns for Botswana (B), Egypt (E), Ghana (G), Kenya (K), Morocco (M), Nigeria (N), and South Africa (S). Monthly data are from 1997:1 to 2009:12. Nominal gross returns denominated in US dollars are used.

The *J*-statistic yields *p*-values of 17%, 19% and 32% respectively for specifications in which risk factors are scaled by world equity portfolio returns, world equity portfolio dividend yields and USA term premia. Thus, for these specifications, the *J*-statistic shows that the data do not reject the model specifications at conventional levels of significance. However, Panel C shows that the data reject the model specification with the Eurodollar deposit rate as the conditioning variable. In that specification, the optimal GMM test for overidentifying restrictions yields a *p*-value of only 0.43 percent. The sup-LM statistic similarly indicates that SDF parameter estimates for this model specification do not pass the stability test.

Figure 3 displays the pricing error plots for the models' estimates. The straight line in each panel is the 45^{o} line, along which all properly priced assets/portfolios should lie. Looking at the plots, it is clear that pricing errors are very large for the specifications in which lagged values of the world market equity portfolio (Panel *a*) and the Eurodollar deposit rates (Panel *c*) are used as the conditioning variables. These two conditioning variables are the least capable of capturing time-varying risk properties of the risk factors.

The pictorial representation therefore clarifies the relatively poor performance of the two specifications as brought out by the *J*-statistic. For these two specifications, aggregate market returns for many of the countries (except South Africa) suffer huge mispricing by the model. In both cases, pricing errors are greatest for Nigeria and Egypt, the two countries whose bilateral exchange rates with the US dollar exhibited the most serious leptokurtosis. Hwang and Satchell (1999) demonstrate that kurtosis is an important factor in modeling emerging market returns. Thus, thick tails in the distribution of stock returns in the two countries could explain their large prediction errors. The existence of large pricing errors in conditional asset pricing is, however, not novel to this study: it has been reported elsewhere in the emerging equity markets studies (Iqbal et al., 2010) as well as in advanced equity markets studies (Fletcher and Kihanda, 2005; Schrimpf et al., 2007). These results are therefore not surprising.

Scaling risk factors by the world equity portfolio dividend yields and USA term premia induces substantial reduction in pricing errors but still do not fully account for the variation in returns in Africa's equity markets. Ghana, whose bilateral foreign exchange rate with the US dollar and US dollar-denominated returns showed significant autocorrelation for up to nine lags, appears to be the most mispriced by the two "well performing" conditioning variables. The most interesting conditioning variable appears to be the cyclical component of USA term premium (UTP). Thus, of the macroeconomic variables investigated, the USA term premium seems to be the most capable of predicting future business cycles in the real economies of the African countries studied. This finding is consistent with those of Schrimpf et al. (2007) and agrees with Harvey (1991), who shows that US instruments have some predictive power over equity returns in foreign markets. The success of the USA term premia as a conditioning variable may be ascribed to the commonly observed phenomenon in which the USA business cycles either lead or coincide with business cycles in other parts of the world.

4.4.2 Further Tests

1. Model Performance with Other Conditioning Variables

To explore the sensitivity of the model's performance to other conditions, I use two additional conditioning variables, namely, the USA index of industrial production and a January dummy. The January dummy is a variable that takes a value of one at the end of each month of January and zero at the end of each of the other calendar months. The January dummy has been used as a conditioning variable in many studies in advanced countries, especially those whose tax years end on December 31. It's appropriateness to Africa's markets, where tax years are varied has not been empirically established. Regarded as one of the best indicators of business cycle changes, the cyclic component of industrial production has gained prominent usage in stochastic discount factor models lately (see, for example, Hodrick and Zhang, 2001, Zhang, 2006 and Iqbal et al., 2010). Use of the USA industrial production index in this study is predicated on Harvey's (1991) observation that USA macroeconomic variables can predict equity returns in other countries. As Dumas and Solnik (1995) explain, the reason why U.S. variables are legitimate instruments for non-U.S. returns is presumably that asset returns are related to business cycles and that the U.S. cycle has led or coincided with other cycles during the period under study.

Table 23 provides results for GMM estimates of the stochastic discount factor model's parameters ($\boldsymbol{b}$) and factor risk premia ($\boldsymbol{\lambda}$). In Panel A, where risk factors are scaled by the cyclical component of USA industrial production, the earlier results that none of the risk factors, scaled or unscaled, contributes significantly to the equity pricing kernel are replicated . In Panel B, where the risk factors are scaled by the January dummy, it is evident that the time-invariant component of the foreign exchange rate contributes to the equity pricing kernel. The parameter estimate is, however, only weakly significant at the 10 percent level (p-value = 9.3%). Similarly, it is seen that the interaction between the January dummy and the world market return yields a statistically significant risk price at the 5 percent level. A possible interpretation of this is that the world market price of risk is different in January than in other months. This result is consistent with Schrimpf et al. (2007) who find the price of the interaction between the January dummy and the domestic market portfolio significant in their investigation of the standard domestic CAPM. In both model specifications, the sup-LM statistics do not reject the hypothesis that SDF parameter estimates are stable.

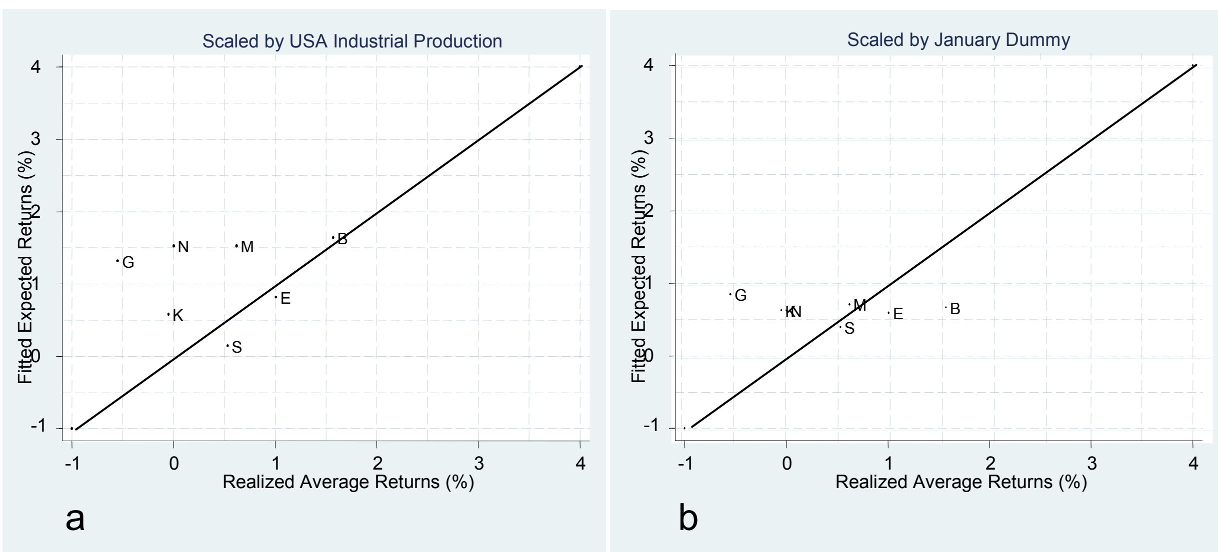
Table 23
Parameter estimates for the SDF model with other conditioning variables

Panel A: Factors Scaled by USA Industrial Production (UIP)					
SDF Parameters	Constant	b_{WLD}	b_{NXR}	$b_{WLD \cdot UIP}$	$b_{NXR \cdot UIP}$
Coefficient	-1.984	3.005	-0.015	-0.024	3.334
t-value	(-0.75)	(1.13)	(-0.36)	(-1.15)	(1.16)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot UIP}$	$\lambda_{NXR \cdot UIP}$
Coefficient	1.012***	-0.008	0.329	-1.438***	-0.005
(*t*-value)	(289.91)	(-1.30)	(0.62)	(3.52)	(-0.61)
Model Tests		J-Statistic	Wald1	Wald2	Sup-LM
Chi-square		10.3124	15.15	2.60	11.831
(*p*-value)		[0.3257]	[0.0044]	[0.2728]	
Panel B: Factors Scaled by January Dummy (JAN)					
SDF Parameters	Constant	b_{WLD}	b_{NXR}	$b_{WLD \cdot JAN}$	$b_{NXR \cdot JAN}$
Coefficient	13.842	-12.636	-0.444*	-0.217	42.53
(*t*-value)	(1.65)	(-1.53)	(-1.69)	(-1.07)	(1.59)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot JAN}$	$\lambda_{NXR \cdot JAN}$
Coefficient	1.009***	0.004	1.250	0.044**	-0.007
(*t*-value)	(297.30)	(0.45)	(1.40)	(2.95)	(-1.21)
Model Tests		J-statistic	Wald1	Wald2	Sup-LM
Chi-square		13.0300	198.6	16.31	12.629
(*p*-value)		[0.1613]	[<0.0001]	[0.0003]	

The table uses monthly nominal gross returns for the period 1997:1 to 2009:12. Returns are denominated in US dollars. The table reports GMM estimates of parameters of the stochastic discount factor model (b) and factor risk premia (λ). The reported *t*-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. *J*-statistic is Hansen's (1982) test for overidentifying restrictions; Wald1 is the joint Wald test that all risk factor pricing parameters equal zero; Wald2 is the joint Wald test that parameters λ_{NXR} and $\lambda_{NXR.CON}$ equal zero (where *CON* is a specific conditioning variable). *WLD* and *NXR* stand for gross monthly returns on MSCI world equity portfolio and monthly nominal foreign exchange rates respectively. The foreign exchange rate factor is the equal-weighted average of binomial exchange rates between the US dollar and currencies of each of the seven African countries. All conditioning variables are lagged one period. Numbers in square brackets are *prob*-values of the test statistics. Sup-LM is Andrews (1993) test of structural stability of the parameters of the stochastic discount factor model: critical values are obtained from Table 1 of Andrews (2003). *, ** and *** indicate statistical significance at 10%, 5% and 1% respectively.

The *J*-statistic is worse (*p*-value of 16.1%) under the January dummy specification than under the USA industrial production specification (*p*-value of 32.6%). The zero-beta rate is 1.2% and 0.9% per month respectively for the USA industrial production and January dummy specifications, again within the range of observed values. The two specifications perform relatively poorly in estimating the factor risk premia, with only the time-varying world market factor appearing significantly priced, at the 1% and 5% levels, respectively for the two specifications. In both cases, neither the time-varying nor time-invariant foreign exchange risk factors appear to be individually priced. However, the Wald test (Wald2) for joint significance reveals that the foreign exchange risk is significantly priced (*p*-value = 0.03%) with the January dummy as the conditioning variable.

Figure 4
Pricing errors for the Stochastic Discount Factor model with other conditioning variables



The figure shows pricing error (monthly percentages) plots for the stochastic discount factor model of equity returns with MSCI world equity portfolio and a foreign exchange variable as risk factors. The foreign exchange rate factor is measured as a weighted average of binomial exchange rates between the US dollar and currencies of each of the seven African countries. The risk factors are scaled by lagged values of USA Industrial Production (Panel a) and the January dummy (Panel b). The cyclical component of the USA industrial production index is extracted using the Hodrick and Prescott (1997) filter. Mean realized returns (horizontal axis) are plotted against mean returns implied by the respective model specification (vertical axis). The test assets are equity market aggregate returns for Botswana (B), Egypt (E), Ghana (G), Kenya (K), Morocco (M), Nigeria (N), and South Africa (S). Monthly data are from 1997:1 to 2009:12. Nominal gross returns denominated in US dollars are used.

Figure 4 presents a visual display of estimation errors. The figure shows that many of the countries' aggregate market equity portfolios, notably Ghana and Nigeria, suffer large mispricing with USA industrial production as the conditioning variable. Pricing errors for all countries, except Ghana and Botswana, are lower with the January dummy as the conditioning variable. The pattern of pricing errors for the January dummy (*JAN*) specification also appears to closely mirror that of the world market equity portfolio dividend yield (*WDY*), depicted in Panel *b* of Figure 3. Thus, although many of these countries' tax years do not end on December 31, there appear to be some weak evidence of abnormal equity return behavior coinciding with the month of January, a fact that could possibly form a basis for equity return prediction in these markets.

2. Time-varying intercept term

Following practice in the literature, the model specification used earlier did not include the conditioning variable as a risk factor. Whether or not this omission causes the model to be misspecified is now investigated. For this purpose, I use the two "well performing" models where risk factors are scaled by the MSCI world equity portfolio dividend yields and the USA term premia, respectively. The resulting estimation outputs are presented in Table 24. For the world dividend yield specification, many of the SDF coefficients turn statistically significant when the time-varying intercept is included. Importantly, the world market equity portfolio dividend yield appears to significantly contribute to the pricing kernels in the investigated stock markets. And the parameters appear to be stable as evidenced by the sup-LM statistic.

The specification in which the USA term premium is used as the conditioning variable yields mixed results. Panel B of Table 24 confirms earlier results that none of the SDF coefficient estimates is significant; specifically, the conditioning variable, used as a risk factor, appears not to significantly contribute to the equity pricing kernel. And, as the sup-LM statistic indicates, the SDF parameters remain stable even when the time-varying intercept term is incorporated. The *J*-statistic for the specification with the time-varying intercept (*p*-value = 39.7%) is better than that without (*p*-value = 31.8%, Table 22, Panel D), implying that time-varying intercept may matter.

Pricing error plots (Figure 5) also reveal that, for many countries, estimation errors tend to be lower for the specification with the time-varying intercept (Panel *b*) than the specification without the time-varying intercept (Panel *d*). However, none of the factor risk premia is statistically significant and both Wald tests fail to reject the hypotheses of zero joint significance of factor risk premia. Importantly, the conditioning variable is not priced. It is therefore difficult to conclude that inclusion of the time-varying intercept improves model performance. Confounding findings of a similar nature have been reported elsewhere by Iqbal et al. (2010).

Table 24
Parameter estimates for the SDF model with time-varying intercept term

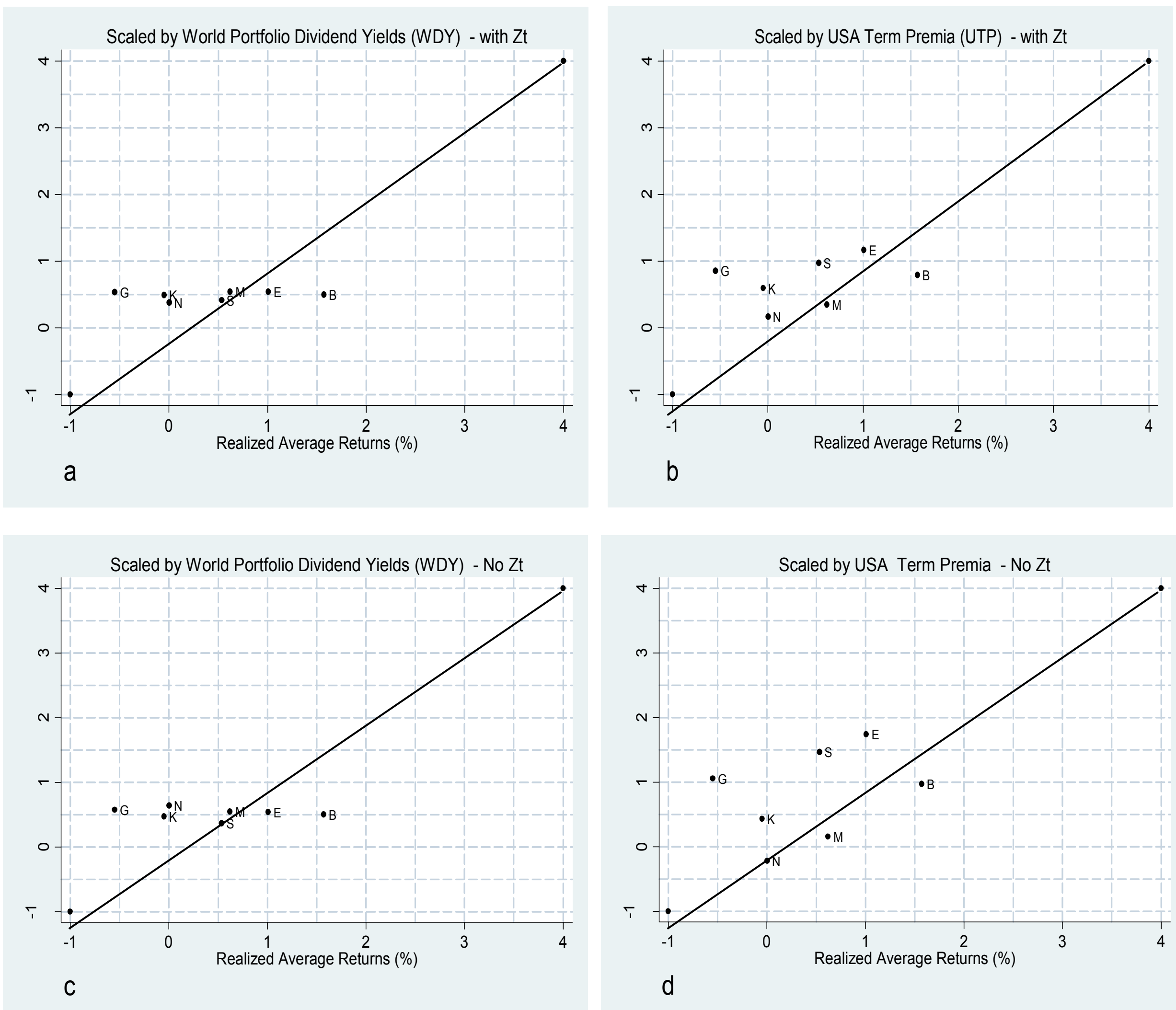
Panel A: Factors Scaled by *MSCI* world equity portfolio Dividend Yield (WDY)						
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{WDY}	$b_{WLD \cdot WDY}$	$b_{NXR \cdot WDY}$
Coefficient	2.744	-1.709	-0.061**	-38.728***	38.940***	1.871***
(t-value)	(0.88)	(-0.55)	(-2.14)	(-2.95)	(2.96)	(3.18)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{WDY}	$\lambda_{WLD \cdot WDY}$	$\lambda_{NXR \cdot WDY}$
Coefficient	1.006***	0.001	0.269	0.049	-0.002	-0.119
(t-value)	(248.87)	(0.21)	(0.98)	(0.19)	(-0.08)	(-1.52)
Model Tests		*J*-Statistic	Wald1	Wald2	Sup-LM	
Chi-square		12.4558	4.92	0.99	1.905	
(p-value)		[0.1320]	[0.2956]	[0.6106]		
Panel B: Factors Scaled by USA Term Premium (UTP)						
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{UTP}	$b_{WLD \cdot UTP}$	$b_{NXR \cdot UTP}$
Coefficient	0.666	0.339	-0.023	6.413	-6.410	0.005
(t-value)	(0.16)	(0.08)	(-0.34)	(0.77)	(-0.77)	(0.02)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{UTP}	$\lambda_{WLD \cdot UTP}$	$\lambda_{NXR \cdot UTP}$
Coefficient	1.009***	0.011	0.737	0.446	-0.279	0.440*
(t-value)	(238.95)	(0.89)	(0.68)	(0.94)	(-0.57)	(1.77)
Model Tests		J-Statistic	Wald1	Wald2	Sup-LM	
Chi-square		8.3869	4.68	1.13	18.208	
(p-value)		[0.3966]	[0.3220]	[0.5689]		

The table uses monthly nominal gross returns for the period 1997:1 to 2009:12. Returns are denominated in US dollars. The table reports GMM estimates of parameters of the stochastic discount factor model (b) and factor risk premia (λ). The reported t-statistics are robust to heteroskedasticity and autocorrelation; the number of lags for the Bartlett kernel was set at 3, which is consistent with the Newey and West (1987) 2-lag kernel. *J*-statistic is Hansen's (1982) test of overidentifying restrictions; Wald1 is the joint Wald test that all factor pricing parameters equal zero; Wald2 is the joint Wald test that parameters λ_{NXR} and $\lambda_{NXR.CON}$ equal zero (where *CON* is a specific conditioning variable). *WLD* and *NXR* stand for gross monthly returns on MSCI world equity portfolio and monthly nominal foreign exchange rates respectively. The foreign exchange rate factor is the equal-weighted average of binomial exchange rates between the US dollar and currencies of each of the seven African countries. All conditioning variables are lagged one period. Numbers in square brackets are *prob*-values of the test statistics. Sup-LM is Andrews (1993) test of structural stability of the parameters of the stochastic discount factor model: critical values are obtained from Table 1 of Andrews (2003). *, ** and *** indicate statistical significance at 10%, 5% and 1% respectively.

3. Dynamic Performance of the SDF Model Specifications

My discussion thus far has concentrated largely on the goodness-of-fit of the SDF model specifications from a cross-sectional perspective. The preceding discussion of pricing errors has also focused largely on risk factor pricing estimates, and their associated covariance measure, at the neglect of the parameter estimates of the stochastic discount factor. The following diagnostic check, advocated by Farnsworth et al. (2002), focuses on the economic magnitudes of pricing errors for particular assets (countries, in this dissertation) from a time series perspective. The

Figure 5
Pricing errors with and without the time-varying intercept



The figure shows pricing error (monthly percentages) plots for the stochastic discount factor model of equity returns with MSCI world equity portfolio and a foreign exchange variable as risk factors. The foreign exchange rate factor is measured as a weighted average of binomial exchange rates between the US dollar and currencies of each of the seven African countries. The risk factors are scaled by lagged values of MSCI world equity portfolio dividend yields – without the time-varying intercept term (Panel a), and with the time-varying intercept term (Panel c); USA term premium without the time-varying intercept term (Panel b), and without the time-varying intercept term (Panel d). Cyclical components of all conditioning variables are extracted using the Hodrick and Prescott (1997) filter. Mean realized returns (horizontal axis) are plotted against mean returns implied by the respective model specification (vertical axis). The test assets are equity market aggregate returns for Botswana (B), Egypt (E), Ghana (G), Kenya (K), Morocco (M), Nigeria (N), and South Africa (S). Monthly data are from January 1997 to December 2009. Nominal gross returns denominated in US dollars are used.

central idea behind the authors' proposal is that if the model is working well, then the SDF's pricing errors, $M_{t+1}R_{t+1} - 1$, should not be predictable using any information available at the investors' disposal at time t, as proxied by the set of conditioning variables, Z_t. In this application, the pricing kernel, M_{t+1}, is defined as

$$M_{t+1} = \hat{b}_0 + \hat{b}_{WLD}R_{WLD} + \hat{b}_{NXR}R_{NXR} + \hat{b}_{WLD \cdot CON}R_{WLD \cdot CON} + \hat{b}_{NXR \cdot CON}R_{NXR \cdot CON} \qquad (65)$$

where *WLD* represents the world market equity portfolio, *NXR* denotes the change in nominal exchange rate and *CON* is a conditioning variable; R_k is the gross return on factor (or interaction between factor and conditioning variable) k at time $t+1$; a circumflex above a coefficient implies that the GMM estimate of the coefficient is used. Farnsworth et al. (2002) demonstrate that the standard deviations of the fitted values of a regression of the model pricing errors, $M_{t+1}R_{t+1} - 1$, on instrumental variables set, Z_t, should explain how well a particular model specification accounts for predictable variation in returns.[54]

Table 25
Dynamic performance of SDF model specifications

	Average	Minimum	Maximum
Risk factors scaled by			
WLR	0.0422	0.0273	0.0716
WDY	0.2051	0.1660	0.3416
UDR	0.1647	0.1286	0.2741
UTP	0.0447	0.0301	0.0548

The table uses monthly nominal gross returns for the period from 1997:1 to 2009:12. Returns are denominated in US dollars. The table reports results of the tests for dynamic model specification performance using a procedure proposed by Farnsworth et al. (2002). The model pricing errors, $M_{t+1}(\hat{b})R_{t+1} - 1$, for each model specification are regressed against the set of the conditioning variables. The standard deviations of the resulting fitted values are then obtained in each case. A low standard deviation indicates that a particular model specification does well in "explaining" the time series variation of the gross returns on a country's aggregate stock market. The conditioning variables used include the gross returns on the MSCI world equity portfolio (WLR), dividend yields on the MSCI world equity portfolio (WDY), Eurodollar deposit rates (UDR) and the USA term premium (UTP). All conditioning variables are lagged one period. Cyclical components of all conditioning variables, except WLR, are extracted using the Hodrick and Prescott (1997) filter.

[54] Fletcher and Kihanda (2005) and Schrimpf et al. (2007) implement a variant of this approach in which the vector of excess asset returns, rather than gross returns is employed in the regressions. Consistent with Farnsworth et al. (2002), they both find that conditional asset pricing models struggle to capture time-varying predictability of stock returns. Differences in results are not expected, however, provided that the same risk-free rate of return is used to compute excess returns for all portfolios/assets.

The smaller the sample standard deviation, the better is the model specification's ability to "explain" return variation. The one strength of this diagnostic check is that standard deviation measure does not place any penalty on model specifications that get the average return wrong. I perform this test for each of the seven countries in the sample; results are presented in Table 25.

In terms of the average standard deviation, Table 25 shows that the specification with WLR as the conditioning variable appears to perform the best in capturing the time series predictability of gross returns on national equity indexes. However, the range of standard deviations is much smaller for the specification with *UTP* as the conditioning variable (0.0247) than the one with *WLR* (0.0443). Thus, it would appear that the distribution of standard deviations is tighter for the specification with *UTP* as the conditioning variable than the one with *WLR*. Indeed, the standard deviation of the averages of the computed standard deviations is only 0.0099 for the UTP specification against 0.0144 for the *WLR* specification. Thus, the *UTP* specification yields smaller standard deviations and, consequently, still proves to be the "best performer". This corroborates the inference drawn from the *J*-statistic. It is worthy of note that one of the specifications with better cross-sectional performance, *WDY*, also turns out to be the "worst performer" from a time-series perspective. This confirms the earlier assertion that the methodology used in this diagnostic check is not biased by the ability of a model specification to correctly capture average returns.

4.4.3 Concluding Remarks

I have used the stochastic discount factor model to investigate whether foreign exchange risk is priced in major stock markets of Africa. A two-factor model with the world market equity portfolio and nominal foreign exchange rate changes as risk factors is investigated. Following Zhang (2006), I invoke the market integration hypothesis which assumes that only global factors are priced and that prices are uniform across the seven countries studied. Different specifications of the two-factor model, in which the factors are scaled by conditioning variables, are examined. Scaling factors helps to capture time variation in risk premia. To approximate the time-varying risk premia, four conditioning variables, namely, the gross return on the MSCI world equity portfolio, MSCI world equity portfolio dividend yield, the Eurodollar deposit rate and the USA term premium are employed. All conditioning variables are lagged one period.

Hansen's (1982) J-statistic rejects the specification with the Eurodollar deposit rate as the conditioning variable but fails to reject the other model specifications. As a robustness check, the dynamic performance of the model specifications used is investigated. Using a method proposed by Farnsworth et al. (2002), I find that the specification that prices average returns best also performs the best in capturing the time series predictability in asset returns. That specification uses the USA term premium as the conditioning variable. This finding appears to support the Harvey's (1991) argument that USA macroeconomic variables can help explain equity returns in other countries.

Empirical results suggest that both the world market and foreign exchange risk factors, whether constant or time-varying, do not contribute significantly to the equity pricing kernel in the studied markets. As a risk factor, however, the world market factor is found to have significant time-varying risk premia. Further, there is strong evidence suggesting that foreign exchange rate changes have significant time-varying risk premia. For most of the model specifications examined, the Wald test for joint significance also rejects the hypothesis of zero foreign exchange rate risk in the seven stock markets. There are several implications of this finding. First, financial managers with exposed cash flows, liabilities, profits and assets must put in place measures to hedge their firms against fluctuations in foreign exchange rates. Second, portfolio managers must also ensure that their holdings in Africa's equity markets are adequately hedged against changes in foreign exchange rates. Importantly, security markets regulatory authorities, and other security markets agencies, in African countries must consider the wisdom of establishing well-functioning foreign exchange derivatives markets. With the exception of South Africa, all the countries studied have no markets for derivative securities.

CHAPTER FIVE

THE DYNAMIC RELATION BETWEEN FOREIGN EXCHANGE RATES AND THE FLOW OF INTERNATIONAL PORTFOLIO CAPITAL

5.1 Introduction

The preceding chapter has established that foreign exchange risk is conditionally priced in equity markets in Africa. The key questions that arise from this finding are: Do fluctuations in foreign exchange rates in Africa contain currency risk pricing information that informs the decisions of foreign portfolio investors? Do international portfolio flows influence movements in foreign exchange rates in a manner that causes them to attract risk premia in Africa's capital markets? Portfolio flows are chosen for this analysis for two reasons. First, portfolio flows, by their nature, are more temporary than foreign direct investments and are therefore likely to pose a bigger challenge to macroeconomic stability if they influence changes in real exchange rates. Second, the literature has concentrated on the relationship between real exchange rates and capital flows in general, leaving a knowledge gap on the role of portfolio flows, in isolation, in real exchange rates changes: filling this knowledge gap is made paramount by the fact that these relationships have not been studied in Africa. Thus, the analysis in this chapter tries to establish whether causality exists between the flow of international portfolio capital and real exchange rates in Africa's major capital markets.

5.2 Monthly Data

In this section, the bivariate behavior between international portfolio flows and changes in real foreign exchange rates is explored. The initial investigation takes the approach of a visual representation of the unconditional co-movement between the two data series. This is accomplished by using plots of the contemporaneous relationships between end-of-month real foreign exchange rate changes and the monthly net foreign portfolio flows, expressed as a percentage of the Gross Domestic Product (GDP), for four African countries whose monthly cash flows data are available. Portfolio flows data are obtained from US Treasury websites.[55]

[55] Portfolio flows data, available only for Egypt, Morocco and South Africa, are obtained from the USA Treasury Department website: http://www.ustreas.gov/tic/s1_other.tic. Data is also available at http://www.treas.gov/tic/afroils_57215.txt for oil-exporting African countries. Because of its relatively better developed financial markets, I assume that the bulk of the reported

Monthly GDP is calculated on the assumption that annual GDP is evenly spread throughout the year. Net foreign portfolio flows at a given point in time are defined as the arithmetic difference between purchases and sales, of foreign (African) bonds and stocks by US residents.

Real exchange rates are preferred because they capture both inflationary expectations and nominal exchange rate changes, both of which, theoretically, have a mutual relationship with international cash flows. However, it is important to point out that results would not differ if nominal exchange rates were used instead. The data reveals that there is a high correlation between nominal and real exchange rate changes: the correlation coefficients for monthly data are 0.984 (Botswana), 0.940 (Egypt), 0.834(Kenya), 0.951 (Morocco), 0.988 (Nigeria) and 0.994 (South Africa). For annual data, the correlation coefficients are: 0.991 (Botswana), 0.886 (Egypt), 0.920 (Kenya), 0.984 (Morocco), 0.993 (Nigeria) and 0.990 (South Africa).

5.2.1 Preliminary Analysis

The analysis begins by looking at some basic characteristics of the portfolio flows data. In particular, I analyze the relative volatility and persistence of international portfolio flows for the four countries: Egypt, Morocco, Nigeria and South Africa. For our purposes here, volatility is measured by the coefficient of variation (CV). A high coefficient of variation indicates high volatility of portfolio flows, and vice versa. Summary statistics for monthly portfolio flows for the four countries are provided in Table 26.

Table 26
Summary statistics for net monthly international portfolio flows

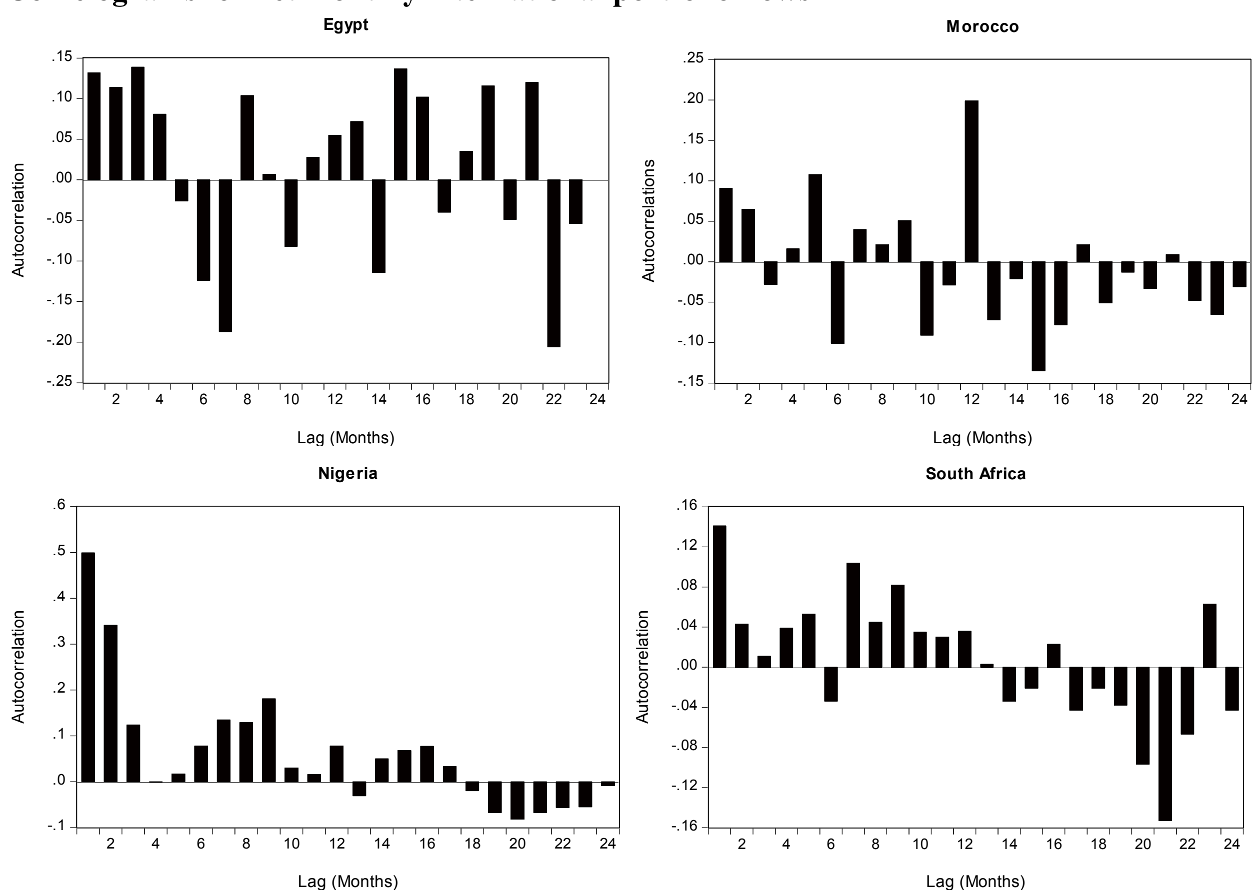
	Mean	Max	Min	Std Dev	CV	Skew	Kurt	Jarque-Bera
Egypt	-0.0207	0.0267	-0.0168	0.5094	24.5862	1.4815	11.9598	578.87***
Morocco	0.0256	0.0121	-0.0032	0.1520	5.9328	3.5871	27.0000	4078.56***
Nigeria	0.6330	0.0752	-0.0457	1.4400	2.2749	2.2847	11.7352	631.69***
South Africa	0.3624	0.1245	-0.0728	1.6186	4.4664	2.8274	28.0648	4291.44***

The table displays summary statistics for net international portfolio flows, as a percentage of GDP, for the period 1997:1 through 2009:12. Mean is in percentage; "Max" and "Min" respectively denote maximum and minimum observed values. "Std Dev" represents standard deviation (in percentage), "CV" the coefficient of variation, "Skew" denotes skewness, "Kurt" is the kurtosis. *** denotes rejection of the null hypothesis of normality at the 1 percent level of significance.

cash flows in this category are in respect of Nigeria and, accordingly, use the data as a proxy for Nigeria's portfolio flows. Froot et al. (2001) have discussed some of the weaknesses of these data.

Some interesting observations can be made from the summary statistics. First, portfolio flows appear not to contribute much to the countries' Gross Domestic Products (GDP): net portfolio flows to GDP are less than 1 percent for all countries. Second, for the 1997-2009 period, moderate to very high levels of portfolio flows volatility are observed. Egypt records the highest volatility in net international portfolio flows, with a variability of over 24 times the period's mean portfolio flows; Nigeria the lowest, with variability of about 2 times the mean portfolio flows. Further, portfolio volatilities for these countries are mostly higher than those recorded in more developed markets: for instance, for the 1976-1992 period, Claessens et al. (1995) find coefficients of variation of 1.22 for Japan, 1.48 for Germany, 1.49 for the USA and 2.35 for the UK; in this study, the lowest recorded CV is 2.27 for Nigeria.

Figure 6
Correlograms for net monthly international portfolio flows



In the literature, the notion of high volatility has been associated with incidents in which international investors get into a market for relatively short periods, largely to take advantage of temporary aberrations from parity conditions, and withdraw their investments at the slightest indication of a change in these conditions. Such investments have been labeled "hot money" in the literature. Third, the distribution of net portfolio flows cannot be described as normal in any of the countries, as attested to by the Jarque and Bera (1987) statistics, which are all significant.

Of more concern to many policy mandarins and researchers is whether capital flows are persistent or transitory in nature. The notion of low persistence has also been associated, in the literature, with hot money. Thus, it is interesting to examine whether net foreign portfolio flows to Africa exhibit persistence over time. A number of measures can be used for this purpose. The autocorrelation function is typically used as the first indicator of persistence. A persistent series is expected to be high and positively autocorrelated, whereas a transitory series should have a low or negative autocorrelation. The autocorrelation functions are presented in Figure 6. The figure demonstrates low persistence in portfolio flows for all the countries. In all cases, the autocorrelation at lag one is low (the maximum value is 0.499 for Nigeria) and the function remains positive for only a few lags thereafter. The autocorrelation functions for Egypt and Morocco change signs at frequent intervals, indicative of very low levels of persistence.

Table 27
Variance ratio statistics for net monthly international portfolio flows

Horizon (Months)	2	3	6	12	18	24	Joint test
Egypt	0.5165*** (-3.03)	0.3388*** (-2.92)	0.2299** (-2.28)	0.1042* (-1.93)	0.0751* (-1.65)	0.0619 (-1.46)	3.0333 [0.014]
Morocco	0.5202** (-2.00)	0.3831* (-1.92)	0.2121* (-1.88)	0.0815* (-1.68)	0.0784 (-1.48)	0.0615 (-1.39)	1.9979 [0.245]
Nigeria	0.6662** (-1.98)	0.5986* (-1.69)	0.3265* (-1.82)	0.1730 (-1.57)	0.1337 (-1.41)	0.0942 (-1.34)	1.9799 [0.254]
South Africa	0.5640** (-2.03)	0.3937** (-2.10)	0.2094** (-2.15)	0.1049** (-2.10)	0.0807** (-2.00)	0.0625* (-1.93)	2.1505 [0.175]

The table reports variance ratios for net portfolio flows using monthly data for the period 1997:1 through 2009:12. The statistics are corrected for bias in variance estimates. In round parentheses are heteroskedasticity-robust *z*-statistics; *p*-ratios are in square braces. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively.

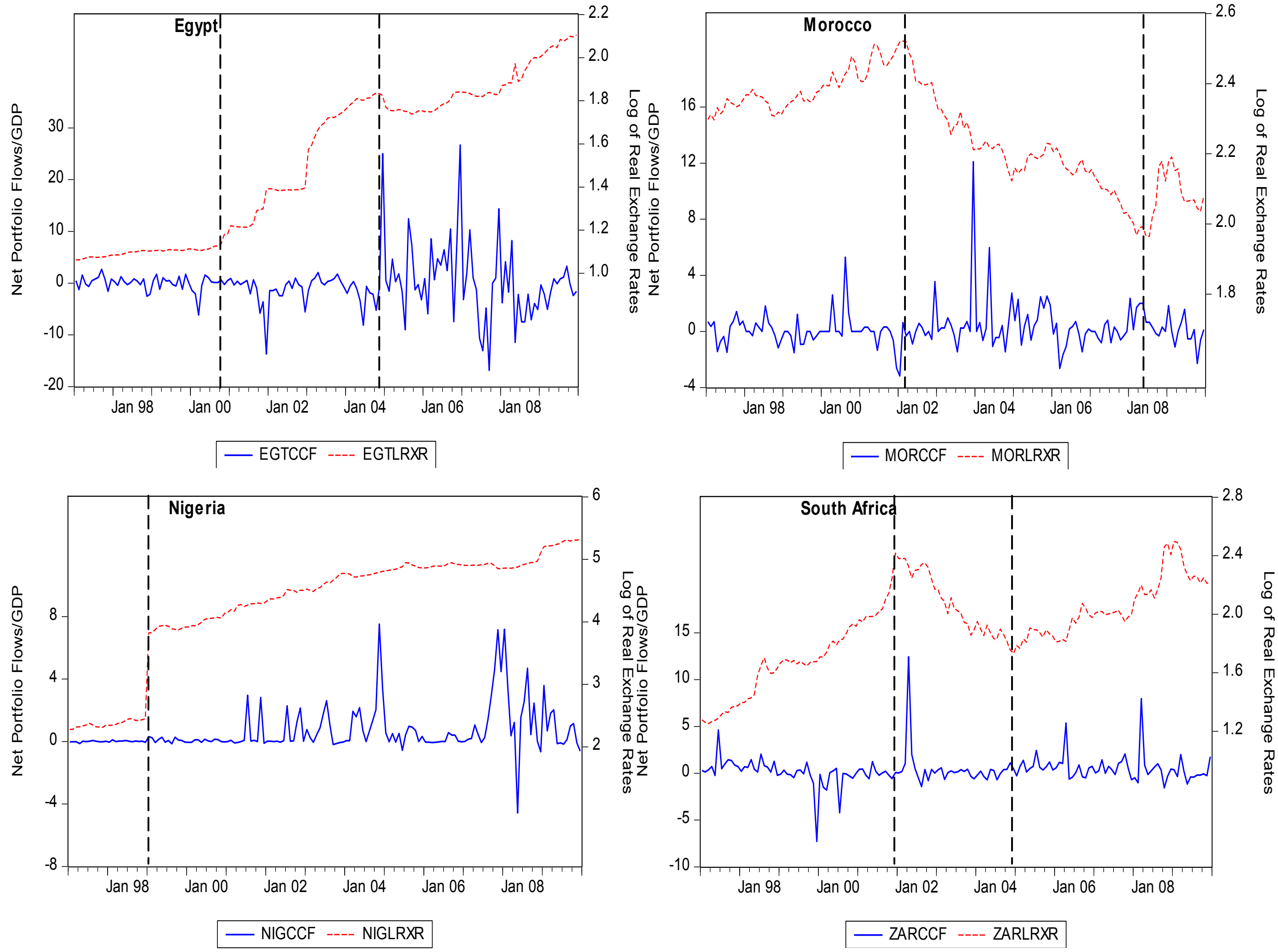
To corroborate the inferences drawn from the autocorrelation function, I perform the variance ratio test (Lo and MacKinlay, 1988, 1989). The form of the variance ratio statistic used here corrects for biasness in the variances, allows for drift and is heteroskedasticity consistent. The statistic compares the variance of monthly flows with the variance of flows measured over q = 2, 3, 6, 12, 18 and 24-month intervals. Results for the variance ratio tests are reported in Table 27. If the net portfolio flows follow a random walk so that they are completely random, each of the variance ratios should be close to 1.0. Numbers higher than 1.0 indicate positive serial correlation in net portfolio flows. Conversely, numbers less than one indicate, on average, negative serial correlation in net portfolio flows. As explained earlier, negative serial correlations in portfolio flows denote lack of persistence.

The table shows that all the point estimates are less than one and diminish strongly with horizon. From these results, the following inferences can be drawn: (i) From the Chow and Denning (1993) maximum $|z|$ joint-test statistic, the hypothesis that net portfolio flows follow a random walk cannot be rejected for all the countries, but Egypt. (ii) Net portfolio flows for all the four countries cannot be described as persistent. Since the standard errors have been adjusted for heteroskedasticity, one cannot attribute this result to changing variance of portfolio flows over time. (iii) Froot et al. (2001) observe that the incidence of high frequency persistence alone leads to variance ratios leveling out as horizons increase. In the current case, the leveling out of variance ratios with increasing horizons can be interpreted to mean that it is much more difficult to predict longer term portfolio flows based only on knowledge of current flows.

Next is a preliminary analysis of the nature of the relationship between net portfolio flows and real exchange rates for each of the four countries under investigation. Figure 7 presents a visual display of the basic relationships, across time, between monthly net portfolio flows, as a proportion of GDP, and real exchange rates, in logs. The figure suggests stationarity for levels of net portfolio flows, an observation that is consistent with existing empirical evidence (Edwards, 1998; Ndung'u and Ngugi, 1999; Froot et al., 2001). Real exchange rates appear non-stationary. For each of the four countries, the figure does not clearly indicate the nature of the relationship that exists between the two variables. However, correlation analysis suggests a weak inverse relationship between real rates of exchange and net foreign portfolio flows for Morocco (–

0.109). For South Africa (+0.002) and Egypt (–0.007) there seems not to be any co-movement between the two variables while results for Nigeria suggest a fairly strong positive relationship (+0.253). Since the direct quotation, in which increases in the exchange rate denote appreciation of the US dollar is used, a positive relationship implies that net portfolio flows increase (inflows outstrip outflows) when the local (African) currencies are weakening, and vice versa.

Figure 7
Contemporaneous co-movements between real exchange rates and net foreign portfolio flows



The surfix *LRXR* denotes natural logarithms of real foreign exchange rates; *CCF* denotes net foreign portfolio flows as a percentage of the Gross Domestic Product (GDP). Thus, *EGTLRXR* series represents natural logarithms of real foreign exchange rates between the Egyptian pound and the US dollar; *EGTCCF* series represents Egypt's net portfolio flows as a percentage of the corresponding Gross Domestic Products, and so on. Because they are very low in absolute terms, net portfolio flows to GDP ratios have been adjusted upwards by different powers of ten: Egypt, 10^3; Morocco, 10^3; Nigeria, 10^2; South Africa, 10^2. The dashed lines in the figure mark the visually estimated possible structural break points for real exchange rates. All observations run from 1997:1 to 2009:12.

A few other important observations emerge from Figure 7. First, portfolio flows appear to be a relatively new phenomenon in Africa's markets. The graphs, particularly those of Egypt and Nigeria, show a trend in which portfolio flows are initially sluggish and then become rather pronounced, both in volume and volatility, in the years after 2003. A similar, even if less pronounced, pattern is observed for net portfolio flows of Morocco and South Africa. For this reason, the study period is partitioned into two: the first sub-period begins January 1997 and ends December 2003; the second sub-period begins January 2004 and ends December 2009. The study then investigates causality for the full sample period as well as separately for two sub-periods.

Second, the graphs also illustrate the existence of structural breaks in the exchange rate series for each of the four countries: Egypt (end of 2000, end of 2003); Morocco (end of 2002, mid 2008); Nigeria (end of 1998) and South Africa (end of 2001 and end of 2004). For Nigeria, one of the devaluation events alluded to in Section 4.2.3 of Chapter 4 is clearly visible at the end of 1998. Structural breaks are also apparent for net portfolio flows of some countries: Egypt (end of 2003); and Nigeria (mid 2001, end of 2003 and mid 2006). In at least one case, structural breaks for portfolio flows and those of real exchange rates coincide, an observation that heightens the interest to empirically establish whether one of these variables causes the other.

Tests for Structural Breaks in Real Foreign Exchange Rates Series

To formally demonstrate the existence of structural breaks in the real exchange rate data series, I conduct the CUSUM tests on parameters from the regression of each series on its own trend and intercept. The CUSUM test, suggested by Brown et al. (1975), is based on the *cu*mulative *sum* of the recursive residuals. It plots the cumulative sum together with 5% critical lines. Parameter instability is found if the cumulative sum goes outside the area between the two critical lines. The CUSUM tests on real exchange rates for all countries indicate a case of variance instability or the presence of structural breaks (see Figure 8). In order to identify the specific points of break in the data series, I conduct a one-step forecast test. Again, the existence (and locations) of structural breaks is detected in each case (see the lower part of Figure 8). The upper portion of the plot (right vertical axis) reports the recursive residuals and standard errors. The lower portion (left vertical axis) shows the probability values for those sample points where the hypothesis of parameter constancy would be rejected at the 5, 10, or 15 percent levels. The points with p-values less than 0.05 correspond to those points where the recursive residuals go outside the two standard error bounds. Graphs of the one-step forecast tests show that many of the countries' exchange rates have multiple structural breaks over the study period.

5.2.2 Unit Root Tests Results

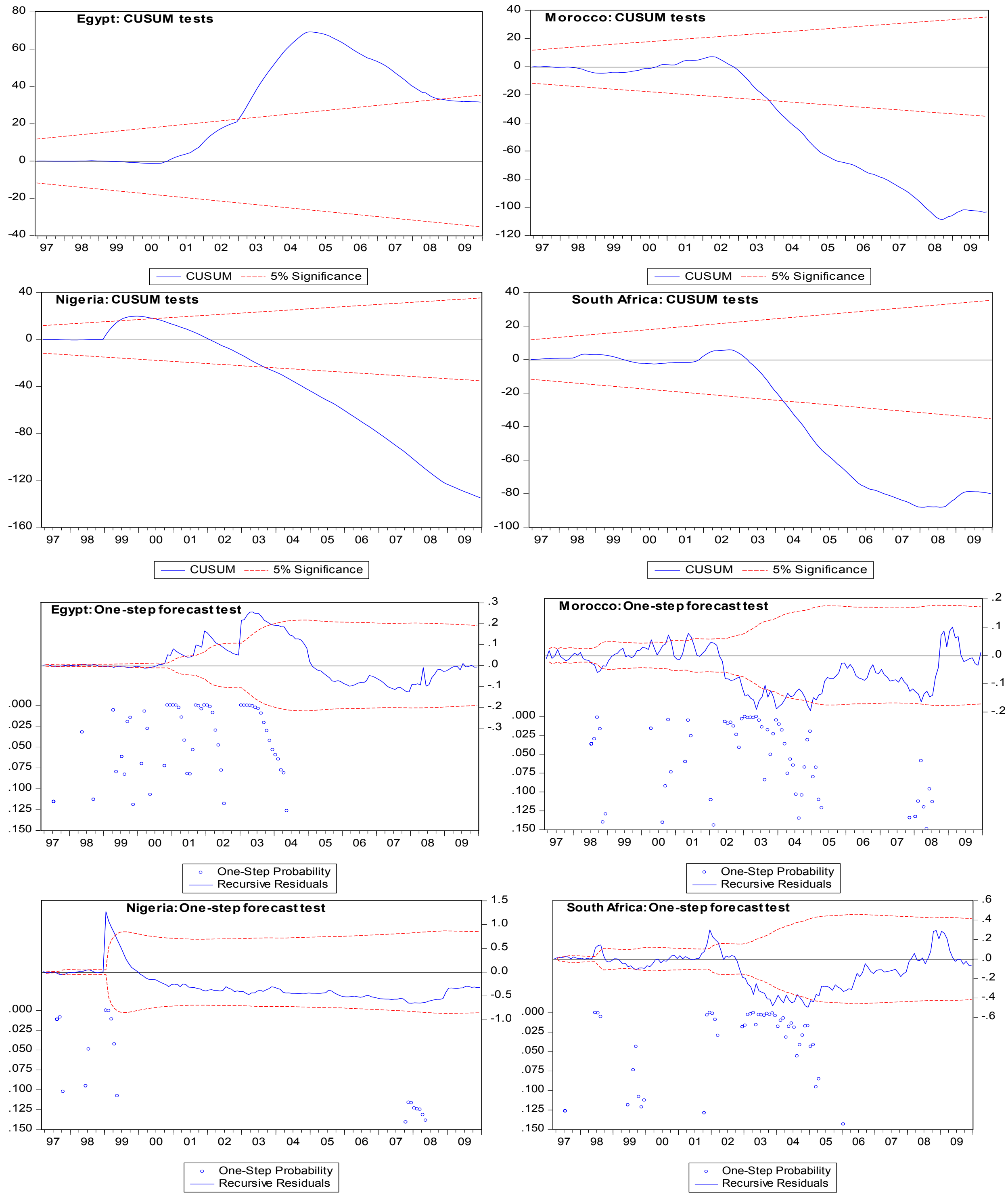
The existence of structural breaks implies that traditional unit root tests may not yield reliable results. Perron (1990) observes that structural breaks (arising from such events as oil shocks and currency crises) in a stationary series may bias the usual unit root tests in favor of accepting the null hypothesis i.e., that there is unit root in the series. He proposes the inclusion of a dummy variable at the point of the structural break. However, Zivot and Andrews (1992) observe that including a dummy variable at the point of observed structural break is tantamount to "pre-testing" the data, a fact that is itself problematic in empirical research. They show that Perron's (1990) method does not adequately deal with the systematic bias in favor of accepting the null in unit roots testing. Accordingly, Zivot and Andrews (1992) propose the following method:

$$\Delta Y_t = \beta_1 + \beta_2 t + \alpha Y_{t-1} + \gamma DU_t(\lambda) + \sum_{i=1}^{k-1} \theta_i \Delta Y_{t-i} + \varepsilon_t \qquad (66)$$

where $\Delta Y_t = Y_t - Y_{t-1}$ is the first difference of the random variable Y at time t; ε_t is the white noise term at time t, $DU_t(\lambda) = 1$ for $t > T\lambda$, otherwise $DU_t(\lambda) = 0$; $\lambda = T_B/T$ represents the location of the structural break. T is the number of observations; T_B, the date when the structural break occurred; and k is the lag length. To capture as much information as possible, this model is tested here within the range $0.03 \leq \lambda \leq 0.97$; the value of T_B for which the ADF t-statistic is minimized is selected. In this study, the Zivot and Andrews (1992) method is justified by the fact that the data exhibit multiple structural breaks as is evident from the foregoing tests.

Results of the unit root tests are presented in Table 28. From Panel A, Zivot and Andrews (1992) tests results show that the hypothesis that a unit root exists in the series cannot be rejected for foreign exchange rate for all countries except Nigeria. On the contrary, the net portfolio flows for all the countries are stationary. Stationarity in capital flows series is not surprising: similar results have been reported elsewhere by Edwards (1998) and Ndung'u and Ngugi (1999).

Figure 8
Structural breaks tests results for monthly real exchange rates series



I also conduct unit root tests using the traditional Augmented Dickey-Fuller method. Results are reported in Panel B of Table 28. The results show that, for all countries, the unit roots hypothesis cannot be rejected in the exchange rate series in levels but is rejected in first differences. Thus, as predicted by theory, exchange rates for the four major African countries are integrated of order one, *I*(1). However, all the portfolio flows appear to be integrated of order zero, *I*(0). Because exchange rates and portfolio flows are not integrated of the same order, it is not expected that a long-run equilibrium exists between them and hence, cointegration tests are not indicated.

Table 28
Unit root tests results

	Egypt	Morocco	Nigeria	South Africa
Panel A: Zivot and Andrews (1990) Unit Root Tests				
Foreign Exchange Rates (Level)	–3.87 [0.46]	–3.19 [0.46]	–13.89*** [0.16]	–3.39 [0.45]
Foreign Exchange Rates (First Difference)	–12.30*** [0.46]	–10.80*** [0.46]	–12.58*** [0.16]	–11.15*** [0.45]
Net Portfolio Flows (Level)	–11.67*** [0.80]	–11.73*** [0.46]	–7.84*** [0.95]	–10.98*** [0.40]
Net Portfolio Flows (First Difference)	–20.94*** [0.80]	–20.83*** [0.46]	–17.55*** [0.95]	–19.64*** [0.40]
Panel B: Augmented Dickey and Fuller (1979) Unit Root Tests				
Foreign Exchange Rates (Level)	–1.68 (0.75)	–2.30 (0.43)	–1.92 (0.64)	–1.76 (0.72)
Foreign Exchange Rates (First Difference)	–12.02*** (0.00)	–10.74*** (0.00)	–12.40*** (0.00)	–10.83*** (0.00)
Net Portfolio Flows (Level)	–10.80*** (0.00)	–11.28*** (0.00)	–7.52*** (0.00)	–10.68*** (0.00)
Net Portfolio Flows (First Difference)	–11.64*** (0.00)	–11.21*** (0.00)	–17.55*** (0.00)	–8.80*** (0.00)

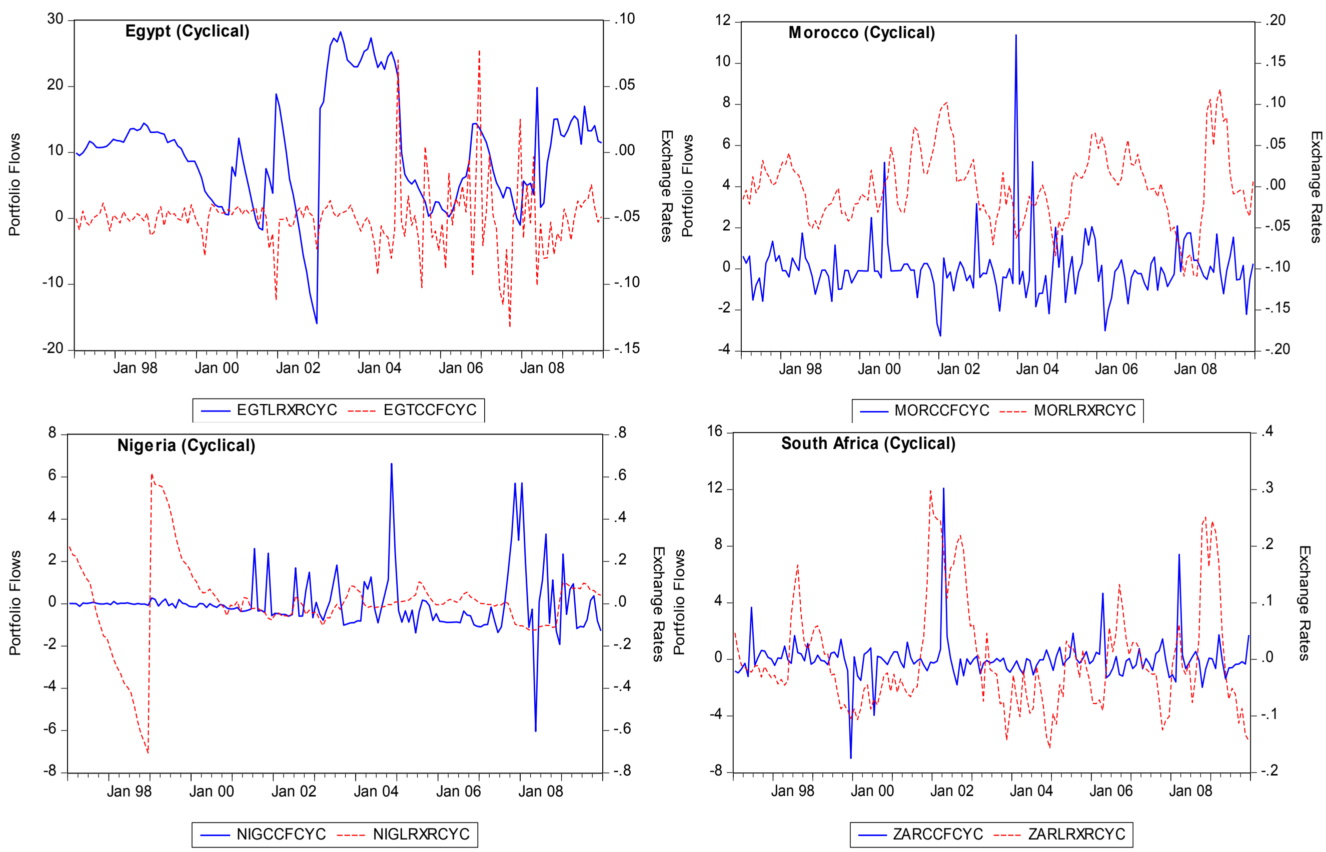
The table shows the results of unit root tests for monthly data (1997:1 to 2009:12). Values of λ for Zivot and Andrews (1992) tests are shown in square braces; *p*-ratios for the Dickey and Fuller (1979) tests are provided in round brackets. *** indicates the rejection of a unit root hypothesis at the 1% level. Asymptotic critical values for Panel A, obtained from Table 2(A) of Zivot and Andrews (1992) are –5.34, –4.80 and –4.58 respectively at the 1%, 5% and 10% significance levels. Asymptotic critical values for Panel B, based on MacKinnon (1996), are –4.018, –3.439 and –3.144 respectively at the 1%, 5% and 10% significance levels.

5.2.3 Granger Causality Tests Results

Despite the existence of structural breaks, traditional Granger causality tests are sufficient for studying the relationship (Granger et al., 2000), in this case between foreign exchange rates and net portfolio flows. As already explained, Granger causality tests require that all the series used be stationary. In this regard, these tests are typically conducted in the literature using first differences of the series involved. This is because many economic variables exhibit the *I*(1) characteristic and the tests usually involve one or more series with the same order of integration.

In the case of the current investigation, foreign exchange rates and portfolio flows are not integrated of the same order, as already observed. Because of this, I follow Edwards (1998) and conduct the causality tests using series detrended through the Hodrick and Prescott (1997) filter. All detrended series exhibit stationarity. The resulting detrended series are displayed in Figure 9. Some observations can be made from the figure. Visual investigation of the graph for Egypt presents a clearly discernible co-movement pattern in the first half of the study period. For Morocco and Nigeria, it is noticeable that the two series tend to move together in the second half of the period. Co-movement is apparent for the entire period in the case of South Africa.

Figure 9
Detrended monthly real foreign exchange rates and net international portfolio flows



The surfix *LRXRCYC* denotes natural logarithms of real foreign exchange rates, detrended; *CCFCYC* denotes net foreign portfolio flows as a percentage of the Gross Domestic Product (GDP), detrended. Thus, *EGTLRXRCYC* series represents cyclical components of natural logarithms of real foreign exchange rates between the Egyptian pound and the US dollar; *EGTCCFCYC* series represents cyclical components of Egypt's net portfolio flows as a percentage of the corresponding Gross Domestic Products, and so on. All observations run from 1997:1 to 2009:12.

Table 29
Granger causality tests results

		F-statistic	p-ratio
Panel A: Full Sample			
Egypt	Real exchange rates does not Granger-cause portfolio flows	0.7040	0.4028
	Portfolio flows does not Granger-cause real exchange rates	0.0074	0.9314
Morocco	Real exchange rates does not Granger-cause portfolio flows	2.0825	0.1511
	Portfolio flows does not Granger-cause real exchange rates	0.0200	0.8878
Nigeria	Real exchange rates does not Granger-cause portfolio flows	0.0903	0.7642
	Portfolio flows does not Granger-cause real exchange rates	0.5379	0.4644
South Africa	Real exchange rates does not Granger-cause portfolio flows	4.3705**	0.0382
	Portfolio flows does not Granger-cause real exchange rates	2.7534*	0.0991
Panel B: January 1997 to December 2003			
Egypt	Real exchange rates does not Granger-cause portfolio flows	1.2595	0.2651
	Portfolio flows does not Granger-cause real exchange rates	3.5267*	0.0640
Morocco	Real exchange rates does not Granger-cause portfolio flows	0.9686	0.3280
	Portfolio flows does not Granger-cause real exchange rates	0.8069	0.3717
Nigeria	Real exchange rates does not Granger-cause portfolio flows	0.0425	0.8372
	Portfolio flows does not Granger-cause real exchange rates	0.0891	0.7661
South Africa	Real exchange rates does not Granger-cause portfolio flows	5.5942**	0.0204
	Portfolio flows does not Granger-cause real exchange rates	1.9074	0.1710
Panel C: January 2004 to December 2009			
Egypt	Real exchange rates does not Granger-cause portfolio flows	2.3804	0.1275
	Portfolio flows does not Granger-cause real exchange rates	0.7361	0.3939
Morocco	Real exchange rates does not Granger-cause portfolio flows	3.2311*	0.0767
	Portfolio flows does not Granger-cause real exchange rates	1.6747	0.2000
Nigeria	Real exchange rates does not Granger-cause portfolio flows	0.3533	0.5542
	Portfolio flows does not Granger-cause real exchange rates	5.6929**	0.0198
South Africa	Real exchange rates does not Granger-cause portfolio flows	0.1006	0.7521
	Portfolio flows does not Granger-cause real exchange rates	1.9632	0.1657

* and ** denotes that the null hypothesis has been rejected at the 10% and 5% levels respectively.

I conduct Granger causality test in levels, using the following bivariate vector autoregression:

$$\begin{aligned} NPF_t &= \beta_0 + \textstyle\sum_{k=1}^{L} \beta_{1k} NPF_{t-k} + \sum_{k=1}^{L} \beta_{2k} RER_{t-k} + \varepsilon_{NPF} \\ RER_t &= \delta_0 + \textstyle\sum_{k=1}^{L} \delta_{1k} RER_{t-k} + \sum_{k=1}^{L} \delta_{2k} NPF_{t-k} + \varepsilon_{RER} \end{aligned} \tag{67}$$

where ε_i (i = NPF, RER) denote the white noise processes and L is the number of lagged terms, chosen using the Schwarz (1978) Bayesian information criteria. The null hypothesis that RER

does not Granger-cause NPF is rejected if β_{2k} are jointly significant; the null hypothesis that NPF does not Granger-cause RER is rejected if δ_{2k} are jointly significant. The null hypotheses are tested using the F-statistic of the Granger causality tests.

Results of the tests are reported in Table 29. Results for the full observation period are shown in Panel A. No evidence of a causal relation is found between net portfolio flows and real exchange rates for all countries except South Africa, which exhibits a bi-directional causation. The first sub-period (January 1997 to December 2003) results, reported in Panel B, provide evidence of a unidirectional causal relation from net portfolio flows to real exchange rates for Egypt, and from real exchange rates to net portfolio flows for South Africa. No evidence of causality is found in respect of the other countries. In Panel C, where tests results for the second sub-period (January 2004 to December 2009) are reported, unidirectional causality is established from real exchange rates to net portfolio flows for Morocco, and from net portfolio flows to real exchange rates for Nigeria. Other countries exhibit no evidence of causality in respect of this sub-period.

Overall, these results do not indicate a clear direction of causality between real exchange rates and net portfolio flows that can be generalized across all the countries investigated. Rather, causality between the two macroeconomic variables is both country-specific and time-varying. The country-specific nature of the relationship can be explained by idiosyncrasies around attractiveness of the specific capital markets to foreign investors, their perceived openness to international capital flows, levels of state control, legal structures and accounting practices. Thus, the South African capital market, which exhibits the highest level of sophistication among the markets investigated, tends to be perceived by investors as more accessible and is therefore much more likely to reflect greater awareness of foreign investors than the other markets. As such, the decisions of international portfolio investors in South Africa's markets are more likely to be guided by changes in macroeconomic fundamentals including fluctuations in foreign exchange rates. Similarly, because of its relatively more flexible foreign exchange management policy, exchange rates fluctuations may draw more closely from market forces as dictated by the flow of cross-border capital. This may explain the finding of bidirectional causality between foreign exchange rates and portfolio flows for South Africa over the full sample period.

Country idiosyncrasies may also explain the time-varying nature of causality between the two variables in some countries. The unidirectional causality from net portfolio flows to real exchange rates found for Egypt and Nigeria for the first and second sub-periods respectively can be attributed to policy changes in which the exchange rate is revalued or devalued in response to fluctuations in international capital flows. This explanation draws from the presentation in Section 1.1.1 (Chapter 1), in which changes in the foreign exchange policy in the sampled countries are briefly documented. The Nigerian currency, the naira, was formally pegged between 1997 and 2009 during which period a pro-rata system of foreign exchange allocation to end-users was in place. Similarly, Egypt changed its exchange rate policy from a managed float in 1997, announced a new exchange rate of Egyptian pound 3.85 per US dollar and introduced a "crawling peg" system, which was in place throughout the first sub-period, when causality has been established. These foreign exchange rate systems typically mandate monetary authorities to make periodic adjustments to the exchange rate if it is deemed to have significantly deviated from the officially acceptable levels/band following pressures from market forces such as international capital flows. Indeed, preliminary analysis of foreign exchange rates data (Section 4.2.3, Chapter 4) clearly present a picture in which monetary authorities of these two countries are actively involved in the foreign exchange market.

5.2.4 Vector Autoregressions (VARs)

The foregoing issues can be clarified by understanding the dynamic interactions between real exchange rate fluctuations and net portfolio flows. The clarification is done by the use of vector autoregressions (VARs). VARs are dynamic equations that examine the inter-relationships between two, or more, economic variables with minimal assumptions about the underlying structure of the economy. A VAR answers the question as to whether variable X predicts changes in variable Y over and above the predictions already provided by lagged values of Y, and vice versa. One might be interested in knowing, for instance, whether foreign portfolio flows provide additional information that can explain changes in real foreign exchange rates beyond information already provided by lagged real exchange rates. Following Froot et al. (2001), I delve into this subject by estimating a system of two equations capturing the joint dynamics of net portfolio flows and real exchange rates, for each country, in the following unrestricted VAR.

$$\begin{bmatrix} NPF_{it} \\ RER_{it} \end{bmatrix} = \begin{bmatrix} \alpha_{10} \\ \alpha_{20} \end{bmatrix} + \begin{bmatrix} \emptyset_{11}(L) & \emptyset_{12}(L) \\ \emptyset_{21}(L) & \emptyset_{22}(L) \end{bmatrix} \cdot \begin{bmatrix} NPF_{i,t-1} \\ RER_{i,t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{it}^{NPF} \\ \varepsilon_{it}^{RER} \end{bmatrix} \tag{68}$$

where NPF_{t-1} represents lagged values of net portfolio flows; RER_{t-1} denotes lagged values of real exchange rates; the error terms ε_{it}^{j} $(j = NPF, RER)$ are assumed to be normally distributed with mean zero and variance-covariance matrix Σ_i, α_{j0} are country-specific constants and $\emptyset_{jk}$ (k = 1, 2) are parameters to be estimated, for all i = 1, … N; t = 1, …, T; L is the lag operator. For each country i, the variance-covariance matrix is defined as:

$$\Sigma_i = \begin{pmatrix} \sigma_{i,NPF}^2 & \rho\sigma_{i,NPF}\sigma_{i,RER} \\ \rho\sigma_{i,NPF}\sigma_{i,RER} & \sigma_{i,RER}^2 \end{pmatrix} \tag{69}$$

where ρ is the correlation coefficient between the two error terms for each country i. The diagonal coefficients $\emptyset_{11}$ and $\emptyset_{22}$ capture the extent to which each variable can be predicted by its own lagged values. The off-diagonal coefficients $\emptyset_{12}$ and $\emptyset_{21}$capture the dynamic interaction between one variable and the other with non-zero values implying causality. Since all the Hodrick-Prescott filtered variables exhibit stationarity, the VAR is estimated in levels. The number of lags (p = 1) is chosen using the Schwarz (1978) Bayesian Information criteria, as reported in Appendix II. Results of the VAR estimation are presented in Table 30.

The table shows that, for all countries, lagged values of net portfolio flows and lagged values of real exchange rates jointly explain not less than two-thirds of the movement in real exchange rates in both the full sample and the sub-periods (see the adjusted R^2 column). However, it is also clear that the bulk of that predictive information comes from the lagged values of real exchange rates themselves. Lagged values of portfolio flows provide predictive information for real exchange rates only for South Africa (consistently) and for Egypt (first sub-period) and Nigeria (second sub-period). Net portfolio flows, on the other hand, cannot be predicted by the two variables: the adjusted coefficients of variation are very low (sometimes negative) in all cases. With few exceptions, lagged portfolio flows also seem not to provide adequate information that can be used by policy makers/ investors to form an opinion about next period's flows. The few exceptions are Nigeria and South Africa (for the full sample, 15.48% and 1.98%, respectively); Egypt (first sub-period, 4.45%) and Nigeria (second sub-period, 17.87%). Similarly, fluctuations in real exchange rates do not appear to be capable of explaining future changes in portfolio flows. Thus, consistent with the Granger-causality findings, there appears not to be strong evidence of causality either from net portfolio flows to real exchange rates or from real exchange rates to net portfolio flows that can be generalized across Africa's capital markets. Instead, the evidence continues suggesting that causality is time-varying and country-dependent.

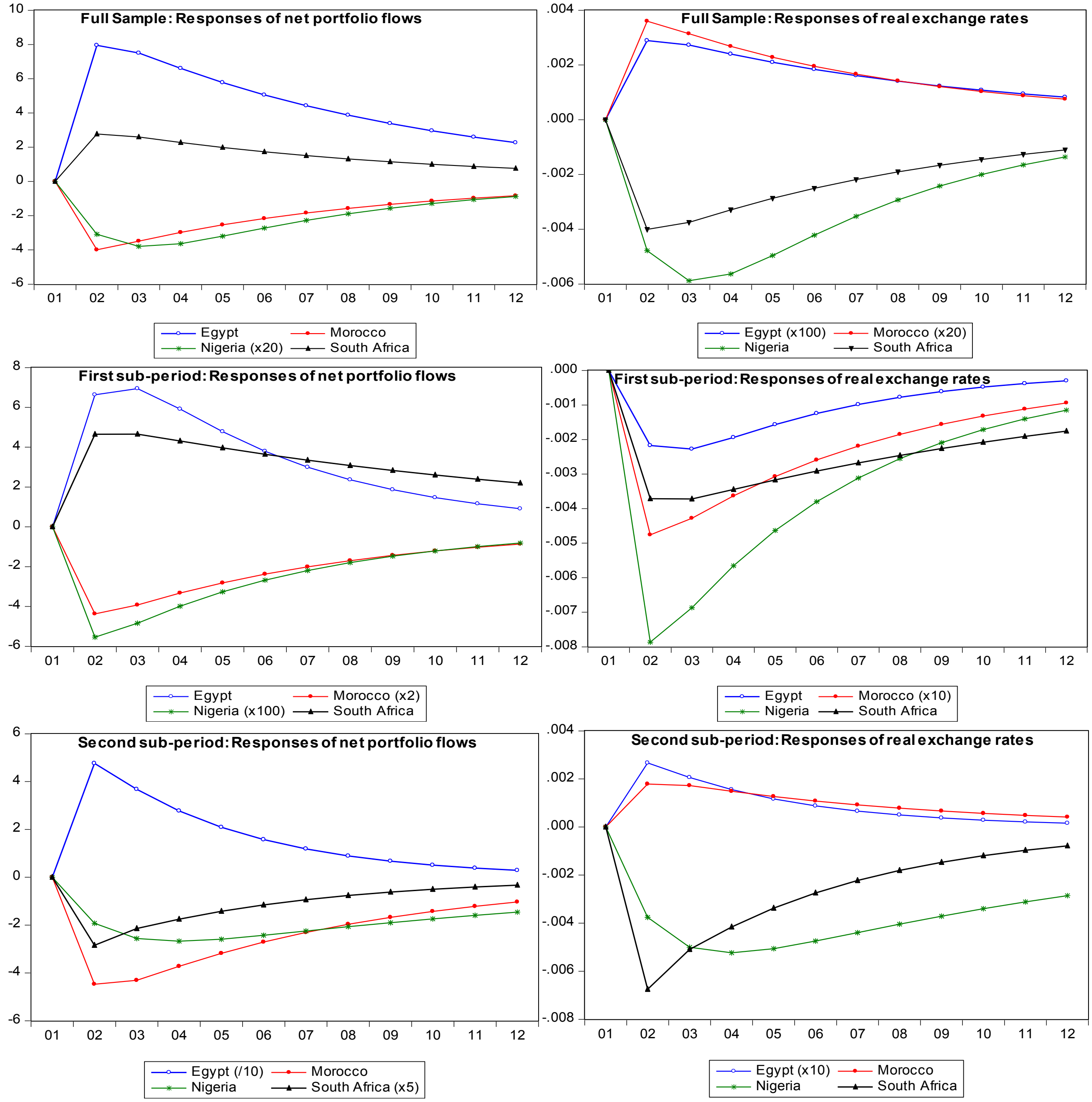
Table 30
Unrestricted Vector Autoregression results for real exchange rates and net portfolio flows

		Constant		NPF(-1)		RER(-1)			Corr
	Dependent variable	Coef	*t*-stat	Coef	*t*-stat	Coef	*t*-stat	Adj R^2	$(\varepsilon_1, \varepsilon_2)$
Panel A: Full sample period									
Egypt	Net portfolio flows	0.0007	(0.00)	0.0685	(0.85)	7.9523	(0.84)	-0.39	-0.145
	Real exchange rates	0.0000	(0.03)	0.0000	(0.09)	0.8747	(22.24)***	76.18	
Morocco	Net portfolio flows	-0.0041	(-0.03)	0.0201	(0.25)	-3.9861	(-1.44)*	0.22	-0.142
	Real exchange rates	0.0001	(0.07)	0.0002	(0.14)	0.8535	(19.83)***	72.35	
Nigeria	Net portfolio flows	-0.0035	(-0.04)	0.4069	(5.47)***	-0.1542	(-0.30)	15.48	0.010
	Real exchange rates	-0.0015	(-0.17)	-0.0048	(-0.73)	0.8224	(18.27)***	68.43	
South Africa	Net portfolio flows	0.0033	(0.03)	0.0504	(0.63)	2.7708	(2.09)**	1.98	-0.010
	Real exchange rates	-0.0012	(-0.31)	-0.0040	(-1.66)**	0.8859	(22.29)***	76.29	
Panel B: First sub-period (January 1997 to December 2003)									
Egypt	Net portfolio flows	-0.0010	(-0.00)	0.2332	(2.16)***	6.6305	(1.12)	4.45	-0.165
	Real exchange rates	0.0002	(0.07)	-0.0022	(-1.88)**	0.8141	(12.86)***	67.20	
Morocco	Net portfolio flows	0.0013	(0.01)	0.0540	(0.34)	-2.1849	(-0.44)	-2.10	-0.243
	Real exchange rates	-0.0006	(-0.25)	-0.0005	(-0.23)	0.8437	(13.13)***	67.58	
Nigeria	Net portfolio flows	-0.0007	(-0.01)	0.0546	(0.49)	-0.0554	(-0.21)	-2.15	0.064
	Real exchange rates	-0.0026	(-0.16)	-0.0079	(-0.30)	0.8190	(13.02)***	67.14	
South Africa	Net portfolio flows	-0.0011	(-0.01)	0.0625	(0.57)	4.6465	(2.37)***	5.42	-0.096
	Real exchange rates	-0.0026	(-0.57)	-0.0037	(-1.38)*	0.9394	(19.48)***	82.38	
Panel C: Second sub-period (January 2004 to December 2009)									
Egypt	Net portfolio flows	0.0196	(0.02)	0.0365	(0.31)	47.6390	(1.54)*	0.76	-0.135
	Real exchange rates	0.0003	(0.15)	0.0003	(0.86)	0.7357	(9.12)***	54.26	
Morocco	Net portfolio flows	0.0092	(0.06)	0.1015	(0.85)	-4.4717	(-1.46)*	1.84	-0.001
	Real exchange rates	0.0002	(0.07)	0.0018	(0.72)	0.8623	(13.60)***	72.55	
Nigeria	Net portfolio flows	0.0076	(0.04)	0.4303	(3.87)***	-1.9228	(-0.59)	17.87	-0.101
	Real exchange rates	0.0002	(0.09)	-0.0038	(-2.39)***	0.9009	(19.67)***	86.12	
South Africa	Net portfolio flows	0.0032	(0.02)	-0.0536	(-0.44)	-0.5685	(-0.32)	-2.54	-0.179
	Real exchange rates	-0.0035	(-0.58)	-0.0068	(-1.40)*	0.8077	(11.52)***	66.22	

"Coef" is the regression coefficient; "*t*-stat" denotes the *t*-statistic and Adj R^2 is the adjusted coefficient of determination, in percentage; Corr $(\varepsilon_1, \varepsilon_2)$ is the correlation coefficient between the error terms from the two equations. NPF(-1) and RER(-1) respectively represent the lagged values of net portfolio flows (as a proportion of GDP) and real exchange rates, in logs.

To further examine the short-run dynamic relations between the two variables in Africa's capital markets, I conduct innovation accounting through impulse response functions and forecast error variance decomposition. Figure 10 reports the results of impulse response experiments.

Figure 10: Impulse responses of one variable to one unit shock in error terms of the other variable



A number of observations can be made from Figure 10. First, a one-unit shock in net portfolio flows has little effect on the cyclical component of bilateral real exchange rates for all the countries. Second, as theory suggests, the little effect exerted on bilateral real exchange rates by a one-unit shock in portfolio flows is largely in the form of an appreciation. This is the case for all the countries in the first sub-period. In the full sample and the second sub-period, Egypt and Morocco, disobey this theoretical regularity. Third, one-unit innovations in bilateral real exchange rates elicit large responses on net portfolio flows for all the countries and particularly for Egypt. For Nigeria, however, portfolio flow responses are subdued in the first sub-period, perhaps a manifestation of the fact that the abnormal behavior of exchange rates arising from the observed devaluation of the nominal exchange rate at the end of 1998 did not send a signal to investors. In the second sub-period, the relative stability in Nigerian naira-US dollar exchange rates appears to surprisingly misinform international investment decisions, with depreciations in the naira leading to diminished inflows. It is also surprising that one-unit shocks in real exchange rates cause a decline in net portfolio flows for Morocco and Nigeria (all periods) and South Africa (second sub-period). The observation that shocks in real exchange rates lead to a reduction in net portfolio flows does not find support in theory. Theoretically, depreciations of the local currency are expected to make local assets less expensive from the perspective of foreigners, who should ideally respond by purchasing more of those assets.

Fourth, shocks in real exchange rates are followed by increments in net portfolio flows in the full sample and the first sub-periods for South Africa, and in all the periods for Egypt. Since an increment in net portfolio flows implies that inflows exceed outflows for that particular period, the implication is that foreign investors to these countries tend to increase their portfolio investments when the US dollar appreciates. This result is consistent with theoretical predictions. In general, results from impulse response functions conform to Granger-causality tests results: it is not possible to make conclusive inferences that can be generalized across the countries, about the nature of causality that exists between the flow of international portfolio capital and foreign exchange rates, using monthly data. Rather, each country experiences a relationship between the two variables that is dependent on country idiosyncrasies and changes in time.

A Digression: Further Evidence on Persistence of Portfolio Flows

To provide further evidence on the persistence of net portfolio flows, I use impulse responses to investigate how one-unit shocks on error terms of net portfolio flows propagate themselves on net portfolio flows over time. As Claessens et al. (1995) point out, if a time series is highly positively autocorrelated, it will take a long time for a shock to die out; if the autocorrelations are low, the shock should vanish quickly. Results, displayed in Table 31, show that the half life (the amount of time it takes for the shock to lose half or more of its initial value) of portfolio flows is approximately one month i.e. it takes a month or less for the effect of a shock on portfolio flows to dissipate. The short memory of portfolio flows makes it reasonable to infer, once again, the absence of persistence in portfolio flows into and out of these countries. This appears to support the presence of the "hot money" phenomenon in these African markets.

Table 31
Impulse responses of net portfolio flows to one unit shocks in their own error terms

Month	1	2	3	4	5	6	7	8	9	10	11	12
Panel A: Full sample												
Egypt	1	0.068	0.005	0.001	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Morocco	1	0.020	0.000	-0.001	-0.001	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Nigeria	1	0.407	0.166	0.069	0.029	0.012	0.006	0.003	0.002	0.001	0.001	0.001
S. Africa	1	0.050	-0.009	-0.011	-0.010	-0.008	-0.007	-0.006	-0.006	-0.005	-0.004	-0.004
Panel B: First sub-period (January 1997 to December 2003)												
Egypt	1	0.233	0.040	-0.006	-0.014	-0.014	-0.011	-0.009	-0.007	-0.006	-0.005	-0.004
Morocco	1	0.054	0.004	0.001	0.001	0.001	0.001	0.001	0.000	0.000	0.000	0.000
Nigeria	1	0.055	0.003	0.001	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
S. Africa	1	0.062	-0.013	-0.018	-0.017	-0.016	-0.015	-0.013	-0.012	-0.011	-0.010	-0.010
Panel C: Second sub-period (January 2004 to December 2009)												
Egypt	1	0.037	0.014	0.010	0.008	0.006	0.004	0.003	0.003	0.002	0.001	0.001
Morocco	1	0.101	0.002	-0.007	-0.007	-0.006	-0.005	-0.005	-0.004	-0.003	-0.003	-0.002
Nigeria	1	0.430	0.192	0.092	0.050	0.031	0.023	0.018	0.016	0.014	0.012	0.011
S. Africa	1	-0.054	0.007	0.003	0.002	0.002	0.001	0.001	0.001	0.001	0.001	0.001

Forecast Error Variance Decomposition

A major shortcoming of the impulse response analysis arises from the high correlation in the residual covariance matrix. If the high correlation is due to omitted variables, it is very difficult to interpret the residuals as a possible measure of autonomous errors (structural shocks). A large residual correlation coefficient does not necessarily imply a structural simultaneous effect nor does it imply incorrectly specified expectations (Juselius, 2006). I have, as a robustness check, attempted to deal with the problem of omitted variables by incorporating into the VAR other variables that are theoretically believed to influence both capital flows and exchange rates. The outcomes of that check are discussed in the next sub-section. In the current on-going discussion, I utilize error variance decomposition techniques to shed light on the extent to which the error variances of the two equations in the vector autoregressions can be attributed to the influence of each of the two endogenous variables. Results are presented in Table 32.

For the full sample, error variances for each of the variables are explained largely by their own innovations. Consistent with theory (see Enders, 2004: 280), each variable explains smaller proportions of its own forecast error variances at longer horizons, save for Morocco (Panel A2). The South African case is worth mentioning: at the twelve-month forecast horizon, the proportion of forecast error variances for portfolio flows explained by innovations in real exchange rates for the full sample is 2.948% (Panel A1) and the proportion of real exchange rates error variance explained by portfolio flow shocks is 5.115% (Panel A2). This explains the finding that the two variables Granger-cause each other in South Africa for the full sample.

Variance error decomposition for the sub-periods presents interesting findings. In the first sub-period, 2.731percent of the one-month horizon forecast error variance for Egypt's real exchange rates are explained by innovations in net portfolio flows (Panel B2). The figure increases steadily to 13.838 percent in the twelve-month horizon. For South Africa, 7.507 percent of the twelve-month horizon error variance of portfolio flows is explained by innovations in real exchange rates (Panel B1). In the second sub-period, 21.534 percent of twelve-month forecast error variance of real exchange rates for Nigeria is explained by shocks on portfolio flows (Panel C2). These findings are consistent with the Granger-causality tests results, which report a significant unidirectional causality from net portfolio flows to real exchange rates in the first sub-period for Egypt, a unidirectional causal relationship from real exchange rates to portfolio flows for South Africa in the first sub-period and a unidirectional causal relation from net portfolio flows to real exchange rates for Nigeria in the second sub-period.

Table 32
Decomposition of forecast error variances for monthly net international portfolio flows and real exchange rates

	Error variance of net portfolio flows explained by:								Error variance of real exchange rates explained by:							
	NPF	RER	NPF	RER	NPF	RER	NPF	RER	NPF	RER	NPF	RER	NPF	RER	NPF	RER
Horizon	Egypt		Morocco		Nigeria		South Africa		Egypt		Morocco		Nigeria		South Africa	
	Panel A1: Full sample								Panel A2: Full sample							
1	100.000	0.000	100.000	0.000	100.000	0.000	100.000	0.000	2.102	97.898	2.002	97.998	0.010	99.990	0.997	99.003
2	99.894	0.106	99.624	0.376	99.984	0.016	99.322	0.678	2.007	97.993	1.851	98.149	0.134	99.866	3.246	96.754
3	99.800	0.200	99.338	0.662	99.962	0.038	98.736	1.264	1.971	98.029	1.801	98.199	0.275	99.725	4.075	95.925
4	99.727	0.273	99.132	0.868	99.941	0.059	98.290	1.710	1.954	98.046	1.777	98.223	0.381	99.619	4.472	95.528
5	99.671	0.329	98.982	1.018	99.924	0.076	97.953	2.047	1.944	98.056	1.763	98.237	0.451	99.549	4.697	95.303
10	99.537	0.463	98.666	1.334	99.892	0.108	97.164	2.836	1.927	98.073	1.741	98.259	0.571	99.429	5.073	94.927
12	99.518	0.482	98.628	1.372	99.889	0.111	97.052	2.948	1.926	98.075	1.739	98.261	0.580	99.420	5.115	94.885
	Panel B1: First sub-period (January 1997 to December 2003)								Panel B2: First sub-period (January 1997 to December 2003)							
1	100.000	0.000	100.000	0.000	100.000	0.000	100.000	0.000	2.731	97.269	5.891	94.109	0.415	99.585	0.911	99.089
2	99.534	0.466	99.927	0.073	99.983	0.017	98.730	1.270	8.032	91.968	6.796	93.204	0.270	99.730	3.475	96.525
3	99.029	0.971	99.868	0.132	99.971	0.029	97.490	2.510	10.728	89.272	7.117	92.883	0.222	99.778	4.470	95.530
4	98.665	1.335	99.825	0.175	99.962	0.038	96.452	3.548	12.103	87.897	7.269	92.731	0.200	99.800	4.959	95.041
5	98.430	1.570	99.795	0.205	99.956	0.044	95.593	4.407	12.842	87.158	7.354	92.646	0.188	99.812	5.244	94.756
10	98.077	1.923	99.733	0.267	99.946	0.054	93.029	6.971	13.785	86.215	7.486	92.514	0.170	99.830	5.768	94.232
12	98.054	1.946	99.726	0.274	99.945	0.055	92.493	7.507	13.838	86.162	7.498	92.502	0.169	99.831	5.842	94.158
	Panel C1: Second sub-period (January 2004 to December 2009)								Panel C2: Second sub-period (January 2004 to December 2009)							
1	100.000	0.000	100.000	0.000	100.000	0.000	100.000	0.000	1.808	98.192	0.000	100.000	1.020	98.980	3.194	96.806
2	98.503	1.497	99.163	0.837	99.938	0.062	99.952	0.048	1.181	98.819	0.417	99.583	7.112	92.888	7.645	92.355
3	97.632	2.368	98.398	1.602	99.834	0.166	99.925	0.075	0.988	99.012	0.613	99.387	11.763	88.237	8.762	91.238
4	97.143	2.857	97.834	2.166	99.721	0.279	99.907	0.093	0.905	99.095	0.712	99.288	14.866	85.134	9.276	90.724
5	96.868	3.132	97.428	2.572	99.615	0.385	99.895	0.105	0.863	99.137	0.767	99.233	16.927	83.073	9.549	90.451
10	96.530	3.470	96.580	3.420	99.271	0.730	99.875	0.125	0.817	99.183	0.855	99.145	20.973	79.027	9.929	90.071
12	96.516	3.484	96.480	3.520	99.197	0.803	99.873	0.127	0.815	99.185	0.863	99.137	21.534	78.466	9.956	90.044

NPF represents net portfolio flows as a proportion of GDP; RER is the bilateral real exchange rates, in logs. All variables in the VAR are Hodrick-Prescott (1997) filtered.

Some puzzling results are also noticeable in the forecast error variance decompositions. The proportion of one-month horizon real exchange rates error variance explained by shocks in net portfolio flows is very high at 5.891 percent for Morocco (Panel B2), indicative of a possible unidirectional causal relation from portfolio flows to real exchange rates. This finding deviates from those of Granger-causality tests, which reports no such relation. To check the robustness of this surprising result, I run the first sub-period VAR for Morocco, in first differences: variance decomposition results (Table A3, Appendix II) are qualitatively similar to those reported here.

5.2.5 Robustness of the VAR Tests Results

Following Siourounis (2004) we now estimate the VAR with interest rate differentials. Unlike Siourounis (2004), however, we enter interest rate differentials in the VAR exogenously and, since our net portfolio flows include both bond and equity flows, we exclude equity return differentials from the model. Interest rate and inflation differentials are both known to have mutual relationships with both exchange rates and cross-border capital flows. Since exchange rates are defined in real terms in our study, including inflation differentials would amount to double counting and may result in spurious regression. Thus, we estimate the following unrestricted interest rate differentials-augmented bivariate VAR.

$$\begin{bmatrix} NPF_{it} \\ RER_{it} \end{bmatrix} = \begin{bmatrix} \alpha_{10} \\ \alpha_{20} \end{bmatrix} + \begin{bmatrix} \emptyset_{11}(L) & \emptyset_{12}(L) \\ \emptyset_{21}(L) & \emptyset_{22}(L) \end{bmatrix} \cdot \begin{bmatrix} NPF_{i,t-1} \\ RER_{i,t-1} \end{bmatrix} + \begin{bmatrix} \delta_{13} \\ \delta_{23} \end{bmatrix} \cdot INTDIF_{it} + \begin{bmatrix} \varepsilon_{it}^{NPF} \\ \varepsilon_{it}^{RER} \end{bmatrix} \tag{70}$$

where NPF_{t-1} represents lagged values of net portfolio flows; RER_{t-1} denotes lagged values of real exchange rates; $INTDIF_t$ are contemporaneous interest rate differentials computed as the arithmetic difference between foreign (African-country) short-term interest rates and USA short-term interest rates. Short-term interest rates are estimated as one-month yields on three-month Treasury bills rates, obtained from IFS. Monthly yields are estimated from annualized percentages using the formula: $r_{f,monthly} = 100 \times \left[\exp(r_{f,annualized}/12) - 1\right]$. As before, lag structures are chosen through the Schwarz (1978) criteria. All variables are Hodrick and Prescott (1987) filtered and exhibit stationarity in levels. VAR estimation results are reported in Table 33.

Table 33
Unrestricted VAR results for robustness check

	Dependent variable	Constant	NPF(-1)	RER(-1)	INTDIF	Adj R^2	Corr $(\varepsilon_1, \varepsilon_2)$
Panel A: Full Sample							
Egypt	Net portfolio flows	-0.001	0.062	13.110	-0.335	-0.30	-0.1315
		(-0.00)	(0.77)	(1.23)	(-1.06)		
	Real exchange rates	0.000	0.000	0.828***	0.003***	76.89	
		(0.03)	(0.26)	(19.02)	(2.37)		
Morocco	Net portfolio flows	-0.004	0.020	-3.601	0.127	-0.07	-0.1414
		(-0.04)	(0.24)	(-1.28)	(0.75)		
	Real exchange rates	0.000	0.000	0.853***	-0.001	72.17	
		(0.07)	(0.14)	(19.41)	(-0.07)		
Nigeria	Net portfolio flows	-0.003	0.403***	-0.187	0.016	15.02	0.0082
		(-0.03)	(5.36)	(-0.36)	(0.43)		
	Real exchange rates	-0.001	-0.005	0.818***	0.002	68.30	
		(-0.17)	(-0.80)	(17.94)	(0.60)		
S. Africa	Net portfolio flows	0.003	0.050	2.749**	0.002	1.33	-0.1009
		(0.03)	(0.62)	(1.82)	(0.03)		
	Real exchange rates	-0.001	-0.004**	0.854***	0.003*	76.47	
		(-0.31)	(-1.76)	(18.98)	(1.48)		
Panel B: First sub-period (January 1997 to December 2003)							
Egypt	Net portfolio flows	-0.001	0.233**	6.748	-0.008	3.24	-0.1695
		(-0.00)	(2.14)	(0.97)	(-0.03)		
	Real exchange rates	0.000	-0.002**	0.731***	0.005**	68.76	
		(0.09)	(-2.07)	(10.14)	(2.23)		
Morocco	Net portfolio flows	-0.003	0.010	-3.013	-0.444*	-0.55	-0.2445
		(-0.02)	(0.06)	(-0.60)	(-1.49)		
	Real exchange rates	-0.001	-0.000	0.844***	0.000	67.17	
		(-0.25)	(-0.21)	(12.97)	(0.09)		
Nigeria	Net portfolio flows	-0.001	0.056	-0.076	0.006	-3.39	0.0625
		(-0.01)	(0.49)	(-0.26)	(0.21)		
	Real exchange rates	-0.003	-0.007	0.800***	0.005	67.00	
		(-0.16)	(-0.27)	(11.88)	(0.82)		
S. Africa	Net portfolio flows	-0.002	0.064	5.074**	-0.034	4.34	-0.0961
		(-0.01)	(0.58)	(2.09)	(-0.30)		
	Real exchange rates	-0.003	-0.004*	0.945***	-0.000	82.16	
		(-0.57)	(-1.36)	(15.86)	(-0.16)		
Panel C: Second sub-period (January 2004 to December 2009)							
Egypt	Net portfolio flows	0.047	0.019	64.494**	-0.691	1.63	-0.0916
		(0.06)	(0.16)	(1.93)	(-1.27)		
	Real exchange rates	0.000	0.000	0.644***	0.003***	58.26	
		(0.09)	(1.21)	(7.68)	(2.74)		
Morocco	Net portfolio flows	0.011	0.068	-4.057*	0.361**	4.89	-0.0121
		(0.08)	(0.57)	(-1.34)	(1.78)		
	Real exchange rates	0.000	0.002	0.864***	0.002	72.21	
		(0.08)	(0.65)	(13.51)	(0.40)		
Nigeria	Net portfolio flows	0.007	0.431***	-2.329	-0.015	16.68	-0.1009
		(0.04)	(3.84)	(-0.58)	(-0.18)		
	Real exchange rates	0.000	-0.004***	0.902***	0.000	85.91	
		(0.09)	(-2.37)	(15.92)	(0.04)		
S. Africa	Net portfolio flows	0.004	-0.056	-0.767	-0.035	-3.95	-0.2008
		(0.02)	(-0.45)	(-0.40)	(-0.28)		
	Real exchange rates	-0.003	-0.008**	0.733***	0.013***	69.36	
		(-0.56)	(-1.68)	(10.20)	(2.82)		

The table reports VAR coefficients and *t*-statistics (in parentheses). NPF(-1) = lagged values of net portfolio flows (as a proportion of GDP); RER(-1) = lagged values of log of real exchange rates; INTDIF = interest rate differentials. Adj R^2 is the adjusted coefficient of determination, in percentage. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels respectively. All variables are Hodrick-Prescott filtered. Observations run from 1997:1 to 2009:12

Consistent with earlier findings, it is important to observe (i) that, in all the countries, values of real exchange rates in one period significantly explain the next period's values, indicative of a high level of serial correlation in the real exchange rates series; (ii) except for South Africa (full sample and first sub-period) and Egypt and Morocco (second sub-period), real exchange rates have virtually no explanatory power on next period's portfolio flows. Similarly, with the exception of South Africa (full sample), Egypt (first sub-period) and Nigeria (second sub-period), lagged values of portfolio flows do not predict contemporaneous real exchange rates; (iii) net portfolio flows seem not to exhibit serial correlation. Clearly, the dynamic relationship between the two variables remains the same under different tests.

Contrary to theoretical predictions, our results do not suggest a strong relationship between interest rate differentials and real exchange rates. Interest rate differentials are found to explain real exchange rates only for Egypt (consistently) and South Africa (full sample and second sub-period). The coefficients of the interest rate differentials are very small in magnitude for all countries; except for Morocco (full sample) and South Africa (first sub-period), they are also positive denoting the tendency of real exchange rates to depreciate with increasing interest rates in these markets. Similarly, the results do not suggest any clear association between interest rates and net portfolio flows of African countries; statistically significant coefficients are reported only for Morocco and only for the sub-periods.

5.2.6 Summary of Findings from Monthly Data Analysis

Overall, the study of monthly data finds the nature of the bivariate relationship between real foreign exchange rates and net portfolio flows in Africa's capital markets to be time-varying and country-dependent. Save for South Africa, where evidence of bidirectional causality seems to hold across the period, net portfolio flows to the rest of the continent appear to be driven largely by international investors' appetite for short-term speculative gains: international portfolio flows to African countries exhibit high volatility and low persistence, characteristics that qualify them to be described generally as "hot money". In generic terms, these findings concur with Sula and Willett (2009) who find that portfolio flows and private loans are more reversible (i.e. less stable) than foreign direct investments in the emerging markets. The authors also demonstrate that volatility of capital flows in normal periods is not a good predictor of the size of their reversal during crises periods.

The situation is particularly interesting for Nigeria, where monetary policy managers seem to attach their foreign exchange rate policy to net portfolio flows in the second part of the study period. In that sub-period, devaluations of the Nigerian naira-US dollar exchange rate, apparently intended to correct deviations from international parity relationships brought about by huge portfolio flows, are noticeable. However, although the direction of causality is not distinct, this study reveals, consistent with findings from other developing regions, that net portfolio flows are generally associated with real exchange rate appreciations in Africa's capital markets (see for example Calvo et al. (1993) and Edwards (1998) for studies of some Latin American markets).

These results do not support portfolio balance theory's predictions that portfolio flows should exert an influence on real exchange rates. Failure to attribute real exchange rate fluctuations to portfolio flows is consistent with Morrissey et al. (2004), who find that neither "permanent" (FDI, remittances, grants) nor "non-permanent" (loans, equities, portfolio investments) capital inflows have a short-run effect on the real exchange rate in Ghana: they ascribe real exchange rate misalignments to changes in trade volume, arising mainly from changes in exports. They, however, find that both permanent and non-permanent capital inflows have a strong and significant appreciation effect on the real exchange rate in the long-run, but that the extent of the appreciation is slightly greater for the permanent inflows than for the non-permanent inflows. In this study, lack of cointegration did not permit the determination of long-run causality. Also, the results do not depart markedly from Ndung'u and Ngugi (1999), who find only a weak feedback from real exchange rate movements to Kenya's capital flows volatility.

5.3 Annual Data

5.3.1 Preliminary Analysis

In this section, I perform preliminary investigation of net international portfolio flows data for individual major African countries and for the group of countries. The number of countries increases to six from the four investigated in the preceding section because annual portfolio flows data are available for Botswana, Egypt, Kenya, Morocco, Nigeria, and South Africa.

Portfolio flows data are not available for Ghana. Summary statistics for annual net portfolio flows and real exchange rates are presented in Table 34.

Table 34
Summary statistics for

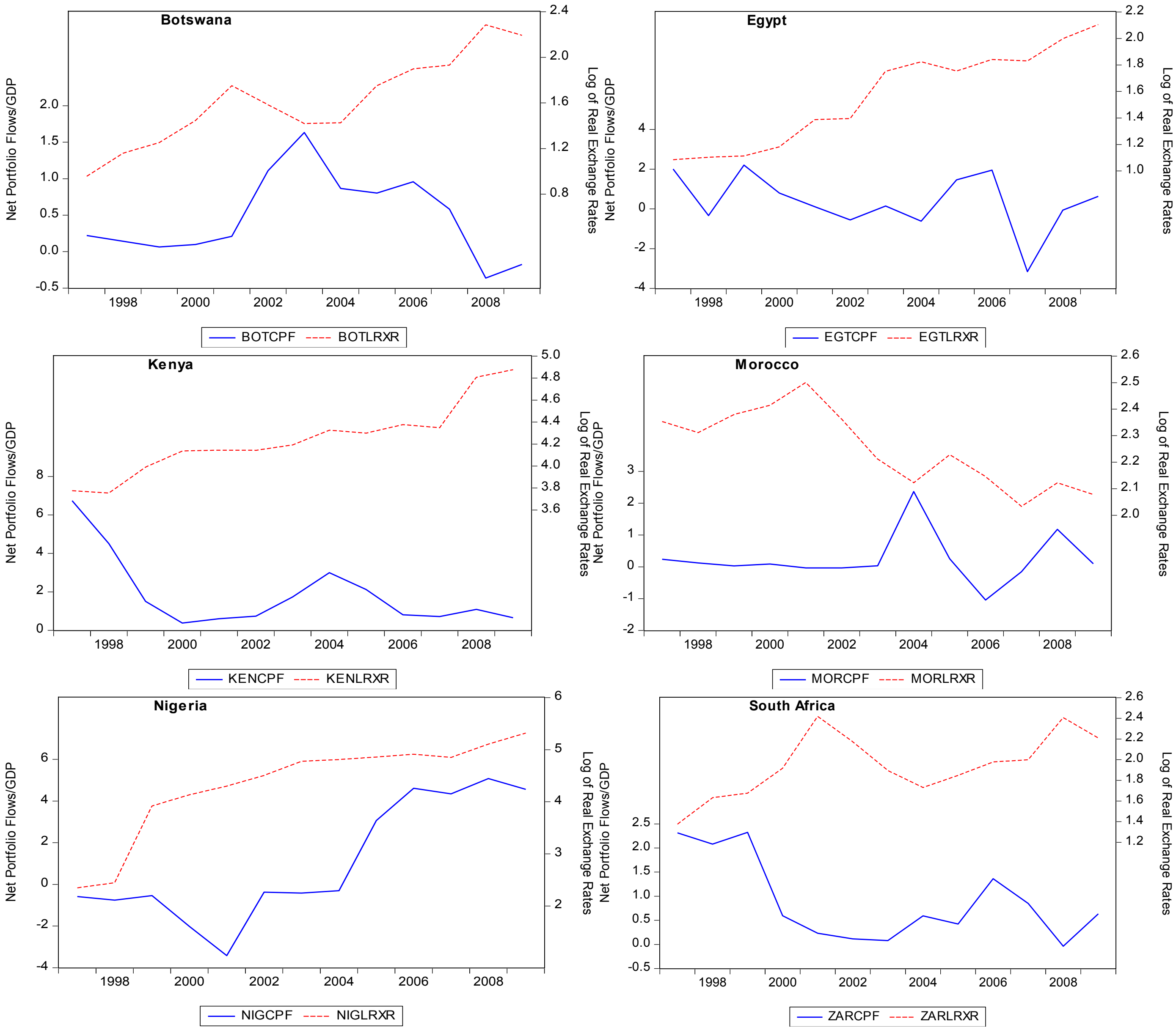
	Mean	Max	Min	Std Dev	CV	Skew	Kurt	Jarque-Bera *LM*-stat	Jarque-Bera *p*-value
Panel A: Real exchange rates									
Botswana	1.6193	2.2845	0.9561	0.3956	0.2443	0.1132	2.1408	0.4276	0.8075
Egypt	1.5668	2.1051	1.0841	0.3672	0.2343	-0.1202	1.5008	1.2487	0.5356
Kenya	4.2438	4.8769	3.7537	0.3311	0.0780	0.4705	2.7977	0.5017	0.7781
Morocco	2.2511	2.5008	2.0333	0.1460	0.0648	0.1042	1.7843	0.8241	0.6623
Nigeria	4.3296	5.3180	2.3483	0.9407	0.2173	-1.2944	3.4289	3.7301	0.1549
South Africa	0.8891	2.3243	-0.0378	0.8534	0.9598	0.7525	2.0768	1.6885	0.4299
Group	2.6589	5.3180	0.9561	1.2697	0.4775	0.6934	2.0012	9.4934	0.0087
Panel B: Net portfolio flows									
Botswana	0.4732	0.0163	-0.0036	0.0057	1.2092	0.4563	2.3602	0.6728	0.7143
Egypt	0.0354	0.0022	-0.0032	0.0014	4.0809	-0.8319	3.7478	1.8022	0.4061
Kenya	0.0038	0.0001	0.0000‡	0.0037	0.9868	1.6111	4.5889	6.9911	0.0303
Morocco	0.0000‡	0.0000‡	-0.0000‡	0.0000‡	3.2645	1.4313	5.5252	7.8924	0.0193
Nigeria	0.0000‡	0.0000‡	-0.0000‡	0.0000‡	2.8291	0.2386	1.6148	1.1627	0.5592
South Africa	0.8891	0.0232	-0.0004	0.8534	0.9598	-0.0100	2.3490	0.2297	0.8915
Group	0.2336	0.0232	-0.0036	0.5329	2.2815	2.5445	9.2410	210.76	0.0000

The table displays summary statistics for log of real exchange rates and net international portfolio flows, as a proportion of GDP, for the period 1997 through 2009. Mean is in percentage; "Max" and "Min" respectively represent maximum and minimum observed values; "Std Dev" represents standard deviation (in percentage), "CV," the coefficient of variation, "Skew" denotes skewness, "Kurt" is the kurtosis and J-B is the Jarque and Bera (1987) statistic. "LM-stat" is the Jarque-Bera Lagrange Multiplier statistic. ‡ denotes that the figure is zero correct to four decimal places.

Table 34 shows that real exchange rates and portfolio flows can be described by the normal distribution for most of the countries. However, because of the very short span of the time series used in this analysis, such an outcome must be interpreted with caution. Indeed, when the data is stacked together in a panel, it is clear that the null hypothesis of normality is easily rejected, at the 1 percent level, by both real exchange rates and net portfolio flows. Thick tails are observed in the distribution of portfolio flows data of many countries as well as for the group. Because of the short time series, an autocorrelation function was not expected to be efficient and was not run for the data. The table also reveals that net annual portfolio flows are generally more volatile than real exchange rates; the most volatility, consistent with monthly data, is recorded for Egypt.

Portfolio flows also constitute a very small proportion of GDP, recording average values less than 1 percent for all countries. The co-movements between net annual portfolio flows and the annual real exchange rates for the six countries are reported in Figure 11.

Figure 11
Trends in annual net international portfolio flows and real foreign exchange rates



The surfix *LRXR* denotes natural logarithms of real foreign exchange rates; *CCF* denotes net foreign portfolio flows as a percentage of the Gross Domestic Product (GDP). Thus, *EGTLRXR* series represents natural logarithms of real foreign exchange rates between the Egyptian pound and the US dollar; *EGTCPF* series represents Egypt's net portfolio flows as a percentage of the corresponding Gross Domestic Products, and so on. Observations run from 1997 to 2009.

From the figure, a negative relationship is observed for most countries. The two variables, however, show a marked positive relationship for Nigeria. These observations are confirmed by correlation analysis: the correlation coefficients between net portfolio flows and real exchange rates for the six countries are –0.201 for Botswana, –0.295 for Egypt, –0.599 for Kenya, –0.215 for Morocco, 0.550 for Nigeria and –0.763 for South Africa. Thus, from a country-by-country perspective, the evidence points to a relationship between net portfolio flows and foreign exchange rates, in which foreign investors tend to purchase more, and/or sell less (i.e., international portfolio inflows to the various African markets tend to outstrip outflows) of financial assets when the US dollar appreciates against the local (African country) currency.

However, when the data is pooled for all countries, the correlation becomes positive (+0.299) and significant at the 1 percent level. Thus, evidence from pooled data is supportive of a linkage between net portfolio flows and exchange rate fluctuations in which net portfolio flows into the African countries capital markets are positively correlated with an appreciating US dollar. The latter observation is consistent with findings of Hau and Rey (2006) that net equity inflows are associated with an appreciating foreign currency. Such a positive relationship is supportive of the traditional view of causality in which foreign exchange rate changes lead foreign capital flows. Empirical evidence, discussed next, sheds more light on this conjecture.

5.3.2 Panel Unit Root Analysis

A major shortcoming of time series studies for individual countries, involving such tests as the ones in section 5.2, is that the powers of the unit root tests may be impaired in finite samples. Weakness in the power of statistical tests is often a worrying econometric issue if the data span is very short, as is the case in the current analysis where the investigation period spans for a period of only thirteen years. To mitigate this shortcoming, this analysis employs the rapidly developing new line of tests based on panel data to exploit the extra power from combining cross-sectional and time series data. As Al-Iriani (2006) notes, using panel data allows us to gain more observations by pooling the time series data across sections leading to higher power for the causality tests. Another positive by-product of using panel data instead of single time series is that the test statistics are asymptotically normally distributed and are approximately normally distributed for the finite sample sizes generally encountered in econometrics. The specific panel methodologies, employed in this section, are reviewed in Appendix III.

Table 35
Results for panel stationarity tests

	Panel Test Statistic			
	Breitung	Im et al.	Fisher ADF	Hadri
Portfolio Flows (Level)	-2.0538** (0.0200)	-2.6628*** (0.0039)	29.7144*** (0.0031)	3.8004*** (0.0001)
Portfolio flows (First Difference)	-2.2071** (0.0137)	-4.0842*** (0.0000)	38.5812*** (0.0001)	1.3338* (0.0911)
Foreign exchange rates (Level)	0.0069 (0.5027)	-0.8689 (0.1925)	21.5304** (0.0431)	5.3291*** (0.0000)
Foreign exchange rates (First Difference)	-2.1488** (0.0158)	3.5485*** (0.0002)	33.5946*** (0.0008)	1.1784 (0.1193)

*, ** and *** indicate the rejection of the null hypothesis at the 1 percent, 5 percent and 10 percent levels, respectively. *P*-values are in parentheses. Annual data for the period 1997 through 2009 are used.

Following practice in the literature, I conduct unit roots tests for the panel using several methods. The unit roots tests are performed with individual (cross-sectional units) intercepts and no trend. However, by its definition, the Breitung (2000) test is performed with individual intercepts and stochastic trends. The null hypothesis for the Breitung (2000), Im et al. (2003) and Fisher ADF tests is that the data exhibit the presence of a panel unit root. Conversely, the null hypothesis for the Hadri (2000) test is that the panel exhibits stationarity. The null hypotheses are rejected if the test statistic is significant. Results of the various panel unit root tests are reported in Table 35.

Excepting the results from the Fisher ADF test of Maddala and Wu (1999), all the tests find that real foreign exchange rates have a panel unit root in levels and are stationary in first differences, i.e., they are integrated of the first order, $I(1)$. On the other hand, net international portfolio flows series appear to exhibit panel stationarity, i.e., they are integrated of order zero, $I(0)$. This is the case for all the tests excepting the Hadri (2000) test which rejects the hypothesis of panel stationarity (recall that the null hypothesis for this test is the presence of stationarity; thus, a significant test statistic implies the presence of a unit root). These results are similar to those obtained in section 5.2.2 for monthly data. They also confirm results from other researchers (see, for example Edwards, 1998) that portfolio flows series generally exhibit stationarity. Since the two series are not integrated of the same order, cointegration tests cannot follow.

5.3.3 Panel Causality Analysis

The direction of causality between net foreign portfolio flows and real exchange rates, in a panel context, is determined next based on the following regressions:

$$\begin{aligned} NPF_{it} &= \beta_{1i} + \textstyle\sum_k \beta_{11ik} NPF_{i,t-k} + \sum_k \beta_{12ik} RER_{i,t-k} + \varepsilon_{1it} \\ RER_{it} &= \beta_{2i} + \textstyle\sum_k \beta_{21ik} RER_{i,t-k} + \sum_k \beta_{22ik} NPF_{i,t-k} + \varepsilon_{2it} \end{aligned} \qquad (71)$$

where NPF_{it} and RER_{it} are, respectively, net portfolio flows and real exchange rates for country i at time t; and k is the lag length, chosen using the Schwarz (1978) Bayesian information criteria. The null hypotheses for the causality tests are that, for all i and k, $\beta_{12ik} = 0$ in the first equation and $\beta_{22ik} = 0$ in the second equation of the VAR system (71).

Table 36 reports results for the Granger causality tests. Since the series are not integrated of the same order, the test is performed by first, using the Hodrick and Prescott (1997) filter to detrend the real exchange rates and net portfolio flows for each country in the panel. The resulting detrended series, which exhibit panel stationarity, are then used in the Granger causality test. This procedure follows Edwards (1998) where it was successfully used to test the dynamic relation between real exchange rates and capital flows for several Latin American countries. Because the Schwarz (1978) Bayesian information criteria puts the optimal lag length for the detrended series at zero, the Hannan and Quinn (1979) information criteria is used to choose the lag length.

Table 36
Results for panel Granger causality tests

	F-statistic	*P*-ratio
Real exchange rates does not Granger-cause net portfolio flows	2.8225**	0.0475
Net portfolio flows does not Granger-cause real exchange rates	1.3839	0.2578
Lag length	3	

** denotes that the null hypothesis is rejected at the 5 percent level of significance. Annual data for 1997 to 2009 are used.

The procedure used in the foregoing causality test implicitly imposes the constraint that the underlying structure is the same for each cross-sectional unit. Thus, it ignores the possibility that each unit has an "individual effect" which translates in practice to its own intercept. The individual effect summarizes the influence of unobserved variables which have a persistent effect on the dependent variable. Since the other right hand side variables are typically correlated with the individual effect, its omission results in inconsistent estimates. Suppose that an individual effect (f_i) exists for cross-sectional unit i. The time-stationary VAR relationship can be expressed in the form (Holtz-Eakin et al., 1989):

$$\begin{aligned} Y_{it} &= \alpha_0 + \textstyle\sum_{j=1}^{L} \alpha_j Y_{i,t-j} + \sum_{j=1}^{L} \delta_j X_{i,t-j} + f_{yi} + u_{it} \\ X_{it} &= \beta_0 + \textstyle\sum_{j=1}^{L} \beta_j X_{i,t-j} + \sum_{j=1}^{L} \gamma_j Y_{i,t-j} + f_{xi} + v_{it} \end{aligned} \tag{72}$$

where Y_{it} and X_{it} ($i = 1, \ldots, N; t = L + 1, \ldots, T$) are the two cointegrated variables; u_{it} and v_{it} are error terms; f_{yi} and f_{xi} are fixed effects for variables Y and X, respectively, unique to each cross-sectional unit, i, in the panel. System (72) cannot be estimated in its current form because the lagged dependent variables are correlated with the individual fixed effects as well as the error terms. The standard solution to this problem in the literature is to difference the data so as to remove the individual fixed effects and then estimate the resulting VAR:

$$\begin{aligned} \Delta Y_{it} &= \alpha_0 + \textstyle\sum_{j=1}^{L} \alpha_j \Delta Y_{i,t-j} + \sum_{j=1}^{L} \delta_j \Delta X_{i,t-j} + \Delta u_{it} \\ \Delta X_{it} &= \beta_0 + \textstyle\sum_{j=1}^{L} \beta_j \Delta X_{i,t-j} + \sum_{j=1}^{L} \gamma_j \Delta Y_{i,t-j} + \Delta v_{it} \end{aligned} \tag{73}$$

where Δ is the difference operator. Holtz-Eakin et al. (1989) point out that this procedure does not eliminate the simultaneity problem as the differenced error terms are correlated with the differenced lagged regressors. Further, Al-Iriani (2006) observes that heteroskedasticity is expected to be present as heterogeneous errors might exist with different cross-sectional units in the panel. Holtz-Eakin et al. (1989) recommend the use of an instrumental variable regression procedure as a way of dealing with these problems. By assuming that the error term, u_i, is uncorrelated with all past values of Y and X, and the individual effects, they demonstrate that one can use the instrument variable vector $[1, Y_{it-2}, \dots, Y_{i1}, X_{t-2}, \dots, X_{i1}]$ in the estimation. Thus, if $L = 1$ for instance, the instrument vector consisting of a vector of ones, and second-lagged values of X_{it} and Y_{it} would suffice. The joint hypotheses, $\delta_j = 0$ and $\gamma_j = 0$ $(j = 1, \dots, L)$, are the null for the absence of causality. The VAR in system (73) is commonly estimated using the Generalized Method of Moments (GMM) estimator of Arellano and Bond (1991), which has been shown to produce efficient and consistent estimators. Results are presented in Table 37.

Table 37
Results of dynamic panel GMM estimation and causality tests

	NPF(-1)		RXR(-1)		J-statistic		Wald statistic	
Dependent variable	Coef	*t*-stat	Coef	*t*-stat	Chi-sq	*p*-value	Chi-sq	*p*-value
Net portfolio flows	0.4581***	15.29	0.9684***	3.07	1.9423	0.7464	9.4131***	0.0022
Real exchange rates	0.0139	1.16	0.8410***	19.78	5.5647	0.2341	1.3408	0.2469

NPF(-1) and RXR(-1) are, respectively, the first lagged net portfolio flows and real exchange rates. "Coef" and "*t*-stat" are, respectively, the estimated coefficient and the *t*-statistic; "Chi-sq" is the *chi* square statistic. The Wald statistic is the test that the coefficients of the lagged real exchange rate (first equation) and lagged net portfolio flows (second equation) are zero. J-statistic is the test of overidentifying restrictions in the GMM estimation. *** denotes that the reported figure is significant at the 1 percent level. Annual data for the period 1997 through 2009 are used.

VAR results in Table 37 show: (i) lagged net portfolio flows have a predictive power on future net portfolio flows. This is at variance with the monthly data analysis which displayed no such characteristics. (ii) Lagged portfolio flows have no predictive power on future real exchange rate movements. (iii) Lagged real exchange rates can predict both portfolio flows and real exchange rates in the next period. The last two findings are consistent with monthly data analysis results.

5.3.4 Summary of Findings from Annual Data Analysis

The results from the two panel causality tests (Tables 36 and 37) demonstrate clear unidirectional causality from real exchange rates to net portfolio flows in Africa. This finding is consistent with Shah and Patnaik (2008) who document results indicating that expectations of currency appreciation attract portfolio flows into India. However, the finding contrasts recent dynamic panel-data evidence of Jongwanich (2010) for nine emerging Asian countries during roughly the same period (2000–2009). His results suggest that the swift rebound of capital inflows into the region could result in an excessive appreciation of the (real) currencies, especially when capital flows are in the form of portfolio investment and bank loans.

The results lend support to the traditional theoretical explanation of causality in which changes in foreign exchange rates are expected to precede changes in international capital flows and, ultimately, changes in the behavior of the domestic capital market. Given the positive correlation between net portfolio flows and real exchange rates, the intuition behind this relationship is clear. A weak Moroccan dirham, for instance, should make dirham-denominated financial assets cheaper, and therefore more attractive, from the point of view of investors foreign to Morocco. If some, or all, of the conditions governing perfect capital markets are violated, as is often the case in practice, the International Fischer Effect[56] would not hold exactly and a depreciation of the Moroccan dirham should encourage inflows of foreign portfolio funds to take advantage of exchange rate induced low-priced Moroccan dirham-denominated financial assets.

The finding in the current study that real exchange rates drive net portfolio flows can be explained as follows. The exposition in section 1.1.1 (Chapter 1) clearly shows that many of the African countries' currencies, although exhibiting various levels of flexibility, cannot be accurately described as freely floating. The Nigerian naira, for instance, has witnessed various episodes of devaluation in the last few years that have more or less distorted its true value. Speculative foreign investors are likely to respond by increasing their purchases of financial assets, in markets in which such distortions occur, to take advantage of the temporary exchange rate-induced return differentials. When the exchange rate begins to stabilize at its new level, a profit-taking stampede follows wherein the speculators quickly close off their net long positions, withdraw their capital and, hopefully, repeat the act in another hapless market. This behavior fits in neatly with the preliminary finding that portfolio flows to African capital markets are non-persistent in nature and therefore largely point to the "hot money" phenomenon in those markets.

[56] The basic message of the Fisher effect is that there exists a relationship between interest rates and the rate of inflation of the form: $i = i' + \alpha + i'\alpha$ where *i* is the nominal rate of interest, *i'* is the real rate of interest and α is the anticipated inflation rate. The international counterpart of this relationship merely recognizes that a similar effect must hold for different foreign currencies: $1 + S_f = (1 + i_D)/(1 + i_F)$, where S_f is the rate of change in the value of foreign currency; i_D is the domestic nominal rate of interest; and i_F is the foreign nominal rate of interest. The theory suggests that foreign currencies with relatively high interest rates will depreciate because the high nominal interest rate would incorporate both relatively high default risk and relatively high anticipated inflation rate. Empirical evidence shows that the IFE does not hold in the short run.

5.4 Conclusions

This chapter has investigated the nature of the relationship between real exchange rates and the flow of portfolio capital in Africa's capital markets from a time series perspective using monthly data of four countries and through panel procedures using annual data of six countries. Analysis of the monthly data documents a time-varying and country-dependent causality. Country idiosyncrasies are invoked to explicate the non-uniformity of the results across countries and through time. On the other hand, annual data analysis documents a clear unidirectional panel causality running from real exchange rates to portfolio flows. This finding that has been attributed to exchange rate policies in the countries, which allow monetary policy managers to intervene in the foreign exchange markets, sometimes causing temporary aberrations in international parity conditions and creating incentives for international arbitrage and speculative activities in the destination countries investigated.

The existence of a relationship between foreign exchange rates and international capital flows is an empirical regularity. What has always been contentious, however, is the direction of causality between the two variables. Froot and Stein (1991) have observed that a ten percent US dollar depreciation is associated with additional Foreign Direct Investment (FDI) inflows of about $5 billion into the US economy. This observation points to a positive relationship similar to the finding from annual data analysis in the current study. Similarly, Frankel and Schmukler (1996) establish a relationship between changes in country fund discounts (caused by investors taking short positions) and changes in the exchange rates. However, the relationship is only partial in nature, indicating that the fall in the discount (following the Mexican peso devaluation, December 1994) was greater than would be expected from the magnitude of the devaluation and the usual pattern associated with exchange rate changes.

In contrast, Edwards (1998) documents a negative relationship between capital inflows and the real exchange rate, in which increases in capital inflows are associated with real exchange rate appreciation while declines in inflows are associated with real exchange rate depreciation. Edwards' finding is not surprising, coming from the Latin American experience in the late 1980s and early 1990s, when international portfolio flows were believed to have been driven largely by interest rate changes in the USA. In that scenario, many researchers (for example Calvo et al., 1993; Chuhan et al., 1993 and Fernandez-Arias, 1994) have argued that international capital flows might have a disruptive effect on the real value of the recipient country's currency. In the developed capital markets, Heimonen (2009) also reports findings suggesting that net equity flows from the euro area to the USA lead to appreciations of the dollar. However, he explains that equity flows affect the exchange rate through order flows driven by equity return differentials between the euro area and the USA. Accordingly, Heimonen (2009) effectively establishes a puzzling relationship between equity returns and foreign exchange rates in which low equity returns in the domestic country, relative to foreign countries, tend to have an appreciation effect on the value of the domestic country's currency. This counterintuitive finding is explained by deviations from the minimum variance portfolio. That is, an increase in the return on domestic-country equities with respect to foreign returns increases the relative share of domestic equities in investors' total wealth and implies a deviation from the minimum variance portfolio of domestic equities. As a result, the investor decreases their holdings of domestic country equities. There is an equity outflow from the country with the excess equity returns which generates an outflow in the domestic exchange market, the effect of which is a depreciation of the currency with excess equity returns.

CHAPTER SIX

CONCLUSIONS, POLICY IMPLICATIONS AND RECOMMENDATIONS

6.1 Introduction

In the early 1990s, most African countries, pushed by conditionalities imposed by the Bretton-Woods institutions for additional budgetary and development support, started liberalizing their economies. The many "bitter pills" swallowed by these countries included: less regulation of import-export activities, abolition of exchange and capital controls, liberalization of interest rates regimes, relaxation of rules regulating foreign participation in domestic capital markets, and adoption of flexible foreign exchange rates. The implementation of these conditionalities has presented serious challenges to macroeconomic stability in many of these countries. In particular, the immediate impact of the adoption of a flexible exchange rate system was, and continues to be, instability in the foreign exchange markets. Such instability is disruptive to the economies of affected countries as it impacts negatively on international investors' perception about the safety of their investment in view of the huge foreign exchange risk that it suggests.

Foreign exchange risk is one of the most widely investigated issues in the International Finance literature. Several studies have sought to establish whether instability in the foreign exchange markets introduces risk that is perceived by investors to be priced in the capital markets. In the advanced economies, conflicting findings have been documented, with early studies, based on unconditional asset pricing models, reporting no evidence that currency risk is priced in capital markets. However, recent evidence provided largely by studies that employ conditional asset pricing models indicates that some success has been recorded in finding foreign exchange risk to be a priced factor in advanced equity markets. In emerging markets outside Africa, relatively few studies have been conducted but their dominant finding is that fluctuations in foreign exchange rates constitute a risk that commands a significant premium in the capital markets. My study has been the first to investigate the pricing of foreign exchange risk in Africa's capital markets.

It was anticipated that the adoption of liberalization policies as advocated by the World Bank and International Monetary Fund (IMF) would accrue tangible macroeconomic benefits to complying countries. Key among the expected benefits was enhanced flows of international money into the capital markets of these countries. However, this "dream" is yet to be realized several years after these policies were implemented: available evidence indicates that Africa still lags behind other developing regions in attracting international capital. In particular, IMF and US Treasury Department data (presented in Chapter 5) show that, for many African countries, net foreign portfolio flows do not constitute a significant portion of the Gross Domestic Product (GDP). It is also evident that, like foreign exchange rates, the volatility of portfolio flows has increased in the years after these policies were adopted. It is this observation that motivated this study to establish whether a linkage exists between net international portfolio flows and foreign exchange rates and whether a scientific explanation exists to which the existence of such a nexus can be ascribed.

In summary, I investigated two major issues in this dissertation. First, I sought to establish whether foreign exchange risk is priced in Africa's capital markets. Because private debt markets, other than bank financing, are relatively poorly developed in the continent, I examined the pricing of foreign exchange risk only in equity markets. In this nexus, I also sought to establish the extent of integration of Africa's equity markets with equity markets in the rest of the world. Second, I investigated the dynamic relationship between real foreign exchange rates and the flow of international portfolio capital in Africa's capital markets; this also involved a study of the persistence of portfolio flows. On all the issues investigated, I have documented interesting findings, which I recapitulate next.

6.2 Foreign Exchange Risk Pricing

When a foreign investment takes place, payments for the assets/securities generally involve a foreign exchange transaction. Frequently, the foreign investor, who buys assets, pays the seller of the assets in the seller's home currency. This requires the investor to accomplish a foreign exchange transaction that converts the investor's currency to the seller's currency. The price or rate of conversion between the two currencies is the exchange rate. Foreign exchange rates are quoted in two ways: the *direct quote*, which gives the rate as the number of units of the home currency required to buy one unit of the foreign currency and the *indirect quote*, which gives the rate as the number of units of foreign currency required to buy one unit of the domestic currency. The latter approach has been used in this study, with the US dollar as the domestic currency.

When they are determined freely in the open market place, foreign exchange rates continually adjust in response to demand for and supply of the two currencies involved. The demand for a country's currency depends on the total value of the country's exports, the inflow of private capital funds and the inflow of foreign governments' capital funds. Conversely, the supply of a nation's currency depends on the total value of its imports, the outflow of government funds and the outflow of private capital funds. Since the forces that drive currency values do not follow a distinct, predictable pattern over time, one would expect adjustments or movements in currency values to be random and unpredictable in nature. For economic agents with assets, liabilities, profits or expected future cash flow streams denominated in foreign currencies, adjustments in currency values may cause a change, for better or for worse, in the future domestic currency values of those items. Such items are said to be exposed to foreign exchange risk.

Formally, foreign exchange risk is defined as the variability in an investment's expected returns attributable directly to adjustments in foreign exchange rates. Foreign exchange risk is said to be priced if investors demand a compensation for bearing it in the capital markets. In this work, the pricing of foreign exchange risk in Africa's capital markets has been investigated using the arbitrage pricing theory (APT)-based unconditional asset pricing theoretical framework as well as a stochastic discount factor (SDF)-based conditional asset pricing model. The two models are reviewed in Section 3.3 of Chapter 3. A summary of the major findings from these tests follows.

6.2.1 The Unconditional Pricing of Currency Risk and Equity Markets Integration

As already explained, the Arbitrage Pricing Theory of Ross (1976) extended to the international environment by Solnik (1983) provides the theoretical framework for the unconditional asset pricing models used in this study. The theory is implemented in two-factor and three-factor model specifications. Each model specification is used to jointly investigate the pricing of foreign exchange risk and equity markets integration. A sample of countries, drawn based on the relative verve of the foreign investors' segments of their capital markets is used in the analysis. The countries are: Botswana, Egypt, Ghana, Kenya, Morocco, Nigeria and South Africa.

Using the *Afro* real exchange rate index, a representative index developed in this study from real exchange rates data of twelve major African currencies against the US dollar (and the euro), both the two-factor foreign exchange risk model and the three-factor model find that foreign exchange risk is not priced in major equity markets of Africa if returns are measured in the US dollar. It is noted that these findings contrast the influential work of Carrieri and Majerbi (2006) who use a similar methodology and, in principle, a similar real exchange rate index (the 'Other Important Trading Partners' index of the US Federal Reserve Bank) to find that currency risk commands a significant premium in emerging equity markets. The findings in this section can be attributed to low participation by international portfolio investors in Africa's equity markets; which implies that the markets are largely driven by domestic investors whose returns are denominated in local currencies and therefore shielded from the influence of changes in foreign exchange rates. Both the two-factor markets segmentation model and the three-factor model find equity markets in Africa to be partially segmented/partially integrated. The findings on segmentation replicate those reported elsewhere (see for example, Choi and Rajan, 1997; Antell and Vaihekoski, 2007).

To check the robustness of these findings to the currency in which returns are measured, I use the euro. The two-factor foreign exchange risk model finds currency risk weakly priced when returns are measured in the euro, a finding that is not surprising given the euro's dominance as an invoicing currency in the African region. The euro has also been found by Kamps (2006) to be increasing in importance as a world currency against the US dollar whose role appears to be slightly diminishing. However, the three-factor model does not find foreign currency risk priced even with the euro as the measurement currency. The latter finding can be explained by the presence of the idiosyncratic local market factor in the model; a factor that apparently represents omitted factors better than the foreign exchange risk factor. Consistent with the case when returns are measured in the US dollar, both the two-factor markets segmentation model and the three-factor model find Africa's equity markets to be partially segmented/ partially integrated.

I also check the robustness of the unconditional asset pricing models to changes in time. For this purpose, I split the study period into two: the time period during which foreign exchange rates and market index returns data exhibited relatively high volatility (January 1997 to June 2003) and the period of relative tranquility in the data (January 2004 to June 2009). The US dollar is used as the measurement currency. Evidence from the two-factor foreign exchange risk model suggests that currency risk factor is not priced in both sub-periods; it is negative in the first sub-period and positive in the second sub-period, demonstrating its time-varying properties. The two factor market segmentation model finds partial segmentation in both sub-periods. For both sub-periods, the three-factor model also finds that currency risk does not command a premium in the equity markets; it suggests partial equity markets segmentation in the first sub-period but appears misspecified in the second sub-period as it finds neither segmentation nor integration.

Finally, I conduct a check on the models' robustness to country composition and change in the definition of local market factor. Thus, I drop Botswana and Ghana and incorporate Mauritius in the analysis. I then replace the *MSCI Barra* Emerging Markets portfolio index with the *MSCI Barra* Africa Frontier Markets index as the local market factor. Consistent with earlier findings, both the two-factor foreign exchange risk and the three-factor models find that currency risk is non-priced when returns are measured in US dollars. The two-factor market segmentation model finds Africa's equity markets to be fully segmented while the three-factor model finds them to be fully integrated! The conflicting finding is traced to the Africa Frontier Markets index to which the unconditional multi-factor models appear non-robust. The index seems not to be able to explain equity returns in Africa as well as does the emerging markets composite portfolio index.

The unconditional asset pricing models are also tested with portfolio-level data for South Africa with returns measured in US dollars. The two-factor model finds that foreign exchange risk commands a significant premium in South Africa's equity market. It also shows that larger firms (by market capitalization) are more exposed to foreign exchange risk than smaller firms. The three-factor model finds currency risk non-priced; it also suggests that South Africa's equity markets are fully integrated with the world equity markets. This finding is consistent with that of Kabundi and Mouchili (2009) who use a different methodology – the Dynamic Factor Model – to find moderate synchronization of the South African stock market with the world common equity market between 1997 and 2006.

6.2.2 The Conditional Pricing of Foreign Exchange Risk

The Stochastic Discount Factor model provides the analytical framework used to investigate the conditional pricing of foreign exchange risk. Because of its ability to extract foreign exchange risk premia estimates, a pricing kernel methodology similar to the one used in Cappiello and

Panigirtzoglou (2008) was initially proposed for this study. I present that methodology in Appendix I. Usage of that model is predicated on the availability of return information from all assets in the economies under investigation. Cappiello and Panigirtzoglou (2008) estimate the model on the assumption that the typical economy can be broadly represented by returns on three asset classes: national equity market indexes, national debt market indexes, and representative money market instruments. For the African countries studied here, it emerged that the debt market sector was still at the nascent stages of its development, the consequence of which was that the study faced paucity of data from which return on debt could be ascertained. Further, available price data for money market instruments were not of the quality whose utilization could provide usable return series: in most cases, interest rates on the money market instruments (obtained from International Financial Statistics database) remained constant for long periods, yielding long series of zero returns.

Thus, only returns on stock markets indexes were available. However, it was felt that equity returns alone could not adequately represent the full spectrum of economic activity in the sampled markets and, hence, were unlikely to yield reliable monthly series of currency premia estimates. As a result, the pricing kernel methodology was replaced by the alternative asset pricing methodology which is presented in section 3.3.2 of Chapter 3. The alternative model, like the one initially proposed, is based on the stochastic discount factor framework and has recently gained prominence in asset pricing investigations in the literature. However, it has the drawback of lacking an in-built mechanism for extracting foreign exchange risk premia for each country studied. Zhang (2006) has employed a variant of the method to find currency risk priced in US equity markets. In this work, I have similarly used the method successfully to find foreign exchange risk priced conditionally in Africa's equity markets. A recap of major findings follows.

The baseline model specifications show that neither the world market factor nor the foreign exchange risk factor significantly explains the pricing kernel for equity securities in Africa. Both being international factors, these results are not surprising since local factors are expected to exert more influence on the pricing kernel in partially segmented equity markets. Estimates of factor risk premia yield zero-beta rates between 0.6% and 1.1% per month, which is found to be reasonable for Africa's money markets. Results also suggest that both the foreign exchange risk factor and world market risk factor are conditionally priced in Africa's equity markets. The specification with Eurodollar deposit rates (UDR) as the conditioning variable is rejected by the data. USA term premium (UTP) is found to be the most interesting conditioning variable by both the dynamic performance methodology of Farnsworth et al. (2002) and pricing error diagnostics.

I conduct robustness checks using a different set of conditioning variables. Results confirm that the two risk factors (world equity market returns and foreign exchange rates) do not significantly contribute to the equity pricing kernel. The alternative specifications find the time-varying world market factor appearing significantly priced but do not find foreign exchange risk factors priced individually. However, the Wald test finds foreign exchange risk factors to be jointly significant. The zero-beta rate still lies within the expected range. Further checks on the model are performed by incorporating the time-varying intercept term, hitherto excluded from the tests: consistent with Iqbal et al. (2010), results are mixed, making it difficult to conclude that inclusion of the time-varying intercept improves/deteriorates the model's performance. This result suggests that the model specification without time-varying intercept, which I have used in baseline tests, is better suited for this particular analysis than the specification with time-varying intercept.

6.3 The Dynamic Relation between Foreign Exchange Rates and Net Portfolio Flows

International finance researchers and monetary policy mandarins across the world are concerned about the impact of international capital flows on the macroeconomic stability of recipient countries. Of particular concern have been the disruptive effects of capital flows driven by push factors, such as low interest rates in source markets. Such flows are typically transitory in nature and are largely motivated by investors' desire for short-term gains. In this category falls "non-permanent" flows such as short-term debt, long-term debt and equity investments. The nature of these flows is such that they move with relative ease across international borders in search of better return prospects. International flows that exhibit this tendency have been labeled as "hot money" in the literature. The inflow of hot money is undesirable from the point of view of the recipient country as it impacts negatively on the country's current account through its perceived influence on real foreign exchange rates. Recognizing this fact, this study sought to establish whether causality exists between real exchange rates and flows of international portfolio capital in Africa's major capital markets. In this nexus, the study also investigated whether portfolio flows to major African destinations exhibited persistence.

From the US Federal Reserve Bank website, I obtained monthly portfolio flows data for four major African countries: Egypt, Morocco, Nigeria and South Africa. These data are used, first, to investigate the volatility of portfolio flows. With volatility measured simply as the coefficient of variation (CV), results suggest that Egypt had the most volatile portfolio flows, Nigeria the lowest, over the study period. Further, African countries' portfolio flows volatilities are largely higher than those recorded in developed markets (see for example Claessens et al., 1995). Next, I investigate the persistence of portfolio flows: results of autocorrelation functions, variance ratio tests and impulse response functions all indicate that portfolio flows to Africa's markets have a very short (less than one month) memory.

Next, I run several tests geared towards establishing the nature of the relationship between real foreign exchange rates and international portfolio flows. First, unit root tests using Zivot and Andrews (1990) and Augmented Dickey and Fuller (1979) procedures find that real exchange rates are integrated of the first order while international portfolio flows are integrated of order zero. This result implies that cointegration analysis, which is not indicated when one of the series is stationary, cannot be conducted. Second, Granger-causality tests, variance decomposition and analysis of impulse responses from monthly data show that causality between the two variables is both country-dependent and time-varying. The findings in this section can be attributed to country idiosyncrasies around attractiveness of the specific capital markets to foreign investors, their perceived openness to international capital flows, legal structures and accounting practices.

Lastly, from the International Financial Statistics database, I obtained annual international portfolio flows data for six countries: Botswana, Egypt, Kenya, Morocco, Nigeria and South Africa. I conduct unit roots and causality tests on the annual data using panel data procedures. From the several unit roots tests, the hypothesis of a panel unit root is rejected for net portfolio flows while it cannot be rejected for real foreign exchange rates. Both series are stationary in first differences. Causality tests using procedures proposed by Granger (1969) and Holtz-Eakin et al. (1989) concur: there is strong unidirectional causality running from real exchange rates to net portfolio flows in Africa's markets.

6.4 Policy Implications

Several policy implications can be drawn from the findings of this study. First, tests results from the unconditional asset pricing model generally show that equity markets in Africa are partially segmented/partially integrated. Whereas studies from other regions of the world also report similar findings, it is important to point out that African economies will find it difficult to realize their full growth potential unless their financial markets are fully integrated with financial markets elsewhere in the world. Closer integration of Africa's financial markets with other markets can be achieved if African countries pursue liberalization policies more meticulously. Bekaert et al. (2004) find that countries that have liberalized their capital accounts, equity markets or banking sectors, should display a closer association between growth opportunities and future real GDP and investment: in the study, a country's growth opportunities is measured by investigating how her industry mix is priced in global capital markets using the price earnings ratios of global industry portfolios.

Further, Neumann et al. (2009) have shown that portfolio flows appear to show little response to capital liberalization while foreign direct investment (FDI) flows show significant increases in volatility for the emerging markets considered in their study. In addition to expanding growth opportunities, increased market integration has the benefit that it encourages more market participants and promotes competition which leads to better price discovery. This benefit is underscored by Pukthuanthong (2009) who finds that firms from low integrated markets enjoy great benefits when they enter into high integrated markets. Thus, African governments should strive to further strengthen their financial markets by adopting policies that encourage growth in the number of listed companies in their bourses, enacting more investment-friendly legal, accounting and reporting regulations, reviewing the rules governing cross-listing and legislation disallowing or restricting foreign investor participation in domestic securities markets.

The second policy issue emanates from the finding that foreign exchange risk is weakly priced unconditionally when returns from equity investments are measured in the euro but not priced when the same returns are measured in the US dollar. This finding draws from the central role played by the euro as the invoicing currency for bilateral trade and investments between African countries and their major trading partner - the European Union. Evidence is also available of the increasing importance of the euro as a world currency and of a slightly diminishing role of the US dollar (Kamps, 2006). An important implication of these findings is that monetary policy managers in Africa should start regarding the euro as an important reserve currency. Thus, in addition to reserves of gold and the US dollar, the euro should increasingly form a key component of foreign currency reserves held by central banks of many of these countries.

Third, Dumas and Solnik (1995) argue that it is natural to test asset pricing models in their conditional form because investors' decisions are generally informed by existing information that affect the performance of securities of their interests. Such information, which include past returns on securities of interest and other market fundamentals, condition investors to behave in a certain way, and cannot be ignored in asset pricing tests. Indeed, unlike tests based on the unconditional asset pricing models, my conditional asset pricing model tests provide strong evidence suggesting that foreign exchange rate changes command significant time-varying risk premia in Africa's equity markets. Thus, international investors keen on incorporating Africa's equity securities into their portfolios would find foreign exchange exposure hedging useful. Yet, with the exception of South Africa, none of the countries studied have developed markets for derivative securities. A policy implication of these findings is that securities markets regulatory bodies and other financial markets agencies in African countries should consider instituting operationally viable markets for derivative securities within their financial jurisdictions, with a view to promoting risk management activities among market players.

An important issue that arose from my data search effort was that of the unavailability of debt markets indices for most of Africa's capital markets. For the few countries with debt indices, such as Egypt and South Africa, the time series were short and available only in low frequencies. Lack of debt markets data is indicative of poor development of the private debt markets outside commercial banking; indeed, it emerged that many of the African bourses had very few or no corporate debt listings and their debt counters were dominated by Treasury issues. Additionally, data available on money market instruments showed a tendency to remain constant for long periods, suggesting that, either interest rates on those securities are not determined through the interaction of the demand for and supply of money, or data capturing and storage is weak.

Several policy implications stem from the issue of lack of data availability: First, African governments should consider instituting policies that expedite the development of debt markets where corporate bonds and other long-term debt instruments can be traded publicly. This would not only enable corporations to access cheaper capital, it would also ease their valuation as information about their debt and other financing arrangements become publicly accessible. This policy recommendation follows from Ojah and Pillay (2009), who find that public debt-issuing firms in emerging markets experience significant reductions in both overall and systematic risks, and incur lower cost of capital following issuance than non-public debt issuers. Their findings suggest that deepening national debt markets can be a fruitful financial market development exercise. Second, to curb price distortion in the money markets, governments should consider moving towards full liberalization of their interest rate regimes so that money demand and supply can play a key role in the determination of interest rates. Third, African governments should boost their data gathering and preservation capacities by developing and funding national institutions which can, in turn, be entrusted with key economic, financial and other data.

Further policy lessons can also be gleaned from portfolio flows studies. The study of monthly data indicates that portfolio flows to African countries are low, volatile and transitory; a finding that puts these flows firmly within the realm of hot money. To mitigate the hot money problem, African governments should consider the following policy proposals. First, portfolio flows could be boosted and stabilized by employing more "sound" monetary policy. Sound monetary policy principles include such attributes as independence, transparency, predictability, rules rather than discretion, and accountability (Shah and Patnaik, 2008). As the duo point out, sound monetary policy frameworks stabilize business cycles, stabilize capital flows, and in turn, are not attenuated by fluctuations in capital flows. Further, Mody and Murshid (2005) observe that countries with better policies have greater success in absorbing foreign inflows. This could be partly because improved policies raise the marginal product of new investments and creates a conducive environment for the diffusion of new technologies and ideas intrinsic to foreign capital; and partly because it reduces the risk of holding domestic assets, which in turn, by discouraging capital outflows, enhances the relationship between capital flows and investment. Second, African governments should embrace the concept of sovereign credit ratings and strive to be rated highly. Gande and Parsley (2004) have shown that sovereign downgrades are strongly associated with outflows of capital from the downgraded country while improvements in a country's sovereign rating are not associated with discernable changes in equity flows.

Third, Gande and Parsley (2004) provide results suggesting that lowering corruption could mitigate some of the perceived negative effects associated with capital flows. To this end, African governments should institute clear measures that send strong signals to investors about their low tolerance to corruption. African countries continue to receive poor corruption ratings: Transparency International (2009) places Ghana ahead of nine African countries surveyed, with a Corruption Perception Index of 3.9 out of a possible 10. Morocco (3.5) is second, followed by Nigeria (2.7) and Kenya (2.1). These ratings are substantially lower than those of countries such as Switzerland (9.0). Fourth, African governments should put in place policies that encourage the growth of their financial markets in order to attract a sustained flow of international capital. Further, regional integration pursuits currently taking root in the continent should be hastened and embraced by all governments. These proposals are informed by Lozovyi and Kudina (2007), who find that underdeveloped financial markets are a factor restricting portfolio inflows to the Commonwealth of Independent States (CIS) and by De Santis (2006), who finds clear evidence that (i) portfolio asset flows are influenced positively by the relative size of the recipient countries' financial markets and (ii) cross-border portfolio flows among euro area countries have increased due to the catalyst effect of the European Monetary Union.

Analysis of monthly portfolio flows data reveals the existence of causality running from net portfolio flows to real exchange rates for three countries, namely, South Africa[57] (full study period), Egypt (January 1997 to December 2003), and Nigeria (January 2004 to December 2009). It is also clear from impulse response functions that real exchange rates largely appreciate in response to shocks in portfolio flows. In response, government agencies overseeing financial markets development should consider instituting policies that encourage the creation of more

[57] It is important to recall that bidirectional causality is found for the full sample in the case of South Africa. However, policy inference, at this point, is made only in respect of causality from portfolio flows to real exchange rates. This is for consistency with other cases in which only unidirectional causality in the direction mentioned is established.

financial instruments. The resulting deepening of financial markets would partly mitigate the effects of portfolio flows on real exchange rates. This proposal is supported by Saborowski (2009) who finds that the exchange rate appreciation effect of foreign direct investments inflows is attenuated when financial and capital markets are larger and more active.

Finally, annual portfolio flows data investigations reveal the existence of strong panel causality from real exchange rates to portfolio flows. Thus, if exchange rates have a tendency to fluctuate wildly and irregularly, portfolio flows would respond similarly. Instability in portfolio flows is not desirable as it can be disruptive to the economy in many ways: for instance, Hau and Rey (2006) have shown that portfolio flows are intertwined with stock market performance. To have stability in portfolio flows, African governments must exercise strict discipline in monetary policy management. In particular, since this study has demonstrated that inflation and interest rate differentials have an influence on foreign exchange rate changes, it is imperative that measures be put in place to ensure their stability if portfolio flows have to remain stable and non-disruptive to Africa's economies.

6.5 Recommendations for Further Research

The question as to whether currency risk is priced in capital markets has been and may remain empirically unresolved for a while. Similarly, the nexus between foreign exchange rates and capital flows has been a controversial one in the literature and debate is likely to continue as more researchers provide further evidence on the issue. For Africa, where academic discourse on both issues has largely began with this work, paucity of data and research capacity may slacken the pace of debate, implying that these issues might remain thorny for longer than necessary. However, as more capital markets are developed, existing markets further expanded, data becomes available at firm and industry levels for many African countries, and research capacity building improved, further investigations of the many outstanding issues not addressed by this study will become more feasible. There are many possible extensions to this study for which future research is desirable. They include, although not necessarily limited to the following:

Firstly, the foreign exchange risk pricing results could be enriched by a larger data set covering more countries in Africa. Secondly, it will be recalled that, for this study, usable firm level data for the unconditional asset pricing model was available only for South Africa. Results from the data showed, contrary to aggregate market data results, that currency risk is priced in South Africa's equity markets with returns measured in the US dollar. Had similar data been available for the other six countries investigated, it would have been possible to conduct a cross-country comparison of unconditional foreign exchange risk pricing in major equity markets in Africa. Additionally, the availability of such data could have made it possible to study conditional currency risk pricing through the more interesting Fama and French (1993) factors. Future studies can endeavor to fill these needs as the data become available.

Thirdly, research has been done in many developing regions to establish the major factors that drive capital flows. My literature search could not trace any such studies conducted in the African region: this can be attributed largely to the unavailability of capital flows data. For similar reasons, and because it was not within the scope of the current work, a similar study has not been conducted here. As an extension to this work and in order to understand capital flows in Africa better, future research can try to establish which of domestic (pull) factors or foreign (push) factors have a bigger influence on capital flows into and out of Africa's capital markets.

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APPENDICES

APPENDIX I: Extracting Foreign Exchange Risk Premia through the SDF Model

A1.1. The SDF Model and International Asset Pricing

Assume, in the spirit of Robotti (2001), that there are $L+1$ countries and a set of $N=n+1$ assets (excluding the risk-free asset with nominal returns denominated in the measurement currency.) These include n risky assets or portfolios of risky assets and the world portfolio of risky assets. Assume further that nominal returns are measured in the $L+1^{st}$ currency, known as the reference or numeraire currency. In this study, the US dollar is used as the numeraire.

Cappiello and Panigirtzoglou (2008) explain that equation (23) (in the text) still holds when (gross) returns are denominated in a foreign currency. Let S_t denote the spot exchange rate, defined as the number of units of the foreign currency per unit of the domestic currency, that is, the indirect method of exchange rate quotation is used. Further, let R^*_{t+1} be the (gross) nominal rate of return on the foreign asset and let M^*_{t+1} represent the foreign investor's pricing kernel. I use a simple numerical illustration to develop the next formula.

Suppose a US-based investor (the domestic investor) has \$1000 to invest in a foreign country, say Kenya. The current spot rate of exchange S_t is KES 50 to the dollar; the anticipated future (end of first period) spot rate of exchange is KES 40 to the dollar and the one-year money market nominal rate of interest is 20% in Kenya. This implies that the (gross) nominal foreign (Kenyan) rate of return is 1.2. It is assumed, for working purposes, that the Uncovered Interest Rate Parity (UIP) relationship holds exactly. The fair (no-arbitrage) money market rate of interest for US dollar-denominated securities can be computed using the following simple procedure.

The US investor converts the \$1000 into KES at the current spot rate, which is KES 50 to the US dollar, obtains KES (1000×50) and invests the proceeds in Kenyan (foreign) money market securities for one year. At the end of the year, the investor has a total of KES $(1.2 \times 1000 \times 50)$. This amount is converted back into US dollars at the spot rate of KES 40 to the dollar to obtain \$1500 [or $(1.2 \times 1000 \times 50) \div 40$]. The no-arbitrage UIP-implied (gross) nominal rate of return in dollar (domestic currency) terms is 1.5 (\$1500 ÷ \$1000).

The illustration demonstrates clearly that the foreign rate of return, R^*_{t+1}, can be related to the domestic rate of return, R_{t+1}, through the spot rates of exchange. In the illustration, the foreign rate of return, 1.2, is the same as the domestic rate of return, 1.5, multiplied by the future spot rate of exchange, KES 40/\$ and divided by the current spot rate of exchange, KES 50/\$. Symbolically,

$$R^*_{t+1} = R_{t+1}\left(\frac{S_{t+1}}{S_t}\right) \tag{A1}$$

With the assumption that the above money market security (and indeed all the *N* assets under consideration) is traded in both the foreign and domestic currencies, the basic valuation equation (23) can be restated, with returns denominated in the foreign currency, as follows:

$$E_t\left(R_{t+1}\frac{S_{t+1}}{S_t}M^*_{t+1}\right) = 1 \quad \Leftrightarrow \qquad E_t(R^*_{t+1}M^*_{t+1}) = 1 \tag{A2}$$

Relating equation (A2) to equation (23), it follows that the following relationship must be true in complete markets:[58]

$$\frac{S_{t+1}}{S_t}M^*_{t+1} = M_{t+1} \tag{A3}$$

Taking natural logarithms on both sides of equation (A3) and rearranging gives

$$ln\left(\frac{S_{t+1}}{S_t}\right) = \ln(M_{t+1}) - \ln(M^*_{t+1}) \tag{A4}$$

Equation (A4) states that a change in the logarithm of the exchange rate equals the difference between the logarithms of the domestic and the foreign pricing kernels. Thus, a decrease in the domestic marginal utility of consumption (that is, a decrease in the pricing kernel) leads to a depreciation in the price of domestic consumer goods relative to foreign consumer goods (Cappiello and Panigirtzoglou, 2008). Stated differently, a decline in the domestic stochastic

[58] Markets are said to be complete if any source of uncertainty can be perfectly hedged using existing hedging instruments. In such a market, the stochastic discount factor or the pricing kernel is unique (See, for example, Cochrane (2000) and Cappiello and Panigirtzoglou (2008) for a detailed discussion).

discount factor (SDF) relative to the foreign SDF is accompanied by a decline in value of the domestic currency.

As already explained, the model in equation (A4) assumes that markets are complete. However, it is known that markets are generally incomplete so that it is possible to find a multiplicity of pricing kernels that are consistent with observed prices (Balduzzi and Robotti, 2001). Indeed, Ferson (2003: 747) remarks that there will be some M_{t+1} that "works" as long as there are no redundant asset returns. Notwithstanding this market reality, Balduzzi and Robotti (2001) and Robotti (2001) demonstrate that choice of the pricing kernel with the lowest variability can solve the problem of multiple admissible pricing kernels. Consequently, equation (A4) should hold, with a unique pricing kernel, for investors operating in markets that are incomplete.

A1.2. Empirical Specification of the Stochastic Discount Factor (SDF) Model

Empirical implementation of the SDF model takes the view of a global investor whose returns are calculated in U.S. dollars. The investor is assumed to be unhedged in exchange rates. Parameter estimation takes the Cappiello and Panigirtzoglou (2008) approach. But, unlike Cappiello and Panigirtzoglou (2008) who use a tripartite framework to price currency risk in three countries, this study follows a bilateral approach with the USA as the domestic country and the US dollar as the reference currency. The bilateral approach has been successfully used by Drobetz et al. (2002) to estimate market risk premia for several emerging market economies with the Swiss Franc as the reference currency. Their study, which employs one-, two- and three-factor models, fails to reject the overall goodness-of-fit of the stochastic discount factor model in each of the three cases. Drobetz et al. (2002) do not include foreign exchange risk in their study.

Robotti (2001) explains that a *normalized* pricing kernel can be constructed by scaling the stochastic discount factor by the risk-free rate. This is achieved with the aid of equation (23) in the text. Let $q_{t+1} \equiv M_{t+1}R_{ft}$. Then, equation (23) can be expressed as $E_t(q_{t+1}) = 1$. Thus, the mean of the normalized pricing kernel is one. As Balduzzi and Robotti (2001) explain, by setting the mean of the normalized pricing kernel equal to one, the need to model the conditional mean of the "usual" pricing kernel, *M*, is obviated. Plugging q_{t+1} into equation (23) yields

$$E_t\left(R_{i,t+1} \cdot \frac{q_{t+1}}{R_{ft}}\right) = 1 \quad \Leftrightarrow \quad E_t\left(R_{i,t+1} q_{t+1}\right) = R_{ft} \tag{A5}$$

Since there are a total of *N* assets under consideration, it is usually convenient to present the relationship in equation (A5) in the form of the following pricing vector.

$$E_t(\boldsymbol{R}_{t+1} q_{t+1}) = \boldsymbol{\iota} R_{ft} \tag{A6}$$

where $\boldsymbol{R}_{t+1}$ is the vector of *N* risky gross real returns and $\boldsymbol{\iota}$ is an *N*-vector of ones. Following Cochrane (2000: 104) the normalized pricing kernel can be expressed as a linear combination of factors (represented by asset returns), in the following form:

$$q_{t+1} = a_t + b_{1t} R_{1,t+1} + b_{2t} R_{2,t+1} + \ldots + b_{kt} R_{k,t+1} = a_t + \boldsymbol{b}_t' \boldsymbol{R}_{t+1} \tag{A7}$$

where $\boldsymbol{b}_t$ is a $(k \times 1)$ vector of factor loadings and, $\boldsymbol{R}_{t+1}$ is a $(k \times 1)$ vector of asset returns. The result in equation (A7) is plugged in equation (A6), and the result rearranged to give:

$$E_t\left[\boldsymbol{R}_{t+1}(a_t + \boldsymbol{b}_t' \boldsymbol{R}_{t+1}) - \boldsymbol{\iota} R_{ft}\right] = 0 \tag{A8}$$

This equation characterizes the population moment conditions, sometimes called orthogonality conditions. The latter term, attributed to Hansen (1982) and popularized by Ferson and Foerster (1994) only becomes meaningful after conditioning information (surrogated by instrument variables with returns orthogonal to the model's residuals) is incorporated in the model.

A1.3. Incorporating Conditioning Information

The unobservable market-wide information set, Ω_t, at the disposal of investors at time *t*, is typically proxied in econometric analysis by a set of predefined instrumental variables, Z_t, assumed to contain time *t* information. The set of instruments comprises of observable economic variables or portfolios of assets whose returns mimic the returns on those variables. Following Cochrane (1996), the instruments set is used to scale the factors so as to allow the parameters to be modeled as linear functions of Z_t:

$$a_t = a_0 + a_1 Z_t\,; \quad b_t = b_0 + b_1 Z_t. \tag{A9}$$

Assuming only one factor, R_{t+1}, and one instrument, Z_t, equation (A7) can now be rewritten with the normalized pricing kernel expressed as a scaled multifactor model with constant coefficients:

$$q_{t+1} = a_0 + a_1 Z_t + (b_0 + b_1 Z_t) R_{t+1} = a_0 + a_1 Z_t + b_0 R_{t+1} + b_1 Z_t R_{t+1} \tag{A10}$$

Thus, instead of a single factor model with time-varying factor weights, there results a three factor model with constant (or fixed) weights (Cochrane, 2000: 135). The procedure in equation (A10) can be easily extended to a multifactor, multi-instrument framework. Because the weights are fixed, the scaled factor model can be tested unconditionally by applying the Law of Iterated Expectations.[59] Now, since $q_{t+1} = a + \boldsymbol{b}'\boldsymbol{R}_{t+1}$, and $\boldsymbol{z}_t$ is a set of s instrumental variables, the first sample moment conditions can be expressed as:

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T} f_t(\theta) = \frac{1}{T}\sum_{t=1}^{T} (a + \boldsymbol{b}'\boldsymbol{R}_{t+1} - 1)\, \boldsymbol{z}_t = \boldsymbol{0}_s \tag{A11}$$

where $\boldsymbol{0}_s$ is an $(s \times 1)$ vector of zeros. The rest of the definitions remain as before. The second sample moment conditions to be estimated are as follows:

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T} f_t(\theta) = \frac{1}{T}\sum_{t=1}^{T} \left\{ \left[\boldsymbol{R}_{t+1}(a + \boldsymbol{b}'\boldsymbol{R}_{t+1}) - \boldsymbol{\iota} R_{ft} \right] \otimes \boldsymbol{z}_t \right\} = \boldsymbol{0}_{ks} \tag{A12}$$

where $\boldsymbol{0}_{ks}$ is an $(ks \times 1)$ vector of zeros; k is the number of factors; $\otimes$ is the Kronecker product (multiply each term in the square bracket by every instrument); $\boldsymbol{0}$ is a vector of zeros. Equation (A11) is the multi-factor system to be tested. Data analysis makes use of the Generalized Method of Moments. In the GMM terminology, $g_T(\theta)$ represents the *sample average* of pricing errors. The subscript T (the sample size) indicates the dependence of this statistic on the sample; θ represents all the parameters to be estimated. With returns measured in the US dollar, the GMM estimates of the coefficients in equation (A12) are used to estimate the pricing kernels for the US investor. Following Cappiello and Panigirtzoglou (2008), estimates of the pricing kernels for the

[59] The law states that taking an expected value using less information of an expected value that is formed on more information, gives the expected value using less information (Cochrane, 2000: 129). For instance, $E[E_t(X)] = E(X)$.

African countries under investigation are then derived from these results using the following stochastic process.[60]

$$ln(M_{t+1}) = -lnR_{ft} - \frac{1}{2}\lambda_t^2 - \lambda_t \Delta W_{t+1} \quad \text{(A13)}$$

where W_{t+1} is the source of uncertainty and the quantity λ_t (the standard deviation of M_{t+1} at time t) is the volatility of the pricing kernel, also referred to as the market price of risk.

A1.4. Estimation of Currency Risk Premia[61]

Consider a one-period risk-free asset with gross return $R_{ft} = 1 + r_{ft}$, where the rate r_{ft} is the one-month yield on the risk-free asset. Assuming its payoff at time $t + 1$ is certain and equal to one unit of consumption in all states of the world, it will satisfy

$$E_t(M_{t+1}) = \frac{1}{R_{ft}} \quad \text{(A14)}$$

Equation (A14) indicates that the pricing kernel is expected to fall according to the risk-free rate. Taking the logarithm of both sides of equation (A14) followed by a Taylor expansion around the term $E_t ln(M_{t+1})$, the following approximation results:[62]

$$E_t ln(M_{t+1}) = -lnR_{ft} - \frac{1}{2}Var_t[ln(M_{t+1})] \quad \text{(A15)}$$

where $Var_t(\cdot)$is the variance operator conditional on the information set Ω_t. The stochastic process for the pricing kernel can then be written as

$$ln(M_{t+1}) = -lnR_{ft} - \frac{1}{2}\lambda_t^2 - \lambda_t \Delta W_{t+1} \quad \text{(A16)}$$

[60] See Appendix B for a detailed discussion.
[61] This discussion is informed largely by Cappiello and Panigirtzoglou (2008)
[62] The assumption of log-normality of the pricing kernel is implied in these calculations. Notice that the higher order terms of the Taylor series have been omitted from the output.

where W_{t+1} is the source of uncertainty and the quantity λ_t (the standard deviation of M_{t+1} at time *t*) is the volatility of the pricing kernel also referred to as the market price of risk.[63] A similar relationship holds for the pricing kernel of the foreign investor:

$$ln(M^*_{t+1}) = -lnR^*_{ft} - \frac{1}{2}(\lambda^*_t)^2 - \lambda^*_t \Delta \mathrm{W}^*_{\mathrm{t+1}} \tag{A17}$$

Substituting equations (A16) and (A17) into equation (A4) yields

$$ln\left(\frac{S_{t+1}}{S_t}\right) = lnR^*_{ft} - lnR_{ft} + \frac{1}{2}[(\lambda^*_t)^2 - \lambda^2_t] - \lambda_t \Delta W_{t+1} + \lambda^*_t \Delta \mathrm{W}^*_{\mathrm{t+1}} \tag{A18}$$

Taking expectations conditional on information available at time *t*, equation (A18) reduces to

$$E_t\left[ln\left(\frac{S_{t+1}}{S_t}\right)\right] = \left(lnR^*_{ft} - lnR_{ft}\right) + \frac{1}{2}[(\lambda^*_t)^2 - \lambda^2_t] \tag{A19}$$

Equation (A19) is uncovered interest rate parity relationship augmented by risk premia. This is the equation that is used to compute the foreign exchange risk premia.[64]

A1.5. Operationalization of the Model

From equation (23), the following relationship holds for the risk-free asset:

$$E_t(M_{t+1}R_{ft}) = 1 \qquad \Rightarrow \qquad E_t(q_{t+1}) = 1 \qquad \Rightarrow \qquad E_t(q_{t+1} - 1) = 0 \tag{A20}$$

This equation together with equations A(11) and (A12) are used to compute the pricing kernels. To illustrate, the pricing kernel for the US investor to South Africa is estimated according to the following system of equations:

[63] The market price of risk is the excess return per unit of volatility of an asset that is perfectly positively correlated with the source of uncertainty, W_t. The source of uncertainty is assumed to be perfectly negatively correlated with the pricing kernel.

[64] Notice that the first term on the right hand side is the interest rate differential and would drive the exchange rate differential in the absence of uncertainty in the economy. If there is uncertainty, risk-averse investors would require compensation for any systematic sources. Thus, a priced currency premium would be captured by the second term.

$$q_{t+1}^{US} - 1 = 0 \qquad \text{(i)}$$

$$R_{M,t+1}^{US} q_{t+1}^{US} - R_{ft}^{US} = 0 \qquad \text{(ii)}$$

$$R_{B,t+1}^{US} q_{t+1}^{US} - R_{ft}^{US} = 0 \qquad \text{(iii)}$$

$$R_{M,t+1}^{ZA} q_{t+1}^{US} - R_{ft}^{US} = 0 \qquad \text{(iv)}$$

$$R_{B,t+1}^{ZA} q_{t+1}^{US} - R_{ft}^{US} = 0 \qquad \text{(v)}$$

$$R_{MM,t+1}^{ZA} q_{t+1}^{US} - R_{ft}^{US} = 0 \qquad \text{(vi)}$$

where $q_{t+1}^{US} = \hat{a} + \hat{b}_1 R_{M,t+1}^{US} + \hat{b}_2 R_{B,t+1}^{US} + \hat{b}_3 R_{M,t+1}^{ZA} + \hat{b}_4 R_{B,t+1}^{ZA} + \hat{b}_5 R_{MM,t+1}$ such that subscripts *M* and *B* respectively denote US dollar returns on equity market and bond market indexes; *MM* is the US dollar return on the South African one-month Treasury bill. The South African investor's pricing kernel is obtained using equation (A4):

$$ln\left(\frac{S_{t+1}}{S_t}\right) = ln\left(\frac{q_{t+1}^{US}/R_{ft}^{US}}{q_{t+1}^{ZA}/R_{ft}^{ZA}}\right) \quad \Rightarrow \quad q_{t+1}^{ZA} = \left(\frac{S_t}{S_{t+1}}\right)\left(\frac{R_{ft}^{ZA}}{R_{ft}^{US}}\right) q_{t+1}^{US} \qquad \text{(A21)}$$

These computations are replicated for the remaining African countries.

APPENDIX II: Some VAR Tests Results for Monthly Data

Table A1
Lag structure choice for unit root tests

Lag	1	2	3	4	5	6	7	8	9	10	11	12	13
Egypt Exchange rates													
AIC	-4.787*	-4.764	-4.758	-4.757	-4.738	-4.719	-4.700	-4.680	-4.659	-4.639	-4.619	-4.599	-4.578
SBC	-4.743*	-4.705	-4.679	-4.656	-4.618	-4.579	-4.538	-4.498	-4.456	-4.414	-4.373	-4.331	-4.288
HQC	-4.766*	-4.740	-4.726	-4.716	-4.689	-4.662	-4.634	-4.606	-4.577	-4.547	-4.519	-4.490	-4.460
Egypt Portfolio Flows													
AIC	6.102*	6.112	6.117	6.135	6.151	6.149	6.136	6.122	6.136	6.154	6.174	6.196	6.217
SBC	6.141*	6.171	6.196	6.235	6.271	6.289	6.298	6.304	6.339	6.379	6.421	6.464	6.507
HQC	6.118*	6.136	6.149	6.176	6.200	6.206	6.202	6.196	6.218	6.246	6.274	6.305	6.334
Morocco Exchange Rates													
AIC	-4.579*	-4.581	-4.563	-4.563	-4.570	-4.555	-4.550	-4.554	-4.533	-4.514	-4.504	-4.484	-4.491
SBC	-4.540*	-4.522	-4.484	-4.463	-4.451	-4.414	-4.388	-4.372	-4.330	-4.289	-4.258	-4.216	-4.201
HQC	-4.056*	-4.557	-4.531	-4.522	-4.522	-4.498	-4.484	-4.480	-4.451	-4.422	-4.404	-4.375	-4.373
Morocco Portfolio Flows													
AIC	3.692*	3.708	3.726	3.737	3.743	3.746	3.756	3.775	3.794	3.799	3.821	3.794	3.796
SBC	3.731*	3.768	3.805	3.837	3.863	3.887	3.917	3.957	3.997	4.024	4.067	4.062	4.086
HQC	3.708*	3.732	3.758	3.778	3.792	3.803	3.821	3.849	3.876	3.891	3.921	3.903	3.914
Nigeria Exchange Rates													
AIC	-1.528*	-1.513	-1.491	-1.473	-1.455	-1.437	-1.417	-1.402	-1.389	-1.372	-1.355	-1.336	-1.321
SBC	-1.489*	-1.451	-1.411	-1.374	-1.335	-1.296	-1.256	-1.220	-1.185	-1.147	-1.109	-1.067	-1.031
HQC	-1.512*	-1.486	-1.459	-1.433	-1.407	-1.380	-1.352	-1.328	-1.306	-1.281	-1.255	-1.227	-1.203
Nigeria Portfolio Flows													
AIC	3.303*	3.307	3.312	3.324	3.336	3.342	3.358	3.379	3.389	3.391	3.412	3.409	3.407
SBC	3.342*	3.366	3.391	3.424	3.456	3.482	3.519	3.561	3.592	3.616	3.659	3.677	3.698
HQC	3.319*	3.3307	3.344	3.364	3.384	3.400	3.424	3.453	3.471	3.482	3.512	3.518	3.525
South Africa Exchange Rates													
AIC	-3.192	-3.195*	-3.180	-3.161	-3.142	-3.123	-3.115	-3.112	-3.097	-3.103	-3.084	-3.076	-3.061
SBC	-3.153*	-3.136	-3.101	-3.061	-3.022	-2.982	-2.954	-2.927	-2.893	-2.878	-2.838	-2.808	-2.771
HQC	-3.176*	-3.171	-3.148	-3.120	-3.093	-3.066	-3.050	-3.038	-3.014	-3.012	-2.984	-2.967	-2.943
South Africa Portfolio Flows													
AIC	3.807*	3.826	3.845	3.864	3.881	3.850	3.856	3.876	3.889	3.908	3.929	3.950	3.971
SBC	3.846*	3.885	3.925	3.963	4.000	3.990	4.017	4.058	4.092	4.133	4.176	4.218	4.261
HQC	3.823*	3.850	3.878	3.904	3.930	3.907	3.921	3.950	3.971	3.999	4.029	4.059	4.088

AIC = Akaike Information Criteria; SBC = Schwatz (Bayesian) Information Critera; HQC = Hannan-Quinn Information Criteria. * denotes the chosen lag length. The SBC chosen lag structure is used in the analysis. Maximum lag length is chosen using the formula: $k_{max} = int[12(T/100)^{0.25}]$ (see Hayashi, 2000).

Table A2
Lag structure choice for vector autoregressions

Lag	1	2	3	4	5	6	7	8	9	10	11	12	13
Panel A: Full Sample													
Egypt													
AIC	1.106*	1.157	1.188	1.228	1.250	1.238	1.204	1.238	1.282	1.320	1.337	1.356	1.370
SBC	1.313*	1.447	1.561	1.684	1.788	1.860	1.908	2.025	2.152	2.273	2.373	2.475	2.572
HQC	1.190*	1.275	1.340	1.413	1.468	1.491	1.490	1.558	1.635	1.707	1.758	1.810	1.859
Morocco													
AIC	-1.072*	-1.027	-0.979	-0.949	-0.960	-0.962	-0.929	-0.899	-0.866	-0.860	-0.852	-0.814	-0.854
SCB	-0.864*	-0.737	-0.606	-0.493	-0.421	-0.340	-0.225	-0.112	0.004	0.093	0.184	0.305	0.348
HQC	-0.987*	-0.909	-0.828	-0.763	-0.741	-0.709	-0.643	-0.579	-0.513	-0.473	-0.431	-0.360	-0.366
Nigeria													
AIC	1.808*	1.857	1.871	1.894	1.948	1.999	2.046	2.089	2.137	2.158	2.203	2.238	2.254
SBC	2.015*	2.147	2.244	2.349	2.486	2.620	2.751	2.877	3.007	3.111	3.239	3.357	3.456
HQC	1.892*	1.975	2.023	2.079	2.167	2.251	2.333	2.409	2.491	2.545	2.624	2.692	2.743
South Africa													
AIC	0.503*	0.526	0.568	0.600	0.604	0.607	0.634	0.668	0.713	0.714	0.762	0.796	0.830
SBC	0.710*	0.816	0.941	1.056	1.143	1.228	1.339	1.455	1.583	1.667	1.798	1.915	2.031
HQC	0.587*	0.644	0.720	0.785	0.823	0.859	0.921	0.988	1.067	1.102	1.183	1.251	1.318
Panel B: First Sub-period													
Egypt													
AIC	-0.321*	-0.254	-0.164	-0.161	-0.094	-0.006	0.076	0.179	0.237	0.329	0.375	0.396	-0.128
SBC	-0.002*	0.192	0.410	0.540	0.734	0.950	1.159	1.390	1.576	1.795	1.969	2.117	1.720
HQC	-0.194*	-0.076	0.064	0.117	0.235	0.374	0.507	0.661	0.769	0.912	1.009	1.080	0.607
Morocco													
AIC	-1.012*	-0.948	-0.851	-0.876	-0.920	-0.909	-0.842	-0.822	-0.793	-0.713	-0.657	-0.733	-0.637
SBC	-0.694*	-0.501	-0.277	-0.175	-0.091	0.047	0.242	0.389	0.546	0.753	0.937	0.988	1.211
HQC	-0.885*	-0.770	-0.622	-0.597	-0.591	-0.529	-0.411	-0.341	-0.260	-0.130	-0.023	-0.048	0.098
Nigeria													
AIC	1.309*	1.392	1.479	1.568	1.534	1.628	1.698	1.798	1.905	1.948	2.046	2.079	2.154
SBC	1.627*	1.838	2.053	2.269	2.362	2.584	2.781	3.009	3.244	3.413	3.640	3.800	4.002
HQC	1.436*	1.569	1.707	1.847	1.863	2.008	2.128	2.280	2.438	2.530	2.680	2.763	2.889
South Africa													
AIC	0.782*	0.872	0.884	0.923	0.920	0.908	0.995	1.043	1.104	1.136	1.235	1.182	1.147
SBC	1.100*	1.318	1.457	1.624	1.749	1.864	2.079	2.254	2.442	2.602	2.828	2.903	2.996
HQC	0.909*	1.049	1.112	1.201	1.249	1.288	1.426	1.524	1.636	1.719	1.869	1.866	1.882

Table A2 Continued

Panel C: Second sub-period													
Egypt													
AIC	1.072*	1.133	1.236	1.319	1.368	1.404	1.363	1.426	1.546	1.599	1.579	1.626	1.731
SBC	1.424*	1.626	1.870	2.094	2.283	2.461	2.560	2.764	3.025	3.219	3.340	3.527	3.773
HQC	1.209*	1.325	1.483	1.621	1.725	1.817	1.831	1.948	2.124	2.231	2.267	2.368	2.528
Morocco													
AIC	-1.504*	-1.410	-1.334	-1.389	-1.332	-1.394	-1.371	-1.380	-1.453	-1.444	-1.399	-1.421	-1.409
SBC	-1.152*	-0.917	-0.700	-0.614	-0.417	-0.338	-0.174	-0.042	0.026	0.176	0.362	0.480	0.634
HQC	-1.367*	-1.218	-1.086	-1.086	-0.975	-0.982	-0.904	-0.857	-0.876	-0.812	-0.711	-0.679	-0.612
Nigeria													
AIC	-0.957	-1.097	-1.106	-1.075	-0.946	-0.831	-1.149	-1.259*	-1.156	-1.126	-1.119	-1.030	-1.147
SBC	-0.605*	-0.604	-0.472	-0.300	-0.031	0.225	0.049	0.079	0.323	0.494	0.641	0.871	0.895
HQC	-0.819	-0.905*	-0.859	-0.772	-0.589	-0.419	-0.681	-0.737	-0.578	-0.494	-0.432	-0.288	-0.350
South Africa													
AIC	0.452*	0.511	0.476	0.497	0.592	0.669	0.460	0.554	0.613	0.538	0.601	0.658	0.654
SBC	0.804*	1.004	1.110	1.271	1.508	1.726	1.657	1.892	2.092	2.158	2.361	2.559	2.696
HQC	0.590*	0.703	0.724	0.799	0.950	1.082	0.928	1.076	1.191	1.171	1.288	1.400	1.451

AIC = Akaike Information Criteria; SBC = Schwatz (Bayesian) Information Critera; HQC = Hannan-Quinn Information Criteria.
* denotes the chosen lag length. The SBC chosen lag structure is used in the analysis: Maximum lag length is chosen using the formula: $k_{max} = int[12(T/100)^{0.25}]$ (see Hayashi, 2000).

Table A3
Forecast error variance decomposition for Morocco (First sub-period)

	Response of net portfolio flows (first differences) to:		Response of (log of) real exchange rates (first differences) to:	
Horizon	NPF	RER	NPF	RER
1	100.00000	0.000000	4.980516	95.01948
2	99.12252	0.877477	6.044138	93.95586
3	99.06923	0.930771	6.071935	93.92806
4	99.04942	0.950584	6.092855	93.90715
5	99.04537	0.954632	6.096649	93.90335
6	99.04435	0.955648	6.097652	93.90235
7	99.04411	0.955888	6.097885	93.90212
8	99.04405	0.955945	6.097941	93.90206
9	99.04404	0.955959	6.097954	93.90205
10	99.04404	0.955962	6.097958	93.90204
11	99.04404	0.955963	6.097958	93.90204
12	99.04404	0.955963	6.097959	93.90204

NPF = net portfolio flows as a proportion of GDP; RER = log of real exchange rates

APPENDIX III: Panel Unit Roots and Cointegration Tests: A Review of Methodologies

Panel unit root tests have been proposed by many researchers (Hadri, 2000; Breitung, 2000; Levin et al., 2002; Im et al., 2003; Carrion-i-Silvestre et al., 2005). As Hadri (2000) explains, the main motivation for testing for stationarity in a panel data instead of single time series is that the power of the test increases with an increase in the number of cross-sections. Levin et al. (2002) develop a panel version of the Augmented Dickey-Fuller (ADF) unit root test that restricts parameters γ_i by keeping them identical across cross-sectional units, but allows the lag order for the first difference terms to vary across cross-sectional units as follows:

$$\Delta y_{it} = \alpha_i + \beta_i t + \gamma_i y_{i,t-1} + \sum_{j=1}^{k} \delta_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \tag{A22}$$

where $t = 1, \ldots, T$ time periods and $i = 1, \ldots, N$ cross-sectional units; k is the number of lags and Δ is the first difference operator. The model in equation (A22), estimated using pooled ordinary least squares (OLS), can also be tested with no trend, i.e., under the restriction that $\beta_i = 0$ for all i, or with no trend and intercept, i.e., $\alpha_i = \beta_i = 0$, for all i. The null hypothesis that there is a unit root, i.e., $\gamma_i = \gamma = 0$, for all i, is tested against the alternative that there is no unit root, i.e., $\gamma_1 = \gamma_2 = \ldots = \gamma < 0$, for all i. The test is based on the statistic, $t_\gamma = \hat{\gamma}/SE(\hat{\gamma})$, where SE is the standard error. To make the pooled OLS regression efficient, the method is implemented with bias correction factors to correct for heterogeneity across cross-sectional elements. The model considered by Breitung (2000) is similar but proffers a different treatment for cross-sectional heterogeneity: appropriate transformations on the variables. A major drawback of these methods is that, under both the null and alternative hypotheses, γ_i is restricted by keeping them identical across cross-sectional units.

Relaxing the assumption of identical first order autoregressive coefficients so as to allow γ to vary under the alternative hypothesis, Im et al. (2003) proposes a test that overcomes some of the weaknesses of the foregoing tests. The test is based on the following data generating process:

$$\Delta y_{it} = \alpha_i + \beta_i y_{i,t-1} + \varepsilon_{it} \tag{A23}$$

where $\alpha_i = (1 - \phi_i)u_i$, $\beta_i = -(1 - \phi_i)$ and Δ is the first difference operator. The null hypothesis of the unit roots is H_0: $\beta_i = 0$ for all $i = 1, \ldots, N$. This is tested against the following alternative hypotheses: H_1: $\beta_i < 0$ for $i = 1, \ldots, N_1$ and H_1: $\beta_i = 0$ for $i = N_1+1, N_1+2, \ldots, N$

Thus the alternative hypothesis allows β_i to differ across cross-sectional elements. The test statistic is constructed in two stages. First, the average of the individual Augmented Dickey–Fuller (ADF) t-statistics for each of the countries in the sample is calculated. Second, the standardized t-bar statistic is calculated as follows:

$$t\text{-bar} = \sqrt{N}\left(t_\gamma - \kappa_t\right)/\sqrt{\upsilon_t} \tag{A24}$$

where $t_\gamma = (1/N)\sum_{i=1}^{N} t_{\gamma_i}$, N is the number of cross-sectional units in the panel, and κ_t and υ_t are, respectively, estimates of the mean and variance of each t_{γ_i}. Im et al. (2003) tabulate exact critical values of κ_t and υ_t for various combinations of N and T. One setback of the t-bar test is that it is not applicable in the presence of cross-sectional dependence in the disturbances. However Im et al. (2003) suggest that, as a remedy to this setback, the data can be adjusted by demeaning. They demonstrate that the t-bar statistic converges to the standard normal distribution in the limit and has better finite sample properties than the earlier test statistics.

Maddala and Wu (1999) criticize the Im et al. (2003) test on the following grounds: First, the test works on the implicit assumption that the panel is balanced, so that the expected value and variance of the t-statistics are the same for all individual units in the panel. Second, the test implicitly restricts its practical application to the same lag length for all panel members. In practice, different lag lengths may be required for different cross-sectional units in the panel. Third, the asymptotic validity of the test depends on the number of units in the panel (N) approaching infinity. However, although this may not seriously hamper the test's suitability for finite sample applications, Maddala and Wu (1999) argue that a test whose validity does not depend on N approaching infinity may be a better alternative. They propose the Fisher test as a way of mitigating these weaknesses. Unlike the Im et al. (2003) test, which obtains the panel statistic by combining test statistics from the various cross-sectional units, the Fisher test combines the significance levels (i.e., p-values) from the same statistics. The Fisher test is non-parametric and has a chi-square distribution with $2N$ degrees of freedom. Using the additive property of the chi-squared variable, Maddala and Wu (1999) derive the following test statistic:

$$\lambda = -2\sum_{i=1}^{N} \ln \pi_i \tag{A25}$$

where π_i is the p-value of the test statistic for cross-sectional unit i. Through simulation studies, Maddala and Wu (1999) show that the test performs better than the Im et al. (2003) test.

Choosing a different approach, Hadri (2000) proposes a test for the null hypothesis of stationarity against the alternative of a unit root. Using a components representation in which an individual time series is written as the sum of a deterministic trend, a random walk and a white noise disturbance term, the null hypothesis of trend stationarity corresponds to the hypothesis that the variance of the random walk equals zero. Under the additional assumptions that the random walk is normal and that the stationary error is normal white noise, Hadri (2000) develops the following one-sided Lagrange Multiplier (LM) statistic to test for the level and trend stationarity hypotheses:

$$LM = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{\frac{1}{T^2}\sum_{t=1}^{T} S_{it}^2}{\hat{\sigma}_\varepsilon^2}\right), \qquad S_{it} = \sum_{j=1}^{t}\varepsilon_{ij} \tag{A26}$$

where $\hat{\sigma}_\varepsilon^2$ is the heteroskedasticity and autocorrelation (HAC) consistent estimate of the long-run variance of disturbance terms. The LM statistic is shown to be normally distributed. Carrion-i-Silvestre et al. (2005) demonstrate that Hadri's (2000) LM statistic can, with suitable modifications, be used to test for stationarity in a panel with structural breaks. The resulting test statistic is written as:

$$Z(\hat{\lambda}) = \frac{\sqrt{N}\left(LM(\hat{\lambda}) - \bar{\xi}\right)}{\bar{\varsigma}}; \quad \bar{\xi} = N^{-1}\sum_{i=1}^{N}\xi_i, \qquad \bar{\varsigma}^2 = N^{-1}\sum_{i=1}^{N}\varsigma_i^2, \tag{A27}$$

where $\xi_i = A\sum_{k=1}^{m_i+1}\left(\lambda_{ik} - \lambda_{i,k-1}\right)^2$; $\varsigma_i^2 = B\sum_{k=1}^{m_i+1}\left(\lambda_{ik} - \lambda_{i,k-1}\right)^4$; $i = 1, \ldots, N$; $m_i \geq 1$. The index $k = 1, \ldots, m_i + 1$ denotes structural breaks. One can allow for at most m_i breaks with the location of the breaks specified as a fixed fraction $\lambda_{ik} \in (0,1)$ of *T*. The constants A and B equal 1/6 and 1/45 respectively when there is no stochastic trend in the model, and 1/15 and 11/6300 respectively when stochastic trend is present. λ is used in equation (A27) to denote the dependence of the test on the dates of the break. The test statistic is shown to be normally distributed with zero mean and unit variance.

Panel cointegration tests can also be broken down into those that test for long-run equilibrium relationships between variables in the absence of structural breaks and those employed in the presence of structural breaks. Two types of cointegration tests have been proposed by Pedroni (1999) to test for long-run equilibrium relationships between variables in panels with no structural breaks. The first type is based on the within-dimension approach, and includes four statistics: panel *v*-statistic, panel rho-statistic, panel PP-statistic (nonparametric), panel ADF-statistic (parametric). These statistics pool the autoregressive coefficients across different members for the unit root tests on the estimated residuals (Lee, 2005). The second category of tests is based on the between-dimension approach and includes three statistics: group rho-statistic, group PP-statistic (nonparametric) and group ADF-statistic (parametric). The tests allow for heterogeneity among individual cross-sectional units in the panel, including heterogeneity in both the long-run cointegrating vectors and in the dynamics. Pedroni (1999) recommends that common time effects be removed by demeaning the data prior to the cointegration tests. The statistics are computed as follows.

Panel *v*-statistic: $Z_v = \left(\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{it-1}^{2}\right)^{-1}$

Panel rho-statistic: $Z_\rho = \left(\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{it-1}^{2}\right)^{-1}\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}\left(\hat{e}_{it-1}\Delta\hat{e}_{it} - \hat{\lambda}_i\right)$

Panel PP-statistic: $Z_t = \left(\hat{\sigma}^2\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{it-1}^{2}\right)^{-1/2}\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}\left(\hat{e}_{it-1}\Delta\hat{e}_{it} - \hat{\lambda}_i\right)$

Panel ADF-statistic: $Z_t^* = \left(\hat{s}^{*^2}\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{it-1}^{*^2}\right)^{-1/2}\sum_{i=1}^{N}\sum_{i=1}^{T}\hat{L}_{11i}^{-2}e_{it-1}^{*}\Delta e_{it}^{*}$

Group rho-statistic: $\tilde{Z}_\rho = \sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{e}_{it-1}^{2}\right)^{-1}\sum_{t=1}^{T}\left(\hat{e}_{it-1}\Delta\hat{e}_{it} - \hat{\lambda}_i\right)$

Group PP-statistic: $\tilde{Z}_t = \sum_{i=1}^{N}(\hat{\sigma}^2 \sum_{t=1}^{T} \hat{e}_{it-1}^2)^{-1/2} \sum_{t=1}^{T}(\hat{e}_{it-1}\Delta\hat{e}_{it} - \hat{\lambda}_i)$

Group ADF-statistic: $\tilde{Z}_t^* = \sum_{i=1}^{N}\left(\hat{\sigma}^2 \sum_{t=1}^{T} \hat{s}_i^2 \hat{e}_{it-1}^{*2}\right)^{-1/2} \sum_{t=1}^{T}(e_{it-1}^* e_{it}^*)$

where, $\hat{e}_{it}$ is the estimated residual and $\hat{L}_{11i}^2$ is the estimated long-run covariance matrix for $\Delta\hat{e}_{it}$. Similarly, $\hat{\sigma}^2$ and $\hat{s}_i^2$ $(\hat{s}_i^{*2})$ are, respectively, the long-run and contemporaneous variances for individual i. The other terms are defined in Pedroni (1999). All seven tests are distributed as being standard normal asymptotically. The panel v-statistic is a one-sided test where large positive values reject the null of no cointegration. The remaining statistics diverge to negative infinitely, which means that large negative values reject the null. Asymptotic and finite sample critical values developed from Monte Carlo simulations are tabulated in Pedroni (1999) and Pedroni (2004), respectively.

Westerlund (2006) proposes a test for the null hypothesis of cointegration that accommodates for structural change of a cointegrated panel regression. The test is able to accommodate for an unknown number of breaks in the deterministic component (i.e., the constant and trend) of the individual cross-sectional regressions, which may be located at different dates for different panel members. Westerlund (2006) assumes that the data generating process (DGP) is given by the following system of equations:

$$y_{it} = z_{it}'\gamma_{ij} + x_{it}'\beta_i + e_{it} \tag{A28}$$

where $e_{it} = r_{it} + u_{it}$, $r_{it} = r_{it-1} + \phi_i u_{it}$ and $x_{it} = x_{it-1} + v_{it}$ is a K-dimensional vector of regressors and z_{it} is a vector of deterministic components. The corresponding vectors of parameters are denoted β_i and γ_{ij}, respectively. The index $j = 1, \dots, M_i + 1$ is used to denote the structural breaks. There can be at most M_i such breaks, or $M_i + 1$ regimes, that are located at the dates $T_{i1}, \dots, T_{iM_i}$, where $T_{i0} = 1$ and $T_{iM_i+1} = T$. Furthermore, the initial value of r_{it} is assumed to be zero, which entails no loss of generality as long as z_{it} includes an individual specific intercept. In the structural break model, the locations of the breaks are specified as a fixed fraction $\lambda_{ij} \in (0,1)$ of T such that $T_{ij} = \lambda_{ij}T$ and $\lambda_{i,j-1} < \lambda_{ik}$ for $k = 1, \dots, M_i + 1$. Both M_i and λ_{ij} are assumed to be unknown.

From the DGP in equation (A28), if $\phi_i = 0$, then r_{it} vanishes under the assumption that $r_{i0} = 0$ in which case x_{it} and y_{it} are cointegrated as u_{it} is assumed to be a stationary process. Thus, the null hypothesis that all the individuals of the panel are cointegrated can be stated as H_0: $\phi_i = 0$ for all $i = 1, \dots, N$. This is tested against the following alternative hypothesis: H_1: $\phi_i \neq 0$ for $i = 1, \dots, N_1$ and H_1: $\phi_i = 0$ for $i = N_1+1, N_1+2, \dots, N$.

This formulation of the alternative hypothesis allows ϕ_i to differ across the cross-sectional units. To obtain the locations of structural breaks, Westerlund (2006) proposes the following method, which globally minimizes the sum of squared residuals from equation (A28):

$$\hat{T}_i = \arg\min_{T_i} \sum_{j=1}^{M_i+1} \sum_{t=T_{i,j-1}+1}^{T_{ij}} \left(y_{it} - z'_{it}\hat{\gamma}_{ij} + x'_{it}\hat{\beta}_i\right)^2 \tag{A29}$$

where $\hat{T}_i = \left(\hat{T}_{i1}, \dots, \hat{T}_{iM_i}\right)'$ is the vector of estimated break points, $\hat{\gamma}_{ij}$ and $\hat{\beta}_i$ are the estimates of the cointegration parameters based on the partition $T_i = \left(T_{i1}, \dots, T_{iM_1}\right)'$ and τ is a trimming parameter such that $\lambda_{ij} - \lambda_{ij-1} > \tau$, which imposes a minimum length for each subsample. After establishing the number of breaks, one proceeds to test the hypothesis of cointegration, using the following panel LM test statistic:

$$Z(M) = \sum_{i=1}^{N} \sum_{j=1}^{M_i+1} \sum_{t=T_{ij-1}+1}^{T_{ij}} \left(T_{ij} - T_{ij-1}\right)^{-2} \widehat{\omega}_{i1.2}^{-2} S_{it}^2 \tag{A30}$$

where $\widehat{\omega}_{i1.2}^2 = \widehat{\omega}_{i11}^2 - \widehat{\omega}'_{i21}\widehat{\Omega}_{i22}^{-1}\widehat{\omega}_{i21}$ and $S_{it} = \sum_{k=T_{ij-1}+1}^{t} \hat{e}_{ik}^*$ is any efficient estimate of e_{it}. The statistic is written as an explicit function of $M = (M_1, \dots, M_N)'$ to denote that it has been constructed for a certain number of breaks for each cross-section and that its asymptotic distribution depends on it.