Essays on Commodity Prices and Financial Market Variables

Evidence from Sub Saharan Africa

A thesis submitted

by

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Abstract

Volatility in international commodity prices is almost accepted as a stylised fact in modern financial markets. The drivers of commodity prices have evolved in addition to traditional global demand and supply factors. The literature suggests a number of other drivers, among them, activities of speculators – the so called "financialisation" of commodities postulate, the role of China and Fed policy. The question of whether exogenous shocks to commodity prices are transmitted through financial markets in Africa is investigated. In addition, the hypothesis of "wealth-transfer" from exporters to importers of commodities when prices fall is tested. Further, commodity prices are tested for their in sample and out of sample predictive ability. Finally recommendations are made for policy makers and financial market players in frontier markets in Africa.

The three essays are organised in Chapters 3 to 5 of the study. In Chapter 3, an ARDL bounds testing approach is adopted and we find that a South Africa-specific commodity index significantly predicts (in-sample) the exchange rate in the short-run. While the long-run relationship is weak and the associated error correction process is slow existence of cointegration of commodity prices and the exchange rate suggests that commodities explain a significant part of terms of trade fluctuations for South Africa. With respect to the structural exchange rate models of the South African Rand, using the Dynamic Ordinary Least squares (DOLS) estimator, we find that commodity prices are significant and consistent explanatory variables of the changes in the nominal exchange rate. The commodity price variable improves the in-sample fit of the structural exchange rate models presented in this chapter and this evidence is robust to the other major Rand cross rates. Further, inclusion of the commodity price variable improves the out-of-sample short horizon forecasting ability of canonical exchange rate models.

In Chapter 4 we employ dynamic econometric modelling techniques to confirm the existence of a strong financial channel through which copper price shocks are transmitted to the Zambian economy. In the short run, changes in the copper price lead changes in all financial market variables. Financial market variables and the price of copper share a long-run equilibrium relationship. Importantly, if this system is out of its long-run equilibrium, short-run corrections back to equilibrium are made by adjustments to the short term interest rate. Fittingly, the copper price-interest rate relationship appears strong in the long run. This result suggests that the policy makers "over-rely" on monetary policy to accommodate shocks from the international price of copper. The exchange rate and equity prices appear weakly exogenous to the system in-sample and out of sample.

In Chapter 5, we confirm existence of a structural break in the price of oil in July 2008. We also show that the financial market time series for Kenya and Nigeria also exhibit a significant structural break around this period. We therefore partitioned our sample to investigate the effect of structural shocks to the financial markets of the two markets. On the whole, we find that the nexus between financial markets and oil prices is much stronger and statistically significant for an oil exporter (Nigeria) and weaker and statistically insignificant for a net oil importer (Kenya) after the 2008 financial crisis. Prior to the 2008 oil price shock, the results are roughly the opposite of the post oil shock period for both countries. Our results highlight that it is important to account for major structural shifts in modelling the impact of oil prices in developing countries (Le and Chang, 2011). The "wealth transfer" argument from net importer to net-exporters exists between Kenya and Nigeria in the short run although it is not robust to sample specification. Finally, we highlight the inherent flaws and limitations of ex-post stabilisation funds in Africa and make a case for market based oil-price hedging instruments. We argue for the adoption of market based hedging instruments given the promising growth in financial markets of developing African countries in spite of several thorny implementation difficulties.

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Dedication

To my late grandmother who inducted me to my first business school in life; believed in me and my God-given talents and prayed always that I would rise and grow to fulfil every bit of God's purpose for my life.

Declaration

I, Xolani NDLOVU, declare that this research report is my own unaided work. It is submitted in partial fulfilment of the requirements of the Doctor of Philosophy in Finance degree at the University of the Witwatersrand, Johannesburg. It has not been submitted before for any degree or examination in this university or any other one.

Xolani NDLOVU **DATE**

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1. Introduction

1.1 Outline of research topic

The research is composed of three related essays analysing dependencies between commodity prices and financial market variables for four sub-Saharan African countries. In the first essay, Africa's best known "commodity currency" is considered. The essay re-visits the exchange rate determination puzzle that remains unsolved since the seminal work of Meese and Rogoff (1983) for the case of South Africa. Emphasis is placed on the role of commodity prices. In South Africa, "it is probably fair to say that the exchange rate of the rand, specifically against the US dollar, is the economic indicator that attracts the most interest among South Africans on an almost daily basis…"(Laubscher, 2016). While most attention is given simply to the domestic drivers of the exchange rate including socio-political developments, it is important to understand how exogenous shocks from international commodity markets are transmitted into the exchange rate. Lessons from the South African Rand which has been on a free float for nearly 20 years under an inflation targeting regime may serve as potential benchmarks for other small commodity exporting economies in the region that are subject to various economic and financial market reforms.

In the second essay we analyse the Zambian case, a Southern African country that heavily depends on export of one primary commodity for export revenues, namely copper. Unlike several studies that consider the impact of shocks to commodity prices to welfare and the macro economy, this essay specifically considers the impact of such shocks on financial market variables. Additionally, the essay departs from other studies that focus on the idiosyncratic relationships between commodity prices and macro variables. Financial market variables and commodity prices are considered in a multivariate setup that enables us to uncover dependencies that exists between financial market variables themselves.

The final essay considers shocks to the price of one of the most important commodities in the world – crude oil. There is a rich strand of literature that explores the impact of shocks to the oil price on the economies of either commodity exporters or importers in isolation. In this essay, we seek to find out if the financial markets of a net oil importing country and impacted differently from that of a net oil exporting country, following a shock to the price of oil. From a portfolio management point of view, there may be diversification benefits if the markets of two economies respond differently to a similar shock. Equally, at least in theory, if a common factor like the oil price induces opposite movements in similar asset classes, risk managers may find it possible to hedge risks across markets. To this end, the third essay addresses the oil price-financial markets nexus empirically, using data from Nigeria (SSA's largest crude oil exporter) and Kenya (a net importer of oil with fairly advanced financial markets).

1.2 **Research Motivation**

Headlines such as "World's Worst Currency Prompts Call for Divine Intervention" - Bloomberg October 16, 2015 or "Nigeria Recession 2016: Economy Slumps as Oil Woes Continue"- IBTimes, August 31, 2016 or "#RandReport: Rand weak, low commodity prices weigh stocks down"- Reuters September 16, 2016 present an invitation to researchers to confirm scientifically if the conjectured commodity prices-financial market variables nexus exists. If so, how financial market players in policy and commerce can better manage them. First, we outline our hunch about what could have happened in the past decade.

The idea that financial markets development accelerates economic growth has been around since the seminal work of Schumpeter (1912). There has been an abundance of attempts to verify this claim for both developed and developing markets. Research has taken several directions related to this topic (see Ngare et al., 2014) for a recent survey. Considerable attention has been devoted to financial reforms in banking and capital markets in Africa (Moyo, et al., 2014). The number of stock markets and economic activity has grown significantly in SSA. There has been support from multilateral financial institutions to accelerate financial markets reforms in some SSA countries.[1] South Africa abandoned a dual exchange rate for a free floating regime in 1995 and adopted an inflation targeting policy framework in 2000. Equally Zambia has been following an inflation targeting policy in recent years.

Further, in 2015 alone, there were 28 initial public offerings (IPOs) across the African markets and 105 between 2011 and 2015, raising over $40 billion. There were 47 Corporate and sovereign/supranational debt issuances in 2015 and 489 issuances between 2011 and 2015 raising a combined $110 billion. The combined market capitalisation of African stock exchanges was well over $1 trillion in 2015 with 23% of that outside South Africa. Until 2007, only South Africa, Tunisia and Egypt had sovereign debt issues in international markets. Since then, 15 countries have since issued Eurobonds denominated in US dollars (PwC, 2016).

Meanwhile, there have been several developments in international commodity markets, some structural and some permanent. The first decade of the 21st century has been characterised by very high commodity prices which led some scholars to suggest that the high prices were a new long-term plateau (Walker, 2012). However the collapse of commodity prices in 2014 has reminded researchers and financial players alike of the cyclicality of commodity prices. There is also the idea of "financialisation" of commodities, an idea that is

[1] A shining example is the establishment of the Lusaka Stock Exchange in Zambia with the preparatory and technical assistance from the International Finance Corporation (IFC) and the World Bank in 1993.

gaining traction with researchers such as Arezki et al. (2014) and Gao and Süss (2015).[2] Activities of speculators would in theory amplify short term volatility of commodity prices and exacerbate deviation of their prices form their intrinsic values.

Some scholars such as Klotz et al. (2014) specifically examine and emphasise the role of China's monetary policy and economic growth in the determination of commodity prices. Frankel (2014) emphasises the role of the Fed's monetary policy in the commodity price cycle. He argues that the loose monetary policy stance by the Fed encouraged a boom in commodity prices. It is clear from the literature that determination of commodity prices is almost completely exogenous to the small exporters in SSA.

Given the twin developments in African financial markets as well as international commodity markets, in this study we investigate how the relationship between these two markets has evolved over the past two decades in small African economies. The relationship - or lack of it - would be particularly interesting for policy makers and for players in financial markets. If the link between the two markets is found to be significant, the implication for policy makers would be to modify their monetary and exchange rate policies to protect their economies from exogenous shocks. Further, such evidence may validate the postulate of a "commodity based industrialisation" agenda for African countries (see Weeks and Mungule, 2013, Morris and Fessehaie, 2014). Equally, there could be lucrative opportunities for traders, hedgers and speculators.

1.3 **Research Objectives**

The study has four main objectives which are:

a) To establish if the "financial market" transmission channel exists for international commodity prices to affect the frontier economies in Sub-Sharan Africa.

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[2] Gao and Süss, 2015 report that financial investors represent roughly 80% of all the investors in commodities. Silvennoinen and Thorp, 2013 argue that commodity prices have become popular as an alternative asset class because of their negative (and sometimes positive) correlation with traditional asset classes such as equities.

b) To ascertain through dynamic econometric modelling methods if there is predictive information contained in international commodity prices with respect to exchange rates in Sub Saharan Africa.

c) To test the hypothesis of "wealth-transfer" between commodity exporters and importers with African financial markets, in the event of a commodity price shock

d) To offer recommendations for policy and financial risk management to policy makers and financial market players

1.4 Contribution to the Literature

The contribution of our study is three-fold. First we provide evidence of the existence of a financial channel through which commodity shocks are transmitted to the economies of small open commodity-dependent economies in Africa. Financial markets have been growing in Africa since the turn of the century underpinned by raft of financial market reforms and increased integration into the global financial markets.[3] Second, we focus our study on four individual countries without generalising their financial markets dynamics. This is important because no country in Africa is the same as another, even if they are from the same geographic region. Our approach enables us to flesh out country-specific recommendations for policy makers and commercial players. Finally, as opposed to several studies that consider idiosyncratic dependencies between commodity prices and a specific financial market variable, we analyse dependencies of financial market variables (such as exchange rates, equity prices and short term money market rates) and commodity prices in a multivariate set-up that allows us to uncover relationships among the financial market variables themselves.

[3] See Ngare et al. (2014), PwC (2016)

1.5 Main Findings

We summarise the main findings of our study as follows:

a) In Chapter 3, in an ARDL bounds testing framework, we find that a South Africa-specific commodity index significantly predicts (in-sample) the exchange rate in the short-run. While the long-run relationship is weak and the associated error correction process slow existence of cointegration of commodity prices and the exchange rate suggests that commodities explain a significant part of terms of trade fluctuations for South Africa. With respect to the structural exchange rate models of the South African Rand such as variants of the Purchasing Power Parity (PPP) and monetary models, we find that commodity prices are significant explanatory variables of the changes in the nominal exchange rate. The commodity price variable improves the in-sample fit of the exchange rate models and this evidence is robust to other major Rand cross rates. Further, inclusion of the commodity price variable improves the out-of-sample forecasting ability of canonical exchange rate models. Overall our results indicate that even with the inclusion of the commodity price variable, monetary models fit the data better than the purchasing power parity models – thus inclusion of commodity prices doesn't seem to resolve the purchasing power parity puzzle.

b) In Chapter 4, we confirm the existence of a strong financial channel through which commodity price shocks are transmitted to the Zambian economy. In the short run, changes in the copper price lead changes in all financial market variables such as short term interest rates, equity prices and the exchange rate. The financial market variables and the price of copper share a long-run equilibrium relationship. This means that deviations from the long run equilibrium in one period are corrected in the following period by adjustment of at least one of the variables in the system. Importantly, if this system is out of its long-run equilibrium, short-run corrections back to equilibrium are made by adjustments to the short term interest rate. Accordingly, the cooper price-short term interest rate relationship appears strong in the long run. This result suggests that the policy makers “over-rely” on monetary policy to accommodate shocks to the international price of copper. The exchange rate and equity prices appear weakly exogenous to the system in-sample and out of sample.

c) In Chapter 5, we confirm existence of a structural break in the price of oil in July 2008. We also show that the financial market time series for Kenya and Nigeria also exhibits a significant structural break around this period. We therefore partitioned our sample to investigate the effect of structural shocks to the financial markets of the two markets pre and post the 2008 crisis. On the whole, we find that the nexus between financial markets and oil prices is much stronger and statistically significant for an oil exporter (Nigeria) and weaker and statistically insignificant for a net oil importer (Kenya) after the 2008 financial crisis. Prior to the 2008 oil price shock, the results are roughly the opposite of the post oil shock period for both countries. Our results highlight that it is important to account for major structural shifts in modelling the impact of oil prices in developing countries (Le and Chang, 2011). The “wealth transfer” argument from net importer to net-exporters exists between Kenya and Nigeria in the short run although it is not robust to sample specification. Finally, we highlight the inherent flaws and limitations of stabilisation funds in Africa and make a case for market based oil-price hedging instruments. We argue for the adoption of market based hedging instruments given the promising growth in financial markets of developing African countries in spite of several thorny implementation difficulties.

1.6 **Outline of study**

The study has six chapters. After the introductory chapter, we present a review of relevant literature. The main findings of the study are presented in Chapters 3-5. The main conclusions are summarised in Chapter 6.

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IGC

Working Paper, November 2013

2. Literature Review

2.1 Introduction

Commodity prices and their impact on developing countries has been the subject of study of open macro-economics for a long time (see Erten and Campo, 2012 for a review). Demand and supply shocks that birth booms and busts in commodity prices traditionally ignite new forms of debate in international economics. There is consensus in literature on the importance of commodity prices for the world economy. Commodity price shocks affect the level of stability of export earnings by developing countries, the cost of inputs to production in industrialized countries, the allocation of world capital flows and rates of national economic growth (Muhanji and Ojah, 2011). Policy makers are concerned with the links between commodity prices and macro-financial goals (whether they be net exporters or importers of primary commodities). Financial market players, such as portfolio managers and traders are also concerned with the substitutability of commodities and financial market securities and therefore design of effective hedging strategies in times of uncertainty. There has been interest too in commodity price cycles and design of counter-cyclical strategies for countries that depend on the export of specialised primary commodities for their export earnings (see Venables, 2016 for a recent review).

Our study is closely linked with a strand of literature that connects commodity prices with macro-financial variables in developing commodity exporting countries. A considerable number of studies on commodity exporting countries analyse dependencies between commodity prices and individual macro financial variables such as exchange rates, equity prices and inflation. There are relatively few studies devoted to the study of developing countries' financial markets relationship with the prices of their major export commodities in comparison to OECD countries like Australia, New Zealand and Canada (Kablan et al, 2016).

Understandably, long and reliable time-series data is hard to find in most developing countries in Africa.

Now we provide a review of studies relevant to relationships between commodity prices and financial market variables.

2.2 Commodities and exchange rates

The link between the exchange rate and commodity prices is well researched in the literature, particularly for the Organisation for Economic Co-operation and Development (OECD) countries such as Australia, New Zealand and Canada. International commodity prices are predominantly denominated in US Dollars. Depreciation of the US Dollar therefore would be expected to be associated with a rise in commodity prices in US Dollars, ceteris paribus. To illustrate this point, Kim et al (2013) consider a simple Purchasing Power Parity (PPP) relation:

$$s = p - p^* \tag{2.1}$$

Where $s, p \ and \ p^*$ are the log of the nominal exchange rate (USD/foreign currency), domestic and foreign price levels respectively. Applying the PPP relation to the commodity markets, p can be thought of as the international price of commodities in USD and p^* as the domestic currency of the commodity exporting country. Equation (2.1) can be re-written as:

$$p = s + p^* \ \ \rightarrow \frac{\partial p}{\partial s} = 1 + \frac{\partial p^*}{\partial s} \tag{2.2}$$

From Equation (2.2), unless the exporting country invoices its commodities in its domestic currency that is $\left(\frac{\partial p^*}{\partial s} = -1\right)$ and thus $\left(\frac{\partial p}{\partial s} = 0\right)$, depreciation of the USD will cause inflation of commodity prices. Akram (2009) also provides evidence that a weaker US Dollar leads to higher commodity prices in a structural Vector Autoregressive model. When the US dollar depreciates, the price of gold should rise to preserve the intrinsic value of gold (Le and Yang 2016). For this reason, the gold price provides a natural hedge for an investor who is long in dollar denominated assets. This specific relationship between the gold price and the US Dollar exchange rate has been investigated recently in literature by a number of scholars who include Sjaastad (2008); Reboredo and Rivera-Castro (2014) and Beckmann et al. (2015).

There is quite a substantial amount of literature focusing on "commodity currencies". Commodity currencies are generally defined as those belonging to countries which rely on export of commodities for export earnings (Chen and Lee, 2014). Because export prices of their commodities represent a source for substantial terms of trade fluctuations, these countries' real exchange rates tend to move in sync with commodity prices (Clements and Fry, 2008).

The commodity currency hypothesis has been explored extensively in literature. Earlier attempts can be found in the work of Edwards (1986), which examined the relationship between real coffee prices and Colombia's real exchange rate. The author concluded then that shocks to coffee prices were partially accommodated by money creation and adjustment of the nominal exchange rate.

The work of Chen and Rogoff (2003) "formalised" the commodity currency hypothesis. Using data from three OECD countries, Australia, New Zealand and Canada, the authors found that US Dollar commodity prices were significant drivers of the real exchange rate for these developed countries. Following the Chen and Rogoff (2003) paper, there have been several related empirical studies on developed economies and Latin American developing countries. These studies include Simpson (2002), Hatzinikolaou and Polasek, (2003), Swift (2004) and more recently Issa et al. (2008), Chen et al. (2010) and Cayen et al. (2010).

Until the empirical work of Cashin et al. (2004), there had surprisingly been no major attention given to the study of the relationship between commodity prices and exchange rates in developing economies. Using a combination of cointegration and Granger causality tests, the authors uncovered evidence of a long-run equilibrium relationship between commodity prices and real exchange rates in one third of the 58 countries that they studied.

Several country-specific studies exist. Bova (2009) finds that the volatility of the copper price and the exchange rate in Zambia are related in an EGARCH framework. Further, in a multivariate structural model, Weeks and Mungule (2013) find that the main source of exchange rate instability are capital account movements which are moderated by foreign exchange holdings of the Bank of Zambia. The main source of foreign exchange reserves is export revenues from copper. Bhundia and Ricci (2004) investigated the South African Rand crises of 1998 and 2001 to found out causes and lessons from the crisis. Among other explanations of the crisis, they find that commodity prices played an important part in the Rand weakness in 1998. That year, global demand for commodities had weakened significantly on the back of the Asian financial crisis putting downward pressure on the market prices of some of South Africa's commodity exports and probably contributed to the large depreciation of the rand in July of that year. In their empirical analysis, they found that that a one percent fall in the real price of commodities exported by South Africa is associated in the long run with a real exchange depreciation of 0.5 percent. Other papers on South Africa include Schaling et al. (2014), MacDonald and Ricci (2002) and Frankel (2007b). Kohlscheen et al. (2016) use daily data and a country-specific commodity price index for 11 commodity exporting countries and conclude that the commodity prices are distinct drivers of exchange rates of commodity exporting countries including South Africa. Hegerty (2013, 2014) finds that shocks to commodity prices exert pressure on the real exchange rates of Western African countries and Latin America respectively.

Studies that connect oil prices and exchange rates (sometimes referred to as oil currencies), generally consider oil separately from other commodities. Prominent papers in this strand of literature include Kilian et al. (2009), and Bodenstein et al. (2011). Lizardo et al. (2010) adds an oil price variable to the monetary model of the exchange rates and find strong links between the oil price and USD exchange rates of major economies. Specifically, an increase in the price of oil significantly weakens the USD against currencies of oil exporters such as Canada, Mexico and Russia. Beckmann et al. (2014) rely on the copula approach to model dependencies between exchange rates and oil prices of a sample of 12 oil-exporting and oil-importing countries. They find that the relationship between exchange rates and oil prices has increased over time and that extreme events for both variables are likely to occur simultaneously. Several other references are provided in Reboredo et al. (2014).

2.3 Commodities and equity prices

The relationship between commodity prices and stock prices is a topic of interest to financial market players because many investment portfolios include commodities as an asset class and stocks (Ildirar and Iscan, 2015). Gold, for example, is considered as an alternative asset class to stocks in Buyuksalvarci, (2010). Unsurprisingly therefore, there is substantial amount of literature about this topic.

In a structural VAR model Chaban (2009) analyses the relationships among the nominal prices of natural resources, nominal returns on equity and nominal exchange rates of three OECD countries: Australia, Canada and New Zealand. The author finds that the portfolio-rebalancing channel of Hau and Rey (2006) is weaker in these countries because commodity prices' flexibility plays a special role in the transmission of shocks by linking equity markets across countries and reducing the need for portfolio rebalancing.[4] A positive supply shock such as increased productivity in the U.S. that affects U.S. equity returns positively is transmitted to commodity-exporting countries through commodity prices.[5]

Correlations between the stock market and commodity markets have been explored extensively by, among others, Lombardi and Ravazzolo (2016). The authors report that correlations between commodity markets and stock markets have increased significantly post 2008, a result that is robust to different correlation specifications. This result is at odds with the proposition of using commodities as a hedge for stock markets because of the actual or perceived negative correlation between the two variables (e.g. Gorton and Rouwenhorst 2006). Other scholars document the effect of increased participation by speculators. Büyüşahin and Robe (2012) report that correlation between the two markets has increased due to increased activity of speculators.

Le and Chang, (2015) use daily data to analyse dependences between stock prices, gold and oil prices in a bounds testing framework. Their study finds that the gold price denominated in Japanese yen has a negative impact on the stock prices in Japan in the long run. The influence of oil on the stock prices is found to be insignificant in the short run.

Rossi, (2012) investigates the link between commodity and equity prices (focusing on their evolution over time) for commodity exporting countries that include Chile, Australia and South Africa. Using a combination of Granger causality tests and forecasting regressions,

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[4] In the portfolio rebalance model, Hay and Rey (2006) provide theoretical micro foundations of the relationship between equity markets and exchange rates. Their model assumes that foreign exchange risk is unhedged; that domestic investors hold a portfolio of domestic and foreign equities and that domestic investors hold equity and foreign exchange risk as a bundle. A portfolio is allowed to deviate due to exogenous shocks such as shocks in equity markets. For example if foreign equities outperform domestic stocks, the share of foreign equities increases in the portfolio, exposing that portfolio to increased exchange rate risk. To rebalance the portfolio risk, domestic portfolio managers find it optimal to withdraw a portion of the foreign equity, in so doing appreciating the domestic currency via increased demand of local currency from inflows from sale of foreign equities. The model therefore implies that stronger equity markets are associated with weaker exchange rates due to the portfolio rebalancing motive.

[5] Increased demand for industrial raw materials in the US increases demand for commodities leading to inflation of commodity prices. Booming commodity prices in USD drive domestic equity returns of commodity exporting countries up and through increased capital inflows appreciate commodity currencies. It is potentially interesting to find out what relationships exist between developing stock exchanges in Africa and major export commodity prices.

the study documents that global commodity price indices are positively correlated with lagged equity values. Further, the study reports that the time series properties of commodity prices have however drastically changed since the 2000s, and correlation between commodity prices and equity markets has strengthened. With respect to forecasting ability, equity market values are found to have significant predictive ability for commodity prices relative to a random walk.

Some studies consider the peculiar relationship of the oil prices and stock markets. The influence of oil prices of stock market values is aptly captured in Jones et al. (2004): "Ideally, stock values reflect the market's best estimate of the future profitability of firms, so the effect of oil price shocks on the stock market is a meaningful and useful measure of their economic impact. Since asset prices are the present discounted value of the future net earnings of firms, both the current and the expected future impacts of an oil price shock should be absorbed fairly quickly into stock prices and returns, without having to wait for those impacts to actually occur." The impact of a shock to the oil price will depend on whether a country is a net importer or exporter of oil.

In the spirit of differentiating analysis between oil importing and exporting countries, Filis et al., (2011) employ a Dynamic Conditional Correlation GARCH in the spirit of Glosten, Jagannathan, and Runkle (1993) method to analyse time-varying correlations between stock market prices and oil prices for a set of oil exporting countries (Canada, Mexico and Brazil) and oil importers (USA, Germany and Netherlands). The study reports that time-varying contemporaneous correlation between oil price and stock markets exists for the two set of countries only for demand side and not supply side shocks to the oil price. The lagged correlation results however show that oil prices exercise a negative effect in all stock markets, regardless of the origin of the oil price shock. In times of financial turmoil therefore, the oil market does not offer a safe haven for investors.

Miller and Ratti (2011) confirm the existence of a long-run relationship between real oil prices and real stock prices for a sample of six OECD countries using a Vector Error Correction framework with additional regressors. The authors suggest that the relationship between the two variables could have changed at the turn of the century. Their results suggest a possible presence of bubbles in stock market prices and/oil prices in recent years.

In a recent study of the GCC economies, Arouri et al. (2011) uncover the existence of substantial return and volatility spill overs between world oil prices and GCC stock markets. Using daily data over the period from June 7, 2005 to February 21, 2010 for the six member countries of the GCC and a generalized VAR-GARCH approach, the study finds that direct transmission of conditional volatility is more apparent from the oil to the stock markets. Chang et al., (2011) also find evidence of volatility spill overs between oil and equity prices in a multivariate GARCH framework. Other contributors for the GCC countries include Narayan and Narayan, (2010), Hammoudeh and Choi (2006) and Mohanty et al., (2011). The latter employed weekly data and found a significant positive association between oil prices and stock market indices for all GCC countries. Choi and Hammoudeh (2010) suggest that commodity traders concurrently look at both stock and commodity markets fluctuations to infer the trend of each market.

Finally, closer to our area of interest, Fowowe, (2013) finds a negative but insignificant effect of oil prices on stock returns in Nigeria using a GARCH-jump models. The author suggests that low liquidity in the stock market, high transaction costs that discourage investment and the few oil-related stocks in the exchange as possible explanations for this counter-intuitive result. There are generally few studies that investigate this important relationship in emerging African stock markets, a gap that our study aims to fill. Other notable contributors to this strand of literature include Jones and Kaul, 1996; Sadorsky, 1999; and Basher and Sadorsky, 2006.

2.4 Commodities and interest rates

The relationship between commodity prices and interest rates can be traced back to the work of Hall (1982), a study that considered the role of commodity prices as an input into the US Federal Reserve monetary policy that is in the form of a monetary policy reaction function. Several related studies followed such as Garner (1985, 1989) who demonstrated that the Consumer Price Index (CPI) and commodity prices are not fully cointegrated. For this reason therefore he argued that monetary policy should not respond to commodity prices as they cannot control them. A counter argument was proposed by Boughton and Branson (1988) who argued that commodity prices frequently preceded turning points in CPI inflation. According to these authors (Boughton and Branson), commodity prices could be interpreted as leading indicators for inflation. Their argument was followed by Furlong (1989) who argues in line with Christiano, et al. (1996) that commodity prices are set in continuous auction markets with efficient information, and thus they can be early indicators of macroeconomic activity and can be used in monetary Vector Auto Regression (VAR) models. Cody and Mills (1991) suggest that the Fed eased monetary policy in the 1980s without taking into account information from commodity prices.

In the 21st century, the commodity prices-interest rate nexus is discussed by Frankel (2008). A report by Cobank (2012) that interviewed Frankel on the relationship of the Fed's quantitative easing policy and commodity prices, Frankel is described as having "studied this dynamic over the course of his academic career, and argues that interest rates should always be kept in mind when assessing the future price for many commodities." In his 2006 paper, Frankel picks up the topic of commodity prices which had fallen out of favour with researchers in the 1980s and 1990s, an era that coincided with low commodity prices. We review Frankel's "cost of carry" theory which is also aptly presented in Frankel (2014).

According to Frankel (2006), real interest rates will affect commodity prices via three channels. Higher short-term interest rates will incentivise commodity producers to increase production so that they can liquidate them and invest in higher yielding interest rates markets. In so doing, supply of commodities would increase leading to decreases in spot prices. Higher US interest rates in the 1980s coincided with a period of low commodity prices in the US. The second channel relates to the activities of speculators. Higher interest rates invite speculators looking for yield to reallocate portfolios from commodity contracts to interest bearing instruments. Finally, lower interest rates lower firm's cost of holding inventories and thus encouraging them to carry inventories contributing to an increase in demand. This "cost of carry" theory has roots in the work of Pindyck (2001) and was revisited by recently by Frankel (2014).

According to Frankel (2014), the cost of carry model can be thought of as a relationship between the short term real interest rate and the spot price of the commodity relative to its expected long-run equilibrium price. The model borrows from Dornbusch (1976) overshooting model for exchange rates. The model is based on two assumptions about expectations of market participants and their decisions whether to hold commodity inventories for another period or sell at current prices and invest in money markets.

Key inputs of the model are:

$s \equiv$ The natural logarithm of the commodity spot price

$p \equiv$ The log of the economy-wide CPI

$q \equiv s - p$ The real price of the commodity

$\bar{q} \equiv$ Log of the long run real price of the commodity

$i \equiv$ Nominal interest rate

cy ≡ Convenience yield from holding inventory[6]

sc ≡ Storage cost (e.g. rental rates for tanks, feed lot rates for cattle)

f ≡ log of forward/futures rate of the same maturity as the interest rate.

$rp \equiv (f - s) - E(\Delta s)$ ≡ risk premium

Market participants' expectations of the evolution of the future real commodity price are governed by whether the current price lies above or below the perceived long-run equilibrium represented as equations (2.3) or (2.4):

$$E[\Delta(s - p)] \equiv E[\Delta q] = -\theta(q - \bar{q}) \tag{2.3}$$

$$E(\Delta s) = -\theta(q - \bar{q}) + E(\Delta p) \tag{2.4}$$

The decision of whether to hold commodity inventories for another period or to sell them at today's period is governed by the expected rate of return. Such an expected rate of return must be equalised such that:

$$E(\Delta s) + c = i \text{ where } c \equiv cy - sc + rp \tag{2.5}$$

Combining equations (2.4) and (2.5), we get:

$$-\theta(q - \bar{q}) + E(\Delta p) + c = i \Rightarrow q - \bar{q} = -\left(\frac{1}{\theta}\right)(i - E(\Delta p) - c) \tag{2.6}$$

The relation implied by equation (2.6) says that the real price of a commodity relative to long-run equilibrium value is inversely proportional to real interest rates. Thus when interest rates rise, commodity prices will fall as investors shift their portfolios out of commodities and commodity prices will fall until such a point that the commodity prices are perceived to lie sufficiently below future equilibria. Equally, when interest rates fall, investors would shift out of money markets into commodities; commodities prices will thus rise until such a point that they are perceived to lie sufficiently above future equilibria. In both cases, the quasi-arbitrage or carry trade condition will be satisfied.

[6] For example, the psychic value of holding a precious commodity like gold or simply the insurance value of holding stock a particular essential input in the event of supply disruption.

A related representation of the relationship is presented by Bryne et al. (2013). Taking an asset pricing approach to commodity prices, they assume that storage costs are constant and that due to arbitrage, the expected return on commodity prices in the next period must equivalent to the real rate of return. Defining:

$E_t CP_{t+1}$ ≡ Expected value of commodity in the next period

r_t ≡ real rate of return for a risk free asset (e.g. US Treasury bill rate)

he risk-neutral rational valuation formula can be represented as:

$$[E_t CP_{t+1} - CP_t] = r_t \tag{2.7}$$

$$\Rightarrow CP_t = \frac{[ECP_{t+1}]}{(1 + r_t)} \tag{2.8}$$

From equation (2.8), it is clear that bond returns and commodities have a negative relationship.

Taking logs such that $\ln(CP_t) = cp_t$ and $R_t = \ln(1 + r) \approx r_t$, we get:

$$cp_t = E_t cp_t + \gamma R_t \tag{2.9}$$

Where $\gamma < 0$, a constant parameter.

Equation (2.9) suggests that commodity prices are a function of their future values and the real interest rate. The negative coefficient of the real interest rate is consistent with the carry trade model of Frankel (2014), Svensson (2008) and Calvo (2008).

If we forward iterate equation (2.9) assuming that the solution is convergent, the commodity prices are given by:

$$cp_t = \gamma E_t \sum_{k=0}^{\infty} R_{t+k} \tag{2.10}$$

Where $R_t = \delta R_{t+1}$, an AR(1) process; therefore $E_t R_{t+k} = \delta^k R_t$ and $\sum_{k=0}^{\infty} \delta^k = (1 - \delta)^{-1}$

From the foregoing relations, the value of commodity prices can be expressed as:

$$cp_t = \gamma \sum_{k=0}^{\infty} \delta^k R_t = \frac{\gamma}{1-\delta} R_t = \alpha R_t = f(R_t) \quad (2.11)$$

Where $f(.)$ is the relationship between log of commodity prices and real interest rates. Incorporating a time varying risk premium (ρ_t) element for risk-averse investors (Frankel, 2008 and Svensson, 2008), the asset pricing model can be written as:

$$[E_t CP_{t+1} - CP_t] = r_t - \rho_t \quad (2.12)$$

Such that:

$$cp_t = f(R_t, \rho_t) \quad (2.13)$$

Estimating the asset pricing equations, Bryne et al. (2013) confirm existence of a negative relationship between interest rates, uncertainty and a common factor in commodity prices using a Factor Augmented VAR (FAVAR) approach.

Another interest rate-commodity price relationship is postulated by Harvey et al. (2017). The authors propose a partial equilibrium model inspired by Borensztein and Reinhart (1994), Frankel (2006), Deaton and Laroque (2003) and Arango et al. (2012). The starting point is a log linear demand function given by:

$$d_t = \alpha y_t - \beta p_t + \gamma + \varepsilon_t^d \quad (2.14)$$

Where d_t is demand, y_t is the log of world income, and p_t is the world price of an internationally traded commodity.

The complementary supply function is given by:

$$s_t = \delta s_{t-1} + \eta r_{t-1} + \theta p_t + \varepsilon_t^s \quad (2.15)$$

Where s_t is the current supply function which depends on the previous period level, and r_t represents the interest rate. In equilibrium, supply must be equal to demand. It can be shown that:

$$p_t = (\beta + \delta)^{-1}[\alpha y_t + \gamma - \delta s_{t-1} + \eta r_{t-1} + \varepsilon_t^d - \varepsilon_t^s] \quad (2.16)$$

Harvey et al. (2017), use a very long dataset spanning 1650 to 2004 to examine relationships among commodity prices, economic growth and interest rates. Using a stationary VAR, the authors unearth evidence that commodity prices Granger cause income and interest rates. They also find that Granger causality runs from interest rates to commodity prices. Specifically they confirm Frankel (2006) argument that loose monetary policy has tended to support higher commodity prices.

In a recent study, Wang and Hu (2015) employ the econophysics technique of multifractal detrended cross-correlation analysis (MF-DXA) developed by Podobnik and Stanley (2008) to examine cross correlations between the US Effective Federal Funds rate and prices of four commodities. Their study confirms significant cross-correlation between commodity prices and the US Effective Federal Funds rate, a result that holds under qualitative analysis of the cross-correlation test, and the quantitative analysis of the MF-DXA. Employing a rolling window technique, they find that the cross-correlations between commodity prices and interest rates are time-varying.

Responding to paucity of literature relating to commodity prices and interest rates in periphery economies, Sujithan et al. (2013) attempt to answer the question whether monetary policy responds to commodity prices using a sample of developed and emerging markets. They use Markov-switching models and demonstrate that commodity markets have an clear impact on the short-term interest rate of all economies in their study – the US, Euro area, Brazil, India and Russia. Moreover, higher volatility in commodity prices is associated with lower short-term interest rates for all the countries under study except Russia. The study, through analysis of impact response functions, found that commodity volatility unit shocks to the interest rate are absorbed faster in emerging than developed markets.

Using quarterly data in a five-variable VAR model for oil prices, a food price index, metals, industrial raw materials and the US real interest rate, Akram (2009) finds that oil prices and industrial raw material prices display overshooting behaviour in response to interest rate shocks. Further, he finds that negative shocks to the world economy leads to lower interest rates and commodity prices. Evidence from this study is at ultra vires to the argument advanced by Frankel (2006, 2014) which suggests that lower US interest rates are associated with higher commodity prices.

Related to this body of literature are studies on inflation hedging using commodity prices (for example Bernanke and Gertler, 1999; Ciner et al., 2013; Batten et al., 2014; Hammoudeh et al., 2015) and the relationship between commodity prices and conduct of monetary policy (for example Hegerty, 2016; Hove et al., 2015; Mallick and Soussa, 2013; Ntantamis and Zhou, 2015).

2.5 Conclusion

Our study follows the work of Bhar and Hammoudeh (2011), Creti et al. (2012) and Bhardwaj & Dunsby (2013) who examine dynamic interrelations among individual industrial commodity prices and commodity-sensitive macro financial variables. Bhar and Hammoudeh (2011) employ four multivariate Markov Switching (MS) models to examine the relationships between copper, oil, gold and silver and three financial variables namely short term interest rates, exchange rates and the world equity index. In all the four MS-models the authors find that relationships between each of the commodity price series and financial market variables in their sample were regime dependent and that macro-financial variables transmit more information to commodity prices than the other way round.

There is clearly a dearth of literature analysing the important dependencies of commodity prices and financial market variables in African markets. African economies which depend on commodities are largely price takers from the international markets; their financial markets are generally too small to exert any measure of influence on commodity prices. As such they are exposed to first round effects of commodity price shocks which are increasingly transmitted via financial markets. Ours is an attempt to shed some light on the understanding of these important relationships which will find traction with decision makers in policy, exporters and financial market players.

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3. The South African Rand, Fundamentals and Commodity Prices

3.1 Introduction

The connection between exchange rates and macro-economic fundamentals is intuitively and theoretically plausible. The literature documents several fundamentals-based theoretical models of exchange rates that worked well in the seventies and early eighties. However, Meese and Rogoff (1983) seminal work demonstrated that several linear structural exchange rate models of the 1970s failed to forecast more accurately than the random walk, a conclusion that has as yet not been overturned by subsequent research attempts.

There is vast exchange rate literature focusing mainly on developed industrialised economies and emerging markets. The periphery countries such as those in sub Saharan Africa have been overlooked (Kablan et al., 2016). The major limitation of research in periphery countries is availability of long, reliable floating exchange rate data. Empirical exchange rate models are mainly concerned with the behaviour of floating exchange rates between countries which are open to trade and have liberalized capital markets, where the currency values are most likely to reflect various market forces.

South Africa is a major commodity exporting economy and the Rand is commonly classified as a "commodity currency" (Cashin et al. 2004). Commodities are transacted in high speed international markets therefore represent an observable and measureable source of terms of trade fluctuations for South Africa (Chen and Rogoff, 2003). Our first objective in this chapter is to investigate the relationship between the exchange rate and commodity prices. Our second objective is to test several popular structural exchange rate models - such as the purchasing power parity (PPP) and variants of the monetary models - using nineteen years of South African floating exchange rate data.

We extend this analysis to investigate if standard structural models adjusted for commodity prices perform better in South Africa, judged by conventional standard of goodness of fit criteria and short horizon forecasts.

As Sub Saharan Africa's largest and most industrialised economy, lessons from South Africa may serve as useful benchmarks for other smaller commodity exporting countries that are pursuing different market liberalisation and industrialisation agendas. For South Africa, understanding the drivers of the exchange rate may help manage exchange rate volatility exacerbated by the country's improved integration into global financial markets.[7]

Speculative flows into debt markets have been pointed as a major source of volatility (Hassan, 2014). Additionally, given the "financialisation" of commodity markets (Gao and Süss, 2015) and the end of the recent commodity price super cycle (Venables, 2016), it is clear that understanding how this asset class interacts with the exchange rate will be helpful to both hedgers, who invest in commodities futures markets for risk management and speculators, who invest in these markets for financial gain. For traders in the Rand, Moosa and Burns, (2012, 2014) argue that a forecast-based currency trading strategy outperforms a random walk model based strategy. Silvennoinen and Thorp, (2013) provide a recent review for the increasing appeal of commodities as an asset class.

This chapter is split into two sections. The first part takes a microscopic view into the nexus between the South African Rand and commodity prices. We employ the ARDL bounds testing approach as pioneered by Pesaran, Shin and Smith (1996), Pesaran and Shin (1998) and Pesaran, Shin & Smith (2001) to investigate if the exchange rate of the Rand and commodity prices are cointegrated. The second part considers structural exchange rate models using bilateral exchange rates of the Rand against the US dollar, British pound sterling and the Euro. The structural exchange rate models are augmented with commodity

[7] AsgiSA, the framework for shared growth to 2014 adopted by the government of South Africa, identifies six binding constraints to be addressed in order to achieve its goals on growth and distribution, and places 'the volatility and level of the currency' at the top of the list

prices and tested for their in-sample and out-of-sample short horizon forecast performance. Next, we provide a brief history of the Rand.

3.2 History of the South African Rand

The South African currency was born through a Commission established in 1956. The Commission recommended giving up the South Africa pound sterling when South Africa's membership of the British Commonwealth and the Sterling area was terminated. The name of the currency derives from the Witwatersrand ("White-waters-ridge"), the ridge where most of South Africa's gold deposits were found and where Johannesburg was built.[8]

We characterise the history of the Rand into three major episodes illustrated in Figure 3.1, namely, the multiple-exchange rate regime era, the managed float era and the new millennium. The period 1960-1995 was characterised by significant attention to stabilisation measures in the foreign exchange markets, mainly as coping mechanisms from political and economic pressures associated with the Apartheid regime.[9] During this period, exchange rate stability was itself an objective of policy, explained in part by the fact that South Africa was a signatory to the Bretton Woods agreement to manage fixed exchange rates.[10] The multiple regime changes are displayed in Table 3.1.

For the first decade, the Rand was pegged to the British pound sterling until the collape of the Bretton Woods system in 1971. It was then pegged to the United States dollar in August 1971.The dollar peg did not last long as the rand deprecated 12.28% against the dollar following currency realignments brought about by the Smithsonian Agreement.

The Rand was re-pegged to sterling briefly between December 1971 and September 1972 after which the dollar peg was re-established and maintained until May 1974. Between June 1974 and May 1975, the South African Reserve Bank introduced a policy of

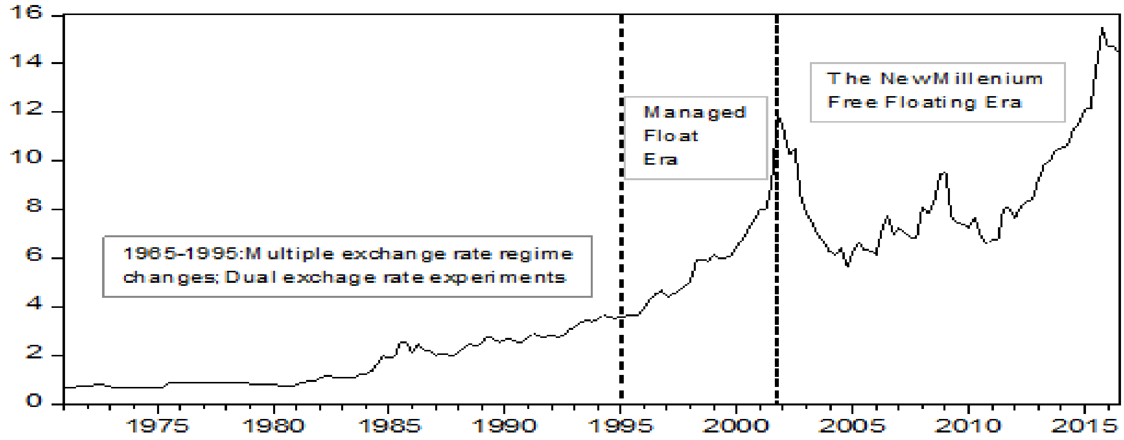
[8] https://www.oanda.com/currency/iso-currency-codes/ZAR

[9] The Apartheid a system of racial segregation in South Africa enforced through legislation by the National Party (NP), the governing party from 1948 to 1994. Under apartheid, the rights, associations, and movements of the majority black inhabitants and other ethnic groups were curtailed, and white minority rule was maintained.

[10] See Wakeford, 2002, Van de Merwe, 1996, De Kock Commission, 1985)

independent managed floating where the value of the rand was pegged to a basket of currencies and the fixed Rand-dollar rate adjusted every few weeks.

Figure 3.1: History of the South African Rand



Source: Author

This policy was abandoned in June 1975 in favour of fixed dollar peg that was maintained until 1979. Jones and Muller (1992) argue that the South African government sought to maintain an overvalued exchange rate in order to import cheap capital goods to support a rapid industrialisation program

A combination of domestic political turmoil and international sanctions imposed on the apartheid regime precipitated massive outflows of non-resident capital. As noted by the De Kock Commission, 1985, adverse global economic developments, notably the devaluation of the USD in 1978 made maintenance of fixed exchange rates difficult.[11] In response, the South African authorities introduced the dual exchange rate system in 1979. Under this reform, there were two exchange rates namely the commercial rand (principal rate applied to all transactions by residents) and the financial rand (a secondary rate applied only to capital transactions by non-residents). The De Kock Commission recommended the financial rand in 1978 as an interim measure and a further step towards the long-term objective of a unified

[11] The De Kock Commission was a Commission of Inquiry into the Monetary System and Monetary Policy in South Africa.

exchange rate. Under the dual exchange rate regime, the financial rand was left to float freely with occasional interventions by the authorities while the commercial rand was pegged to the US-dollar, subject to intervention by the authorities.

Guided by the De Kock Commission, the financial rand was discontinued in February 1983 and the exchange rates were unified to a managed float. In Garner (1994), several economic improvements were cited by the Finance Minister for this decision. The key improvements in the economy included a reduction in the current account deficit in 1982, forecast of a current account surplus in 1984, substantial improvements in capital account inflows and loan financing and a 3.6 billion rand in foreign reserves in 1982. In line with this policy, the authorities introduced a forward exchange rate market although it remained heavily regulated by the Reserve Bank (see Van der Merwe, 1996 and Aron et al, 1997, 2000). Following the abolition of the dual exchange rate however, the rand depreciated sharply against the US dollar. The depreciation was in part due to weakening of the gold price and increased divestment from South Africa.

Table 3.1: Exchange rate regimes in South Africa

Dates	Exchange Rate Regime
Feb 1961-July 1971	Fixed Exchange rate regime: rand pegged to the British pound
August 1971-November 1971	Fixed Exchange rate regime: rand pegged to the US Dollar
December 1971-September 1972	Fixed Exchange rate regime: rand pegged to the British pound
October 1972-May 1974	Fixed Exchange rate regime: rand pegged to the US Dollar
June 1974-May 1975	Crawling peg regime: rand pegged to basket of currencies
June 1975-May 1979	Fixed Exchange rate regime: rand pegged to the US Dollar
June 1979-Jan 1983	Dual exchange rate regime: crawling peg for commercial Rand; free floating financial rand
Feb 1983-Aug 1985	Unified exchange rate: managed float.
September 1985-Feb 1995	Dual exchange rate regime: managed float for commercial rand; free float for financial rand
March 1995-Jan 2000	Unified exchange rate: Managed float rand
Feb 2000-present	Unified exchange rate: free floating Rand with inflation targeting policy

Source: Author

At the end of August 1985, a number of international banks refused to roll-over short term loans to South African borrowers as a result of imposition of UN economic sanctions. South Africa subsequently defaulted on its foreign debt obligations by declaring a moratorium (also known as debt standstill) on more than half of its international obligations (Ayogu & Dezhbakhsh, 2008). As a response to the crisis, the dual exchange rate regime was reinstated, essentially in the same form as one that existed between 1979 and 1983. The only difference was that the commercial and financial rates were allowed to float although the commercial rand was subject to more intervention by the authorities. The depreciation of the rand in 1985 due to economic sanctions was arrested briefly between 1986 and 1988 due to the rising gold price. The subsequent fall in the gold price in 1988 precipitated a crisis in the South African mining industry in May 1988 leading to a sharp depreciation into the 1990s. The political crisis escalated in the early 90s, until 1993 when the government agreed to share power with the African National Congress (ANC) for five years after the first all-race election. After the political reconciliation of 1994 and subsequent removal of economic sanctions, the commercial and financial rand exchange rates were unified on 12 March 1995 and the dual exchange rate regime was discontinued.

For a period of eleven months after the abolition of the dual exchange rate system the unified rand was stable at around R 3.60 to the US dollar. This period was followed by a sharp sell-off with the rand falling 20% by June 1996. The contagion arising from the Asian crisis of 1997 also hit the Rand which plummeted by over 20% in real terms in 1998 although it regained its composure through 1999 trading in a broad range between R5.50 and R6.40 to the dollar. Official interventions from the Reserve Bank continued during this period to stabilise the exchange rate and also as a part of broader reforms to integrate the country into the global economy. Mboweni, (2004) notes that reserve Bank grew is open position on its forward book in the foreign exchange market. This intervention however resulted in a large open position against the Reserve Bank with its attendant negative effects on foreign investments and international markets' assessment of the economy, (see Ayogu & Dezhbakhsh, 2008).

The current free-floating regime of the Rand began in 2000 when the Reserve Bank adopted inflation targeting as a framework for monetary policy. Adoption of inflation targeting means that the Reserve Bank overtly focuses on a target inflation benchmark as the variable with interest rate as a policy instrument. For efficacy and credibility of the Central bank, inflation targeting precludes pre-commitment to an exchange rate target (Masson, et al 1998). Accordingly, the Central Bank ceased its forward book in the forex market in February 2004 (IMF, 2004; Mboweni, 2004). Thus while the Reserve Bank occasionally participates in the foreign exchange markets in South Africa, mainly for reserve accumulation, the current exchange rate regime is closest to a free float. This regime and increased integration to the global financial markets implies that the pricing of the currency can be recognised with economic forces. The rand has continued to weaken in the new regime and has become one of the most liquid emerging market economies – trading as a proxy for emerging market currencies in terms of expressing both negative and positive sentiments (Kissi, 2013).

The new millennium was associated with several events of note with respect to the Rand although the general weak trend has continued through 2014. On the positive, the country's integration to the global markets increased significantly, indicated by extensive trade in USD/ZAR. According to Kissi (2013), the Rand is the only BRICS currency that utilises the CLS settlement system, making it more appealing to international investors and promoting its basis as an international currency. The Bank of International Settlement (BIS) (2013) survey notes that trade in USD/ZAR accounted for almost 1.1% of global trade of which 55% of daily turnover occurred in London between non-South Africans. This statistic indicates the potential difficulties that the Reserve Bank may face in attempts to significantly influence the exchange rate consistently.

On the negative, the currency has continued to depreciate against the dollar and other majors for a number of well documented reasons. In 2001, the currency lost 40% against the US dollar over a relatively short period, promoting the government to appoint the Myburgh Commission of Inquiry (2002) which investigated the factors behind the rapid depreciation. [12] In December 2001, the currency reached a record low at R13.84/US dollar although the currency appreciated 75% between this period and September 2004 (Hodge, 2005). The 2008 global financial crisis also affected the Rand, having it fall to multi year lows of R11.86 in October 2008.

This period was followed by several months of recovery until the second quarter of 2011. From this period the Rand has weakened sharply from levels of around R6.75 to close 2014 at R11.6. A host factors have been cited as responsible for the collapse of the exchange rate. Concern over South Africa's worsening current account is cited by van der Merwe (2012) alongside labour market disturbances across the mining sector that resulted in the so called "Marikana Massacre." Other factors cited for the weakness include a rating downgrade, weakening growth, large external deficits in exposure to the Chinese economic growth slowdown. According to a Morgan Stanley research note in 2013, the Rand was classified with the Brazilian Real, Indonesian Rupiah, Indian rupee and Turkish lira as the "Fragile Five" or troubled emerging market currencies. Speculation around unwinding of the stimulus package by the US Federal Reserve Bank, the end of the commodity prices super cycle, Chinese growth concerns and European sovereign debt crisis combined to make the Rand the worst performing currency of the 16 major currencies tracked by Bloomberg (Bonorchis, 2013).

[12] The commission noted factors such as strength of the US dollar against other majors; higher relative inflation in South Africa, contagion effects from the Argentina crisis and economic problems in neighbouring Zimbabwe as factors behind the rapid collapse.

At the time of writing, the rand has set an all-time low at R17.83/US dollar on 10 January 2016 largely blamed on the President Jacob Zuma's surprise firing of Finance Minister Nhlanhla Nene and the subsequent appointment of little known David van Rooyen as his replacement.

3.3 Exchange rates and commodity prices

The correlation of exchange rates and economies' major source of export revenues has its roots in economic theory (See Chen and Rogoff, 2003, MacDonald and Ricci, 2002 Blundell-Wignall et al., 1993, Broda ,2004; Cashin et al., 2004 and more recently Chen et al. 2010; Issa et al., 2008; Cayen et al., 2010 and Ferraro et al, 2015). There are at least two theoretical links between commodity prices and exchange rates.

The first link comes from macro-economic and trade theory arguments discussed in Clements and Fry (2006) and Chen and Rogoff (2003). To illustrate this link between the exchange rate and commodity prices, consider a small open commodity exporting economy with a tradable and non-tradable goods sector. An increase in the price of the exported commodity in world markets would affect the demand for non-traded goods through its effect on wages – a channel similar to the well documented Balassa-Samuelson (1964) effect. Assuming that prices of non-tradable goods are sticky (conceivable in the case of South Africa), the exchange rate instead of prices would have to adjust to preserve efficient resource allocation.[13] Thus, a positive terms of trade shock such as a boom in commodity markets leads to an appreciation of the exchange rate in an environment of nominal price rigidities in the non-tradable sector.

The second channel works through the "portfolio balance" class of models (see Frankel, 1993). In this class of models, it is assumed that asset holders allocate portfolios in shares that are well defined functions of their expected rates of return. The portfolio balance

[13] That is, the price of the domestic currency against foreign currency, instead of prices of goods and services would have to adjust upwards to maintain efficient relative price between traded and non-traded goods.

approach treats domestic and foreign assets as imperfect substitutes; thus the exchange rate is a function of demand and supply of foreign and domestic assets and not just money. Under the assumption of perfect capital mobility, the exchange rate must adjust instantly to equilibrate the international demand for national assets. For a commodity exporting economy, a boom in the price of the exported commodities in international markets would typically lead to an excess supply of dollars and accumulation of foreign reserves, increasing pressure on the relative demand of their domestic currencies. To equilibrate the demand for the domestic currency, the price of the domestic currency would have to appreciate in terms of the foreign currency.[14]

To summarise, depending on whether a country is a net commodity exporter or (importer), commodity price rises may be expansionary (or serve as tax) and therefore cause and appreciation (or depreciation) of the domestic nominal exchange rate.

3.4 The Rand and commodity prices: the ARDL bounds testing approach

3.4.1 Empirical model (ARDL)

We employ the Autoregressive Distributed Lag (ARDL) procedure to test the short-run and long-run relationships between the Rand and commodity prices. The ARDL model is chosen in order to address typical problems associated with time series modelling, namely a short sample period, the order of integration problem, serial correlation and endogeneity of regressors.

The ARDL approach was developed and popularised by Pesaran, Shin and Smith (1996), Pesaran and Shin (1998) and Pesaran, Shin & Smith (2001). The ARDL approach has several advantages and strengths over earlier cointegration methods like the Engle Granger (1987) and Johansen (1995) methods. First, the method is applicable whether the variables

[14] See also Edwards(1994); Isard,(2007); Ricci et al.(2008), Coudert et al (2015), Chaban (2009)

involved are purely stationary $I(0)$ or difference stationary $I(1)$, a mixed or mutually cointegrated. Thus, the approach overcomes the requirement for variables to be integrated of the same order as in the Johansen (1995) approach. OLS estimates of the long run coefficients of the ARDL models are consistent provided that the lag structure of the model is identified. Valid inferences on long-run parameters can thus be made using standard asymptotic theory (Pesaran and Shin, 2008). Secondly, the method is suited for small samples because the approach eliminates the need for pre-testing for unit roots. Pesaran, Smith and Akiyama, (1998) show that commonly used standard unit root tests can be problematic in distinguishing the null of a unit root from the stationary alternatives in short sample periods. Thirdly, the ARDL method addresses the problem of serial correlation and endogeneity of regressors by appropriate augmentation of the order of regressors. Harris and Sollis (2003) show that long-run relationships estimated using the ARDL approach will be unbiased even when some of the regressors are endogenous in the model.

Implementation of the ARDL procedure involves two steps. The first step involves testing for the existence of a long-run relationship between the relevant variables in the context of an error correction framework. The test is achieved by computing the F-statistic with respect to the significance of the lagged levels of the variables in the error correction form of the underlying ARDL. The asymptotic distribution for this F-statistic is however non-standard. Pesaran et al (1996, 2001) provide a different test statistic to be used specifically for an ARDL model. They compute two sets of asymptotic critical values for the two polar cases: an upper bound assuming that all the regressors are $I(1)$ and a lower bound assuming that all the regressors are $I(0)$. The new critical values are tabulated from an extensive set of stochastic simulations under differing assumptions for the number of regressors, inclusion of an intercept and deterministic trends. If the calculated F-statistic lies outside the critical value bounds, then the researcher can make conclusions without any need to know whether the underlying variables are $I(0)$ or $I(1)$ or fractionally integrated. Specifically, one can conclude that a long-run relationship exists among variables if the computed F-statistic lies above the upper bound and that no long-run relationship exists if the F-statics lies below the lower bound. If the computed F-statistic lies inside the critical value bounds, inference is inconclusive and knowledge of the order of integration of the underlying variables becomes necessary. In that case, Banmani-Okkooee and Nasir (2004) show that an efficient way of establishing a long-run relationship is applying the ECM version of the ARDL model - if the error correction term is negative and significant, the researcher can conclude that cointegration exists (see also Bahmani-Oskooee, 2001 and Pahlavani et al, 2005).

The second step of the ARDL approach is to estimate the coefficients of the long-run relations and make inferences about their value. This step is valid only if the F-tests in the first stage do not reject the existence of a long-run relationship between the variables – thus the regressors can be considered as the "long-run forcing" variables explaining the dependent variable.

The ARDL model can therefore be thought of as a re-parametrization of the Vector Auto regression (VAR) model.

Starting with a Var (p) model:

$$z_t = \alpha + \sum_{i=1}^{p} \phi_i z_{t-i} + \varepsilon_t \tag{3.1}$$

Where z represents a vector of variables

Assuming that the elements of z are at most $I(1)$, that is, they don't have explosive roots, equation (1) can be written as a simple vector error correction model:

$$\Delta z_t = \alpha + \Pi z_{t-1} + \sum_{i=1}^{p-1} \Gamma \Delta z_{t-i} + \varepsilon_t \tag{3.2}$$

Where $\Delta \equiv 1 - L$ is the difference operator; $\Pi = -(I_{k+1} - \sum_{i=1}^{p} \Phi_i)$ and $\Gamma = -\sum_{j=i+1}^{p} \phi_j$, $i = 1, \dots, p-1$ are the $(k+1)x(k+1)$ matrices of the long-run multipliers and short run dynamic coefficients.

Assuming that there is only one long-run relationship among the variables, Pesaran *et al* (2001) partition z_t into a dependent variable y_t and a set of "forcing" variables x_t. The matrices $\alpha, \beta, \Gamma$and the long run multiplier Πcan also be partitioned as follows:

$$\alpha = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix}, \quad \Gamma_i = \begin{bmatrix} \gamma_{11,i} & \gamma_{12,i} \\ \gamma_{21,i} & \gamma_{22,i} \end{bmatrix}, \Pi = \begin{bmatrix} \pi_{11} & \pi_{12} \\ \pi_{21} & \Pi_{22} \end{bmatrix}$$

The key assumption that x_t is a long-run forcing variable of y_t implies that the vector $\pi_{21} = 0$, that is, there is no feedback from the level of y_t on Δx_t. Thus, the conditional model for Δy_t and Δx_t can be written as:

$$\Delta y_t = \alpha_1 + \pi_{11} y_{t-1} + \pi_{12} x_{t-1} + \sum_{i=1}^{p-1} \gamma_{11,i} \Delta y_{t-i} + \sum_{i=1}^{p-1} \gamma_{12,i} \Delta x_{t-i} + \varepsilon_{1t} \qquad (3.3)$$

$$\Delta x_t = \alpha_2 + \Pi_{22} x_{t-1} \, \pi_{12} x_{t-1} + \sum_{i=1}^{p-1} \gamma_{21,i} \Delta y_{t-i} + \sum_{i=1}^{p-1} \gamma_{22,i} \Delta x_{t-i} + \varepsilon_{2t} \qquad (3.4)$$

Under the standard assumptions about the error terms, (3.3) can be written as a long-run and short run error correction model respectively as follows:

$$y_t = \delta_1 + \phi y_{t-1} + \theta x_{t-1} + \sum_{i=1}^{p-1} \upsilon_i y_{t-i} + \sum_{i=1}^{p-1} \varphi_i x_{t-i} + \omega_t \qquad (3.5)$$

$$\Delta y_t = \delta_1 + \phi y_{t-1} + \theta x_{t-1} + \sum_{i=1}^{p-1} \upsilon_i \Delta y_{t-i} + \sum_{i=1}^{p-1} \varphi_i \Delta x_{t-i} + \gamma ECT_{t-1} + \omega_t \qquad (3.6)$$

Equation (3.6) is the “unrestricted” error correction model (ECM). For a long-run relationship to exist among level variables, the parameters ϕ and θ should be non-zero. The null hypothesis of no cointegration is therefore the joint hypothesis that $\phi = \theta = 0$ via a bound test.

The choice of an ARDL model is selected using the Akaike Information Criterion (AIC) or Schwartz Bayesian Criterion (SBC). The chosen model is then estimated by Ordinary Least Squares (OLS). We check the chosen models for serial auto-correlation to confirm that the resulting regression results are not spurious. The bounds test is then conducted to ascertain the existence of long-run relationships between the two variables. Inferences are then made from the resulting long and short run coefficients.

As observed by Ouattare (2004) it is important to note that the variables included in the ARDL model should at most be $I(1)$. If the variables are $I(2)$, then the critical values provided in Pesaran *et al* (1997, 2001) which are computed based on I(0) and I(1) variables are no longer valid. For this reason, we conduct unit root tests.

3.4.2 Stability tests

To check the stability of the estimated coefficients, we apply the cumulative sum of recursive residuals (CUSUM) and cumulative square of recursive residuals (COSUMSQ) tests proposed by Brown, Durbin, and Evans (1975). The latter test is based on “recursive residuals”- defined as uncorrelated sums with zero mean and a constant variance. The stability tests are used when there is a possibility of structural breaks. The null hypothesis is that the coefficient vector is the same in every period against the alternative that it is not. The CUSUM test plots the cumulative sums together with the 5% critical lines.

The CUSUM test statistic is computed as:

$$W_t = \sum_{r=k+1}^{t} \frac{w_r}{s} \qquad T = K+1, \dots, T \tag{3.6}$$

Where w is the recursive residual and s is the standard error of the regression fitted to all T sample points. If the b (coefficients vector) remains constant from period to period, then $E[W_t] = 0$, but if β changes , W_t will tend to diverge from the zero mean values line. The quantum of departure from the zero mean line is examined with reference to the 5% significance lines. The distance between the significance lines increases with t.

The 5% significance lines can be found by connecting the points given by:

$$\left[k \pm 0.98(T-K)^{\frac{1}{2}}\right] \; and \; \left[T, \pm 3x0.948(T-K)^{\frac{1}{2}}\right] \tag{3.7}$$

Parameter instability occurs if the W_t moves outside the area between the two critical lines.

The CUSUM sum of squares is based on a test statistic computed as:

$$S_t = \sum_{r=k+1}^{t} \left(\frac{W_r^2}{\sum_{r=k+1}^{T} W_r^2}\right) \tag{3.8}$$

The expected value of S under the hypothesis of parameter constancy is given by: $E|S_t] = (t-k)/(t-k)$ which goes from zero at $t = k$ to unity at $t = T$. The significance of departure of S form its expected value is examined by reference to two parallel lines around the mean. Movement outside the parallel lines suggest parameter or variance instability.

3.4.3 Model Specification

We estimate a bivariate model of the United States nominal Dollar/South African Rand (USDZAR) exchange rate featuring the United States Dollar nominal non-fuel commodity index between 1996 and 2014. We choose to focus our study form 1996 to 2014 to examine the behaviour of the floating exchange rate. In the spirit of Gregory and Hansen (1996a), Chen and Rogoff (2003), Cashin *et al* (2004) and more recently Bodart *et al* (2012), we specify the model as:

$$s_t = \alpha + \beta com_t + \varepsilon_t \tag{3.9}$$

Where s_t is the log of the nominal USD/ZAR exchange rate and com_t the log of the nominal non-fuel commodity price index and ε_t is i.i.d error term. The objective of the model is to ascertain the connection between the nominal exchange rate and the commodity price fundamental. Recalling the notoriously poor performance of fundamentals-based models as detailed by Frankel and Rose (1995), we seek to isolate the interactions between the exchange rate (and its lags) and commodity prices in a model that allows the variables to affect each other over time while at the same time addressing the endogeneity problem of the data generation process. Accordingly, we specify the unrestricted ECM version of (3.9) as:

$$\Delta s_t = \alpha + \sum_{t=1}^{p} \beta_{0,i} \Delta s_{t-i} + \sum_{i=0}^{p} \beta_{1,i} \, \Delta com_{t-i} + \gamma_0 s_{t-1} + \gamma_1 com_{t-1} + \mu_t \tag{3.10}$$

Employing the Pesaran *et al* (2001) bounds test for no cointegration, the null for cointegration to be tested is $\gamma_0 = \gamma_1 = 0$ against the alternative of $\gamma_0 \neq \gamma_1 \neq 0$.

3.4.4 Results of the ARDL model

3.4.4.1 Data and preliminary analysis

We use the nominal monthly United States dollar/South African Rand bilateral nominal exchange rate (s_t) obtained from the International Finance Corporation (IFS) database for the period 1996-2014.[15] Here the exchange rate is quoted as South African Rand per United States dollar such that an increase in the exchange rate implies depreciation of the Rand and vice versa. For the commodity prices (com_t), we construct a South Africa-specific commodity price index based on four major export commodities, namely gold, platinum, iron ore and coal following Cashin etal, (2002). There is significant co-movement in the prices of the four commodities in-line with the findings of Pindyck and Rotemberg, (1990). The price data are obtained from the IMF database while the share of major commodity exports was extracted from the UN Comtrade database (see data appendix A for the specific computation of the commodity index). All data are transformed into natural logarithms for ease of interpretation and are illustrated in Figure 3.2.[16]

The data series exhibit several structural breaks over the sample period. In this study we employ the methods developed by Bai (1997) and Bai and Perron (1998, 2003a) to ascertain multiple unknown break points. The advantage of this approach is that it allows us to endogenously estimate break points without a priori knowledge of the break dates. Specifically, we employ the "Global Bai-Perron L Breaks vs. None" method and allow for 5 break dates and 15% trimming percentage. We report the results in Table 3.2.

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[15] www.ifs.org.

[16] We estimated our model using real variables and the results do not change much. Results are available on request.

Figure 3.2: Data Plots

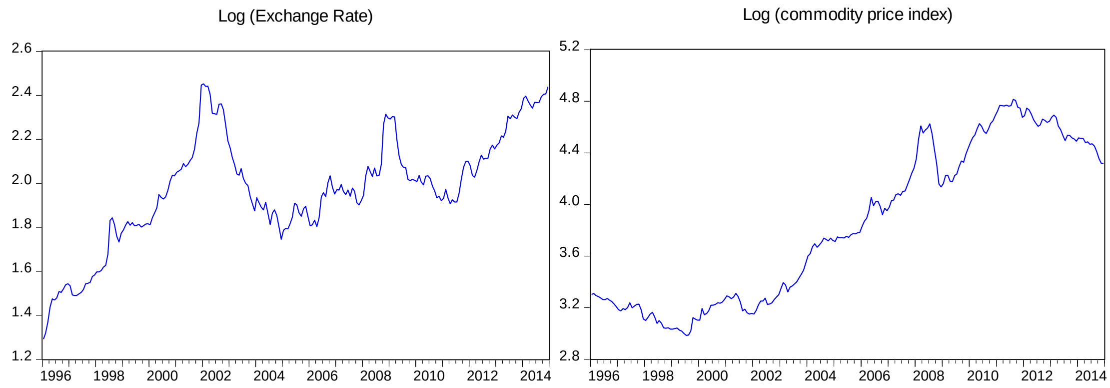


Table 3.2: Structural break tests

Break test options: Trimming 0.15, Max. Breaks 5

	Exchange Rate	**Commodity Price**	**Model**
Schwarz criterion selected breaks:	4	4	4
LWZ criterion selected breaks:	4	4	4
Estimated break dates:	2000M5; 2003M07; 2008M02 2012M03	1998M1; 2003M07; 2006M05 2012M04	2000M11; 2004M02; 2008M08; 2012M03

The Schwarz and LWZ criteria indicate at most 4 structural breaks in both the Rand and commodity price variables. Tests on the bivariate model also indicate 4 structural breaks. We tested the model with date dummy variables following Pahlavani et al (2005) and tested the dummy variables for significance.[17] The results indicate that only two break dummies are statistically significant with 95% confidence, namely, January 1998 and March 2012. We note that the two major structural breaks roughly coincide with two crises that have plagued South Africa and Asia, namely the Asian crisis of 1998 (see Bundia and Ricci 2005 for a detailed account) and the combination of the platinum strikes and Marikana massacre in

[17] See Bai, and Perron (1998, 2003a) for details.

2012.[18] Accordingly, we include two dummy variables (D_1 and D_2) to control for the two breaks in our ARDL estimation.[19]

The basic descriptive statistics are illustrated in Table 3.3.

Table 3.3: Basic descriptive statistics

	Log(Nom Exch)	**Log(Nom Comm)**
Mean	1.9723	4.7273
Median	1.9706	4.6180
Std. Dev.	0.2497	0.3206
Skewness	-0.2729	0.2048
Kurtosis	2.8576	1.5601
Jarque-Bera	3.0226	21.2897
Probability	0.2206	0.0000
Observations	228	228

The variables appear to be non-stationary although it is not possible to ascertain the order of integration from visual inspection of Figure 3.2. The nominal exchange rate appears to be highly kurtotic with negative skewness. It is noteworthy that the commodity prices have positive skewness alongside a negative skewness for the exchange rate. This property suggests that extreme movements in the commodity prices are associated with depreciation rather than appreciation of the exchange rate.

The ARDL model requires that the variables be at most $I(1)$ as the bounds test's critical values are invalid for variables of integration of order higher than $I(1)$. For this reason we test and present the unit root tests in Table 3.4. The results confirm that both variables are difference stationary $I(1)$. We proceed to estimate the ARDL model for the two asset prices.

[18] Mpofu and Peters (2016) note the Marikana massacre on 16 August 2012 led to the Rand/US dollar, Rand/British Pound and Rand/Euro depreciating by 1.86 percent, 2.21 percent and 2.31 percent respectively from the date before the strike began. See also Breckenridge (2014), Wehmhoerner (2012). On account of the 1998 Asian crises: between end-April and end-August in 1998, the rand depreciated by 28 percent in nominal terms against the U.S. dollar. This was accompanied by increases of around 700 basis points in short-term interest rates and long term bond yields, while sovereign U.S. dollar-denominated bond spreads increased by about 400 basis points. At the same time share prices fell by 40 percent and output contracted during the third quarter of 1998 (quarter on-quarter).

[19] We estimated the model without the dummy variables. The results are significantly different as measured by stability diagnostics and the rejection of the cointegration null.

3.4.4.2 *Estimation results of the ARDL model*

We present the results of the long run estimation in Table 3.5, the short run error correction model in Table 3.6 and the bounds test for cointegration in Table 3.7. One of the most important aspects of the ARDL approach is the choice of the lag order. We follow Pesaran and Smith (1998) and choose the lag order of the model based on the Schwartz Bayesian Information Criterion (SBIC). The authors argue that the SBIC method tends to define a more parsimonious model in the case of short samples. Accordingly, our choice of the SBIC is motivated by our small sample size. However, we include the model chosen by the Akaike's Information Criteria (AIC) for robustness. For these information criteria, the optimal number of lags for each of the variables in the ARDL are specified as (3, 1) and (2, 1) under the AIC and SBIC respectively.

Table 3.4: Unit root tests

Variable	**ADF TEST STATISTIC**		**PHILIPS PERRON TEST**	
Levels	**Constant with trend**	**Constant without trend**	**Constant with trend**	**Constant without trend**
Log (Nom exch)	-2.5096	-1.5857	-2.3960	-1.5293
Log(Nom Comm)	-1.8772	-0.7633	-1.8127	-0.7388
First Differences				
ΔLog (Nom exch)	-10.5214	-10.9841	-10.4536	-11.0558
ΔLog(Nom Comm)	-9.9844	-10.5672	-10.0271	-10.6113

Notes: The Critical values for rejection are -4.0296, -3.4444 and -3.1471 at a significant level of 1%, 5% and 10% respectively for models with a constant and linear trend and -3.4812, -2.8830, -2.5787 at a significant level of 1%, 5% and 10% respectively for models without a linear trend. The optimal lag for the ADF test was chosen based on the Schwartz Information Criterion and the truncation parameter for the PP test was selected using the Newey-West truncation method.

3.4.4.3 *The dynamic relationship between the Rand and commodity prices*

Our results provide support for the nexus between commodity prices and the nominal exchange rate of the Rand in South Africa. When we control for structural breaks using dummy variables (Dummy_1 and Dummy_2), the ARDL model indicates that the two asset prices are cointegrated (Table 3.5). The F-test reading of 6.4947 is greater than the upper bound critical value (5.7300) under the SBIC model. The null of no cointegration is rejected at 10% level of significance under the AIC model and is inconclusive at 5% level or better.

Table 3.5: Bounds test for cointegration

Null Hypothesis: No long-run relationships exist

	Lag length Selection Criteria			
	AIC		**SBIC**	
$F-Statistic$	5.3868		6.4947	
Significance	**I0 Bound**	**I1 Bound**	**I0 Bound**	**I1 Bound**
5%	4.9400	5.7300	4.9400	5.7300

These results lead us to examine an alternative way of establishing cointegration suggested by Kremers *et al* (1992), Bannerjee et al (1998) and Bahnani-Oskooee (2001). These authors demonstrate that a negative and significant lagged error term within the bounds testing approach is suggests cointegration among the relevant time series variables.

Table 3.6: Estimated short-run error correction model

Dependent variable Δs_t

	Lag length Selection Criteria	
	AIC	**SBIC**
Δs_{t-1}	0.3065***[4.8078]	0.2702*** [4.4402]
Δs_{t-2}	-0.1235**[-1.9660]	
ΔCom	-0.3804***[-4.8386]	-0.3685*** [-4.6714]
$Dummy_1$	-0.0000*[-0.0003]	-0.0002 [-0.0075]
$Dummy_2$	0.0221[0.6757]	0.0277[0.8452]
ECT	-0.0461***[-3.3957]	-0.0562***[-4.2184]
R^2	0.9820	0.9821
DW-Stat	1.9443	1.8769
F-stat	1697.4360***	2012.2201***

*Notes: *, **, *** indicate statistical significance at 10%, 5% and 10% level of significance respectively.*

Turning to the results in Table 3.6, although the error correction term coefficient is small , it is clear that it is significantly different from zero and is negatively signed under both the AIC and SBIC models. These results corroborate the bounds test suggesting that the two variables are cointegrated. Our estimations suggest that when there is an exogenous shock to the system and the exchange rate is above (or below) the equilibrium level, approximately 0.05% adjustment will be achieved in the first month under both models. The speed of adjustment of the exchange rate to the shocks in the commodity markets appears to be very slow. This result may indicate that while the commodity price variable matters in the exchange rate equation, its influence in the short run is relatively muted at a monthly frequency. The negative sign of the error correction term is important and suggests that the error correction process is stable and convergent.

The short-run model also shows that the coefficient of Δcom has the expected sign and is statistically significant (-0.38 and -0.37 under AIC and SBIC respectively). Therefore a 10% increase in commodity prices is associated with approximately 3.7% appreciation in the value of the nominal exchange rate. The R^2 reading is quite large for both the AIC and SBC models; the F-statistic is also statistically significant suggesting that the model fits the data well. This result lends support to the notion that the value of the South African Rand moves in in line with commodity prices.

While our results point to the presence of a cointegration relationship between the two nominal prices, the long-run impact of the commodity price with respect to the nominal exchange rate (Table 3.7) is not only very small (0.06) but statistically insignificant at all conventional levels.[20] We conclude that while our results show equilibrium between the two variables, the long run relationship is weak and the correction process slow. The relationship is however significant and strong in the short run. We suspect that factors such as structural changes in the South African economy and integration with global financial markets have varying effects on the relationship of the two variables in the long run. These results motivate the substance of the next section, where we test the standard fundamental models of the exchange rate, augmented by the commodity prices variable.

[20] We obtain very similar results when we test the model using the real effective exchange rate and the real commodity prices(i.e. nominal commodity prices deflated by the Manufacturing Value Index [MUV])

Table 3.7: Estimated long-run coefficients

Dependent Variable s_t

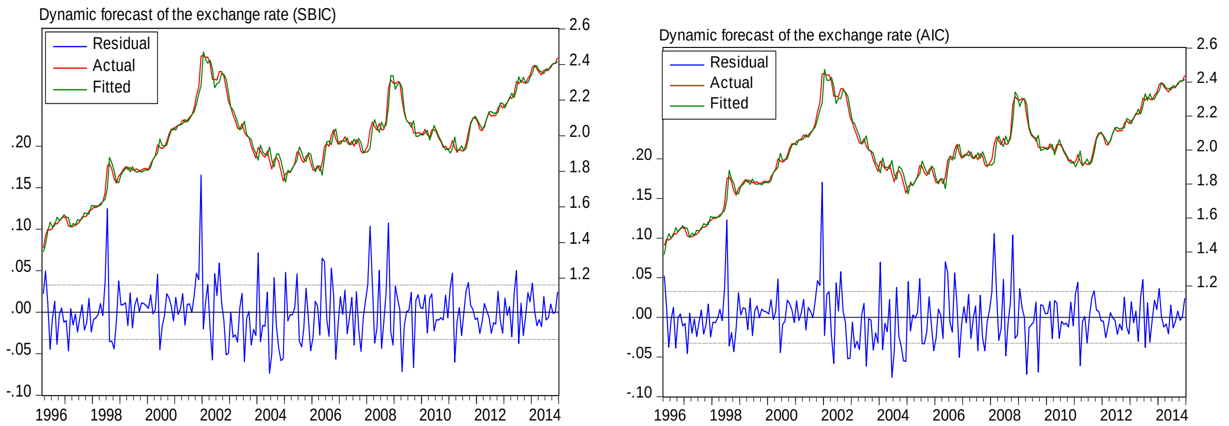
	Lag length Selection Criteria	
	AIC	**SBIC**
Com	0.0580 [0.3072]	0.0610[0.3614]
D_1	0.4265***[2.9398]	0.4181*** [3.9175]
D_1	0.3532***[2.5506]	3.3170*** [3.3170]
Constant	1.3663 [1.5472]	1.3475*[1.6931]

*Notes: *, **, *** indicate statistical significance at 10%, 5% and 10% level of significance respectively.*

3.4.4.4 Stability Tests Results

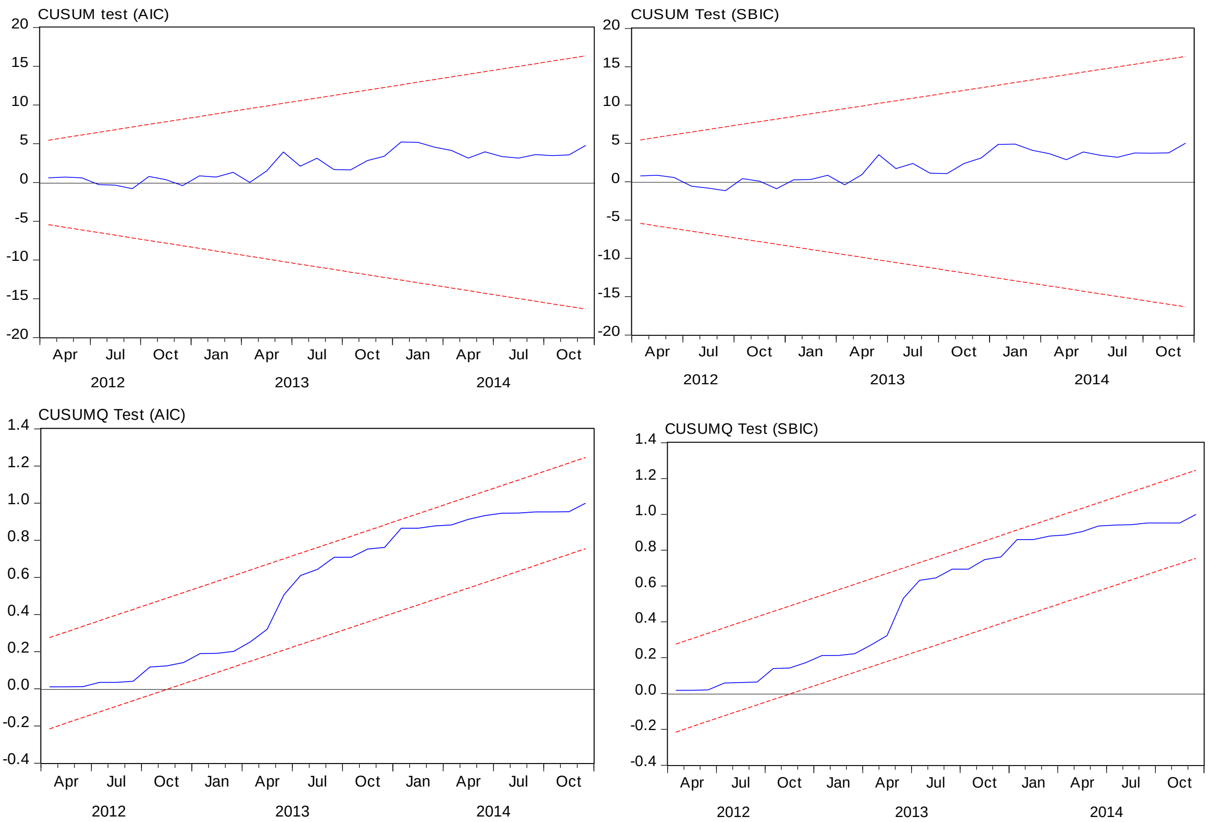
According to Pesaran and Pesaran (1997), the stability of the coefficients of the ARDL model must be investigated. First we consider the dynamic forecast of our model against the actual variables over the sample period. A look at the plots in Figure 3.3 suggests that our model fits the data quite well. The residuals appear to be stable over the sample period as well. The model also passes the tests of autocorrelation and heteroscedasticity (see Table 3.8).

Figure 3.3: Plots of actual and forecast values of the exchange rate



Additionally, we examine the graphical representation of the CUSUM and CUSUMQ following Brown, Durbin, and Evans (1975), Bahnani-Oskooee (2001) and Pahlavani (2005) in Figure 3.4.

Figure 3.4: Plots of CUSUM and CUSUMQ statistics for coefficient stability



The null hypothesis (that is the regression equation is correctly specified) cannot be rejected if the plot of the CUSUM and CUSUMQ remain within the 5% critical band over the sample period. It is clear from this figure that the plots of both CUSUM and CUSUMQ are within the 5% critical boundaries and confirm the stability of the coefficients of our model under both the AIC and SBIC information criteria.

Table 3.8: Tests of autocorrelation and heteroscedasticity

Breusch-Godfrey Serial Correlation LM Test :*Null Hypothesis: no residual autocorrelations*		
	AIC	**SBIC**
F-statistic	1.3434[0.1962]	1.5460[0.1100]
Heteroscedasticity Test: Harvey :*Null: no heteroscedasticity in residuals*		
F-statistic	1.6807[0.1149]	2.0429[0.1612]

Notes: p-values in [ ] parenthesis

3.5 Commodity Prices and Structural Exchange Rate Models

In this section, we incorporate the commodity price fundamental in testing standard structural exchange rate models. The extensive literature on exchange rate economics shows that identifying the elusive connection between exchange rates and economic fundamentals is no easy task (Obstfeld and Rogoff, 2000).[21] In the context of South Africa, evidence in favour of the standard monetary model of exchange rate determination is mixed (see de Bruyn et al., 2012; Hassan and Simeone, 2011 and Moll, 1999, 2000). Evidence in favour of PPP is rare and mixed at best. Using recent floating era data, support has been found for PPP but only after allowing for non-linearity in the data generation process (Larceda et al., 2010). Other modifications to the PPP model include allowing for half-life definitions (Mokeona et al., 2009a) and long memory (Mokoena et al., 2009c).

We argue that inclusion of the commodity price variable is a plausible modification of the standard exchange rate models. In the spirit of Chen and Rogoff (2003), we conjecture that the commodity price variable may be a potential omitted variable in the standard models for South Africa (which depends on the export of its primary commodities for a significant portion of its foreign exchange revenues).[22] Moreover, there is reasonably long market determined exchange rate time series data available since South Africa abandoned the dual exchange rate regime in favour of a floating exchange rate in 1995.[23]

Specifically, this section intends to answer the following questions about the Rand: 1) in the context of standard structural models, does the commodity price fundamental help explain movements in the Rand? 2) Without the commodity price fundamental, do standard

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[21] The difficulties manifest themselves in the exchange rate disconnects puzzle - the missing link between exchange rate and a menu of economic and financial fundamentals that theory suggests should drive exchange rates.
[22] The UN Comrade data shows that exports of gold, platinum, coal and iron ore accounted for an average of 35% of South African exports between 2005 and 2014. See Data Appendix for details.
[23] Further, new evidence in the search for answers to the exchange rate disconnect puzzle suggests that forecast based trading strategies outperform the random-walk based trading strategies (see Moosa and Burns, 2012, 2015).

exchange rate models fit the data better? 3) Does inclusion of the commodity price fundamental improve the out-of-sample forecasting ability of the standard models?

3.5.1 Model specifications

The starting point in our analysis concerns the standard workhorse of exchange rate determination models – the Purchasing Power Parity (PPP) model. According to the PPP hypothesis, the nominal exchange rate is the ratio of the countries' price levels, according to the law of one price. Specifically, we follow the assumption that the exchange rate reflects the ratio of purchasing power between countries. Providing for transport costs and other factors, we follow Chen (2002) to specify two commodity-price augmented variants of the PPP model:[24]

PPP1Model: $$s_t = a - \beta_1 p_t^{com} + \beta_2 (p_t - p_t^*) + \varepsilon_t \tag{3.11}$$

PPP2 Model: $$s_t = a - \beta_1 p_t^{com} + \beta_2 p_t - \beta_3 p_t^* + \varepsilon_t \tag{3.12}$$

All variables are logarithms. s_t is the nominal exchange rate quoted as units of domestic currency/foreign currency such that an upward movement represents depreciation of the home currency. p^{com} is the log of South Africa-specific commodity price index; $p\ and\ p^*$ are the domestic and foreign CPIs respectively and ε_t is a stationary disturbance.[25] The commodity price variable enters the equations with a negative sign since we expect an upward movement of commodity prices to appreciate the exchange rate.

The PPP model is central to exchange rate modelling. Additional restrictions can be imposed on the PPP relation to build the popular monetary class models of the exchange rate. Assuming money market equilibrium, that is, that the log of real money demand depends linearly on the log of real income and nominal interest rates, it can be shown that:

[24] We augment the structural exchange rate models by the commodity prices, which measure the terms of trade shocks for South Africa – a predominantly commodity exporting economy.
[25] We follow the Casellian version of the PPP model that measures relative prices in terms of CPIs.

$$m_t - p_t = \beta_y y_t - \beta_i i_t + \varepsilon_t \qquad (3.13)$$

Where m_t is the domestic money stock; y_t is domestic and foreign income and i_t is domestic short-term interest rate; ε_t represents a stationary disturbance.

If we have an identical equation for a foreign country, extending the PPP model, the relative CPIs would cancel and the exchange would become a function of relative money stocks, interest rate differentials and relative real income for the two countries as follows:

$$s_t = \alpha + (m_t^* - m_t) - \beta_y (y_t^* - y_t) + \beta_i (i_t^* - i_t) + \varepsilon_t \qquad (3.14)$$

In equation (3.14), the nominal interest rate differentials reflect inflation risk premia. With increasing domestic inflation, investors would sell domestic currency and invest in the bonds market exerting downward pressure on the domestic exchange rate. Assuming that the uncovered interest parity (UIP) condition holds, we have:

$$i_t^* - i_t = E_t(s_{t+1} - s_t) \qquad (3.15)$$

Equation (3.15) reflects international markets equilibrium and assumes that domestic and foreign assets are perfect substitutes. Following Killian (1999) and Chen (2003), the UIP condition can be incorporated into equation (3.14) to give the first order stochastic differential equation for the exchange rate. The exchange rate can be expressed as the expected present value of relative money stock and relative real income. Assuming that relative money stock and real income follow a driftless random walk, we obtain:

$$s_t = \alpha + (m_t^* - m_t) - \beta_y (y_t^* - y_t) + \varepsilon_t \qquad (3.16)$$

Equations (3.14) and (3.16) are two variants of the flexible price monetary model of the exchange rate. In this study we test these two specifications with and without the inclusion of the commodity price index as follows:[26]

MM1 Model: $$s_t = a - \beta_1 p_t^{com} + \beta_2(m_t - m_t^*) - \beta_3(y_t - y_t^*) + \varepsilon_t \quad (3.17)$$

MM2 Model: $$s_t = a - \beta_1 p_t^{com} + \beta_2(m_t - m_t^*) - \beta_3(y_t - y_t^*) + \beta_4(i_t - i_t^*) + \varepsilon_t \quad (3.18)$$

Coefficients β_2, β_3 and β_4 represent elasticity with respect to money stock, money demand and interest rates respectively.

3.5.2 Data characteristics

We test the structural models discussed in section 3.5.1 using the USD/ZAR bilateral exchange rate and for robustness, we test the performance of the models against other major Rand cross rates namely the Pound-Sterling/Rand (GBP/ZAR) and the Euro-Rand (EUR/ZAR). In all cases the exchange rates are measured as monthly averages in Rand per base currency. The money supply variable is measured as M1 in all cases except for the UK where we use M0 in USD billions. Inflation is measured as CPI in all cases while we use the three month Treasury-bill rate in percent/annum in all cases. The real GDP numbers are measured in billions of USD. Given that the data is measured in quarterly intervals, we use linear interpolation to obtain monthly numbers as suggested by Sjuib, (2009).[27] With the exception of the GDP series from the World Bank database, all the other data was extracted from the IFS database. For commodity prices, we construct a South Africa-specific production-weighted commodity price index based on four major export commodities

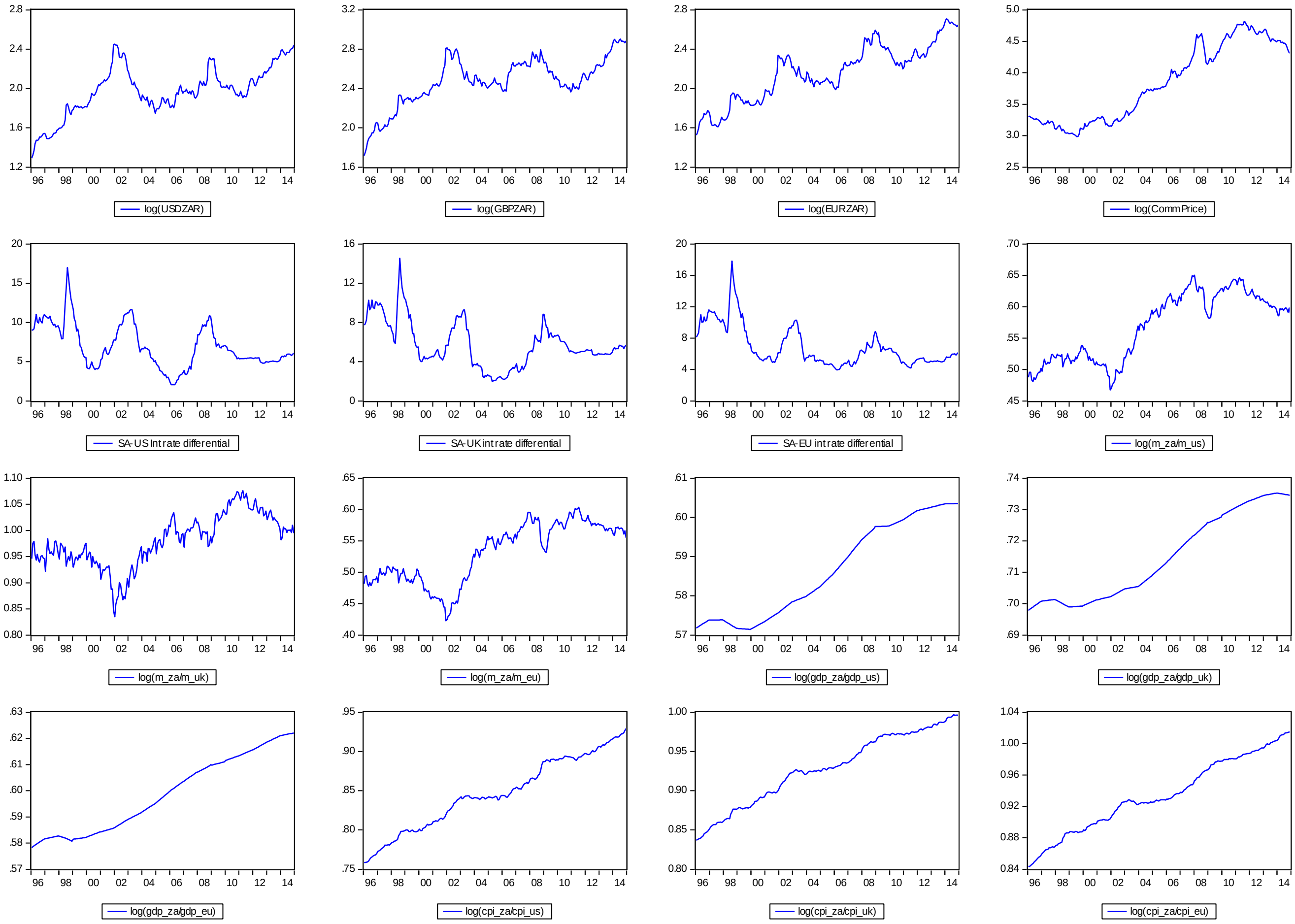
[26] Details on these models are provided in Bilson (1978), Frankel (1976), and MacDonald and Taylor (1994)
[27] See also Kodongo and Ojah(2012)

following Cashin et al. (2004).[28] Figure 3.5 illustrates the graphical representation of the data series. All the variables appear to be non-stationary.

In Tables 3.9-3.12 we report the augmented Dickey Fuller (ADF) test (Dickey and Fuller, 1979, Dickey and Fuller, 1981) and Philips-Perron test (Phillips and Perron, 1988) for unit root. Both tests suggest that all the variables are $I(1)$.

Additionally, we test the four different specifications for cointegration (Table 3.13). Except for the trivariate PPP model for EUR variables in Model 2, the UK version of Model 3 and 4, cointegration is detected in all cases.

Figure 3.5: Data plots



[28] This is the same price index used in the first section of the chapter.

Table 3.9: Representative Tests for unit root (South Africa Variables)

Null hypothesis: variables have unit root or non-stationary

Variable	**ADF TEST STATISTIC**		**PHILIPS PERRON TEST**	
Levels	**Constant with trend**	**Constant without trend**	**Constant with trend**	**Constant without trend**
Log (Nominal USD exch-rate)	-2.5096	-1.5857	-2.3960	-1.5293
Log (Nominal GBP exch-rate)	-2.2170	-2.4965	-2.0685	-2.6037
Log (Nominal EUR exch-rate)	-3.0562	-1.5360	-2.8913	-1.5482
Log(Commodity price Index)	-1.8772	-0.7633	-1.8127	-0.7388
Log(Money Supply)	-1.4895	-1.2959	-1.6869	-1.2936
Log(CPI)	-2.3069	-1.5973	-2.3859	-1.4283
Log (Real GDP)	-2.0936	-2.0682	-0.8908	-0.8097
Interest Rate	-3.5610	-3.2684	-3.1285	-2.7288
First Differences				
ΔLog (Nominal USD exch-rate)	-10.5214	-10.9841	-10.4536	-11.0558
ΔLog (Nominal GBP exch-rate)	-12.6666	-12.0289	-12.6666	-12.0289
ΔLog (Nominal EUR exch-rate)	-12.2599	-11.7054	-12.1278	-11.6480
ΔLog(Commodity price Index)	-9.9844	-10.5672	-10.0271	-10.6113
ΔLog(CPI)	-10.7911	-11.2256	-11.0640	-11.4644
ΔLog(Money Supply)	-13.7374	-14.7410	-13.7309	-14.7305
ΔLog (Real GDP)	-3.3455	-3.1858	-3.0396	-2.9585
ΔInterest Rate	-9.9786	-9.1913	-9.7857	-8.8630

Table 3.10: Representative Tests for unit root (USA Variables)

Variable	**ADF TEST STATISTIC**		**PHILIPS PERRON TEST**	
Levels	**Constant with trend**	**Constant without trend**	**Constant with trend**	**Constant without trend**
Log(Money Supply)	0.1074	-2.0931	-0.0223	4.0642
Log(CPI)	-2.2015	-1.0725	-3.0385	-1.1246
Log (Real GDP)	-2.5569	-1.6123	-1.3796	-2.3869
Interest Rate	-3.7775	-2.2350	-2.7475	-1.2614
First Differences				
ΔLog(CPI)	-8.9037	-11.3815	-7.6475	-8.0917
ΔLog(Money Supply)	-4.6923	-5.7754	-14.9500	-14.9006
ΔLog (Real GDP)	-3.2518	-3.8392	-4.1988	-3.5106
ΔInterest Rate	-7.2661	-3.7328	-13.7077	-12.1440

3.5.3 The Dynamic OLS (DOLS) Estimator

We employ Stock and Watson's (1993) Dynamic OLS (*DOLS)* to estimate the cointegrating vectors in our model. The DOLS method possesses some advantages of alternative methods of estimating cointegrating systems.

Table 3.11: Representative Tests for unit root (UK Variables)

Null hypothesis: variables have unit root or non-stationary

Variable	**ADF TEST STATISTIC**		**PHILIPS PERRON TEST**	
Levels	**Constant with trend**	**Constant without trend**	**Constant with trend**	**Constant without trend**
Log(Money Supply)	-1.8581	-1.0052	-2.7413	-1.0758
Log(CPI)	-1.7994	1.0694	-2.3292	1.1885
Log (Real GDP)	-2.4571	-1.7683	-2.4606	-1.7648
Interest Rate	-2.6283	-1.0251	-2.7849	-0.7514
First Differences				
ΔLog(CPI)	-2.1774	-5.0694	-17.4920	-29.2864
ΔLog(Money Supply)	-15.1755	-3.7170	-28.4711	-16.3631
ΔLog (Real GDP)	-10.4051	-15.4543	-14.7313	-154543
ΔInterest Rate	-6.6963	-7.2957	-6.6964	-7.3529

Table 3.12: Representative Tests for unit root (EU Variables)

Null hypothesis: variables have unit root or non-stationary

Variable	**ADF TEST STATISTIC**		**PHILIPS PERRON TEST**	
Levels	Constant with trend	Constant without trend	Constant with trend	Constant without trend
Log(Money Supply)	-2.2914	-0.5721	-2.2546	-0.6027
Log(CPI)	-0.8850	-1.3852	-1.1751	-1.3120
Log (Real GDP)	-1.4590	-2.4482	-1.7271	-3.1061
Interest Rate	-2.6933	-1.5035	-2.7769	-1.5533
First Differences				
ΔLog(CPI)	-5.2349	-11.2582	-12.7853	-11.5327
ΔLog(Money Supply)	-4.3134	-13.9563	-10.6794	-13.9124
ΔLog (Real GDP)	-3.5413	-4.1217	-5.7297	-6.3582
ΔInterest Rate	-5.2759	-9.6377	-7.0663	-10.0864

Notes: The Critical values for rejection are -4.0296, -3.4444 and -3.1471 at a significant level of 1%, 5% and 10% respectively for models with a constant and linear trend and -3.4812, -2.8830, -2.5787 at a significant level of 1%, 5% and 10% respectively for models without a linear trend. The optimal lag for the ADF test was chosen based on the Schwartz Information Criterion and the truncation parameter for the PP test was selected using the Newey-West truncation method.

The Johansen maximum likelihood estimator, for example is known to be plagued by small sample bias and producing widely dispersed estimates (Chen, 2002). Further, the Johansen maximum likelihood approach, being a full information approach, is vulnerable to the problem that parameter estimates in one equation may be affected by misspecification in other equations. The DOLS method addresses this problem by its design as a robust single equation method which has been shown to have the same asymptotic optimality properties as the Johansen distribution (Al-Azam and Hawdon, 1999; Masih and Masih 1996a). The DOLS methodology overcomes the regressors' endogeneity problems associated with simple OLS regressions by the inclusion of leads and lags of first differences of the regressors, and for serially correlated errors by a General Least Squares (GLS) procedure.

Table 3.13: Johansen Test for Cointegration

Base Country	**Number of Cointegrating relationships** Trace Statistic	Eigenvalue Statistic
Model 1: log(Exch-rate); log(Commodity Price); log(CPI_SA/CPI_foreign)		
USA	1	1
UK	1	1
EU	1	1
Model 2: log(Exch-rate); log(Commodity Price); log(CPI_SA) log(CPI_foreign)		
USA	2	2
UK	2	1
EU	0	0
Model 3: log(Exch-rate); log(Comm Price); log(Mon Supply_SA/Mon Supply_foreign); log(GDP_SA/GDP_foreign)		
USA	1	1
UK	0	0
EU	0	1
Model 4: log(Exch-rate); log(Comm Price); log(Mon Supply_SA/Mon Supply_foreign); log(GDP_SA/GDP_foreign); Interest differentials		
USA	2	1
UK	0	0
EU	0	1

The DOLS procedure basically involves regressing any cointegrated I (1) variables on other I (1) variables, any I (0) variables and leads and lags of the first differences of any I(1) variables. It the context of structural models the model is represented in the following econometric specification:

$$s_t = \beta_0 + \beta_1 f_t + \sum_{j=-p}^{p} \delta_j \Delta f_{t-j} + u_t \tag{3.19}$$

Where s_t is the exchange rate, f_t is the matrix of fundamental explanatory variables

β_1 is the cointegrating vector; that is, represents the long-run cumulative multipliers or, alternatively, the long-run effect of a change in f on s_t and $(-)p$ and p are the lag lead lengths respectively. AS pointed out earlier, the lags and leads of Δf are added to the DOLS model for the purpose of making its stochastic error term independent of all past innovations in stochastic regressors. We also employ the heteroscedasticity consistent covariance (HAC) method proposed by Newey and West (1987) to address the problem of heteroscedasticity and autocorrelation in the regression errors. Finally, we carry out unit root tests on the residuals to ascertain whether our estimations are spurious.[29]

3.5.4 Dynamic OLS Estimation Results

We report the Stock-Watson DOLS parameter estimates with all variables appearing in levels for the USD/ZAR, GBP/ZAR and EUR/ZAR models in Tables 3.14-3.16. We report the long run elasticities of the level regressors on the nominal exchange rate with asymptotic standard errors in parenthesis. We show the estimates of the models with and without the commodity prices variable.

The commodity price variable appears with the correct anticipated sign and is significant in all but the monetary models based on the GBP and EUR. On average, the commodity price variable enters all models with a negative coefficients of -0.56, -0.36 and -0.25 in the dollar, GBP and EUR based models respectively. The data fits the PPP dollar based models fairly well as well as the other Rand crosses; all appearing with correct signs and significant. In all cases the adjusted R^2 reading improves notably by inclusion of the commodity price variable, suggesting an improvement of fit. Turning to the stability of the models, inclusion of the commodity price variable further improves the ADF statistic for stationarity of residuals across all PPP and monetary models. These findings are robust to the other Rand crosses as well.

[29] Choi *et. al.*, (2008) point that a regression is technically called a spurious regression when its stochastic error is unit-root nonstationary

Table 3.14 Estimating the Cointegrating relationships under Dynamic OLS [Base: US]

	PPP1	PPP1 + Comm Price	PPP2	PPP2 + Comm Price	MM1	MM1 + Comm Price	MM2	MM2 + Comm Price
Log(Com Price) [-]		-0.3611*** (0.0412)		-0.5602*** (0.0623)		-0.6621*** (0.1311)		-0.6702*** (0.1106)
Log(CPI_ZA/CPI_US)[+]	5.0525** (0.5826)	8.9000*** (0.6203)						
Log(CPI_ZA) [+]			1.4903*** (0.2102)	1.4121*** (0.1811)				
Log(CPI_US) [-]			-2.5002*** (0.5842)	-0.4921 (0.6602)				
Log(M1_ZA/M1_US)[+]					-2.6711*** (0.9222)	-1.5603*** (0.7831)	-2.9001*** (0.9622)	-1.5911* (0.8912)
Log(GDP_ZA/GDP_US)[-]					-14.3502* (7.7522)	22.3444** (8.8625)	-13.6712** (8.8902)	23.0400*** (8.1323)
Tbill_ZA-Tbill_US [-]							-0.0003 (0.0002)	-0.0000 (0.0002)
Residual ADF Test	-3.8333	-5.1301	-3.9725	-4.9123	-3.6122	-4.2014	-3.1152	-4.1902
Adjusted R^2	0.8112	0.8925	0.8233	0.9232	0.8642	0.9002	0.8531	0.9002
No. of Observations	221	221	221	221	221	221	221	221

Notes: The following notes are applicable to the results in tables 3.15 and 3.16.

a) *The output shows the estimated models with and without the commodity prices variable. The expected signs of the coefficients are shown in parenthesis[ ] against the variables on the tables*
b) *The Stock-Watson (1993) dynamic OLS (DOLS) methodology is used to obtain super consistent estimators of the cointegrating vectors with approximate asymptotic standard errors reported in parenthesis ( ) against the coefficient estimates.*
c) **, **, *** denotes statistical significance at 10%, 5% and 1% level respectively.*
d) *A South-Africa specific commodity prices, constructed according to Cashin et al (2003) is the production weighted average of the top four commodity exports from South Africa, priced in USD. See data appendix A for details.*
e) *The models were estimated up to $j = \pm 3$ lags of each dependent variable to orthogonalise.*
f) *The residuals ADF test statistic is shown for each model. The null is hypothesis is: residuals have unit root against the alternative that the residuals are stationary. Non-*

stationarity of residuals suggests that the model may be spurious. The Critical values for rejection are -3.4812, -2.8830, -2.5787 at a significance level of 1%, 5% and 10% respectively.

Table 3.15: Estimating the Cointegrating relationships under Dynamic OLS [Base: UK]

	PPP1	PPP1 + Comm Price	PPP2	PPP2 + Comm Price	MM1	MM1 + Comm Price	MM2	MM2 + Comm Price
Log(Com Price) [-]		-0.4811***		-0.4910***		-0.1200		-0.3532***
		(0.0403)		(0.0535)		(0.1539)		(0.0919)
Log(CPI_ZA/CPI_UK)[+]	4.0221***	9.5700***						
	(0.7722)	(0.7325)						
Log(CPI_ZA) [+]			2.0602***	2.3433***				
			(0.4346)	(0.3302)				
Log(CPI_UK) [-]			-3.6323***	-2.5232				
			(0.9402)	(0.7719)				
Log(M1_ZA/M1_UK)[+]					-4.3218***	-4.4002***	-4.1301***	-3.8638**
					(0.3922)	(0.6211)	(0.3644)	(0.3407)
Log(GDP_ZA/GDP_UK)[-]					-5.0609	1.8018	3.9645	23.1811***
					(5.8037)	(7.6344)	(4.1826)	(5.4624)
Tbill_ZA-Tbill_UK [-]							-0.0202***	-0.0536***
							(0.0041)	(0.0038)
Residual ADF Test	-2.7604	-4.2225	-3.5711	-4.2928	-3.2624	-3.4108	-3.9039	-4.4489
Adjusted R^2	0.7511	0.8905	0.8145	0.8912	0.9045	0.9022	0.9402	0.9642
No. of Observations	221	221	221	221	221	221	221	221

Turning to the monetary models of the Rand however, we note that these models generally perform poorly across all the base currencies, judged by the size and direction of coefficients. While the money supply variable is consistently significant across all models, it enters the models against our a priori expectations with a negative sign. The same observation applies to the output variable which, in addition to a positive sign, appears in most models with an unreasonably large coefficient and is not always significant. The inclusion of the commodity price variable does not to change the signs in all cases. On the interest rate differential, the exchange rate is consistently unresponsive across all bases except the GBP. The estimated elasticity of the GBP exchange rate, while it appears significant with a correct sign, is quantitatively very small. This finding suggests that South Africa is an unlikely destination of dollar and Euro carry trades.[30] Hassan (2014) demonstrates that most of the carry trade turnover in South Africa is between the Japanese Yen and the Rand. The stability tests indicate that the monetary models (without the commodity price variable) may be spurious regressions at 10% level or better. This is particularly so for MM2 for the USD and EUR bases.

To summarise, we note that the inclusion of the commodity price variable increases the performance of the structural models of the Rand. The size of the estimated coefficients of the commodity price variable is close to the earlier findings from the bivariate ARDL model as well as other authors such as MacDonald and Ricci (2003). These findings support the view that the South African exchange rate is a commodity currency and that the commodity price variable is potentially an omitted variable in canonical structural exchange rate models for commodity exporting economies. Indeed, while the link between exchange rates and fundamentals remains an elusive one, future theoretical models of exchange rates stand to benefit from inclusion of this important fundamental.[31]

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[30] A carry trade is a class of currency speculation strategies designed to profit from a favourable interest-rate deferential, when the high-interest currency does not depreciate substantially (as to erode the interest carry.) relative to the low-interest currency. The simplest way to implement the carry trade is to borrow in the low-interest currency (the funding currency), buy the high-interest currency (the target currency) in the spot market, deposit the proceeds or buy fixed-income securities denominated in the target currency, and finally convert the terminal payoff back into the funding currency facing the exchange rate risk. This is the conventional (textbook) understanding of the carry trade. But it can also be implemented through the derivatives market, for example selling the currency forward when it is at a significant forward premium, or using currency options to hedge the exchange rate risk component (Hassan 2014).

[31] Similar observations are made for the other OECD countries namely, Canada, Australia and New Zealand by Chen (2002).

Table 3.16: Estimating the Cointegrating relationships under Dynamic OLS [Base: EU]

	PPP1	PPP1 + Comm Price	PPP2	PPP2 + Comm Price	MM1	MM1 + Comm Price	MM2	MM2 + Comm Price
Log(Com Price) [-]		-0.2900*** (0.0523)		-0.3131*** (0.0838)		-0.2123** (0.1138)		-0.1911 (0.1238)
Log(CPI_ZA/CPI_EU)[+]	5.4102*** (0.4625)	8.1824*** (0.7111)						
Log(CPI_ZA) [+]			1.7702*** (0.3924)	1.6138*** (0.5009)				
Log(CPI_EU) [-]			-3.3441** (1.2135)	-1.1705 (1.7833)				
Log(M1_ZA/M1_EU)[+]					-4.4611*** (1.1632)	-3.6618*** (1.1509)	-4.3005*** (1.2219)	--3.6107*** (1.1827)
Log(GDP_ZA/GDP_EU)[-]					22.4622 (9.2704)	34.1728 (9.9035)	21.7222 (9.5037)	32.7504*** (10.0344)
Tbill_ZA-Tbill_EU [-]							0.0000 (0.0000)	0.0000 (0.0000)
Residual ADF Test	-2.9801	-3.8425	-3.5200	-3.7245	-3.2342	-2.7608	-2.7615	-2.8212
Adjusted R^2	0.8801	0.9225	0.9049	0.9100	0.8524	0.8617	0.8609	08638
No. of Observations	221	221	221	221	221	221	221	221

3.5.5 Exchange rate predictability from fundamentals

One of the most contested areas of research in international monetary economics is that of exchange rate predictability from macro-economic fundamentals. Since the Meese and Rogoff (1983) work on exchange rate and fundamentals, the empirical failure of structural exchange rate models to outperform the random walk model in short horizon forecasts has become, as Moosa and Burns (2014) observe, “an undisputed fact of life”. There are several subsequent research projects - led by Frankel and Rose (1995) - that have lent currency to this conclusion who question the “value of further time series modelling of exchange rates at high or medium frequencies using macro-economic models”. Evans and Lyons (2004) further note that the Meese and Rogoff puzzle is the "most researched puzzle in macroeconomics", echoing Abhyankar et al (2005) who describe it as a "major puzzle in international finance". Bacchetta and van Wincoop (2006) conclude that the notoriously poor performance of existing macro-exchange rate models is most likely the major weakness of international macro-economics.

While the Meese-Rogoff conclusion has been remarkably resilient over the years, there have been some researchers who claim that it is possible to outperform the naïve random walk model, as judged by the root mean square error and other similar metrics. This class of models has been put forward prominently by Mark (1995), Mark and Sul, (2001), Chinn and Meese, (1995) and MacDonald, (1999). These models have however been challenged by several scholars. Killan (1999), Berkowitz and Giorgianni (2001) and Berben and Dijk (1998) question the robustness and the inference procedures used in these models. Further criticism has been levelled on the models' assumption of a stable cointegrating relationship.[32]

There are several reasons that have been put forward explaining the poor performance of macro-exchange rate models. In their 1983 paper, Meese and Rogoff point out that the failure of fundamentals based models to out-forecast the random walk model may stem from simultaneous equation bias, sampling error, stochastic movements in the true underlying parameters or misspecification. The authors suggest four sources of misspecification, namely uncovered interest parity, proxies for inflationary expectations, goods markets specification and the money demand function. Harvey (2006) corroborates the claim of misspecification of the monetary model – he points out that the monetary model is a stock rather than a flow model. Imposition of proportionality and symmetry restrictions have also been blamed as sources of misspecification by Neely and Sano (2002) and Tawadros, (2001). Chen (2002)

[32] See also Cheung et al (2005) who show that a wide range of models rate not successful in predicting exchange rates.

suggests that the problem with fundamentals based models may stem from potential omitted variable bias. For most countries, while it is difficult to conjecture what the potential missing variable may be, a potential omitted variable for commodity-dependent economies may stem from commodity prices. Chen, (2002) notes that the commodity price fundamental is not only identifiable for such economies but also “easily quantifiable”.

Following our earlier attempt to assess the commodity-augmented exchange rate models’ performance based on several goodness of fit criteria and direction of causality implied by economic theory, we attempt in this section to assess the predictive power of such models in short horizon forecasts of the Rand. While other scholars have done extensive work on statistical evaluation of forecast models (such as Moosa and Burns, 2014), our objective here is to evaluate whether the addition of the commodity price fundamental improves the performance of standard macro-models.

3.5.5.1 Forecasting approach

Our forecasting approach in this section is based on the error correction approach. As an empirical regularity, the time path of cointegrated variables is influenced by the error correction mechanism, that is, the extent of any deviation from long-run equilibrium. Christoffersen and Diebold (1998) show why inclusion of the error correction term improves the forecasting power of exchange rate models. The error correction term measures the extent of deviation of a system of variables from their long-run path and the error correction coefficient measures the rate at which this deviation is corrected.

In a recent study, Moosa and Vaz (2016) evaluate the forecast performance of error correction based models against their first difference equivalents over short horizons. They conclude that the ECM based framework has no forecast superiority over what is offered by the distributed lag structure of dependent and explanatory variables. The authors further argue against imposition of theory-based restrictions in exchange rate models.[33] Our approach follows this view and does not impose theory-based restrictions on cointegrating vectors but rather allows them to be determined by the data.[34] This approach is closer to the Meese-Rogoff, (1983) rolling regression framework. Additionally, our objective is to evaluate the forecast performance of models before and after inclusion of the commodity price variable. We therefore allow the coefficients in the standard exchange rate models to adjust to reflect the effect of inclusion (or exclusion) of this additional variable. We test two monetary exchange rate models specified earlier and their corresponding commodity augmented versions. We also test the PPP version that incorporates the price symmetry (PPP1).

We follow Faust et al (2001) to specify a basic unconstrained short horizon k forecast error correction model (ECM) of the following form:

$$\Delta s_{t+k} = s_{t+k} - s_t = \alpha_k + \emptyset_k z_t + \nu_{\mathrm{t}} \tag{3.20}$$

Where $z_t = \beta f_t - s_t$ and α_k and $\emptyset_k$ are estimated coefficients. f_t is the fundamental value of the exchange rate suggested by the exchange rate model and β is the cointegrating vector as in equation (3.19) and ν_{t} is a stationary disturbance. The intuition in the ECM is that if the exchange rate is cointegrated with fundamentals as implied by the structural models and assuming that the exchange rate is the variable that does the correction back to equilibrium, then deviations from the "fundamental" value of the exchange rate should help predict future exchange rate.[35] For equation (3.20) to be a valid ECM, $\emptyset_k$ must be negative and statistically significant; that is, if the current exchange rate is below its equilibrium value implied by

[33] Moosa and Vaz (2016) argue that imposition of theory-imposed restrictions presumes that the underlying theories are valid, without empirically testing them. Because the theories are far from being perfect, imposing theory-based restriction is therefore a "hazardous endeavour".

[34] This approach may be suitable in the South African context. The theoretical restrictions are based on theories developed for major industrialised countries with relatively longer floating exchange rate period.

[35] For a comprehensive discussion of the ECM forecasting model of the exchange rate see also Christoffersen and Diebold (1998) and Moosa and Vaz (2016).

fundamentals, the exchange should appreciate in the next period to correct the error. For our estimation, we rely on the value of β estimated from data.

For our three models under investigation, the fundamental value of the exchange rate (f_t) is predicted by the following relations:

$$f_t = p_t - p_t^* \tag{3.21}$$

$$f_t = (m_t - m_t^*) - (y_t - y_t^*) \tag{3.22}$$

$$f_t = (m_t - m_t^*) - (y_t - y_t^*) + (i_t - i_t^*) \tag{3.23}$$

Equation (3.20) is the basic error correction version of the standard macro-exchange rate models. Augmenting equation (3.20) with the commodity prices variable and substituting z_t with the error correction term yields:

$$s_{t+k} - s_t = \alpha_k + \emptyset_k[(\beta f_t - \beta_{1,k} p_t^{com}) - s_t]) + \nu_{\mathrm{t}} \tag{3.24}$$

Therefore a test of the null $\emptyset_k = 0$ agaisnt $\emptyset_k < 0$ is a test of exchange rate predictability from fundamentals. We evaluate each of the out of sample forecast for the standard models with and without the commodity price variable.

To carry out recursive out-of-sample estimates we follow Mark (1995), Faust et al (2001), Chen (2002) and Moosa and Burns (2014). We estimate the model in equations (3.20) and (3.24) over part of the sample period, $t = 1,2,\dots,m$ and then a k period ahead forecast generated for the point in time $m + k$. To generate the next forecast, this procedure is repeated and the models are estimated over the period , $t = 1,2,\dots,m+1$ so that the new forecast generated is for the point in time $m + k + 1$. It is clear from this procedure that the estimated forecast log exchange rate $(s_{t+k} - s_t) \equiv \hat{s}_{m+k}$. Using this notation therefore, this estimation procedure is repeated until $\hat{s}_n$ by estimating the models of the period $t = 1,2, \ldots, n - k$ where n is the total sample size.

The forecast log exchange rate therefore, pre and post augmentation with the commodity variable can be represented as:

$$\hat{s}_{m+k} = \hat{\alpha}_k + \widehat{\emptyset}_k z_{m+k} \tag{3.25}$$

$$\hat{s}_{m+k} = \hat{\alpha}_k - \hat{\beta}_k p_{m+k}^{com} + \widehat{\emptyset}_k z_{m+k} \tag{3.26}$$

Where $\hat{\alpha}_k$ and $\widehat{\emptyset}_k$ are the estimated values of α_k and $\emptyset_k$ respectively. The forecast level of the exchange rate is therefore given by:

$$\hat{S}_{m+k} = \exp(\hat{s}_{m+k}) \tag{3.27}$$

The recursive regression procedure is preferred for its forecasting superiority as argued by Moosa and Burns (2014). This methodology ensures that all information is available at the time of the forecast. Similar arguments are made by Nordhaus (1987), Marcelino (2002) and Marcelino et al (2001).

3.5.5.2 Forecast Evaluation

Evaluation of forecast accuracy of various models is achieved through use of an array of tests that evaluate the magnitude of the forecast error. From equation (3.24), once the forecast of the exchange rate $\hat{S}_t$ is known for the period $t = m + k, \ldots, n$, the forecast value can be compared with the actual series S_t to calculate various quantitative measures of forecast accuracy. One such measure of forecast accuracy, employed in our analysis is the Root Mean Squared Error (RMSE) calculated as:[36]

$$\text{RMSE} = \sqrt{\frac{1}{n-k-k+1} \sum_{t=m+k}^{n} \left(\frac{(\hat{S}_t - S_t)}{S_t}\right)^2} \qquad (3.28)$$

The RMSE depends on the scale of the dependent variable. It is used to compare forecast performance of the same series across different models. The basis of the evaluation is that the smaller the error, the better the forecasting ability of that model.[37]

To evaluate the forecast accuracy of the exchange rate models relative to the random walk model, we report Theil's inequality coefficient (U). The U coefficient reaches the lower boundary $U = 0$ for perfect forecasts and assumes a value of 1 when the exchange rate models being evaluated deliver forecast with same standard error as the naïve random walk model. The coefficient increases monotonically as the standard error forecasting of the random walk model improves relative to that of the exchange rate models. In sum, the random walk model outperforms the standard exchange rate models if $U > 1$. Theil's U coefficient is calculated as the RMSE ratio of the estimated model to that of the random walk as:

$$U = \frac{\sqrt{\frac{1}{n-m-k+1} \sum_{t=m+k}^{n} \left(\frac{(\hat{S}_t - S_t)}{S_t}\right)^2}}{\sqrt{\frac{1}{n-m-k+1} \sum_{t=m+k}^{n} \left(\frac{S_t - S_t}{S_t}\right)^2}} \qquad (3.29)$$

[36] The Meese-Rogoff (1983) paper and many subsequent studies used the RMSE to evaluate forecast accuracy of the standard exchange rate models against the random walk model without testing for the statistical significance of the difference. Other papers however have attempted use the Diebold and Mariano (1995) test. Corraddi, et al (2001) provide the conditions under which the Diebold and Mariano test can be used for tests of cointegrated systems. Clark and McCracken (2001) propose an array of tests for forecast comparisons for nested linear models. More recently, Moosa and Burns(2014) revisit the AGS test suggested by Ashley et al (1980)

[37] See also Chen and Yang (2004) for a discussion on stand-alone and relative forecast measures.

3.5.5.3 Forecasting performance of commodity-price augmented exchange rate models

We report the forecast performance of the three specifications of the exchange rate models, pre and post augmentation by the commodity price in Table 3.17.[38] The literature is replete with controversy on the tests for robustness of the different measures of forecast accuracy.[39] Our analysis of forecast accuracy in this section is illustrative only – we don't evaluate the statistical significance of the forecast measures against the random walk model. We therefore evaluate performance of the exchange rate models with and without the commodity prices variable using the simple measures of RMSE and the U-coefficient without assessment of their statistical significance.[40]

Results from the PPP model with price symmetry indicate that the forecast error improves with forecast horizon. The U-statistics also indicate that the model improves in its performance relative to the random walk as the forecast horizon increases. This picture is consistent with the commodity-price augmented version of the PPP model. Comparing the performance of the models pre and post augmentation indicates that addition of the commodity price variable does improve the forecast accuracy of the model. This result is important in our subsequent assessment of the monetary models as the PPP model is an important building block in the flexible price monetary models.

The link between the exchange rate and monetary variables appears to be stronger than CPI.[41] Notably however, in contrast to the PPP model, the forecast accuracy of the monetary models appears to diminish with an increase in the forecast horizon judged by the

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[38] We report findings only for the USD/ZAR exchange rate.

[39] The forecast accuracy criteria, for example, have been shown to suffer from small-sample size bias in coefficient estimates and in asymptotic standard errors in error correction frameworks (see Killian, 1997, Berkowitz and Giorgianni (1997 and Mark and Sul, 2001). Alternative methods to correct for sample bias such as bootstrapping have also been proposed but the robustness of findings of long-horizon forecast has been challenged (See Moosa, 2013). Moosa and Burns (2014) challenge the very notion of the "puzzle" with respect to the Meese and Rogoff paper that judged the performance of the standard exchange rate models against the random walk using the RMSE and related criteria. They propose alternative measures of forecast accuracy such as direction accuracy and profitability. Other scholars propose introduction of dynamics to the exchange rate models to improve their forecasting power (see Taylor 1995, Cheung et al, 2005, Tawadros, 2001, Hwang, 2001 and Aarle et al 2000).

[40] Chen, (2002) adopts a similar approach but uses only the Theil U statistic as a measure of forecast accuracy.

[41] Mark and Sul (2001) find a similar relationship for Australia and New Zealand.

size of both the RMSE and the Theil coefficient. Incorporating the commodity prices variable fails to overturn this pattern. In terms of improving forecast accuracy, commodity prices improve the forecast performance of the MM1 model, however that doesn't appear to be the case for model MM2. This observation is important in the context of Rapach and Wohar (2002) who argue that augmented structural exchange rate models can be viewed as "weaker' versions of the original theoretical models.[42] Without testing for statistical significance of the forecast evaluation criteria, the monetary models overwhelmingly outperform the naïve random walk model at all forecast horizons.

Table 3.17: Forecast Evaluation

	PPP1	**PPP1 + Comm Price**	**MM1**	**MM1 + Comm Price**	**MM2**	**MM2 + Comm Price**
One Month						
RMSE	0.0483	0.0460	0.0183	0.0183	0.0178	0.0178
U	1.0009	0.9005	0.4440	0.4396	0.4226	0.4231
Six Months						
RMSE	0.0459	0.0451	0.0239	0.0238	0.0241	0.0241
U	0.9307	0.9153	0.5186	0.5163	0.5218	0.5220
Twelve Months						
RMSE	0.0416	0.0404	0.0225	0.0224	0.0226	0.0226
U	0.9250	0.8975	0.5247	0.5225	0.5271	0.5273

3.6 Discussion and conclusions

The objective of the first empirical section of this chapter was to establish the nexus between the Rand and commodity prices between 1996 and 2014. The choice of the sample period was motivated by the objective of analysing the relationship during the floating regime of the South African exchange rate. We follow the techniques of Bai and Perron (1998, 2003a) to endogenously determine structural break dates within the time series data. The stability of our

[42] In their particular case, augmentation meant incorporating a linear trend in the cointegrating vector that allowed for a deterministic Balassa–Samuelson effect in real exchange rates.

model improves remarkably with the inclusion of the dummy variables to control for the two significant structural breaks suggested by our tests.

We allow the lags to be included in the ARDL model to be determined by both AIC and SBIC information criteria. The question whether the two variables are cointegrated is answered by the bounds test as well as examination of the significance and sign of the error correction term. The results of the bounds test reject the null of no cointegration at 5% and 10% levels of significance under the SBIC and AIC models respectively. Further, examination of the error correction terms shows that the coefficient is negative and significant. We conclude therefore that the two asset prices share a long-run relationship. The error correction term, which indicates the speed of adjustment which restores the dynamic relationship to equilibrium, is quite small. The long run elasticity is not statistically different from zero suggesting weak cointegration and slow error correction process. The model passes tests of structural stability as indicated by the CUSUM and CUSUMQ statistics and the residuals are also robust to autocorrelation and heteroscedasticity.

Commodity prices are therefore an important explanatory variable of the USD/Rand exchange rate.

In the second empirical section of the chapter we present evidence indicating that not only are commodity prices consistent explanatory variables of the bilateral USD/ZAR exchange rate but that they are also significant explanatory variables of the major Rand crosses as well. We use post-float data for South Africa and find evidence in favour of a long-run relationship in the commodity-price augmented PPP and monetary models for South Africa, supporting findings from other scholars such as Lacerda et al (2007) and Mokoena et al (2009a, 2009b and 2009c).

Evidence in favour of the canonical structural exchange rate models is at best mixed in the existing empirical literature. Judged by conventional goodness of fit criteria and signs of estimated coefficients, the models generally perform poorly despite their theoretical appeal. Moreover, since Meese and Rogoff (1983)'s seminal work, subsequent research has failed to convincingly overturn their conclusion on the forecast superiority of the naïve random walk model over standard exchange rate models. Several explanations have been put forward for what some authors have described as the "major puzzle in international finance" with Chen (2002) suggesting that the reasons for failure of theoretical exchange rate models may be due to "potential omitted variable bias".[43] While this argument has intuitive appeal, questions have been raised over augmentation of the exchange rate models with for example, trend specification, stock prices and commodity prices. The problems associated with endogeneity, observability and measurability of the potential omitted variables has led to some authors to suggest that the augmented models can be viewed as weaker versions of the canonical models (see Rapach and Wohar, 2002 and de Bruyn et al, 2012). However, to the extent that the data can help players in the exchange rate markets to improve performance, both from a policy and commerce stand point, we argue that there is merit in further attempts to solve the exchange rate determination puzzle, especially for an emerging market like South Africa.

Our analysis indicates that a 10% rise in commodity prices is associated with between 3.6% and 6.7% appreciation of the USD/ZAR exchange rate. The GBP and EUR crosses appreciate by between 2.9% and 4.9% respectively.[44] In terms of the type of models, while addition of the commodity price variable improves the in-sample fit of the PPP models, evidence of such improvement, measured in terms of expected a priori signs suggested by economic theory is mixed for monetary models. De Bruyne et al (2012) find similar problems using 101 years of South African data.

We further test the predictive prowess of the standard exchange rate models in the spirit of the Meese and Rogoff paper. We find that the commodity price variable does indeed

[43] See Abhyankar et al (2005)

[44] The fit of the models with the UK and EUR bases improves marginally when we use the GBP and EUR denominated commodity price models.

improve the forecast accuracy of the standard models although the results are not robust to model specification. The forecast accuracy test results suggest a tighter fit between the exchange rate and monetary variables (monetary model) than relative prices (PPP). While the forecasting results are presented for illustrative purposes only, without tests for statistical significance, they nevertheless invite further investigations using inference procedures that are robust to potential biases associated with small samples.

While the value of empirical time series modelling of exchange rates using macro-economic variables has been called to question (see Frankel and Rose, 1995), our study makes a contribution to the important debate of exchange rate determination in South Africa, using newer dataset of the floating Rand. Recent studies (see Moosa and Burns, 2012, 2015) argue that improving forecast accuracy of models of financial asset prices is valuable in terms of improving corporate performance. The authors, for example, demonstrate that a forecasting based currency trading strategy outperforms a simple carry trade strategy based on the random walk model. Results from South Africa, a bell-weather emerging commodity exporting economy with relatively developed financial markets, offers important lessons for other commodity-dependent African economies that are in the process of liberalising their financial markets. Using higher frequency data, improved forecast accuracy measures such as direction accuracy and profitability measures to test theoretical models against floating exchange rate data is a promising area of future research for commodity exporting emerging economies.

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3.8 Data Appendix

Construction of the South Africa Specific commodity Index

The country specific commodity price is constructed following Deaton and Miller (1996) and Cashin, et al. (2004).The nominal country specific commodity index p_t^{com} is constructed as a geometrically weighted index of the nominal prices of 4 individual commodity exports where for South Africa:

$$p_t^{com} = exp\left[\sum_{k=1}^{K}(W_k\,(lnP_k))\right] \qquad (A1)$$

Where $W_k = \left[\left(P_{jk}Q_{jk}\right)/\left(\sum_k P_{jk}\,Q_{jk}\right)\right]$

P_k is the dollar world price of commodity k (taken from the IFS database)

W_k is the weighting item which is the value of exports of commodity k in the total value of all K commodity exports for the constant period j;and

Q is the quantity of exports of commodity k (taken from the UN COMTRADE database.

The four major commodity exports considered in the computation of the index are gold, platinum, coal and iron ore. We calculate the 2005-2014 average total value of primary commodity exports; the four individual main commodity weights are calculated by dividing the 2005-2014 average value of each individual commodity export by the 2005-2014 average total value of primary commodity exports. Table A2 indicates the ten year aggregates and the averages for each commodity export. We show the calculation of the individual commodity weights on Table A1. All commodity weights are gross export weights as found in the UN COMTRADE data provided by the UN Statistical Department. Once the commodity export weights are calculated, these weights are held fixed over the sample period and are used to weight the individual (US dollar-based) price indices of the same individual commodities—taken from the IMF's IFS—to form, a geometric weighted-average index of (US dollar-based) nominal commodity-export prices (base 2010M06 = 100).

Table A1: Weighting of the individual commodities in the commodity index

	%age of total exports	**Weight in the index**
Platinum	12	0.34
Gold	10	0.29
Coal	7	0.20
Iron	6	0.17
Aggregate	35	

Table A2 Breakdown of exports by major commodity

Year	**Total Exports (USDb)**	**Platinum**		**Gold**		**Coal**		**Iron Ore**		**Aggregate**
		Exports (US$b)	**%age of exports**	**Exports (US$b)**	**%age of exports**	**Exports (US$b)**	**%age of exports**	**Exports (US$b)**	**%age of exports**	
2005	46,99	5.32	11%	4.25	9%	3.27	7%	0.942	2%	**29%**
2006	52,60	8.01	15%	5.10	10%	3.13	6%	1.16	2%	**33%**
2007	64,02	9.82	15%	9.47	15%	3.37	5%	1.60	2%	**38%**
2008	73,97	9.80	13%	5.88	8%	4.76	6%	2.40	3%	**31%**
2009	53,86	6.77	13%	6.26	12%	4.20	8%	3.14	6%	**38%**
2010	71,48	9.33	13%	8.52	12%	5.54	8%	5.46	8%	**40%**
2011	92,98	10.99	12%	10.37	11%	7.52	8%	9.01	10%	**41%**
2012	98,87	7.93	8%	8.66	9%	6.79	7%	7.75	8%	**31%**
2013	95,11	8.41	9%	6.61	7%	5.83	6%	8.46	9%	**31%**
2014	90,61	6.50	7%	4.73	5%	5.19	6%	6.74	7%	**26%**
Average			**12%**		**10%**		**7%**		**6%**	**35%**

Source: UN Comtrade database; author's own computations

4. Copper Prices and Financial Markets in Zambia

4.1 Introduction

The impact of copper prices on the Zambian economy has been a subject of considerable research in the development economics literature in the past three decades. Research themes include a general interest in the impact of commodity prices on small developing economies (for example McBean, 1966, Coppock, 1962, Nziramasanga and Obidegwu, 1980); investigation of the resource curse and Dutch disease narratives (for example Crain, 2010); mismanagement of mineral dependence (Auty, 1991); impact of copper prices on the exchange rate (examples include Cashin et al, 2004, Bova, 2009, Mungule 2004, Chipili, 2010, Weeks etal, 2007, IMF 2009), and more recently, the commodity based industrialisation narrative (Weeks and Mungule, 2013, Morris and Fessehaie, 2014).

It is not difficult to see why economists and policy makers in Zambia are concerned with the behaviour of international copper prices versus various economic variables and indicators. Copper and cobalt account for 78% of the country's exports and are clearly the major source of foreign currency earnings (Central Statistics Office, 2015). The near collapse of the sector in the late 90s which was a result of a combination of low copper prices and mismanagement of copper revenues by the Government has been followed by rejuvenation in the past decade. The dynamics of the international copper market therefore raises concerns for policy makers in three areas of macro-economic management namely fiscal policy, monetary policy and exchange rate policy (Weeks, 2013).

Other issues of concern about Zambia have centred on issues of the country's comparative advantages. The abundance of copper raises questions associated "dependence" problem on copper. The World Bank (2001) raised concerns that countries that have this problem may have their growth stagnate in the future as they cannot join the new economy of knowledge and technology. Other issues have centred around the harmful effects that the volatility of copper prices in the international markets may have on the local economic cycle including the Dutch disease, Meller and Simpasa (2011).

In April 2005, Zambia achieved one of “the landmarks in the history of Zambia” (Minister of Finance Budget speech, 2006) by reaching the completion point for the Highly Indebted Poor Countries (HIPC) initiative. Mungule and Weeks (2013) observe that the Zambian Kwacha started appreciating after it became public news that the country had qualified for debt cancellation and this trend continued thereafter. This period coincided with a strong recovery of copper prices and significant inflows into the Zambian capital markets. The economic reforms of the 1990s had given birth to the Lusaka Stock Exchange (LUSE) in 1994. Thus, in addition to the traditional money markets, the economy had an alternative investment destination in the equities exchange. This chapter investigates the short-run and long-run association between copper prices and other financial market variables in Zambia, namely the nominal exchange rate, equity prices and short term interest rates in a cointegrating vector error correction model with a focus on the post debt cancelation period, 2004 to 2014.

In spite of the wide interest in the impact of copper prices on the Zambian economy, we were unable to find previous research that considers the simultaneous interactions of copper prices and financial market variables considered in this study. This chapter attempts to fill this gap. We employ a Vector Error Correction (VECM) framework that treats all variables in the model as a priori endogenous. This approach imposes a minimal set of theoretical restrictions on the model being tested thereby allowing a close to pure statistical analysis of the variables under consideration. Further, the VECM framework overcomes Sim’s (1980) critique that some exogeneity assumptions for some variables in simultaneous equation models are not backed by fully developed theories. The framework allows the variables in the system to interact with themselves and with each other without having to impose a theoretical structure, except that the relationships are assumed to be linear. To examine the out-of-sample performance of the model, we analyse impulse response functions and forecast error variance decompositions.

Now we briefly outline the results of the study. We find that all the variables in the model are integrated I(1) processes; the Johansen cointegration method indicates one cointegrating relationship among the variables. The VECM error correction estimates suggest when the system is out of its long-run equilibrium, the adjustment of the system back to long run equilibrium is borne by corrections to the short term interest rates. In the short run, exchange rates and equity prices appear weakly exogenous to the system – a result confirmed by our out-of-sample analysis. Further, we show that changes in copper prices lead changes in the other financial market prices in the short term. In the long-run, the relationship between copper prices and short-term rates appears strong and significant. The model passes tests of misspecification and structural stability.

The results are of interest to policy makers, financial market participants in Zambia and international organisations such as the IMF.

The remainder of the essay is organised as follows. In the next section we consider the theoretical connections between commodity prices and financial market variables. Data and econometric methodology are presented in Section 4.3, empirical findings in Section 4.4 and Section 4.5 concludes.

4.2 Commodity Prices and Financial Markets

Theoretical links between commodity prices and macro-financial variables are well documented in the literature. Here we provide a brief discussion of theoretical channels through which commodity price shocks can be transmitted to financial markets.

The linkage between exchange rates and commodity prices is well documented in the economic literature (see Clements and Fry, 2008, Chen and Rogoff, 2003, Chen, 2002, and Ndlovu, 2011) for a comprehensive review of literature on the relationship between commodity prices and commodity currencies). Literature discusses two channels namely the trade in goods channel and the "portfolio balance" class of models.

Under the trades in goods channel, consider a small open commodity exporting economy with tradable and non-tradable goods sectors. An increase in the price of the exported commodity in world markets would affect the demand for non-traded goods through its effect on wages – a channel similar to the well documented Balassa-Samuelson effect.[45] Assuming that prices of non-tradable goods are sticky, the exchange rate instead of prices would have to adjust to preserve efficient resource allocation. Thus, a positive terms of trade shock such as a boom in commodity markets eventually leads to an appreciation of the exchange rate in an environment of nominal price rigidities in the non-tradable sector.

Under the portfolio balance model, the exchange rate is treated as a function of demand and supply of national assets; domestic and foreign assets are treated as perfect substitutes. For a commodity exporting economy, a boom in the price of the exported commodities in international markets would typically lead to an excess supply of dollars and accumulation of foreign reserves, increasing pressure in the relative demand of their domestic currencies. To equilibrate the demand for the domestic currency, the price of the domestic currency would have to appreciate in terms of the foreign currency.

The link between commodity prices and stock markets comes from the input costs and activities of financial investors. A positive shock in the price of strategic commodities like oil

[45] The Balassa-Samuelson effect owes its name to two economists Balassa (1964) and Samuelson (1964). The model posits that faster productivity in tradable versus non-tradable goods in a given economy compared to international counterparties would eventually raise the price level and therefore the real exchange rate. The model assumes that labour is an important factor of production and is fully mobile across the tradable and no-tradable goods sectors. A rise in productivity of tradable goods will raise wages in the tradable sector. Since labour is assumed to be perfectly mobile across the two sectors, the wages in the non-tradable sector would also rise. Producers in the non-tradable sector would have to raise prices to match higher labour costs since the rise in wages is not matched by increased productivity.

for example can be expected to feed into the production costs (through the spill-over of higher energy costs, see Lombardi et al. 2012) in a net oil importing economy. Assuming that prices are sticky in the goods market to accommodate the oil shock, an increase in the price of a key production input would most likely negatively affect investor expectations of future earnings and therefore stock market prices. While correlations between commodity prices and stock markets have not been strong until the 2008 financial crisis (Lombardi and Ravazzolo, 2016, forthcoming), increased financialisation of commodity markets justifies another link between the two asset markets. Gao and Süss (2015) report that 80% of all investors in commodity markets are financial investors. Investment in commodities provides an alternative to stock markets. Cassassus and Higuera (2011) provide evidence that oil prices are good predictors of equity prices. In a recent study, Lombardi and Ravazzolo, (2016, forthcoming) shed more light on this link. The authors indicate that joint modelling of the evolution of prices in the two asset markets is valuable from an asset allocation point of view.

The connection between commodity prices and short-term interest rates (a monetary policy variable) is well developed in literature. Frankel (2008) presents three theoretical channels through which interests rates can interact with prices of storable commodities. The first link relates to the activities of financial speculators in commodity markets. All things being equal, low real interest rates are likely to encourage speculators to shift from treasury bills investments towards financial markets for commodities.[46] The second link comes from the extraction decision of commodity producers. To arrive at the decision to extract the commodity "now or later", a producer has to consider a number of factors including cost of storage and risk and opportunity cost of holding commodity inventories. An environment of high real interest rates is likely to encourage producers to extract the commodities now (which are non-interest earning), liquidate them and earn high interest from the proceeds of

[46] See also Gao and Süss (2015) for a discussion on the financialisation of commodities

the sale. Thus high money market rates are likely to increase supply of commodities and exert downward pressure on their prices. The third link comes from the cost of carrying inventories. For a primary commodity producer or processor of the raw materials, the decision to carry inventory is likely going to be affected by costs of storage including real interest rates. If the interest rates are high, it is likely that everyone would like to hold lower inventories, thus lowering demand and commodity prices, ceteris paribus. These theoretical links have been tested in the literature (see Frankel, 2008; Grubber and Vigfusson, 2012, Akram, 2009). In commodity exporting countries, policy makers may introduce countercyclical measures against commodity price declines resulting in increased interest rate volatility (Hegerty, 2016).

The previous discussion on the relationship between commodity prices, interest and exchange rates suggests that interest rates and exchange rates would move in opposite directions in response to a commodity price shock. The relationship departs from one implied by the traditional workhorse open macroeconomic models such as the Mundell-Fleming (MF) model (see Mundell 1962, 1963, Fleming 1962, Dornbusch 1976, Frankel and Rose 1995, Taylor 2002). The MF model, based on the assumptions of perfect capital mobility and interest rate parity implies that interest rates and nominal exchange rate move in the same direction. According to Makin (2013), the traditional open macro-economic models "neglect" the commodity currency phenomenon. In a commodity exporting economy with a "commodity currency", prolonged commodity price booms which result in persistent appreciation of the exchange rate may result in fall in national output. Persistent appreciation of the exchange rate is disruptive to production in the non-commodity exporting sectors of the economy in the long run. For this reason, an optimal monetary policy response should be one that limits the appreciation of the exchange rate, that is, a contractionary policy stance. We now test these conjectured links for Zambia.

4.3 Data and Econometric Methodology

4.3.1 Data Description and Measurement

The study uses monthly data for the sample period January 2004 to December 2014 for copper prices (p_c) and three financial market variables namely, the Zambian Kwacha exchange rate (*NEER)*, the three month Treasury bill rate (*TBILL*) and the Lusaka Stock Exchange Index (*LSEI*). Copper is an industrial metal traded on the London Mercantile Exchange (LME) and constitutes nearly 70% of Zambia's exports with main destinations being Switzerland (61.1%) and China (25.6%).[47] The monthly average spot price is used in this study. The data series was extracted from the IMF database and is measured in USD/Metric tonne.[48] The Zambian Kwacha exchange rate is measured as the Nominal Effective Exchange Rate Index (NEER, 2010=100) published by the IMF. The nominal effective exchange rate (NEER) of the Zambian Kwacha is calculated as the geometric weighted average of a basket of bilateral nominal exchange rates taking into account Zambia's largest international trade partners. An increase in the index (above 100) indicates appreciation of the Zambian Kwacha relative to the United States dollar. Nominal variables are preferred in order to capture the asset price bubbles in the data. The equity index is the Lusaka Stock Exchange main board index. The index (2010=100) was extracted from the CEIC database and measures the monthly average value of the equities index.[49] The measure of short term interest rate is the three month Treasury bill rate published by the Bank of Zambia measured in percent, per annum. The 91 day Treasury bill rate implicitly reflects the stance of monetary policy in Zambia (see Fundanga 2009). To enable us to interpret the estimated coefficients as elasticities, all the data series with the exception of the Treasury bill rate are transformed into natural logarithms.

[47] Republic of Zambia Central Statistical Office, 2015
[48] www.imf.org
[49] http://webcdm.ceicdata.com

Table 4.1: Data descriptive statistics

	Copper Price	*Equity Index*	*T-bill Rate*	*Exchange Rate*
Mean	6458.8521	102.3998	12.0781	106.1978
Median	7049.9751	104.4350	12.0900	100.6750
Maximum	9880.9400	210.6600	19.1300	150.2000
Minimum	2421.4804	15.5800	2.8800	78.4500
Std. Dev.	1957.5931	52.0989	3.4834	16.2338
Skewness	-0.6383	0.1372	-0.1516	1.0421
Kurtosis	2.2504	2.3419	2.3896	3.3090
Jarque-Bera	12.05604	2.7955	2.5545	24.4186
Observations	132	132	132	132

Summary descriptive statics are provided in Table 4.1. The statistics clearly show that with the exception of copper prices, financial market variables in Zambia exhibit large kurtosis indicating extreme market movements on either direction (gains or losses), consistent with the observation of Bhar and Hammoudeh (2011). Kurtosis on the exchange rate is higher than 3, indicating that the Kwacha may be mispriced in relation to other currencies. Weeks and Mungule (2013) attribute mispricing of the Kwacha to the volatility smoothing activities of the Bank of Zambia. The Jarque-Bera statistic rejects equality of distribution hypothesis in the data. This behaviour of variables under study may be as a result of non-linearity in business cycle fluctuations (Tulley and Lucey, 2007).

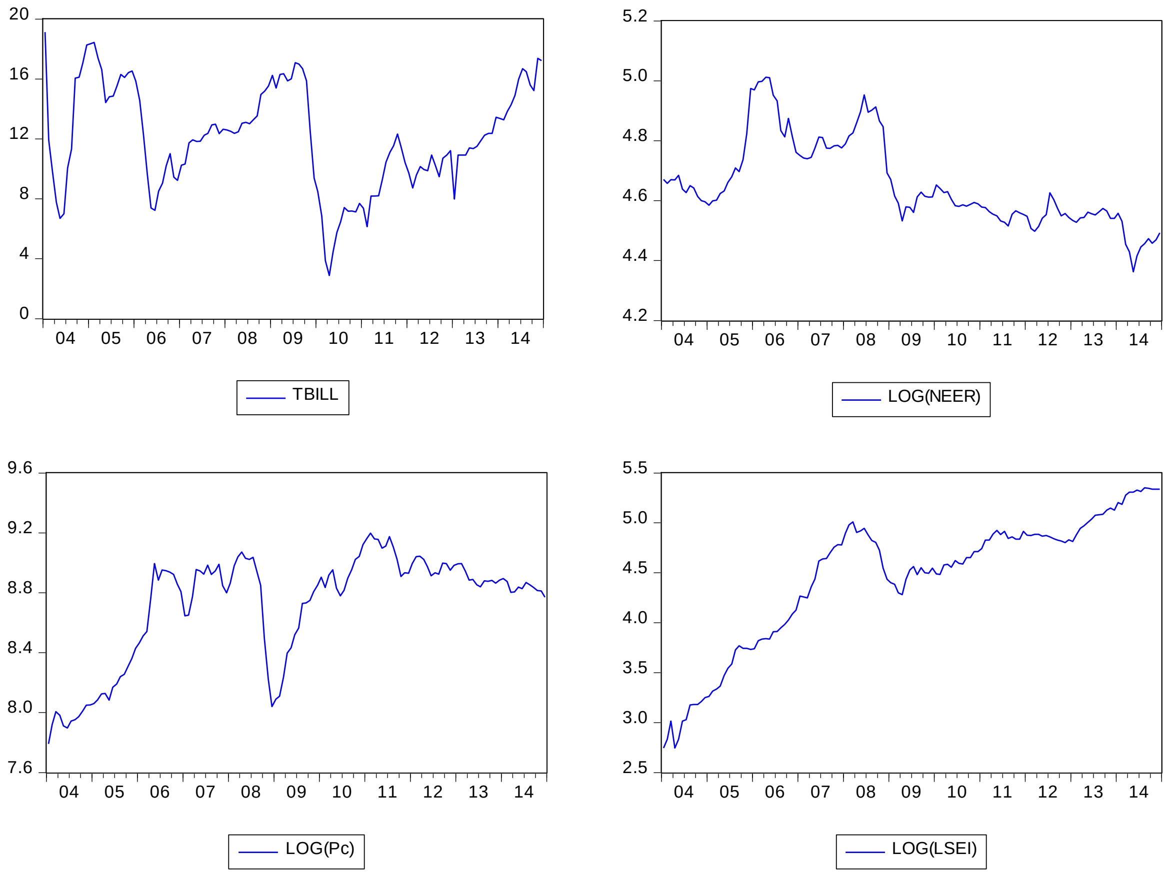
The data plots in Figure 4.1 show the relevant financial time series and the copper price. The series are possibly non-stationary and cointegrated. We formally carry out tests for unit root.

4.3.2 Unit Root Tests

Unit root processes are not uncommon in financial time series analysis. A unit root or a nonstationary time series will have a time-varying mean or a time-varying variance or both. Most economic time series exhibit trending behaviour or non-stationarity in their mean. Prominent examples include asset prices, exchange rates and macro-economic aggregates like Gross Domestic Product (GDP).

There are several methods of examining the stationarity properties of time series data that have now become standard in econometrics. We employ the Augmented Dickey Fuller (ADF) tests (Dickey and Fuller, 1979, Dickey and Fuller, 1981) to examine the stationarity properties of the data. Following Doldado, Jenkinson and Sosvilla-Rivero (1990), the ADF test is implemented starting with the least restrictive of plausible models which includes a trend and a constant. An alternative model including only the constant is tested as well. A similar test is conducted on first differences of the time series data. The Phillips-Perron (Phillips and Perron, 1988) method is added to test the robustness of the ADF test results. The test results can be seen in Table 4.2.

Figure 4.1: Plots of copper prices and financial market variables



The results indicate that the null hypothesis cannot be rejected on all level variables irrespective of the constant and trend assumptions. However the same tests performed on first differences suggest rejection of the null hypothesis at all levels of significance and thus verify that the variables are $I(1)$ processes. The Philips-Perron tests also confirm the ADF results indicating that the series in the model are best described as difference stationary. This implies that the series have long memories, that is, random shocks to the series have persistent effects into the distant future. All the series are therefore treated as first-difference stationary in our model.

4.3.3 Test for deterministic trends

The results of a VAR/VECM specification and the related asymptotic inferences on Granger causality or neutrality are sensitive to the presence of deterministic trends. Stock and Watson (1989) note that finite distributions of tests using time series data not only depend on unit root but on deterministic trends as well. Having ascertained the stationarity properties of the data series, we proceed to test for the presence of deterministic trends in the data. We follow the methodology developed by Stock and Watson (1989).

Table 4.2: Tests for Unit Root

Variable	**ADF TEST STATISTIC**		**PHILIPS PERRON TEST**	
Levels	**Constant with trend**	**Constant without trend**	**Constant with trend**	**Constant without trend**
Copper Prices (p_c)	-2.4895	-2.4710	-2.2758	-2.5553
Exchange rate *(NEER)*	-2.5891	-2.7831	-2.3652	-2.5082
Interest Rate *(TBILL)*	-2.0926	-2.1040	-2.8905	-2.9835
Equity Index *(LSEI)*	-2.5572	-2.4925	2.4944	-2.1538
First Differences				
$\Delta(p_c)$	-7.2606	-7.2158	-7.2246	-7.2276
Δ*(NEER)*	-9.2398	-9.1613	-9.4167	-9.3651
Δ*(TBILL)*	-8.5749	-8.6262	-8.5375	-8.5895
Δ*(LSEI)*	-9.2035	-8.8437	-9.6176	-9.3489

Notes: The Critical values for rejection are -4.0296, -3.4444 and -3.1471 at a significant level of 1%, 5% and 10% respectively for models with a constant and linear trend and -3.4812, -2.8830, -2.5787 at a significant level of 1%, 5% and 10% respectively for models without a linear trend. The optimal lag for the ADF test was chosen based on the Schwartz Information Criterion and the truncation parameter for the PP test was selected using the Newey-West truncation method.

To ascertain the order of the deterministic trends of the data, the first step involves checking for trends in levels. The level time series data is regressed against sufficient lags of itself, and a constant.[50] The second step involves rerunning the regression with a deterministic trend added. The test is essentially a $t-$test for statistical significance of the constant/drift term and the trend term in the equations. The test is repeated for first differences of the data. A statistically significant $t-$value indicates the presence of a drift term and/or constant. This simple test is conducted for all variables to be tested in the VECM model. The test results are presented in Table 4.3.

The test results indicate that the level data series have a constant term but no deterministic trend. The first differences, with the exception of equity prices, have neither a drift term nor a linear time trend. The level data series are therefore best described as having unit root and a drift term and that the first differences have neither drift nor time trend.

Table 4.3: Test for deterministic trends

Variable	Lag Length	Constant (t-statistic)	Trend(t-statistic)
Copper Prices (p_c)	2	2.8508***	0.9464
Exchange rate *(NEER)*	2	3.1164***	0.2571
Interest Rate *(TBILL)*	2	2.8195***	0.5710
Equity Index *(LSEI)*	2	4.1156***	1.0658
$\Delta(p_c)$	1	0.5141	0.8762
Δ*(NEER)*	2	0.2871	0.1947
Δ*(TBILL)*	1	1.1091	1.1271
Δ*(LSEI)*	1	3.0773*	2.1016*

*Notes: the reported t-statistics are for the coefficient of a constant and a coefficient of a trend component included in the autoregressive regression of a variable in levels and first differences. The optimal lag lengths are selected according to the AIC information criterion. *, ** and *** indicate significance at 10%, 5% and 1% levels of significance respectively. Δ is the first difference operator.*

4.3.4 Test for Structural Stability

In testing the relationships between economic time series, it is important to test if the variables are subject to similar processes. Structural change in time series obtains from

[50] Sufficient here is defined as the optimum number of lags as determined by the AIC criterion for each of the data series. Additional lags of up to 8 were tested for each series with no material change in the results. Stock and Watson (1989) carry their tests on six lags on output, money growth, inflation and interest rates.

external forces such as significant economic crises, changes in policies such as exchange rate regimes and wars. For our time series, we suspect that the 2008 sub-prime mortgage crisis may have caused a structural change in the variables under study. Chen, Rogoff and Rossi (2010) emphasise the importance of controlling for structural breaks in analysing the causality between exchange rates and fundamentals. Mirer (1988) argues that the data generation process in time series can only be the same if the relationship is stable.

Enders (2010) suggests that if it is reasonable for a researcher to suspect a structural break, it is straightforward to implement a Chow (1960) test. To implement the Chow test, we identify June 2009 as the break date. [51] The VECM model is then fit for the pre-break and post-break data. If the estimates of the models (pre and post-break) are sufficiently different from each other, then it is possible that there is a structural break in the data.

The first sub-sample therefore starts on January 2004 to June 2009 and the second sub-sample starts from July 2009 to December 2014. We estimate the VECM for the full sample period and for the two subsamples and use the F-test to determine if the models from the two subsamples are significantly different from each other. The null hypothesis is that the coefficients of the model from both subsamples are the same and thus lead to a restricted regression. The alternative hypothesis is that the model estimates from the two subsamples are different, thus relaxing the restriction imposed on the null hypothesis. Denoting the sum of squared residuals by RSS for the unrestricted model and RSS_1 and RSS_2 for the first and second subsamples respectively, it can be shown that the Chow test follows an F distribution such that:

$$F = \frac{(RSS - RSS_1 - RSS_2)/n}{(RSS_1 + RSS_2)/(T - 2n)} \tag{4.1}$$

[51] It is reasonable to expect the Zambian economy to have fully assimilated the effects of the October 2008 collapse of Lehman Brothers investment bank in the United States, the subsequent sell-off of emerging market assets and policy responses by the authorities in Zambia.

where n = number of parameters estimated and the number of degrees of freedom are $(n, T - 2n)$.

The test says that if the estimates of the model for the two subsamples are the same, then

$RSS_1 + RSS_2 = RSS$ and therefore F must be equal to zero. However the larger the calculated value of F, the more restrictive is the assumption that the coefficients are equal (Enders 2010:104). The results of the test can be found on Table 4.4.

Table 4.4: Test of Structural Stability

	RSS Full Sample 2004m1-2014m12	*RSS_1* Sub-sample 1 2004m1-2009m6	*RSS_2* Sub-sample 2 2009m7-2014m12	*F-Statistic*
Copper Prices (p_c)	0.5919	0.4425	0.1271	1.1974
Exchange rate *(NEER)*	0.1299	0.0921	0.0372	0.1370
Interest Rate *(TBILL)*	1.7701	0.5777	1.0786	2.0967
Equity Index *(LSEI)*	0.3261	0.2064	0.0995	2.0115

Notes: The null hypothesis is that the coefficients of the structural relation are the same in Sub-samples 1 and 2, that is, there are no structural breaks in the series. The alternative hypothesis is that the corresponding coefficients of the two sub-samples are different. If the F-statistic exceeds the critical value, then the null hypothesis of non-existence of a structural break is rejected. The critical values for rejection of the null hypothesis are 3.48 and 2.45 at 1% and 5% level of significance respectively

The results suggest that there is no structural break in our sample period at the 5% level of significance or better. We conclude therefore that the relations among the variables are stable over the whole period and proceed to estimate the VECM over the full sample period.

4.3.5 The Econometric Procedure

The literature points to several frameworks that have been developed for estimation and inference in cointegrating systems. The popular methods can be found in Gonzalo (1994). The estimation methods include ordinary least squares (OLS) by Engle and Granger (1987), non-linear least squares (NLS) by Stock (1987), the principal components (PC) method by

Stock and Watson (1988), canonical correlations (CC) by Bossaerts (1988), instrumental variables (IV) by Hansen and Philips (1990), spectral regression (SR) by Philips (1991) and maximum likelihood in a fully specified error correction model (MLECM) by Johansen (1988).

Much of empirical work involving single equation error correction methods was proposed by Engle and Granger (1987) and is based on simple OLS regression of level variables. Inferences can also be made in a straightforward manner using test statistics with appropriate asymptotic distributions. The method has a number of limitations including its inability to estimate models with more than one cointegrating vector. Gonzalo (1994) concludes that the best way to estimate cointegrated systems is full system estimation by maximum likelihood, incorporating all prior knowledge about presence of unit roots. The maximum likelihood procedure developed by Johansen (1988, 1991, 1995) ensures that the estimation results of an error correction specification are symmetrically distributed, median unbiased and asymptotically efficient. Further, the hypothesis tests may be conducted on the cointegrating matrix and the loading matrix using standard asymptotic chi-squared tests. Gonzalo, (1994) finds that the maximum likelihood approach outperforms the other methods in his study even when the errors are non-normally distributed or when the dynamics at play in the data are unknown. Other recent studies that confirm the efficiency of the Johansen maximum likelihood method include Hubrich, Lütkepohl and Saikkonen (2001). Accordingly, the Johansen ML method is used to estimate the VECM in this study.

Johansen (1995) shows that relationships among non-stationary time series variables that are cointegrated of rank $r,$ can represented as a multivariate Vector Error Correction model (VECM). An unrestricted VAR model of order p with $(kx1)$ endogenous variables all integrated of order one, forced by a vector of $(kx1)$ independent Gaussian errors can be formulated as a VECM of the form:

$$\Delta y_t = \mathbf{\Pi} y_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta y_{t-i} + \gamma D_t + e_t \qquad (4.2)$$

Where y_t is a $(k\ x\ 1)$ vector of dependent (endogenous) variables; D_t is a vector containing constants and dummies; $\mathbf{\Pi}$ is a (kxk) matrix, decomposed as $\alpha\beta'$ with matrices α and β, dimensioned (kxr) capturing the speed of adjustments and long-run relations respectively; $\mathbf{\Gamma}_i(i = 1, \dots, p-1)$ are (kxk) parameter matrices capturing the short run dynamics among the variables and e_t is a $(k\ x\ 1)$ vector of disturbances, has mean 0 and covariance matrix $\mathbf{\Omega}$ and is independent, identically distributed (i.i.d.) over time.

If the matrix $\mathbf{\Pi}$ in equation (2) has rank $\mathbf{0} \leq \boldsymbol{r} < \boldsymbol{k}$ where r is the number of linearly independent cointegrating vectors, then there exists a linear combination of y_t that is stationary.

Substituting $\alpha\beta'$ into (4.2), the VECM can be re-written as:

$$\Delta y_t = \boldsymbol{\alpha\beta}' y_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta y_{t-i} + \gamma D_t + e_t \qquad (4.3)$$

The Johansen (1995) estimation approach enables us to conduct hypotheses tests on the cointegrating matrix β and the loading matrix α using standard asymptotic chi-squared tests.

In the VECM framework the characteristics of the data series determine the exact specification of the model. This framework has been the workhorse of time series econometrics since Sim's (1980) seminal article. The approach treats all variables in the model as a priori endogenous. This procedure has at least two distinct advantages. First, the framework imposes a minimal set of theoretical restrictions on the model being tested thereby allowing close to pure statistical analysis of variables under consideration. Second, the VECM framework overcomes Sim's (1980) critique that some exogeneity assumptions for some variables in simultaneous equation models are not backed by fully developed theories. Theoretical models developed using industrialised economies assumptions often perform dismally when tested using developing economies data.[52] To answer the questions of this study, we argue that the VECM framework is suitable for analysing dependencies among financial assets and copper prices in Zambia as it allows the variables in the system to interact with themselves and with each other without having to impose a theoretical structure, except that the relationships are assumed to be linear.

Employing the Johansen (1995) maximum likelihood method, the estimates of the $\boldsymbol{\beta}$ matrix equal the matrix of eigenvectors corresponding to r largest eigenvalues (or characteristic roots) of a residual matrix. To obtain the number of distinct cointegrating vectors, one has to check the significance of the resulting eigenvalues. A cointegration test for the number of eigenvalues that are insignificantly different from unity is done using two test statistics called the trace and maximum eigenvalue statistics. We use the trace and maximum Eigen value statistics to determine the rank (r) of the cointegrating matrix $\mathbf{\Pi}$. The two tests can be written in terms of Eigen values (λ_i) (see Johansen 1995 for details). For a sample size (T):

$$\lambda_{trace}[\mathcal{H}_1(r)|\mathcal{H}_0] = -T\sum_{i=r+1}^{p}\ln(1-\hat{\lambda}_i) \tag{4.4}$$

Where $\hat{\lambda}_i$ is the estimated eigenvalue and the null hypothesis is $\lambda_i = 0$, so only the first r eigenvalues are non-zero.

$$\lambda_{max}[\mathcal{H}_1(r+1)|\mathcal{H}_1(r)] = -T\ln(1-\hat{\lambda}_i) \tag{4.5}$$

For $r = 0, 1, 2.., p - 2, p - 1$. The null is that there exists r cointegrating vectors against the alternative of $r + 1$ vectors.

For the short-run dynamics, we conduct Granger (1969) causality tests. While the existence of cointegration among the variables can be interpreted as long-run causality, the cointegrating relation does not indicate direction of causality. [53] A VECM derived from long-run cointegrating vectors can be used to test Granger causality and therefore whether an endogenous variable can be treated as exogenous (that is, determined outside the system) within a sample period in the short-run (Masih and Masih, 1996).

In addition to the foregoing "within-sample" analysis, we conduct "out-of-sample" analysis of the four variables using Impulse Response Functions (IFR) and Forecast Error Variance Decompositions (FEVDs).

Impulse-response functions (IRFs) measure the dynamic marginal effects of each shock on all of the variables in a VAR system over time. Defined more formally, IRFs gives the j^{th} period response when the system is subjected to a one standard deviation shock. The general representation of the IRF is widely available and can be found in the seminal works of Sims (1980), Bernanke (1986), and Sims (1986).

Forecast error variance decompositions (FEVD), separate the variation in an endogenous variable into the component shocks to the VAR. In effect, variance decompositions indicate the importance of each random innovation in influencing the variables in a VAR/VECM system. FEVDs can be termed out of sample causality tests because they indicate the proportion of movements in a sequence due to a variable's own shocks versus shocks to the other variables (Enders, 2010:314). If shocks in y_{2t} explain none of the forecast error variance in y_{1t} at all forecast horizons, then variable y_{1t} can be said to be exogenous to the system and therefore y_{1t} evolves independently of the y_{2t} shocks. On the other hand if y_{2t} shocks explain all the forecast error variance in y_{1t}, at all forecast horizons, then the variable y_{1t} is said to be entirely endogenous to the system. Sims (1980) notes that a

[53] See Granger et al, 2000

variable that is optimally forecast from its own lagged values will have a forecast error variance accounted for by its own disturbances. Most economic variables will typically have their own shocks explain almost all of their forecast error variance in short horizons and only smaller proportions in longer horizons. Thus, while the IRFs show the effects of shocks on the adjustment path of the variables in the system, FEVDs measure the contribution of each type of shock to the forecast error variance. Both analyses help in analysing how shocks to variables reverberate through the system.

IRFs and FEVDs are obtained from the moving average representation of the VAR. This way, the analysis overcomes the problems associated with the restrictive assumption that the shocks that occur to the system are independent. By construction, the residual errors from the equations in a VAR framework are usually serially uncorrelated and therefore the shocks can be treated as independent. However it is possible to have contemporaneous correlations across errors of different equations and in this case it is possible that a shock in one variable is accompanied by a shock in another variable. In this case, setting all the other errors to zero may provide a misleading picture of the dynamic relationships among variables.

The innovations from the equations in a VAR can be orthogonalised through the decomposition such that y_{2t} does not have a contemporaneous effect on y_{1t} and the resulting covariance matrix of innovations is diagonal and the IRFs can be interpreted. The decomposition forces a potentially important asymmetry to the system and separates the estimated residuals from the reduced form representation of the structural model into orthogonal shocks by restrictions imposed on the basis of an arbitrary ordering of the variables. The decomposition implies that a one unit shock to the first variable causes the value of the first variable to increase without causing a contemporaneous increase to the other variables in the system. In an autoregressive system, the decomposition implies that the first variable responds only to its own exogenous shocks, the second variable responds to the first and the second variable's exogenous shocks and so on.

In the spirit of Sim's argument against imposing "incredible identifying restrictions" to our model, we use the Choleski decomposition which has been shown to impose a minimal set of identification assumptions (Enders, 2010).

Under the Choleski decomposition, the results of IRFs and FEVDs have been shown to be sensitive to the ordering of variables (Kamas and Joyce, 1993). The importance of ordering depends mainly on the correlation of residuals. Enders (2010:310) shows that if the correlation coefficient is found to be zero among residuals, ordering of variables becomes irrelevant.

We follow Gordon and King (1982) who suggest that the variables that respond most to current events such as interest rates and nominal exchange rate be placed last in the ordering so that their values reflect contemporaneous realisation of the variables of a higher order. Thus the innovations in our model were orthogonalised in the following order: $Pc \rightarrow NEER \rightarrow LSEI \rightarrow TBILL$. This analysis is conducted over a 24 month period.

There are a number of researchers that analyse the impact of copper prices to the Zambian economy. Most of them however are concerned with the link between copper prices and the exchange rate. Notable contributions include single cointegration models of Cashin et al (2004), Bova (2009), generalised autoregressive conditional heteroscedasticity model (GARCH) model of Chipili (2010) and the multivariable structural models of Weeks and Mungule (2013). The evidence of whether the Zambian Kwacha is a "commodity currency" is mixed.

4.4 Empirical Findings

4.4.1 Lag-length selection

The accuracy of a VECM specification depends on the choice of lag lengths in the model. Several authors including Braun and Mittnik (1993), Hafer and Sheehan (1989) and Lütkepohl (1993) demonstrate that estimates of a VAR whose lag length differs from the *true* lag length are inaccurate as are the resulting impulse response functions and variance decompositions. Overfitting a lag order that is, selecting a higher lag length order than the true lag length, causes an increase in the mean square forecast errors of a VAR while under fitting the lag length generates autocorrelated errors.

Several length selection criteria have been developed in the literature to assist researchers to fit models of correct lag order. The most popular ones include the Akaike's (1974) information criterion (AIC), Schwarz Bayesian information criterion (SBIC), and the Hannan and Quinn information criterion (HQIC) and the Log Likelihood (LR) method. In this study, the appropriate lag lengths for the estimations are chosen based on the log likelihood (LR) criterion, preferred by Sims (1980).[54]

Table 4.5 : Lag length selection

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-262.9845	NA	0.0010	4.4494	4.5426	4.4874
1	408.7962	1287.5800	1.80e-08	-6.4799	-6.0153*	-6.2912
2	437.9863	54.0018	1.45e-08*	-6.6997*	-5.8635	-6.3601*
3	448.8770	19.4216	1.58e-08	-6.6146	-5.4067	-6.1240
4	462.9525	24.1629*	1.64e-08	-6.5825	-5.0029	-5.9410
5	469.2454	10.3833	1.94e-08	-6.4207	-4.4695	-5.6283
6	488.6550	30.7319	1.85e-08	-6.4775	-4.1546	-5.5342
7	499.8967	17.0497	2.03e-08	-6.3982	-3.7037	-5.3039
8	515.1246	22.0805	2.10e-08	-6.3854	-3.3191	-5.1401
9	536.1299	29.0573	1.98e-08	-6.4688	-3.0309	-5.0726
10	554.0419	23.5840	1.98e-08	-6.5006	-2.6911	-4.9536
11	559.5883	6.9330	2.45e-08	-6.3264	-2.1452	-4.6284
12	582.2713	26.8416	2.30e-08	-6.4378	-1.8849	-4.5889

** indicates lag order selected by the criterion; LR: sequential modified LR test statistic (each test at 5% level); FPE: Final prediction error; AIC: Akaike information criterion; SC: Schwarz information criterion; HQ: Hannan-Quinn information criterion*

54 The lag length selected by the LR criterion eliminates all serial correlation in the residuals in our model.

In our tests for lag length we specify 12 maximum lags to control for any seasonality in our monthly data. The Log Likelihood (LR) information criterion in Table 4.5 indicates the optimum lag as $p = 4$.

4.4.2 The Johansen Test of Cointegration

The results of the Johansen (1995) test for cointegration are displayed in Table 4.6. Both the trace and maximum eigenvalue statistics are significant and reject the null hypothesis of no cointegration at the 5% level of significance and indicate the presence of at most one cointegrating vector.

Table 4.6: Johansen test for cointegration

Hypothesized No. of CE(s)	Trace Statistic	Maximum Eigenvalue Statistic	Critical Values (5%) Trace	Critical Values (5%) Maximum Eigenvalue
$r = 0$	64.8145***	41.3712	47.8561***	27.5843
$r \leq 1$	23.4433	14.1921	29.7970	21.1316
$r \leq 2$	9.2511	7.2128	15.4947	14.2646
$r \leq 3$	2.0383	2.0383	3.8414	3.8414

*Notes: *** denotes significance at 5% level*

4.4.3 Diagnostic Tests

In econometrics modelling, it is standard practice to check the VECM results robustness to residual autocorrelation, heteroscedasticity and normality. This section presents results of stability test, test of residuals normality, autocorrelation and homoscedasticity.

The normality tests of residuals rely on skewness and kurtosis. The Jarque-Bera statistic and chi-squared values for skewness and kurtosis are reported in Table 4.7. The results suggest rejection of the null of residual normality suggesting that most errors are skewed, kurtotic and non-normal. According to Lutkepohl (2011) normality in the distribution of residuals is a desirable but not a necessary condition for the validity of many of the statistical procedures related to VECM and VARs. Juselius, (2003:76; 2006) note that not passing for symmetry and no excess kurtosis in residuals, at least asymptotically has no implications on the validity of either tests or estimators in VECMs.

We perform the Lagrange-multiplier (LM) tests for residual autocorrelation. The Ljung–Box Q-statistics are a popular catch-all test for autocorrelation (Enders, 2010:436). The results in Table 4.7 suggest rejection of the null of residual autocorrelation. Our model is therefore correctly specified with sufficient lags. Equally, the ARCH test indicates that the residuals are homoscedastic. In summary, the model passes the key diagnostic, stability and stationarity tests with the exception of non-normality caused by excess kurtosis in the time series variables. Inferences made from our estimations are therefore valid.

Table 4.7: Diagnostic tests

	LM-Stat(8)	*ARCH* (χ^2)	*Normality*	*Skewness*	*Excess Kurtosis*
Panel A: Single Equation tests					
T-bill	1.0844	0.8691	22.8046	0.0169	5.0675
	[0.3797]	[0.6055]	[0.0000]	[0.0100]	
Copper Price	1.2222	1.3392	66.0665	-0.7290	6.2032
	[0.2932]	[0.2494]	[0.0000]	[0.0155]	
Equity Index	1.0115	0.7428	7.7484	0.0212	4.1777
	[0.4319]	[0.7448]	[0.0274]	[0.0588]	
Exchange Rate	0.6678	1.3508	29.0447	0.3050	5.2521
	[0.7187]	[0.1801]	[0.0000]	[0.0005]	
Panel B: Model tests					
	10.7018	260	76.6020	4.8875	71.7152
	[0.8275]	[0.3225]	[0.0000]	[0.2990]	[0.0000]

Notes: P-values in [ ]

4.4.4 Short-run Dynamics

Granger Causality/Block Exogeneity Wald tests are presented in Table 4.8. While the tests indicate presence of and direction of Granger causality, they do not provide information on the strength of the causal chain among variables. The Wald tests can be interpreted as "within-sample" tests as they do not indicate the variables' exogeneity or endogeneity outside the sample period (Masih and Masih, 1996).

The short run dynamics indicate that changes in copper prices precede changes in the three financial market variables. Among the financial market variables themselves, equity prices Granger cause short term interest rates. This result reflects the superiority of information efficiency of copper market over the other financial markets in Zambia. Copper is transacted in highly efficient auction markets and its price reflects demand and supply shocks rapidly. For a country that depends heavily on copper for foreign currency, price signals in the copper market are leading indicators of future direction of balance of payments and impact on the wider macro-economy in the short run.

In the literature, the short-run link between the copper price and the exchange rate has been found by Bova (2009) in an EGARCH framework. The author highlights that the volatility in the copper market is related to volatility of the exchange rate in the short run. With respect to equity prices, a positive shock to copper prices reflects an improvement in the future dollar incomes of copper exporters and the government (via rents like tax and royalties). For a copper dependent economy, investors such as fund managers and speculators can be expected to adjust their expectations of future incomes of companies and therefore bid for their shares thereby inflating equity prices. This asset-price argument can be extended to the yields on Treasury bills. Our results indicate that changes in equity prices precede changes in the money market rates – we surmise that this mechanism operates via the asset substitution mechanism. Short term interest rates seem to “react” to changes in the equities market. For a portfolio manager with a choice of only two assets namely equities and interest bearing instruments (quite plausible for Zambia); money market rates have to adjust to a change in the equity prices to maintain the same asset weights in a given portfolio.

Short term interest rates appear to be unrelated to the exchange rate in the short-run. This result indicates that monetary policy is ineffective in moderating exchange rate volatility in the short run.

Table 4.8: Granger causality tests

	Independent Variable			
	Δ(ln_cu)	Δ(ln_luse)	Δ(tbill)	Δ(ln_zmw)
Dependent variable	χ^2 Statistic			
Δ(ln_cu)	-	3.4069	2.2090	1.6291
Δ(ln_luse)	6.1850***	-	3.4268	0.2260
Δ(ln_tbill)	3.7885***	5.4845*	-	4.7880
Δ(ln_zmw)	5.9470**	0.0462	0.4193	-

*Notes: *, ** and *** indicate significance at 10%, 5% and 1% respectively. A significant χ^2 Statistic implies that the independent variable Granger causes the dependent variable.*

4.4.5 The Long-Run Dynamics

The presence of at least one cointegrating vector implies that the vector can be generically identified in the long-run through a normalisation process (Boswijk, 1996, Hunter and Ali, 2014). While the literature generally regards normalisation as innocuous, Bowswijk (1996) suggests that empirical normalisation requires that further rank conditions be satisfied. Burke and Hunter (2005) suggest that any identification procedure should preclude normalisation on variables that are either long-run excluded or weakly exogenous to the system.[55] We follow this reasoning in this paper and therefore, before imposing and identification restrictions, we test the variables in our model for long run exclusion (LE) and weak exogeneity (WE).

The LE tests are conducted by imposing a zero restriction on the elements of the long-run matrix β. If this hypothesis cannot be rejected at the 5% level, the cointegrating vector cannot be normalised on this variable. The WE tests are carried out by imposing zero restrictions on the elements of the adjustment matrix α. Failure to reject the zero restriction null implies that the variable is weakly exogenous to the system, that is, it drives the system instead of adjusting to it. The results of the LE and WE tests are presented in Table 4.9.

[55] In our case where we have a single cointegrating vector, normalization implies rendering the model to be related to a particular dependent variable. It makes sense that for the dependent variable to be endogenous to the system, that is, it cannot be excluded from the system in the short and long run.

Table 4.9: Long-run Exclusion (LE) and Weak Exogeneity (WE) tests

	Log(Copper price)	Log(Equity Index)	T-bill	Log(Exchange rate)
Panel A: LE Tests				
$\chi^2(1)$	23.4290	13.8456	26.0057	4.2475
p-value	0.0000***	0.0001**	0.0000***	0.0393**
Panel B: WE Tests				
$\chi^2(1)$	5.8844	0.0687	17.1194	2.1113
p-value	0.0952*	0.7932	0.0000***	0.1462

*Notes: *, ** and *** indicate significance at 10%, 5% and 1% respectively.*

From Table 4.9 Panel A, it is clear that all four variables cannot be excluded from the long-run equation as they are statistically different from zero. From Panel B however, with the exception of the copper price and the short term interest rate, the proposition that equity prices and the exchange rate are weakly exogenous cannot be rejected. This result implies that the long-run equation cannot be normalised on the exchange rate nor equity prices. Therefore the cointegrating vector is normalised on the short term interest rate as follows (with standard errors in ( ) and t-statistics in [ ]):

$$\begin{array}{lllll} TBILL = & -16.4510(p_c)\ *** & +\ 6.8300(NEER)\ ** & +\ 6.5581\ (LSEI)\ *** + 94.4840 \\ & (1.8381) & (2.7242) & (1.1329) \\ & [8.9497] & [-2.5072] & [-5.7884] \end{array} \tag{4.6}$$

*Notes: *, ** and *** indicate significance at 10%, 5% and 1% respectively.*

By inspection of equation (4.6), we observe that all the long run parameters are statistically significant at 5% level or better. In the long run, short term interest rates are associated with movements of -17%, 6.8% and 6.6% to a 1% change in the copper price, exchange rate and equity prices respectively. Moreover, the short run adjustment parameters indicate that the short term rate is the “slave” of the system, that is, it bears the burden of short run endogenous adjustment to bring back the system to its long-run equilibrium.

This result is important as it reflects the importance of the conduct of monetary policy in bringing the system to long run equilibrium in Zambia.[56] While the ultimate monetary policy target is inflation in Zambia, the BoZ has paid attention to exchange rate volatility in recent years. The BoZ estimates the country's marginal propensity to import at over 40% suggesting a remarkable pass-through rate of the exchange rate changes to the general price level. The actions of the BoZ in achieving exchange rate stability through monetary policy explain why the money market rate is the "forcing" variable of the exchange rate to the long run trend.

This specific mechanism, (discussed by Weeks, 2013) functions through the actions of domestic branches of international banks. By increasing the yield on domestic government securities through open market operations, for example, the BoZ narrows the spread between Kwacha and dollar denominated securities. The domestic branches of international banks tend to shift from dollar denominated assets in favour of domestic securities markets, thus appreciating the domestic exchange rate. This mechanism works the same way (by lowering yield on domestic bonds) in the event that the BoZ wants to achieve a depreciation of the exchange rate.[57]

4.4.6 Out of sample analysis

In our out-of-sample analysis, we consider impulse response functions and forecast error variance decompositions. The stability checks in Table 4.7 do not indicate that the model is misspecified. Therefore IRFs and FEVDs are valid and have known interpretations.

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[56] Until 2012, the Bank of Zambia (BoZ) pursued a Monetary Aggregate Targeting (MAT) framework - reserve money was the operating target while broad money was the intermediate target, aimed at controlling inflation, the ultimate target. To achieve these objectives, the BoZ conducted Open Market Operations (OMOs) and regular auctions of Government securities. In April 2012, the policy was changed to interest rate targeting – a policy rate based on the overnight interbank rate was introduced. Although the policy rate has been set based on the interbank rate, bank lending rates have been shown to track closely the 91-day Treasury bill rate (Simpasa et al 2014).

[57] It is not difficult to see why the Central Bank will be concerned about the effect of a volatile exchange rate in Zambia. Weeks (2013) argues that interventions in the exchange rate markets "socialises hedging" foreign exchange risks in a market with relatively shallow hedging products and dominated by a few market participants.

4.4.6.1 Impulse Response Functions

Our discussion of the impulse response functions (IRFs) centres on the responses of the financial market variables to shocks in the copper price, their own shocks and shocks to the other financial market prices. The results are displayed in Figure 4.2.

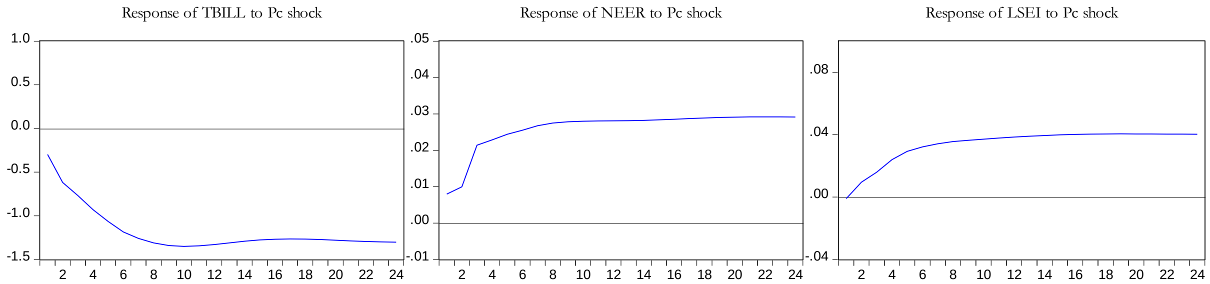
The response of the money market rates to a positive innovation in the copper price is a contemporaneous fall and appears permanent. This result confirms the findings from the in-sample long run equation. The players in the Zambian market must pay attention to the price of copper as unexpected positive shocks tend to be followed by lower interest rates.[58] Following an appreciation of the exchange rate, short term interest rates tend to rise contemporaneously before reverting to steady state in the third month. The interest rates then fall before correcting to steady state on the ninth month. This result indicates the tendency of the BoZ to lower yields on Kwacha denominated debt instruments to moderate the volatility of the exchange rate through the actions of domestic branches of international banks discussed earlier. A positive shock to the short-term interest rate is followed by a contemporaneous rise that stays constant for about four months before it falls gradually for another six months. Following a positive shock in the equity prices, the short-term interest rate falls contemporaneously before rising with a lag of two months, peaking on the sixth month before correcting steadily to the fourteenth month. This result is consistent with short-term Granger causality result. The "reaction" of the money market rates to a positive shock in the equities price most likely reflects the inefficiencies in the Zambian financial markets – one would expect a faster response in the developed financial markets.

The response of the exchange rate to a shock in the interest rate seems transitory; appreciating the exchange rate with a one month lag and correcting steadily by the twelfth month. This result likely reflects leveraging activities by investors – high interest rates attract

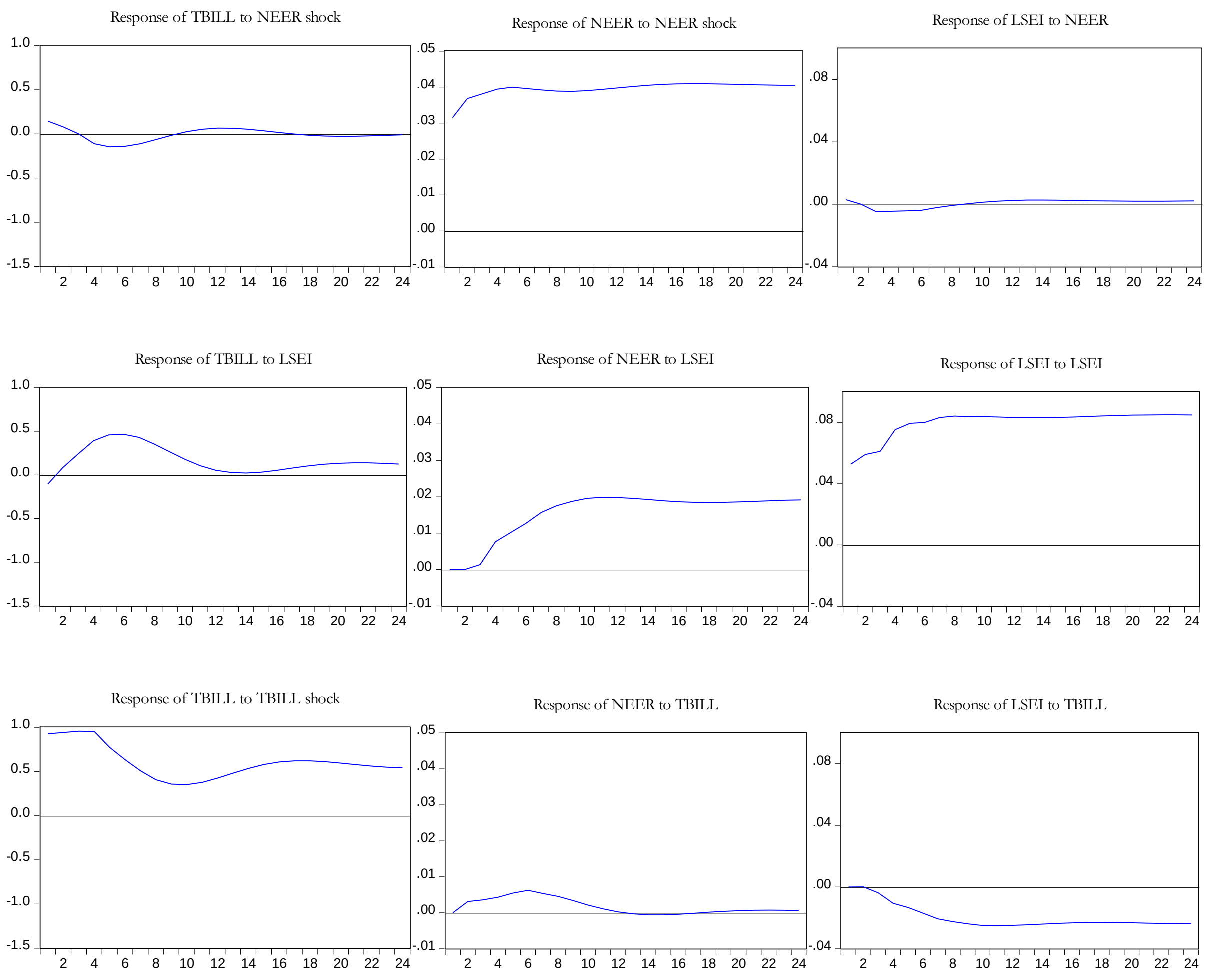
[58] Higher copper prices tend to be followed by appreciation of the exchange rate. The pass-through from the exchange rate tends to adjust inflation and interest rate expectations downwards. See https://www.lusakatimes.com/2015/11/03/bank-of-zambia-hikes-interest-rates-to-15-5-full-statement/

foreign portfolio flows into the money market, exerting upward pressure on the exchange rate (see Muhanga and Soteli, 2009). A positive shock in the copper price leads to a steady and permanent appreciation of the exchange over the two year period, possibly through its effects on the current account movements. This result supports the fairly conventional theoretical views of open economy macroeconomics. A boom in the copper price improves the country's current account leading to an appreciation of the exchange rate (see Krugman 1983a). Equity market shocks appear to appreciate the exchange rate with a two months lag. This observation supports evidence of robust participation by foreigners on the Lusaka Stock Exchange.[59] The portfolio flows in turn increase the foreign exchange order flows on the Zambian Kwacha in the manner described by Hau and Rey (2006). By comparison however, while appreciation of the exchange rate places downward pressure on the equity prices (possibly by increasing the incentive for profit taking and locking in foreign exchange gains by foreign investors), the effect appears to be statistically indifferent from zero. Basher et al (2012) report a similar result using the trade weighted US index and an emerging market equity index.

Figure 4.2: Impulse response functions



[59] Foreign participation on the stock exchange was 21% of total turnover in 2011 (http://www.luse.co.zm/wp-content/uploads/2012/08/Year-End-Statistics-2012.pdf)



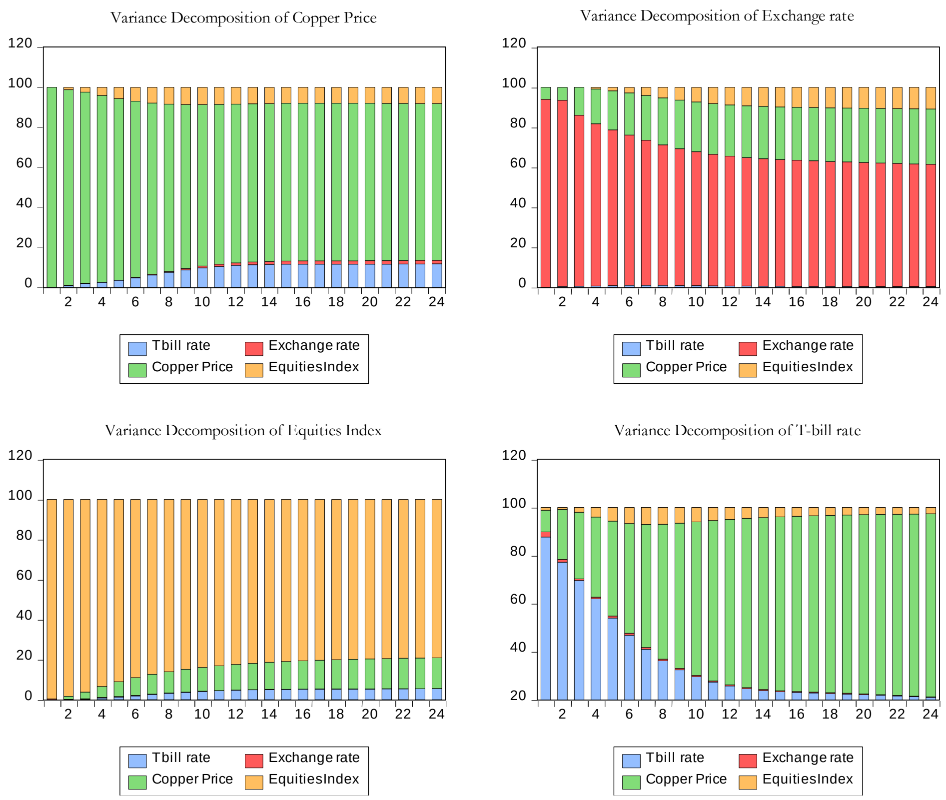
The reaction of the equity index to its own shock does not appear to correct over the experiment period. This result likely reflects inefficiencies in the stock exchange – one would expect shocks in efficient markets to correct rapidly. Stock market prices fall following a shock in the short term interest rates. Equity prices initially rise to a shock in the copper price and thereafter stay at the elevated level indicating the permanent effect of the shock and indicating that the two variables share a long-run equilibrium relationship.

4.4.6.2 Variance decomposition

Following the discussion of the impulse response functions, we now turn to the results of the variance decompositions. The forecast error variance decompositions are displayed in Figure 4.3 for the four variables in our study.

As can be seen from Figure 4.3, the exchange rate, equity prices and the copper price dominate the system in the short run, supporting our in-sample findings of weak exogeneity in the previous sub sections. By the sixth month for example, 75%, 88% and 89% of the variation in the exchange rate, copper price and equity index are explained by their own innovations respectively. By comparison however, the short term interest appears as the least weakly exogenous variable, supporting the proposition that the variable is a "slave" to the system responsible for correction to long-run equilibrium. The copper price dominates the short term rates in the long run – explaining 76% of the variation by the 24th month. Overall, the picture depicted by FEVDs supports the results of the IRFs on the influential role of the copper price in the determination of financial market prices in Zambia; the remote role played by the exchange rate in the equity price and interest rate equation and the long run bi-directional causality between the equities and money markets.

Figure 4.3: Variance Decompositions



4.5 Conclusions and recommendations

We showed that there exists a stable long-run relationship between copper prices and financial market variables in Zambia in a Vector Error Correction Model (VECM). The correction to the long-run equilibrium is borne mainly by short term interest rates suggesting heavy reliance on monetary policy by the policy makers to steady financial markets in the event of exogenous shocks. In-sample short run Granger causality runs mainly, as expected from the copper prices to the financial markets although changes in the equities market lead changes in short term interest rates. The latter may reflect asset substitution activities of financial market players in Zambia- evidence that development of the equities an alternative investment class to the traditional Government securities. The out-of-sample analysis of impulse response functions and variance decomposition largely confirms our in-sample findings, emphasising the endogeneity of short term interest in the short and long run.

Our results raise at least two issues for policy in Zambia. First is the vulnerability of the financial markets (and the macro-economy) to the external shocks to the copper price and secondly the “over-reliance” on monetary policy to absorb the shocks from the copper market. While the country has had rich pickings from the financial market reforms of the 1990s (see Weeks and Mungule, 2013), it has to be borne in mind that the “commodity super-cycle” helped accentuate the period of remarkable economic growth in the last decade.[60] The threat is raised by commodity slump that started in 2014 (viewed by some scholars as the end of the commodity super-cycle, e.g. Goldberg, 2015; Bershidsky, 2015). Bauer and Mihalyi, (2015) list Zambia among the ten least prepared countries for the burst of the commodity super-cycle. Another relevant concern is how long the commodity price slumps will last. The previous commodity price super-cycles were on average between 15-28 years apart (see Le

[60] Zambia's economy experienced strong growth in the last decade, with real GDP growth in 2005-13 more than 6% per year. See http://www.theodora.com/wfbcurrent/zambia/zambia_economy.html

Billon and Good, 2015). In light of our findings we now make some policy recommendations for the Zambian authorities.

First, a set of counter-cyclical policies are required during periods of a commodity price boom. Reliance on tight monetary policy in the event of a commodity price downturn may have the unintended consequences of increasing domestic borrowing costs for the Government at a time when borrowing is required to finance a fiscal deficit.[61] Counter-cyclical measures include revenue and stabilisation funds and adoption of fiscal rules that avoid over-optimistic revenue forecasting and over-expenditure by Government (Frankel, 2011). Such reforms have been adopted in Chile for example.[62] Monetary authorities may consider anchoring policy on price indices tied to the main export prices rather than consumer price index (CPI).[63] Stabilisation funds and sovereign wealth funds built during commodity prices super cycles need to be considered (Varangis et al 2005). Stabilisation funds, built to smoothen revenue in the short-medium term typically involves the government depositing funds in an offshore account in periods of excess revenue for withdrawal when the commodity prices are below a pre-determined reference price (Bauer, 2014). Chile provides a case study with its Pension Reserve and Social and Economic Stabilization Fund.[64]

In the long-term the economic diversification agenda should be accelerated – with the upshot of broadening the portfolio of economic activities that reduces vulnerability of export revenues to copper. A well-developed non-resource based tradable sector may take advantage of Kwacha devaluation in periods of commodity price slumps to increase competitiveness in global markets (Le Billon and Good, 2015). Some practical tools that may be adopted to help the diversification agenda include implementing measured trade restrictions to support local

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[61] Zambia's marginal propensity to import is estimated at over 40% (see Weeks and Mungule, 2013). This means that an exchange rate deprecation of 10% could potentially lead to economy-wide price increases of over 40%.
[62] In Chile, a new structural budget rule requires that the copper price and estimated output be estimated by an independent panel of experts. The two variables of copper are the most important inputs into the state budget.
[63] See Frankel (2006) for a robust proposal for small commodity exporters to peg policy in the export price index.
[64] See http://www.swfinstitute.org/fund/chile.php

manufacturing, improving access to financial markets and market liberalisation.[65] Other forms of diversification may include promotion of sectors with backward and forward linkages with the copper mining industry. Backward linkages involve use of local inputs and local suppliers by mining firms to promote growth of local industry. While such initiatives' success is mixed in low income countries, more flexible approaches such as raising the capability of local firms by mining companies are promising for the future (Sutton, 2014). Forward linkages, (sometimes known as beneficiation) involve further processing of the raw primary commodity for local use or export. Another policy alternative is to harvest copper revenues to invest in sectors that are completely unrelated to copper mining, either through development banks or industrial policy (Venables, 2016). Malaysia offers an example of success of this policy (see Yusof, 2011).

Other policy choices for Zambia include strengthening the governance and transparency in the management of copper revenues. These improvements may include a formal program of raising citizen awareness on the implications of commodity price cycles and related revenue management decisions.[66] The authorities also need to improve management of the "resource for infrastructure" deals particularly with China.[67] Such deals accelerate the diversification program and help transform the sub-soil assets into surface assets at the price of foregoing concurrent consumption. These policy options while they come with their limitations are necessary to reduce the vulnerability of export revenues to the shocks in the copper market and by extension, over reliance on monetary policy as a response instrument to copper price shocks.

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[65] Opening up financial markets and financial liberalisation enables producers access to international financial markets and alternative investments abroad.

[66] As suggested by Venables (2016), Malaysia and Botswana have seen management of citizens expectations improve economic performance and

[67] Halland et al, (2014) note that some of these deals are barter deals and form parts of bilateral trade agreements. Bräutigam and Gallagher (2014) estimate that, between 2000 and 2011, China committed $53 billion to Africa in funding of "resource for infrastructure" type deals.

For financial market players, it is valuable to watch the copper price as a signal of the future direction of interest rates. Equity prices may also signal future direction of short term interest rates.

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5. Oil Price Shocks in Nigeria and Kenya Financial Markets

5.1 Introduction

The behaviour of oil prices has generated interest among researchers in academia, policy makers and financial market participants since the 1973 oil price shock and the subsequent economic recession – the longest of post-World War II (Brown and Yücel, 2002). The works of early researchers like Pierce and Enzler (1974), Rasche and Tatom (1977), Mork and Hall (1980), Hamilton (1983) and Darby (1982) shows that the economic recessions of the 1970s and 1980s were preceded by high oil prices. Since then, there have been numerous attempts to model the oil-macro-economy relationship (e.g. Hamilton, 1983, 1996, 2003, Killian, 2008, 2009, Killian and Vigfusson, 2011a, 2011b, 2013). There have been yet more studies investigating the oil-financial markets relationship (see Blanchard and Gali, 2007, Herrera and Pesavento, 2009, Bernanke, 2006, Brown and Sarkozy 2009, Chisholm, 2014 and Froggatt and Lahn, 2010). The literature shows that despite several attempts to understand how oil prices impact the economy and financial markets, these interactions differ across countries and time (Miller and Ratti, 2009, Bastianin et al., 2016, Chen et al., 2016). The volatility of oil prices and its effects on the economies and financial markets in developing countries provides a compelling reason for researchers and policy makers to understand the dynamics of the oil market with the object of managing the shocks better. This is the aim of this study in the context of Nigeria (a net oil exporter) and Kenya (a net oil importer).

Early studies on oil and financial markets concerned developed countries (as major consumers of oil) and Middle East oil producers. Recent developments in the world such as the emergence of China and India as economic giants, has seen a shift of focus to the emerging world. According to the US Energy Information Agency (EIA), China became the world's largest net oil importer in 2014. The EIA projects further that China, India and the

Middle East will account for 60% of a 30% increase in global energy demand to 2035.[68] The EIA also notes that oil consumption has been growing in Africa over the past decade and is forecast to remain at similar levels in the next two decades (EIA, 2014). As oil consumption increases however, so does sensitivity of the relevant economies to oil price shocks (Pershin et al. 2016).

The question that arises is how the growth in consumption of oil would impact sensitivity of financial markets across net exporters and net importers of oil in Sub-Saharan Africa. More generally, the volatility of commodity prices is an issue of concern in Sub-Saharan countries. Sub-Saharan Africa's (SSA) three largest economies – Nigeria, South Africa and Angola depend heavily on commodities for export earnings. The World Bank predicts that SSA economic growth will trail population growth and will be slowest in years.[69] One of the major sources of the deceleration in economic growth is the collapse in commodity prices.

Our study seeks to answer the following pertinent questions about the oil price and financial markets in an SSA setting. First, how do financial market variables (specifically, exchange rates, stock markets and interest rates) interact with the oil price for a net oil exporter and net oil importer? Does empirical evidence support the theoretical prediction of "wealth transfer" between consumers and producers of oil in the case of Nigeria and Kenya? Secondly, we investigate the question of whether this relationship changed after the 2008 oil price shock as argued by authors like Reboredo (2012) and Hacihasanoglu et al (2013) and Chen et al. (2016). And finally we will answer the question of whether there are opportunities to hedge the oil price risk using market based methods. In the next subsections, we discuss the theoretical channels through which oil price shocks are transmitted to financial markets. Further, we provide a brief discussion of the institutional profiles of Nigeria and Kenya.

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[68] See (http://www.worldenergyoutlook.org/pressmedia/quotes/12/).

[69] See http://www.worldbank.org/en/region/afr/overview

5.2 Oil prices and the financial channel

In this sub-section we discuss the theoretical connections between oil prices and financial market variables. While the literature documents other channels through which oil prices are transmitted into the wide economy (see Khan and Ahmed, 2011, Tang et al., 2010 Alom, 2011, Jones et al., 2004 and others), we restrict our discussion to financial market channels.

There are two distinct channels through which oil price shocks can be transmitted into the foreign exchange market. The first is the terms of trade channel. It is important to note that, in theory, the impact of an oil price shock would impact the terms of trade differently, depending on whether a country is a net exporter or importer of oil. Several scholars investigate this transmission channel, for example, Corden and Neary (1982), Amano and van Norden (1998) and Chen and Rogoff (2003). Rising oil prices would lead to a deterioration of the trade balance for a net oil importing country, leading to depreciation of the exchange rate and vice versa for an oil exporting country. Buetzer et al. (2012) link rising oil prices with the Dutch disease phenomenon in net oil exporting countries. Rising oil prices may lead to inflation in the prices of goods in the non-tradable sector, leading to an appreciation of the exchange rate (see Backus and Crucini, 2000).

The second channel is called the "wealth transfer effect" channel credited to the early work of Golub (1983) and Krugman (1983). The basic premise of the theory is that a shock in the oil prices represents a shift in purchasing power from one country to another depending on whether they are a net importer or exporter of oil. The mechanism operates via the current account and portfolio reallocation. For example, because oil prices are denominated in dollars, when the USD depreciates, oil exporting countries have to raise the dollar price of oil in order to maintain the purchasing power (in local currency) of their oil revenues. On the other hand, net oil importers may have to run down their USD reserves in order to pay for a higher oil import bill. The effect of rising oil prices therefore would be to reduce domestic aggregate demand in an oil consuming country, while the converse is expected in an oil exporter (Galesi and Lombardi, 2009). The net effect of a positive oil price shock therefore would be appreciation of the nominal exchange rate of net exporters and depreciation for net oil importers (see Ding and Vo, 2012, Buetzer et al., 2012; Fratzscher et al., 2014; Rasmussen and Roitman, 2011). Our study seeks to investigate if empirical evidence supports this proposition in the context of Sub-Saharan Africa.

There have been several attempts to find the empirical links between exchange rates and oil prices with mixed and inconclusive results (see Reboredo and Rivera-Castro, 2013 for a recent survey of literature). Other studies include Bénassy-Quéré et al. (2007), Kilian et al. (2009), and Bodenstein et al. (2011).

Further, oil prices can affect stock prices in several ways, depending on whether the country is a net oil exporter or importer. The price of a stock at any one point in time is the equal to the present value of expected future cash flows (Huang et al., 1996, Mohanty et al., 2011). The oil price enters the stock-price equation directly by affecting future cash flows of firms and indirectly via discount rate (sees Basher at al., 2012). At firm level, rising oil prices increase cost of doing business; in the absence of complete substitution effect between factors of production for net oil consuming firms, future earnings will be reduced (Sadorsky, 1999, Maghyereh, 2004). Additionally, if the firms cannot fully pass the increased cost of production to consumers, their profits and future dividends which are the key drivers of stock prices will fall (Le and Chang, 2011).

At a macro-economic level, in an oil importing country, the effect of rising oil prices may affect consumer expenditure via four complementary channels namely the (1) discretionary income effect, (2) uncertainty effect, (3) precautionary savings effect and (4) operating costs effect (see Killian, 2010, Chuku, et al. 2010). Moreover, rising oil prices act as an inflation tax to both consumers and producers in an oil-consuming economy. Consumers' disposable incomes affected by shocks in the oil price would in turn affect their demand for firms' products, negatively for oil consumers and positively for oil producers (Apergis and Miller, 2009, Miller and Ratti, 2009). For producers increasing uncertainty in oil prices can lead to higher "hurdle rates" on investments adversely affecting stock prices (Mohanty et al., 2011). Rising oil prices can also be perceived as inflationary by monetary policy authorities leading them to increase interest rates in a net oil importing country. Higher interest rates adversely affect the values of stock prices via the discount rate (Huang et al. 1996, Miller and Ratti, 2009).

While there is a sizeable literature on the relationship between oil prices and stock markets, there is no consensus among researchers (Fowowe, 2013). Some prominent contributors to this strand of literature include Jones and Kaul (1996), Sardosky (1999), Basher and Sadorsky (2006), Boyer and Filio (2007), Hammoudeh and Aleisa (2004), Huang et al. (2005), Sadorsky (2001) and more recently, Basher et al. (2012) and Fowowe (2013).

The channels through which oil prices enter money and bond markets have evolved over time. Early thinkers like Fried and Schulze (1975) and Dohner (1981) discuss the income transfer and aggregate demand channel. Through this channel, consumer demand in oil producing countries rises in the event of a positive oil shock, a rise that more than offsets the fall in consumer demand in oil importing countries. The net effect is that demand for goods from oil importing countries falls and the world supply of savings increases, exerting a downward pressure on world interest rates. The downward pressure on real interest rates is assumed to more than offset the rise in interest rates from oil consuming countries trying to smooth their consumption (see Brown and Yücel, 2002). A related channel, the real balance effect was emphasised by Pierce and Enzler (1974), discussed in Mork (1994) and Brown and Yücel (2002). Through this channel, an increase in oil prices increases domestic prices and therefore money demand for a net oil importing country. Failure by monetary authorities to increase the money supply to meet money demand puts an upward pressure on interest rates.

Another oil price-interest rate nexus is discussed by Kang et al. (2014). The oil price affects long-term interest rates indirectly through investors' demand for bonds. Unanticipated shocks in the price of oil alter the discretionary income and precautionary savings of investors in both oil importing and oil exporting countries and affect the bond and money markets through the demand for bonds and treasury bills by investors. In their study, the authors find that shocks in the crude oil market account for 30.6% long run variation in real returns for a broad based U.S. bond index with average maturity of five years (a result that held for different corporate and government bonds). On the short term interest rates, oil market specific shocks were found to explain 31.2% of the variation in the real 30-day Treasury-bill return in the long run – a result that underscores the connection between monetary policy and oil prices.

Literature on the connection of oil prices and money and bond markets is scant in comparison to that on exchange rates and equity prices. Despite the sheer size of debt markets, (outstripping the equities market in the US) Kang et al. (2014) also highlight that little attention has been devoted to understanding the link between oil prices and debt markets. Bernanke et al. (1997) address the connection between monetary policy and oil price shocks. Killian and Lewis (2011) find little evidence of systematic policy responses to oil price shocks in oil consuming countries and argue that oil price shocks have different causes.

Our study investigates these conjectured links in a comparative study between Kenya and Nigeria.

5.3 Oil and financial markets in Kenya and Nigeria

Financial markets have grown in leaps and bounds in some Sub Saharan countries in recent years. Relevant growth statistics in recent years are displayed in Table 5.1.

Table 5.1: Oil and financial markets in Kenya and Nigeria

KENYA	2000	2008	2015	Post-Crisis Growth
GDP (USDbn)†	12.7100	35.9002	63.3020	76.0233%
Stock Market Cap (USDbn) †	1.2601	10.8522	26.4000	143.0251%
Stock Market Cap/GDP*	10.0001%	30.3200%	42.111%	38.2351%
Government debt securities (USDbn)*	2.1302	5.6225	14.5732	159.0251%
Government debt securities/GDP	17.0200%	16.1250%	23.0236%	47.02145%
USD Exchange Rate‡	77.9512	78.1221	102.2022	31.3520%
Net Oil imports bill as %age of imports†	22.2202%	27.2311%	25.9800%	-5.1202%
NIGERIA				
GDP (USDbn)†	46.3911	208.0728	481.0712	131.0251%
Stock Market Cap (USDbn)†	2.3702	48.0618	49.9723	4.2301%
Stock Market Cap/GDP	5.2100%	23.2301%	10.2101%	-55.2312%
Government debt securities (USDbn)τ	8.2002	17.0609	44.4044	160.0023%
Government debt Securities/GDP	18.0325%	8.0112%	9.0125%	13.2359%
USD Exchange Rate‡	109.5018	136.0033	199.0012	46.4588%
Oil Prices (Median USD/Barrel)†	28.3009	91.8542	52.8025	-43.3658%
Net oil exports as %age of total exports†	99.6408%	91.7416%	90.8506%	-1.0069%

†Source: The World Bank
*Source: Nairobi Securities Exchange https://www.nse.co.ke/
‡ Source: Thomson Reuters Eikon
τ Source: Debt Management Office Nigeria https://www.dmo.gov.ng/debt-profile/domestic-debts/debt-stock

Stock market capitalisation was 42% and 10% of GDP and with post 2008 crisis growth rates of 4% and 143% in Nigeria and Kenya respectively as at end of 2015. Growth in debt markets measured by stock of outstanding Government securities posted robust growth rates of roughly 160% in both countries. Exchange rates against the US Dollar depreciated 31% and 46% in Kenya and Nigeria respectively while the oil price softened 43% as at end of 2015 compared to its post crisis median of $91.85/barrel. In terms of the importance of crude oil in both countries, it is clear that Nigeria still depends on oil exports for 90% of its export revenues while approximately one quarter of Kenya's import bill goes to crude oil.

In the next section we outline the empirical strategy used to analyse the impact of oil shocks in Kenya (net oil importer) and Nigeria (net oil exporter) on financial market variables namely, exchange rates, equity prices and short term interest rates.

5.4 Econometric Methodology

To analyse the dynamic association of financial market variables and oil prices, a Vector Autoregressive (VAR) framework is used. VARs have been used extensively in the literature as a method of summarising dynamic interrelationships among macro-economic variables. The framework allows the characteristics of the data series to determine the exact specification of the model.

VARs have been the workhorse of empirical macroeconomics since Sims (1980) seminal articles. Vector Autoregressive models are set up such that current values of a variable set are partly explained by past values of the variables involved – thus they describe the joint generation of variables involved (Lutkepohl, 2011). In this framework all variables in the model are treated as a priori endogenous. This approach has at least two distinct advantages. First, the framework imposes a minimal set of theoretical restrictions on the model being tested thereby allowing close to pure statistical analysis of variables under consideration.[70] Second, the VAR framework overcomes Sims' (1980) critique that some exogeneity assumptions for some variables in simultaneous equation models are ad hoc and not backed by fully developed theories. The framework allows the variables in the system to interact with themselves and with each other without having to impose a theoretical structure on the estimates. Once they are correctly specified, VARs can be used to simulate the response over time of any variable in a system to either its own disturbance or a disturbance from other variables in the set-up.

Financial markets in small open economies in Sub-Saharan Africa have been growing in the past two decades, both as destinations for foreign portfolio flows chasing emerging market yields and as effective ways of raising capital for the private and public sector

[70] One major assumption of VARs that we adopt in our study is that the relations among variables under study are linear. Chen et al. (2016) investigate this assumption using oil prices and exchange rate data for 16 OECD countries and find very little evidence of non-linear relations between the two variables.

enterprises. While this growth has been laudable, exogenous shocks from international markets have had remarkable effects on the financial markets and participants in developing countries. One such source of shocks has been the price of crude oil, which enters the economy either as a major source of foreign exchange or a major consumer of foreign currency through the oil import bill. The VAR method is chosen here to analyse and isolate the impact of oil prices in these markets. We rely on the Structural VAR (SVAR) modelling approach which allows us to identify structural shocks grounded in economic theory. Several studies on the interaction of oil prices and macro-financial variables use the VAR methodology. These include Hamilton (1983), Mork, (1989), Bernanke et al (1997). More recently Basher et al., (2012) use a six variable structural VAR model to investigate the relationships between oil prices and several macro-economic variables namely emerging stock market prices and exchange rates. Their work borrows from Killian (2009) who uses a three variable VAR model with a recursive identification scheme to analyse relationships between oil prices and exchange rates. Brahmasrene et al. (2014), Pershin et al. (2016), Chen et al., (2016) follow a similar methodology in analysing the nexus between oil and exchange rates. Ratti and Vespignani (2016) use a global factor augmented (GFAVEC) model in investigating relationship between oil prices, global industrial production, prices, central bank policy interest rates and monetary aggregates.

In this study, we focus on impulse response functions (IRFs) and forecast error variance decompositions (FEVDs). Several studies employ this strategy to analyse the response of one variable to a shock to a specific variable. Brahmasrene et al. (2014) and Pershin et al. (2016) follow a similar methodology. Jones and Kaul (1996) analyse IRFs and FEVDs to analyse repose of stock prices to oil price shocks in the US and Canada. Apergis and Miller (2009) analyse the impact of oil price shocks on international stock market prices of developed countries while Basher and Sadorsky (2006) and Cong et al. (2008) analyse short run dynamics between oil and stock prices using IRFs and FEVDs.

As a starting point, suppose financial market variables and oil prices evolve according to the following VAR relation:

$$Ay_t = v + B(L)y_{t-1} + U_t \tag{5.1}$$

Where $y_t = [y_{1t} \dots y_{kt}]'$ is a column vector of observations on the current values of all variables in the model; v is a column vector of deterministic constant terms. Constant matrix A is a kxk matrix that represents the contemporaneous values of y_t while $B(L)$ is a matrix polynomial. The fundamental sources of uncertainty in this model are represented by the i.i.d random variables U_t, where:

$$E(U_t U_t') = I \tag{5.2}$$

Where I denotes an identity matrix. Given the values of A and $B(L)$, impulse responses to shocks in U can be computed from the moving average (MA) representation of the system of the form:[71]

$$y_t = C(L)U_t = \sum_{s=0}^{\infty} C_s U_{t-s} \tag{5.3}$$

Where

$$C(L) = A^{-1}[I - B(L)]^{-1} \tag{5.4}$$

In practice, it is impossible to observe and estimate the vector of disturbances U_t directly. It is possible however to estimate the reduced form version of equation (5.1) which is obtained by multiplying both sides of equation (5.1) by the inverse of matrix A. The reduced form VAR then takes the form:

[71] The general representation of the IRF is widely available and can be found in the seminal works of Sims (1980), Bernanke (1986), and Sims (1986).

$$y_t = \mu + \bar{B}(L)y_{t-1} + V_t \tag{5.5}$$

Where $$\bar{B}(L) = A^{-1}B(L) \tag{5.6}$$

$$V_t = A^{-1}U_t \tag{5.7}$$

The re-arrangements of the model imply that the reduced form vectors and the underlying shocks can be represented by the following relation:

$$E(V_t V_t') = A^{-1}(A^{-1}) = D \tag{5.8}$$

The moving average representation implied by the reduced form VAR in equation (5.5) can be written as:

$$y_t = \bar{C}(L)\mu + \bar{C}(L)V_t \tag{5.9}$$

Where $$\bar{C}(L) = [I - \bar{B}(L)]^{-1} \tag{5.10}$$

The underlying structural model can be recovered by imposing theoretical restrictions on the contemporaneous structure of the model. Estimation of the reduced VAR model gives us the estimates of the variance-covariance matrix of the residuals: $\Sigma_v = E[V_t V_t']$. Substituting this relation in equation (5.8) yields:

$$\Sigma_v = E[V_t V_t'] = E[A^{-1}U_t U_t'(A')^{-1}] = A^{-1}\Sigma_U(A')^{-1} \tag{5.11}$$

Our intention in this study is not to test a particular economic or a specific model, we follow Basher et al (2012) and impose restrictions on our model based on different economic models and ad-hoc reasoning. Given K variables in our model, we normalise the diagonal elements to 1. The structural equation then requires imposing a minimum of $K(K-1)/2$ restrictions on A. We impose only short run restrictions in our model. Following Christiano et al. (2007), and Basher et al. (2012), short run restrictions have been shown to perform "reasonably well".

Given our interest in impulse response functions and forecast error variance decompositions, we estimate the VAR in levels and cointegration is not tested. This approach, which follows Sims et al. (1990), can be justified as follows. First, Doan (1992) and Sims (1990) emphasise that the goal of a VAR model is to examine the interrelationships among variables and not parameter estimates. Both authors argue against differencing even if the variables are non-stationary as differencing may throw away important information about the co-movement among variables being investigated. Sims et al. (1990) show that although unit roots may characterise the data, the traditional standard asymptotic tests are still valid even if the VAR is estimated in levels – IRFS and FEVDs rely only on consistent parameter estimates. Secondly, our analysis focuses on the short-run dynamics of the model, that is, responses to short run constraints imposed on the model. Even if the constraint of cointegration is excluded in our model, the short run is still valid.

As shown above, shocks enter the system through non-zero residuals. From the reduced form of the VAR in equation (5), shocks enter through the residual vector $V_t = (V_{1t} \ldots V_{kt})'$. A non-zero component of V_t corresponds to similar changes on the left hand side variables that in turn induce further changes on the other variables in the next periods. It is possible to study the marginal effect of V_t on the system by inverting the VAR representation and considering an equivalent moving average representation (see Lutkepohl, 2011). The disturbances V_t can be viewed as one-step forecasts and are also called forecast error impulse responses and the corresponding moving average representation is called a Wold MA representation (Lutkepohl, 2005).

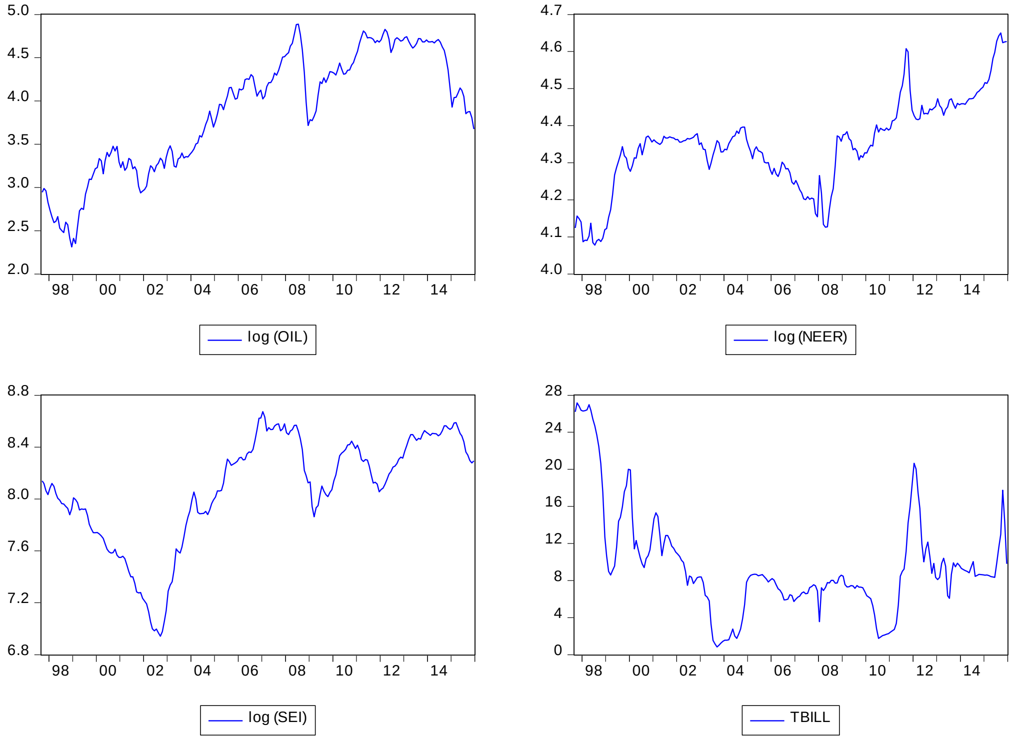
Forecast error variance decompositions (FEVD), separate the variation in an endogenous variable into the component shocks to the VAR. In effect, variance decompositions indicate the importance of each random innovation in influencing the variables in a VAR system. If shocks in in say variable y_{2t} explain none of the forecast error variance in variable y_{1t} at all forecast horizons, then variable y_{1t} can be said to be exogenous to the system and therefore y_{1t} evolves independently of the y_{2t} shocks. On the other hand if y_{2t} shocks explain all the forecast error variance in y_{1t}, at all forecast horizons, then the variable y_{1t} is said to be entirely endogenous to the system. Sims (1980) notes that a variable that is optimally forecast from its own lagged values will have a forecast error variance accounted for by its own disturbances. Most economic variables will typically have their own shocks explain almost all of their forecast error variance in short horizons and only smaller proportions in longer horizons. Thus, while the IRFs show the effects of shocks on the adjustment path of the variables in the system, FEVDs measure the contribution of each type of shock to the forecast error variance. Both analyses help in analysing how shocks to variables reverberate through the system.

5.5 Data Description and Measurement

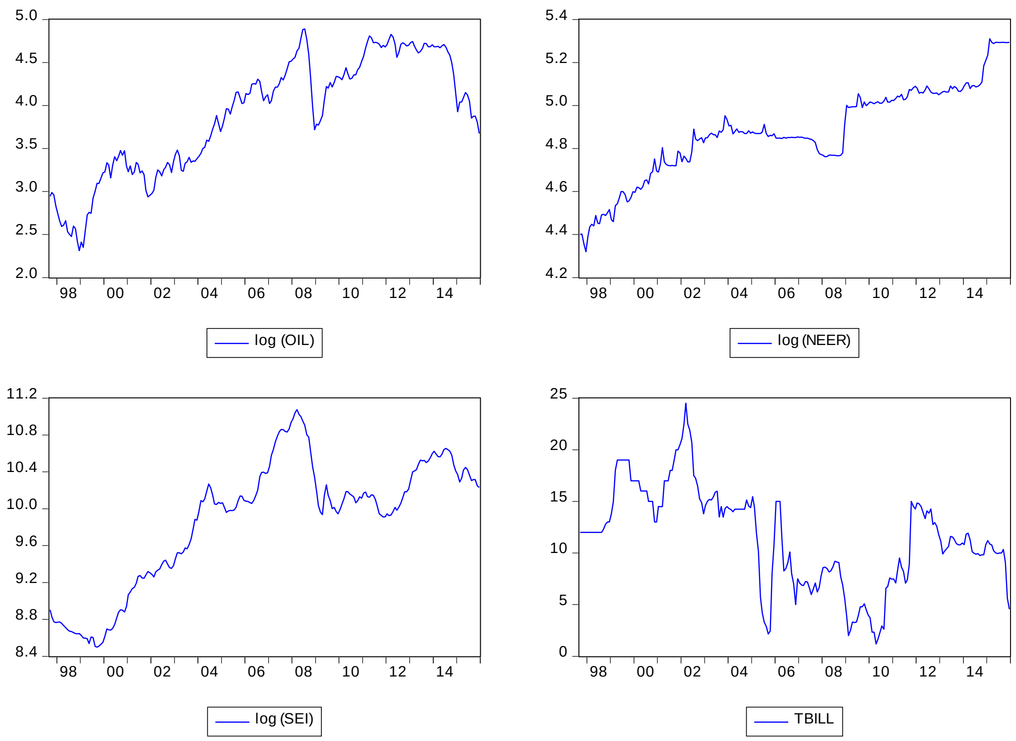
We use monthly nominal data from September 1997 to December 2015. For oil prices (OIL), measured in USD/barrel, we follow Wlazlowski, Hagströmer, and Giulietti (2011) among many others and use Brent-Europe closing monthly spot prices obtained from Thomson Reuters. The Brent–Europe bench-mark is one of the most accurate indicators of world oil prices because it is one of the most traded oils in the current global oil market, and due to this liquidity, it follows the evolution of the global oil prices fairly accurately. For the nominal exchange rate series, we use monthly closing spot price data obtained from Thomson Reuters. For both Kenya and Nigeria, the nominal exchange rate (NEER) is measured as units of domestic currency per USD, so that an increase in the exchange rate represents depreciation of the domestic currency and vice versa. The exchange rate series were sourced from the Thomson Reuters database. The stock market index (SEI) is the closing value of the Nigeria Stock Exchange and the Nairobi Stock Exchange all-share indices for Nigeria and Kenya respectively sourced from Global Financial Data.

Figure 5.1: Data Plots

Panel A: Kenya



Panel B: Nigeria



Finally, we proxy short term interest rates by the 90-day Treasury bill rate (TBILL), for both countries, obtained from Global Financial Data, measured in percent per annum.[72] All data series except the TBILL are transformed into logs, such that the vector of variables of interest can be expressed as: $y_t = [OIL; NEER, SEI, TBILL)$.

The data series are graphed and presented in Figure 5.1 while the descriptive statistics are presented in Table 5.2. Panels A and B represent Kenya and Nigeria respectively. All series appear to be non-normal and kurtotic indicating a tendency for extreme movements. Clearly, all series appear to be non-stationary processes with the exception of, potentially the TBILL rates. We test the series for unit root using the Augmented Dickey-Fuller (ADF) and the Philips Perron test and report the results on Table 5.3. The results show that all series are I(1) processes. The oil price exhibits a potential structural break in mid-2008 associated with the drop in price from $144/barrel to $44. We formally perform the Chow break point test (Chow, 1960) for this break in June 2008 and the results in Table 5.4 show that the null of no structural break in 2008:6 is rejected at 5% at all conventional levels.

We treat the structural break in two ways and compare the results. First we include a break dummy to control for the break in 2008 and estimate our model using full samples for both countries. This approach is in the spirit of authors like Basher et al. (2012) and Chen et al. (2016). Secondly we split the sample into two sub-samples, pre and post the 2008 oil price shock. We surmise that this approach will help reveal whether the relationship between oil and financial market variables in developing markets in Africa changed significantly after the 2008 shock. Pershin et al. (2016) and Hacihasanoglu (2013) follow a similar approach.

[72] We also interpret the TBILL rate as proxy to the policy variables for each country following Frankel (2008) and several others.

Table 5.2: Descriptive Statistics

	Log (NEER_KE)	Log (NEER_NG)	Log(OIL)	Log (SEI_KE)	Log (SEI_NG)	TBILL_KE	TBILL_NG
Mean	4.3467	4.8747	3.8842	8.0739	9.8634	9.6195	11.4393
Median	4.3551	4.8694	4.0405	8.1271	10.0619	8.5065	11.7950
Maximum	4.6501	5.3104	4.8888	8.6726	11.0744	27.1627	24.5000
Minimum	4.0775	4.3202	2.3120	6.9428	8.4981	0.8312	1.2000
Std. Dev.	0.1200	0.2129	0.6873	0.4223	0.6932	5.6683	4.8150
Skewness	-0.1009	-0.3062	-0.3997	-0.8644	-0.5247	1.3417	-0.0176
Kurtosis	3.2099	2.8191	2.0352	3.0185	2.2136	4.9988	2.5616
Jarque-Bera	0.7711	3.7394	14.3906	27.3878	15.7643	102.6402	1.7725
Observations	220	220	220	220	220	220	220

Table 5.3: Representative tests for unit root (Full sample)

Variable	ADF TEST STATISTIC		PHILIPS PERRON TEST	
Levels	**Constant + trend**	**Constant without trend**	**Constant + trend**	**Constant without trend**
Kenya				
Log(NEER)	-2.2816	-1.3481	-2.1119	-1.3128
Log(SEI)	-2.1658	-1.3669	-2.1747	-1.3708
TBILL	-3.7093	-3.7901	-3.0007	-3.2280
ΔLog(NEER)	-11.4244	-11.4400	-11.4823	-11.498
ΔLog(SEI)	-8.3921	-8.4009	-8.2772	-8.2842
ΔTBILL	-9.0382	-9.0058	-8.9721	-8.9652
Nigeria				
Log(NEER)	-2.5558	-1.2007	-2.5558	-1.1724
Log(SEI)	-1.5190	-1.6566	-1.0396	-1.3270
TBILL	-3.0410	-2.3397	-2.7980	-2.1481
Log(OIL)	-1.3740	-1.62323	-0.9691	-1.5620
ΔLog(NEER)	-12.2764	-12.2879	-14.1059	-14.1360
ΔLog(SEI)	-7.9687	-7.9201	-7.8917	-7.8493
ΔTBILL	-10.6821	-10.6909	-10.7539	-10.7624
ΔLog(OIL)	-9.8163	-9.7515	-9.7401	-9.6932

Notes: The Critical values for rejection are -4.0296, -3.4444 and -3.1471 at a significant level of 1%, 5% and 10% respectively for models with a constant and linear trend and -3.4812, -2.8830, -2.5787 at a significant level of 1%, 5% and 10% respectively for models without a linear trend. The optimal lag for the ADF test was chosen based on the Schwartz Information Criterion and the truncation parameter for the PP test was selected using the Newey-West truncation method.

Table 5.4: Chow Breakpoint test

Break date: 2008M6 (Pershin et al. 2016).

Null Hypothesis: No breaks at specified breakpoints

Variable	F-Statistic	Log Likelihood Ratio	Wald Statistic
OIL	138.9650***	108.4914***	138.9650***

*Notes: *, ** and *** denotes significance at 10%, 5% and 1% level respectively*

5.6 Empirical Findings

In the previous section, we outlined our estimation strategy in the presence of a structural break. In this section we present our estimation results. It is generally known in the literature that VAR estimations are sensitive to the choice of lag length. It is thus important to pay special attention to the selection of appropriate lag length. In this study we choose the lag lengths based on the Log Likelihood Ratio (LR) method, preferred by Sims (1980). The detailed discussion of the test is covered in Lütkepohl (1991) and is based on the equation:

$$LR = (T - m)(log|\Omega_{l-1}| - log|\Omega_l|) \sim \chi^2(k^2) \tag{5.11}$$

Where m is the number of parameters per equation under the alternative and T is the number of observations. The test essentially compares the modified LR statistics to the 5% critical values starting from the maximum lag, and decreasing the lag one at a time until it gets the first rejection. The test hypothesis is that the coefficients on lag l are jointly zero using the χ^2 statistics with k^2 degrees of freedom. The lag lengths tests can be found in Table 5.5. We present here lag length choice results for the full samples and sub samples for each country.

Table 5.5: Lag-length selection (LR Tests)

a) Kenya

	Full Sample(1997:09-2015:12)		**Sub-sample 1(1997:09-2008:06)**		**Sub-sample 2(2008:07-2015:12)**	
Lag	LogL	LR	LogL	LR	LogL	LR
0	-687.6031	NA	-316.7576	NA	-126.5760	NA
1	768.1203	2842.7810	481.0965	1530.3100	331.3101	864.8958
2	852.6954	161.9692	523.3151	78.2081	379.7746	87.2362*
3	863.2840	19.8786	536.9232	24.3162	393.0491	22.7141
4	887.7365	44.9832*	555.7125	32.3421*	404.1012	17.9290
5	898.5419	19.4702	567.6351	19.7408	419.4945	23.6030
6	909.7267	19.7316	580.9385	21.1547	432.9927	19.4974
7	918.4253	15.0173	586.2513	8.0997	444.3377	15.3787
8	926.2586	13.2280	600.0748	20.1688	454.2253	12.5243

** indicates lag order selected by the criterion*

b) Nigeria

	Full Sample(1997:09-2015:12)		Sub-sample 1(1997:09-2008:06)		Sub-sample 2(2008:07-2015:12)	
Lag	LogL	LR	LogL	LR	LogL	LR
0	-765.3293	NA	-340.4589	NA	-166.7706	NA
1	699.7008	2860.9550	432.9435	1483.4110	338.5788	954.5489
2	779.3755	152.5847	462.9281	55.5451	380.3119	75.1196*
3	796.0855	31.3707*	480.9249	32.1582	384.7228	7.5474
4	808.6942	23.1951	489.0039	13.9064	400.9265	26.2861
5	818.4100	17.5067	494.7451	9.5059	410.2520	14.2991
6	832.9804	25.7044	513.2236	29.3838*	424.6037	20.7301
7	838.7460	9.9538	523.7484	16.0461	435.9229	15.3438
8	849.2284	17.7013	536.1074	18.0319	450.2104	18.0976

** indicates lag order selected by the criterion*

Having specified the correct lag lengths, before we conduct structural analysis of the VAR model, it is important the VAR model be correctly specified and stable. Inferences based on impulse response analysis and forecast error variance decomposition require that the underlying VAR be correctly specified, that is, the residuals should be white noise processes. McCallum (1993) suggests that estimation of SVAR in levels is only appropriate if the errors from each VAR are serially uncorrelated and stationary [also see Parrado, 2001, Khan and Ahmed, 2011 and Enders (2004) p. 270].

We present the results of these tests in Table 5.6. The tests for stability based on Lütkepohl (2005) and Hamilton (1994) in testing if the inverse roots of the AR characteristic polynomial of the VAR lie within the unit circle. These results are displayed in the appendix. We include twelve lags in our test to account for any possible seasonality in our variables (see Ratti and Vespignani, 2016). The results clearly indicate that the null of no autocorrelation of residuals cannot be rejected at 5% level or better. Additionally, all the inverse roots of the AR characteristic polynomial lie within the unit circle indicating stability for all the underlying VAR models. We thus proceed to the main results of our model.

Table 5.6: Tests for residual autocorrelation

Null Hypothesis: no serial correlation at lag order h (redo with 1 lags to cater for seasonality)

	Kenya			**Nigeria**		
	Full Sample	Sub-sample1	Sub-sample2	Full Sample	Sub-sample1	Sub-sample 2
Lag	LM-Stat	LM-Stat	LM-Stat	LM-Stat	LM-Stat	LM-Stat
1	17.9520	23.2821	18.2952	26.2350*	16.7280	7.4323
2	16.8397	19.5849	19.2511	21.6212	13.7336	16.4155
3	25.0023	13.1192	18.2919	19.1446	12.2741	19.1319
4	24.6550*	16.5551	13.9717	16.6812	9.6166	8.19891
5	16.1851	13.9991	18.0913	21.5069	16.3911	18.5956
6	7.9536	13.9374	15.7006	10.7491	18.0375	16.0638
7	14.8626	18.7895	8.2312	10.4412	13.7601	26.2813
8	9.6690	15.0586	17.0545	12.0446	11.9463	11.9895
9	19.9214	23.1646	23.7921	13.7854	11.6413	19.9657
10	11.5582	19.4909	10.7410	12.1080	11.4238	13.9114
11	25.7231*	9.9039	15.9285	17.5931	6.5838	9.8970
12	17.8999	21.7772	12.0822	10.1221	10.4527	12.6452

*Notes: *, ** and *** denotes significance at 10%, 5% and 1% level respectively*

5.6.1 Identification of structural shocks

The main model presented in Section 2 needs to be identified by imposing a minimum of $K(K-1)/2$ exclusion restrictions. To restate, the model uses monthly data for four nominal variables namely the Brent crude oil price (OIL), nominal exchange rate (NEER), all share stock market indices (SEI) and the 90 day Treasury bill rates (TBILL). Except for the TBILL rate all variables are transformed to natural logarithms. Thus the vector of variables can be expressed as follows, (with variables in logs expressed in small letters) $y_t = (oil, TBILL, sei, neer)$

The short-run restrictions imposed on the model are based on different economic models and adhoc reasoning. The following restrictions are imposed on A:

$$Ay_t = \begin{bmatrix} a_{1,1} & 0 & 0 & 0 \\ a_{2,1} & a_{2,2} & 0 & 0 \\ a_{3,1} & a_{3,2} & a_{3,3} & 0 \\ a_{4,1} & a_{4,2} & a_{4,3} & a_{4,4} \end{bmatrix} \begin{bmatrix} oil \\ TBILL \\ sei \\ neer \end{bmatrix} \tag{5.12}$$

The restrictions imposed on A may be justified as follows. We treat the oil price as contemporaneously exogenous to all the variables in the system on the grounds of information delay. Given that oil is traded in global markets, it is unlikely that the financial market variables of small developing economies in Africa would influence the price of oil within one month. This argument is also made by Killian (2009) and Kilian and Park (2009). Recently Chen et al. (2016) treat the oil price as exogenous in a four variable SVAR model that includes global oil production, global economic activity and global oil production. A similar treatment of the oil price can also be found in Ratti and Vesignani (2016).

Short term interest rates are allowed to respond contemporaneously to the global oil price shock within one month. Rising oil prices are expected to feed into inflation expectations for a net oil importer like Kenya and therefore an upward adjustment of short term rates while the result would presumably be the opposite for Nigeria. See Basher et al. (2012) for their justification of contemporaneous relationship between oil price shock and interest rates.

Stock market prices are allowed to respond contemporaneously to shocks in the oil price and short term interest rates. In small open economies like Kenya and Nigeria, one would expect financial markets to adjust quickly to shocks in macro-economic news namely, interest rates and changes the global oil price and inflation expectations within one month. We assume that foreign investors do not change their investment decisions within one month based on an exchange rate shock.[73]

Finally, exchange rates are assumed to respond contemporaneously to all variables in the model. This treatment of exchange rates can be justified on the grounds of the forward looking nature of exchange rates on asset prices. Moreover, information that influences the value of exchange rates is available on a daily basis and the shocks to the oil price, short term

[73] We relax this restriction on the exchange rate in an alternative specification of the model and our main results do not change at all.

interest rates and stock market prices can reasonably be expected to affect the exchange rates within one month. A similar restriction can be found in Gordon and Leper (1994), Björnland (2009).

5.6.2 Impulse response functions

The main objective of this paper is to trace out the impact of oil price shocks on the financial market variables in two African markets. Given the data characteristics discussed in section 3, the samples for the two countries under study were divided into two. As such, we estimate our model for the full sample (FS) and two sub-samples [subsample 1 (SS1) for the period1997M09-2008M06 and subsample 2 (SS2) for the period 2008M07-2015M12] for each country. We follow Koop et al. (1996), Pesaran and Shin (1998), Khan and Ahmed (2011) and Pershin et al. (2016) in employing generalised impulse response functions (GIRFs) for our analysis. GIRFs are superior to the orthogonalised impulse response functions proposed by Sims (1980) in that they are not sensitive to the ordering of variables (Galesi and Lombardi, 2009).[74] Our intention is to trace the responsiveness of the macro-financial variables to shocks in the oil price. For this reason, we only trace out the response of the financial market variables to a positive one period standard deviation disturbance in the oil price for a twenty-four month horizon using the full sample and two sub samples. In the next subsections, we will discuss in detail the GIRFs of each financial market variable.

5.6.2.1 Oil Price Shock and exchange rates

The GIRFs of exchange rates to a one-period oil price shock displayed in Figure 5.2. The two dashed lines represent 95% confidence interval bands. For Kenya, against our intuition, the

[74] GIRFs are unique and fully take account of the historical patterns of correlations observed amongst the different shocks in a VAR system (Pesaran and Shin, 1998). We specified various orderings of the variables and compared the resulting GIRFs. The results show no significant differences.

exchange rate contemporaneously appreciates for two months in the full sample and SS2 although both responses are not statistically significant. The Kenyan exchange rate however depreciates after an oil price shock in SS1 and peaks around the third month and is statistically significant. The conflicting picture between the two subsamples indicates that the relationship between the two variables changed significantly after the 2008 oil price shock. Specifically, the relationship looked stronger and significant in the first subsample while it looks soft and insignificant in the second sub-sample. In comparison, the Nigerian Naira experienced a much stronger depreciation in SS2 than SS1. The response of the Naira in the full sample is statistically significant although softer than SS2. The relationship between oil prices and the Nigerian exchange rate looks stronger after the 2008 shock - in stark contrast to the Kenyan shilling. The Nigerian results are similar to the findings of Reboredo (2012) and Hacihasanoglu et al (2013). The Kenyan case in sub SS1 is in line with our intuition and cost-push inflation literature. Higher oil prices can be expected to feed into inflationary expectations and thus should be associated with a depreciation of the exchange rate all things being equal. The 2008 oil price shock may have motivated the Kenyan authorities to institute counter-cyclical policies to protect the exchange rate from the oil price shocks.

Figure 5.2: GIRFs Oil Prices and Exchange rates

<table>
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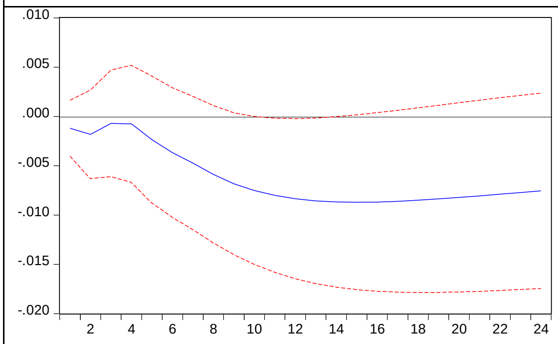
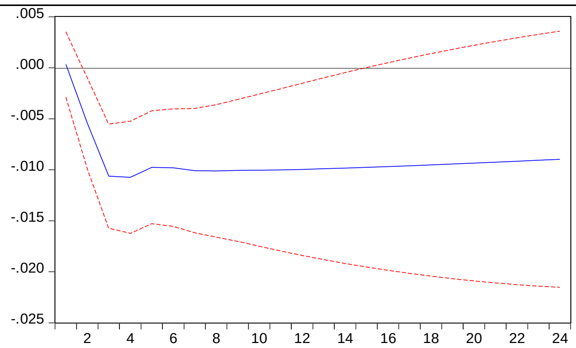
<tr><td colspan="2">Exchange rates (Full Sample)</td></tr>
<tr><td>
</td><td>
</td></tr>
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<table>
<tr><td></td><td></td></tr>
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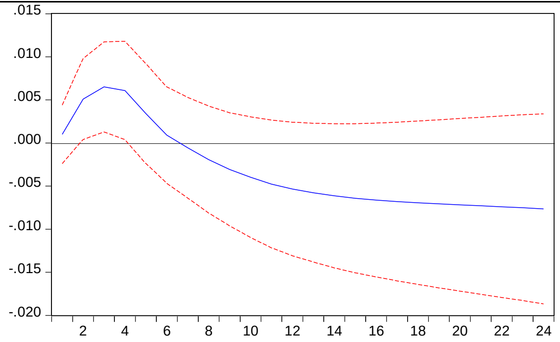
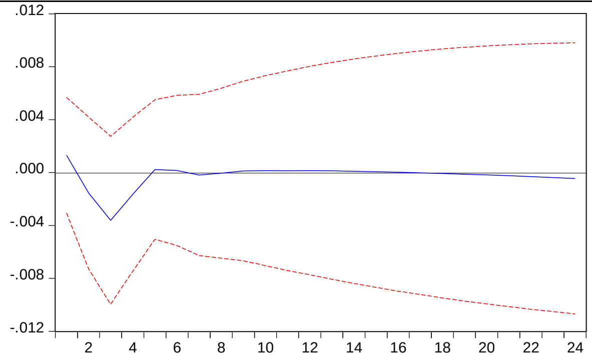
<tr><td>
</td><td>
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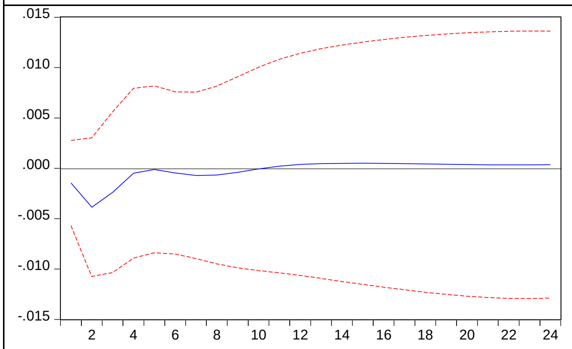
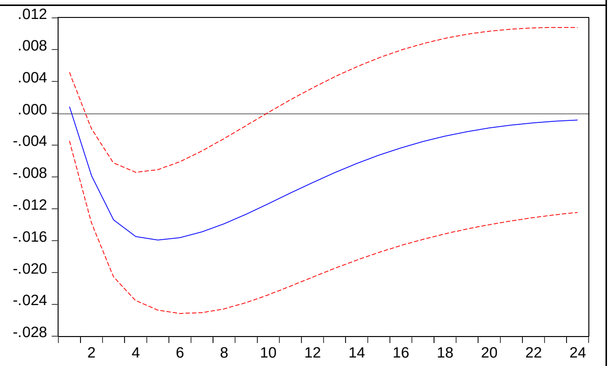
<tr><td colspan="2">Exchange rates(Sub-sample 2: 2008M07-2015M12</td></tr>
<tr><td>
</td><td>
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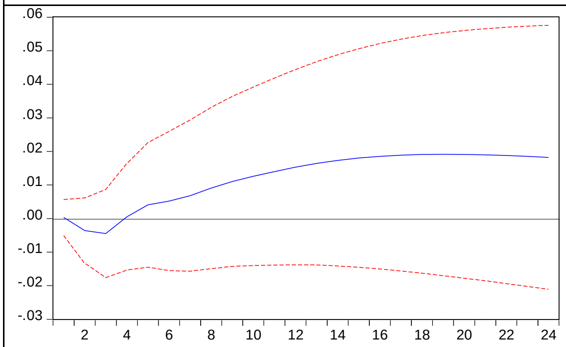
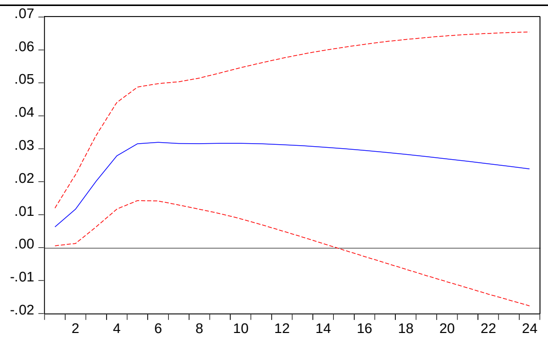
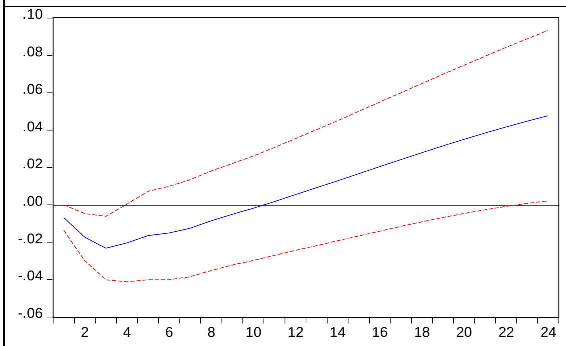
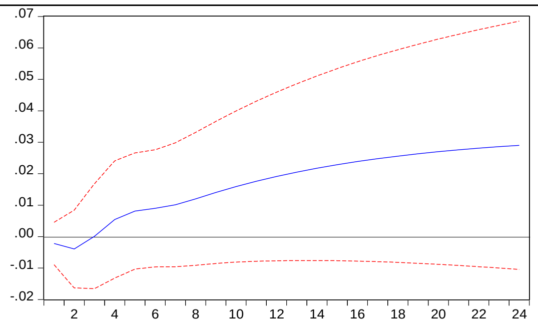
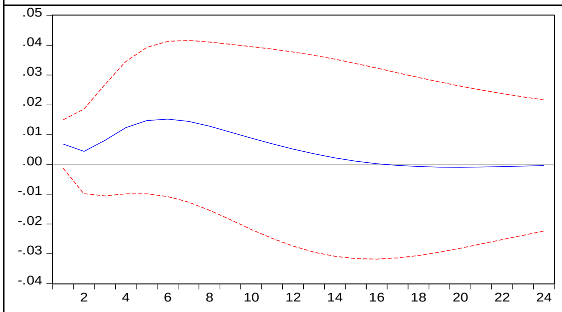
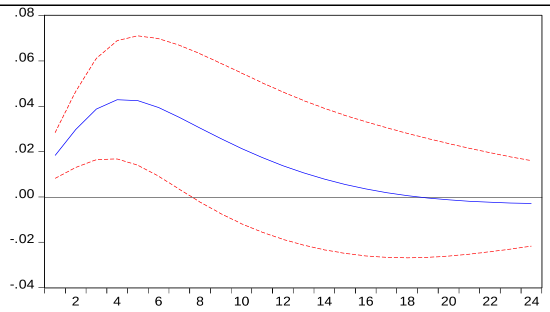
5.6.2.2 Oil price shock and equity prices

In Figure 5.3, we display GIRFs of equity prices to an oil price shock. Kenyan equity prices fall for the first three months before they start rising for the rest of the twenty four months for the full sample and SS1. The result in SS1 is statistically significant for the first three months. The responses are plausible for an economy like Kenya which is a net oil importer. A rise in oil imports would increase cost of production for firms at a micro-economic level and in an economy with sticky consumer prices (which is plausible for Kenya); the future profits for firms fall, negatively affecting stock market prices. This relationship between oil and stock prices however breaks down after 2008. Equity prices appear to increase instead with an oil price shock although this result is not statistically significant. In comparison, equity prices rise contemporaneously to an oil price shock in Nigeria for the full sample and SS2. Both results are statistically significant, consistent with theoretical prediction that improved expected oil revenues would positively affect expectations of firms' profitability and stock prices. Consistent with findings on the exchange rate, the relationship between oil prices appears much stronger after 2008 for Nigeria than Kenya.

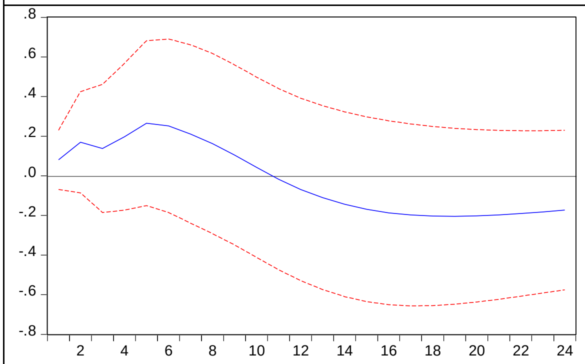
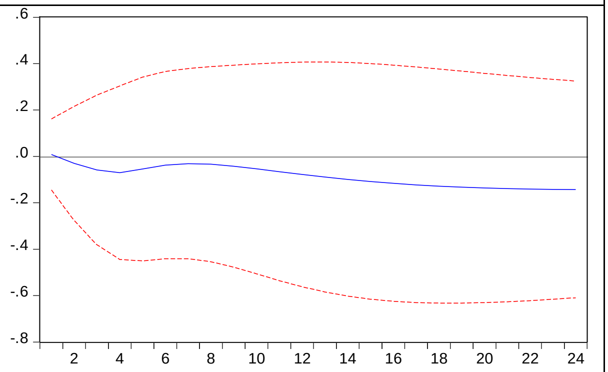
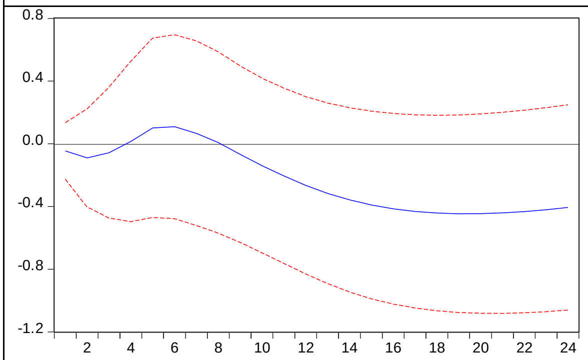
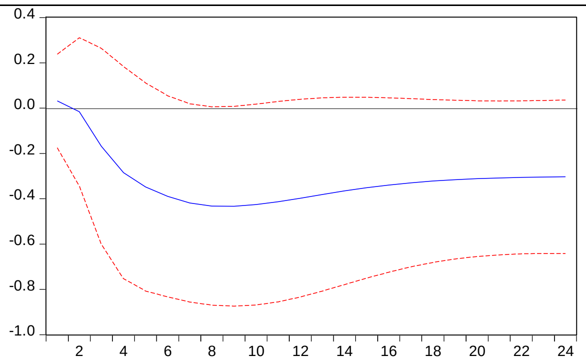
Figure 5.3: GIRFs Oil Prices and Equity Prices

Kenya	**Nigeria**
Equity Prices (Full Sample)	
	
Equity Prices(Sub-sample 1: 1997M09-2008M06	
	
Equity Prices (Sub-sample 2: 2008M07-2015M12	
	

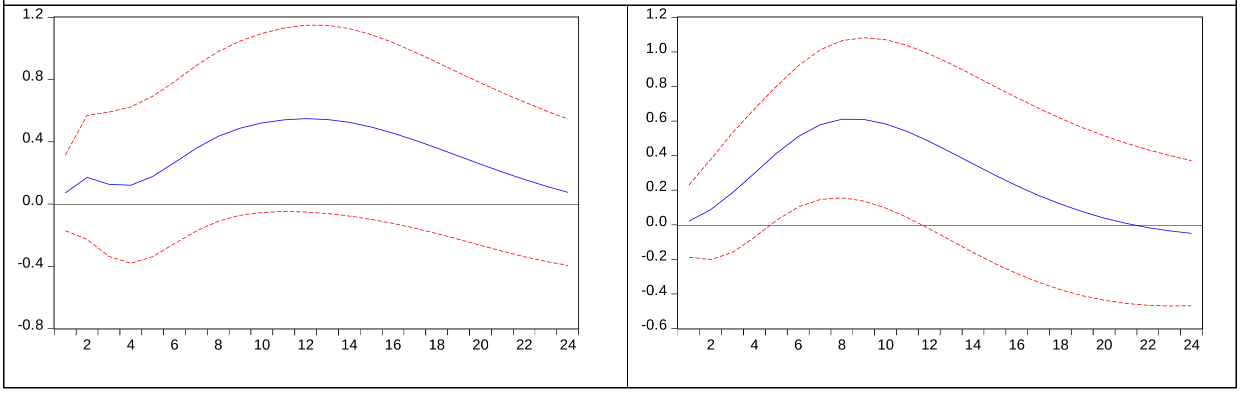
5.6.2.3 *Oil Price shock and interest rates*

The GIRFs of the short term money market rates to an oil price shock are displayed in Figure 5.4. The responses for Kenya look roughly similar for the full sample and SS2, rising mildly in the first month, falling briefly on the second month before the increase accelerates on the third month. For SS1, short term interest rates fall contemporaneously to oil price shock in the first month before they start rising on the second month, falling permanently on the sixth month. The rise in short term rates support the money-demand theory – that is the increase in oil prices increase precautionary demand for money (see Kang et al. 2014). In the absence of immediate increase in money supply by the central bank, short term interest rates rise. All the responses are however not statistically significant.

Figure 5.4: GIRFs Oil Prices and Interest rates

Kenya	**Nigeria**
Treasury bill Rates (Full Sample)	
	
Treasury Bill rates(Sub-sample 1: 1997M09-2008M06	
	

Treasury bill rates(Sub-sample 2: 2008M07-2015M12



For Nigeria, interest rates appear unresponsive to an oil shock in the full sample, staying very close to zero (the original state). In the first sub sample, interest rates fall sharply in the first seven months and stay low throughout the twenty four months horizon. The response in the second subsample is a complete opposite of the first subsample. Interest rates rise contemporaneously with an oil price shock, peaking roughly on the 8^{th} month before they fall back to the original values on month twenty one. Only the responses on the second subsample are statistically significant in line with findings on the exchange rates and equity prices for Nigeria. Once again, the relationship between the oil price and interest rates appears to be much stronger after the 2008 oil price shock for Nigeria. We surmise that improved oil revenues from oil revenues increase demand for short term Government securities pushing money market rates up in Nigeria.

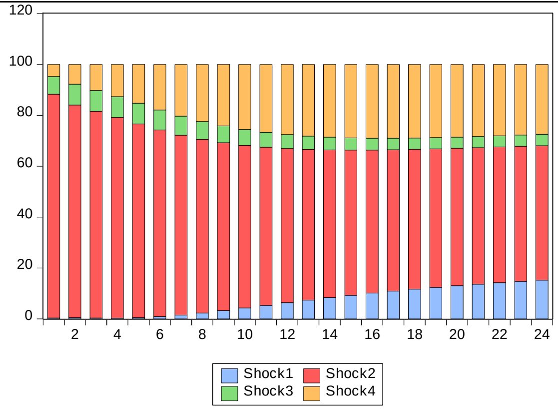
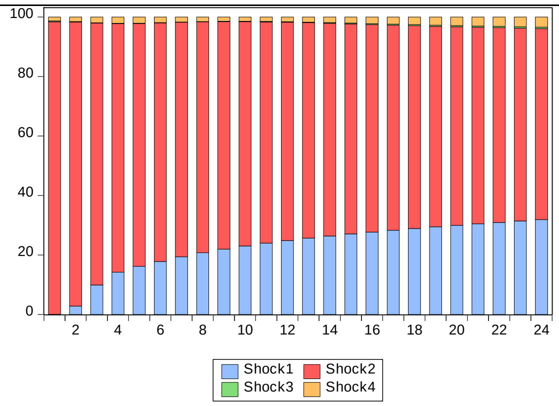
5.6.3 Variance Decomposition

The structural forecast error variance decomposition enables the researcher to identify the relative importance of each dependent variable in explaining the variations in the explanatory variables (Khan and Ahmed, 2011). We display the results of the structural forecast error variance decomposition (FEVDs) for three samples namely the full sample, sub sample 1(SS1) and sub sample 2 (SS2).

5.6.3.1 *Oil Price and exchange rates*

The FEVDs of exchange rates are displayed in Figure 5.5. Consistent with GIRFs for Kenya, it is clear that while oil prices influenced the exchange rate prior to the 2008 structural break, oil price disappears completely as an explicator of the variation of the Kenyan shilling. Up to roughly 16% variation in the Kenyan shilling is explained by oil prices in SS1. The picture is completely the opposite for Nigeria, however. Broadly speaking, the oil price does not enter the exchange rate equation at all in the first subsample – the exchange rate variation is largely explained by itself. The oil price explains roughly 40% and 25% variation in the Nigerian Naira in months 10 and 24 respectively. This result indicates that while the oil price enters the exchange rate equation with a one month lag, it is a persistent explanatory variable of the variation in the Nigerian exchange rate.

Figure 5.5: FEVDs Oil prices and exchange rates

<table>
<tr><td>Kenya</td><td>Nigeria</td></tr>
<tr><td colspan="2">Exchange rates (Full Sample)</td></tr>
<tr><td>
</td><td>
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<tr><td colspan="2">Exchange rates(Sub-sample 1: 1997M09-2008M06</td></tr>
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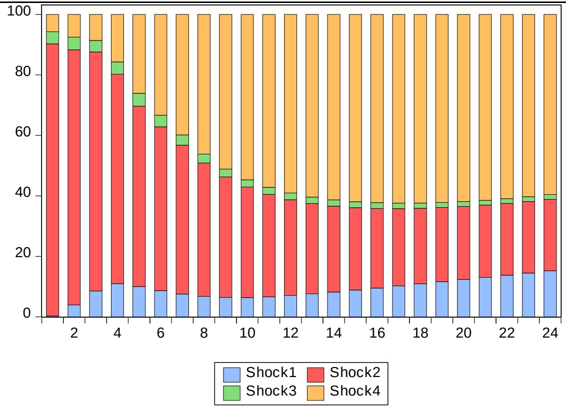
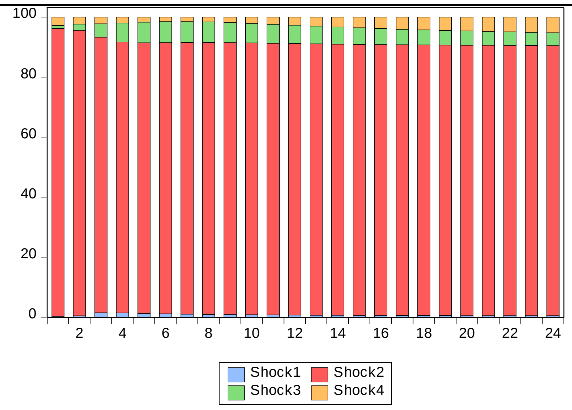
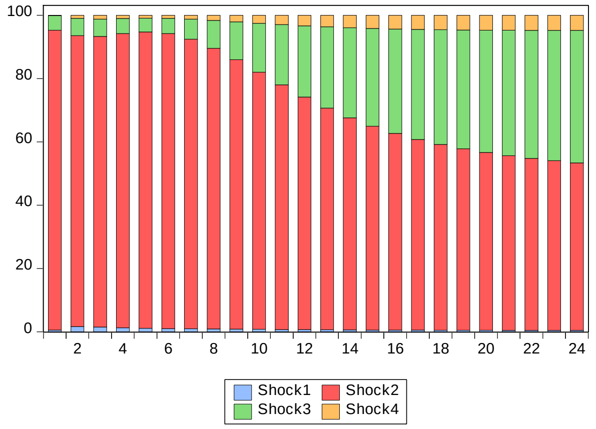
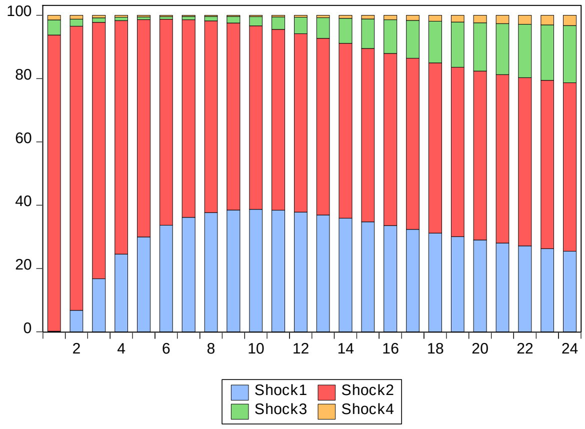




Exchange rates(Sub-sample 2: 2008M07-2015M12





Note: Shock 1=Log(Oil Price); Shock2=Log(Exchange rate) Shock 3= Log(Equity Index); Shock4= Tbill rate

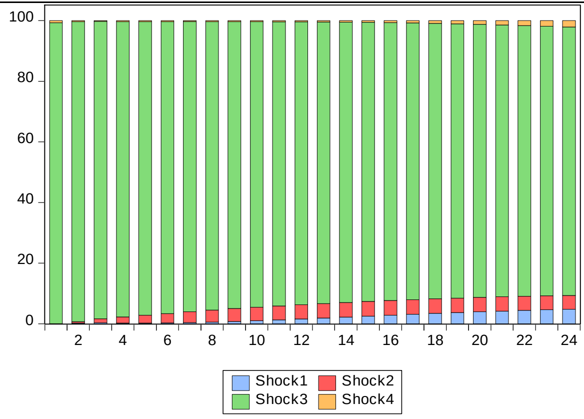
Note: Shock 1=Log(Oil Price); Shock2=Log(Exchange rate) Shock 3= Log(Equity Index); Shock4= Tbill rate

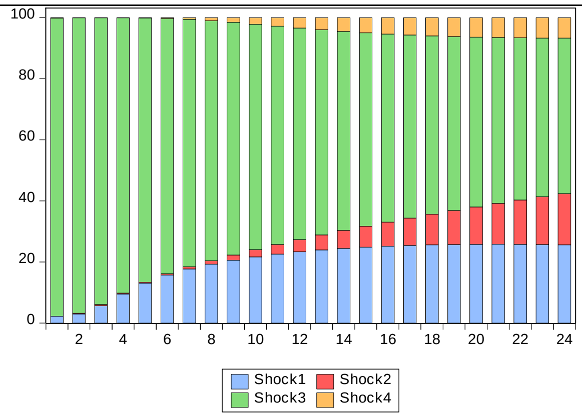
5.6.3.2 Oil Prices and equity prices

The structural FEVDs for the equity prices across the full sample and two subsamples are displayed in Figure 5.6. For Kenya, the importance of the oil price in explaining the variation in equity prices decreases in SS2 and so does the exchange rate. Short term interest rates however increase in their contribution to the variation in equity prices and are persistent throughout the twenty-four months horizon. The result indicates in part that equity prices' sensitivity to externally determined variables (oil price and exchange rate) decreased markedly after 2008; sensitivity to domestic variables instead increased as indicated by the contribution of the domestic short term money market rate increased. This result supports the idea that the Kenyan economy is better protected from external shocks post the 2008 crisis.

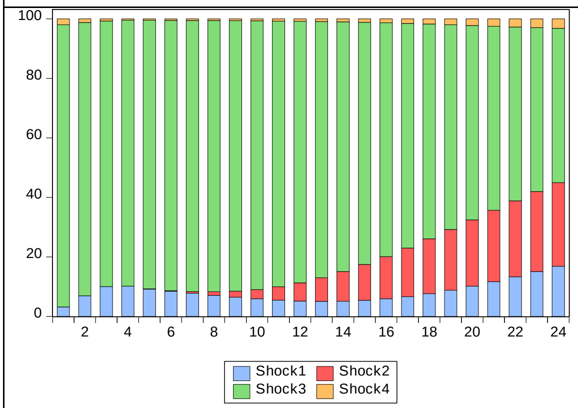
Figure 5.6: FEVDs Oil Prices and Equity Prices

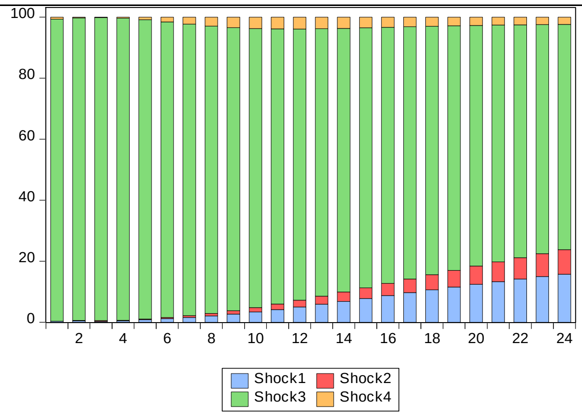
Kenya	**Nigeria**
Equity Prices (Full Sample)	



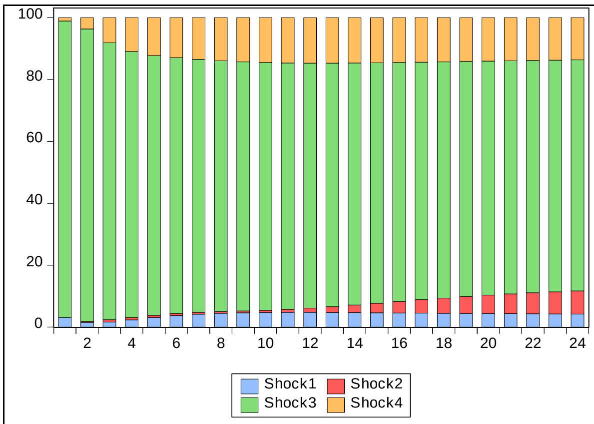
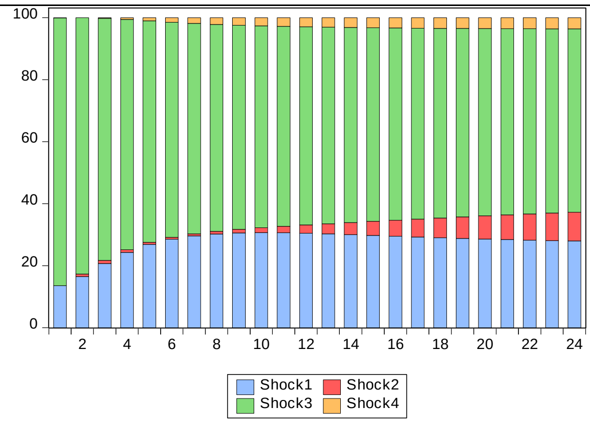


Equity Prices(Sub-sample 1: 1997M09-2008M06



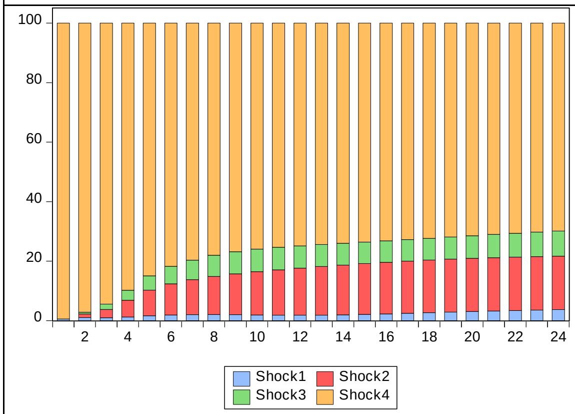
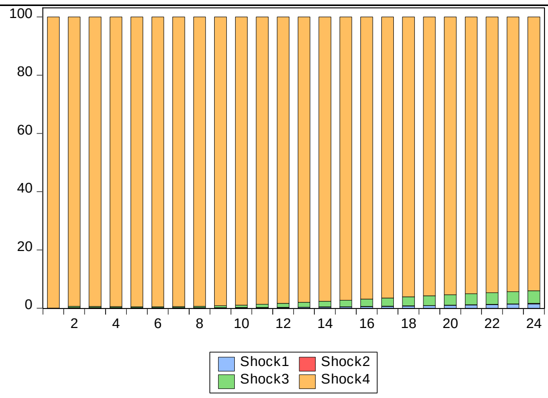


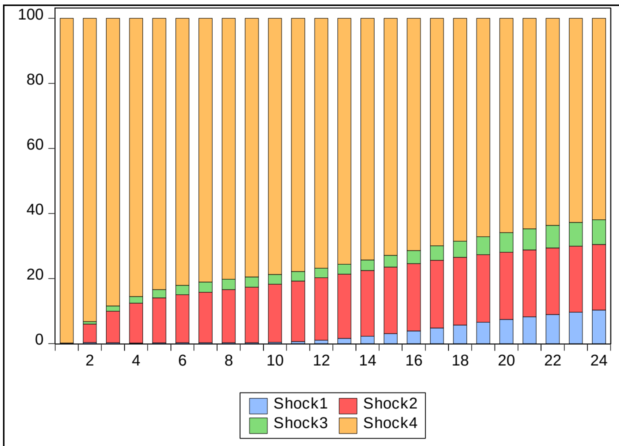
Equity Prices (Sub-sample 2: 2008M07-2015M12

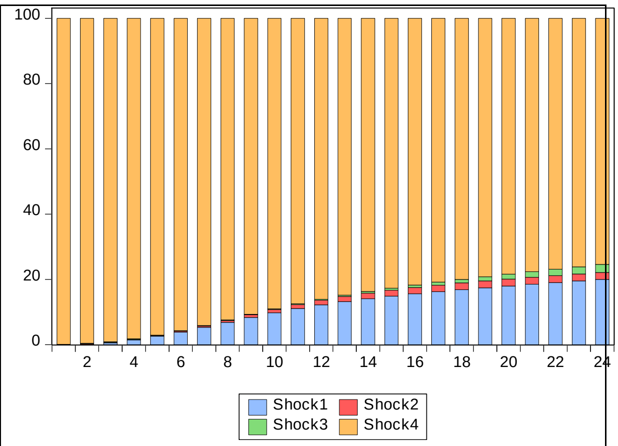
 *Note: Shock 1=Log(Oil Price);* *Shock2=Log(Exchange rate) Shock 3=* *Log(Equity Index); Shock4= Tbill rate*	 *Note: Shock 1=Log(Oil Price);* *Shock2=Log(Exchange rate) Shock 3=* *Log(Equity Index); Shock4= Tbill rate*

For Nigeria however, the contribution of the oil price the variation in equity prices increases significantly in subsample 2, a result that confirms the GIRFs. Whereas oil price enters the equity price equation with a five months lag in subsample 1; it explains between 17% and 28% variation in equity prices over the twenty-four months horizon.

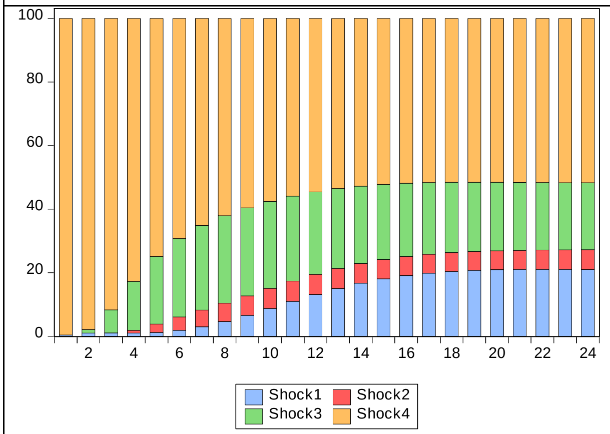
Figure 5.7: FEVDs Oil Price and interest rates

Kenya	**Nigeria**
Treasury bill Rates (Full Sample)	
	
Treasury Bill rates(Sub-sample 1: 1997M09-2008M06	



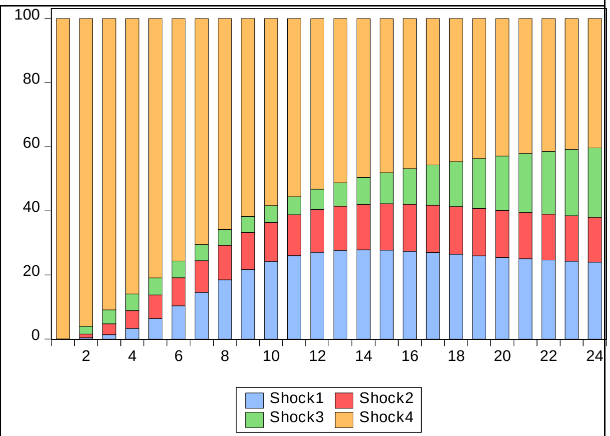


Treasury bill rates(Sub-sample 2: 2008M07-2015M12



Note: Shock 1=Log(Oil Price);

Shock2=Log(Exchange rate) Shock 3= Log(Equity Index); Shock4= Tbill rate



Note: Shock 1=Log(Oil Price);

Shock2=Log(Exchange rate) Shock 3= Log(Equity Index); Shock4= Tbill rate

5.6.3.3 Oil Price and interest rates

FEVDs for short term interest rate for the two countries are displayed in Figure 5.7. Both countries exhibit a similar pattern with respect to the oil price contribution to the variation in the short term interest rate. There is a dramatic and persistent increase in the contribution of oil prices to the variation in the short term interest rates in sub sample 2 for both countries. Equity prices and exchange rate also enter the interest rate equation noticeably for both countries in the second subsample. In the case of a net oil importer like Kenya, a positive shock in the price of oil may lead authorities to increase interest rates in response to expected higher cost-push inflation. For Nigeria, improved oil revenues to the economy are equally likely to increase demand-pull inflation thus eliciting a contractionary monetary policy response as well.

5.6.4 Hedging Oil Price Risk in Africa

While the volatility of the price of oil affects firms at a micro-economic level, the literature provides evidence that governments bear the adverse effects in two main ways. For net oil exporting countries like Nigeria, volatility in the price of oil affects their revenues and thus their fiscal position.[75] For net oil importing countries that rely on oil, a major factor of production, volatility of the oil price causes inflation which may be costly. Some net oil importers often try to smooth domestic oil-led price fluctuations from their fiscal positions. For either set of countries, volatility in oil prices has social, political and economic costs (see Daniel, 2001). While governments across the world have tried several oil price risk management strategies, such as savings and stabilisation funds, the IMF observes that they are inherently flawed and effectively never take the oil-price risk away from government. In this section of our paper we make a case for market based strategies given the promising growth in financial markets in Africa.

The volatility in the price of commodities (that includes oil) has increased significantly in recent years. Triki and Affes (2011) note that in addition to the traditional demand and supply factors in commodities, there have been increases in speculation and a shift to "just-in-time" inventory management systems. More frequent economic busts and emergence of the BRICS as major consumers of commodities have increased demand side

[75] For example, Nigeria increased the 2016 budget deficit forecast by 36 percent after oil prices fell below the $38 per barrel benchmark set by Fiscal authorities. See http://www.bloomberg.com/news/articles/2016-01-21/nigeria-widens-2016-budget-gap-forecast-after-oil-price-plunge

shocks of crude oil.[76] Bierne et al. (2013) report that China's excessive growth adds a premium to the price of oil that increases over time. Supply-side shocks to the oil price emanate from supply disruptions from major oil producers. According to the U.S Energy Information Administration (EIA), the major suppliers of the world's crude oil are countries located in regions that historically have been plagued by political unrest. Major political events in recent years such as the Arab Spring, political events in Nigeria, Venezuela and Libya have been major sources of supply side shocks.[77]

In addition to supply and demand-side shocks, such as ones discussed above, speculation by large hedge funds and other traders have "accelerated and amplified the number and magnitude of price swings" that deviate from market fundamentals (Triki and Affes 2011).[78] While other scholars like Hamilton (2009a, b) find a very limited role of financial speculators in the oil price changes on 2007-8, the consensus in the literature is that there is very little variability in the oil price changes that can be explained by the role of speculators (e.g. Fattouh, et. Al., 2013, Smith, 2009 and Knittel and Pindyck, 2016). The shift in inventory management systems to "just-in-time" (JIT) inventory management in the oil industry also contributes to oil price volatility. The JIT system means that demand changes shift oil prices rather than inventory (Triki and Effes 2011).

Oil price management methods can be classified into two: ex ante and ex post revenue smoothing methods. Doing nothing about the oil price risk is akin to speculating on the oil price – a position that African countries are ill-equipped to handle. Doing nothing about the oil price risk makes planning difficult – making wrong bets about the future spot price of oil in making budgets for example maybe find the government ill-equipped if the assumptions

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[76] For example, a Reuters report of 26 August 2015 notes that oil prices fell on account of concerns over the economic health of China, see http://www.reuters.com/article/us-global-oil-idUSKCN0SD03T20151019

[77] See https://www.eia.gov/finance/markets/spot_prices.cfm

[78] Speculation here is defined as *the purchase (or sale) of an oil-related asset with the expectation that the price of the asset will rise (or fall) to create the opportunity for a capital gain* (Knittel and Pindyck 2016)

turn out to be wrong. In some cases, the government may have to cut expenditure or raise other revenue (in the absence of other financing options). Daniel (2001) argues that such actions by government may make fiscal policy pro-cyclical putting a heavy burden on the private sector and the poor.

5.6.4.1 Ex-post revenue smoothing funds

The revenue smoothing stabilisation oil funds have been around in African countries for some since the 1990s. These include the Oil Reserve Fund (Libya, established 1995), the Revenue Regulation Fund (Algeria, established 2000), Oil Revenue Stabilisation Fund (Sudan, established 2002), the Excess Crude Fund (Nigeria, established 2004) and the Reserve Fund for Oil (Angola, established 2004). Most stabilisation funds are tailored for exporting countries. The purpose of the funds is to smooth revenues and consumption but not to influence the price of oil. Evidence of the success of stabilisation funds is at best mixed (Devlin and Titman, 2004). There are several reasons why the success of stabilisation funds by themselves is limited.

The success of stabilisation funds depends on whether oil prices changes are temporary or permanent and on the assumption that oil prices have a well-defined time-invariant equilibrium value to which they revert (Daniel, 2001). The stabilisation funds in Africa however are generally small and their objective of smoothing consumption and revenue is likely to fail in the event of large and long lasting oil price shocks (Deaton and Miller 1995). According to Le Beloon and Good (2015), the commodity prices slumps have averaged 15-28 years. Triki and Faye (2011) report that only 8 out of 15 Sovereign Wealth Funds in Africa have stabilisation mandates.

Another challenge with stabilisation funds in Africa entails setting up of transfer rules that will trigger transfers to and from the fund. By setting up a benchmark reference spot price for oil, governments assume that the international oil spot price has a well-defined time-invariant "equilibrium" value to which it returns (Daniel, 2001). This assumption is flawed. Daniel (2001) notes that funds generally do not stabilise government finances by themselves, unless accompanied by other policies such as expenditure restraints. Additionally, stabilisation funds create more problems for fiscal policy authorities such as inefficiency in the management of public funds, duplications and overlaps and may encourage poor governance. Indeed, most African stabilisation funds suffer from poor design and governance (Triki and Faye 2011).

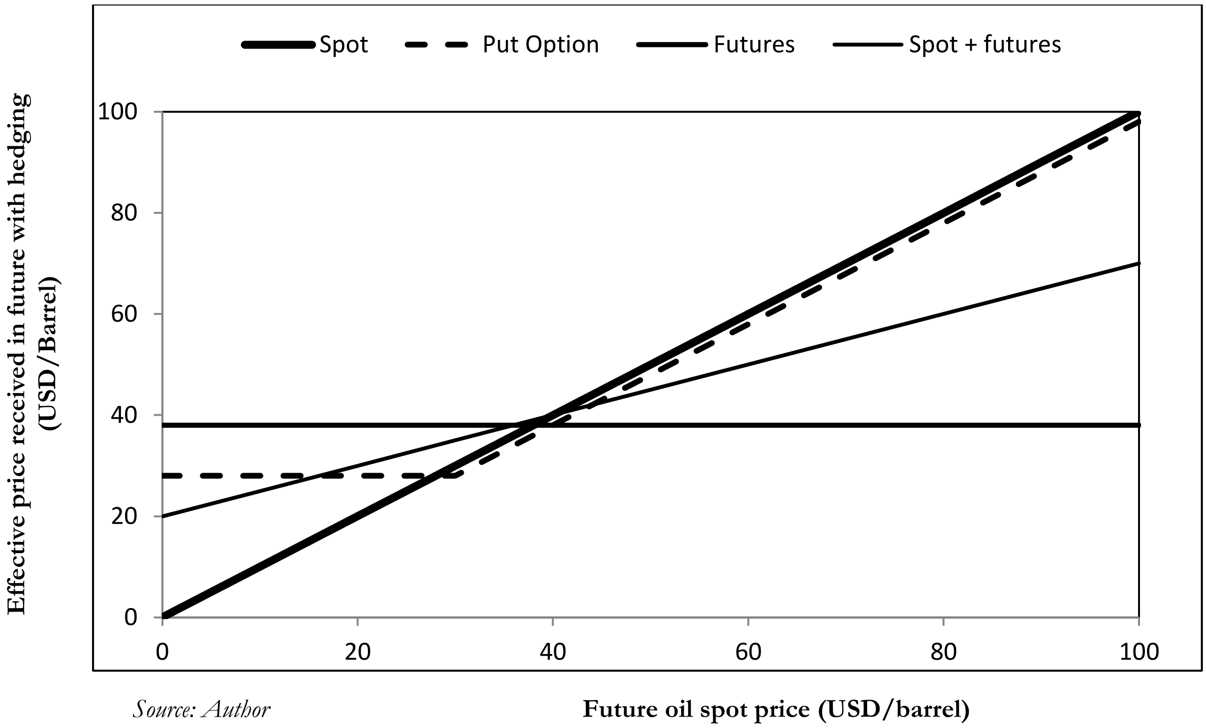
In lieu of setting up pure stabilisation funds governments may opt to create foreign assets/accounts that they can run down when the international price of oil goes down. Because the accounts are generally small, another route will be for the government to borrow to finance its fiscal gap. Most African countries however seldom have the credit worthiness to access the international debt market. Further, during oil price downswings, oil dependent economies are likely to be least attractive to lenders. Even if borrowing was possible, it is potentially politically difficult to accumulate a corresponding surplus to repay the debts when the oil price reverses in most African countries, thus creating solvency problems (Daniel, 2001). In light of these difficulties with stabilisation funds, African governments should consider market based hedging strategies discussed in the next section.

5.6.4.2 Ex-ante market based hedging solutions

African economies (net exporters and importers of oil) are exposed to large oil price risk which they are often times ill-equipped to manage. One way of managing the risk is to transfer the risk to those who can bear it using the oil risk markets. The case for market based oil price risk management strategies has been around for some time although there has been minimal uptake (see Claessens and Duncan, 1993, Engel and Mellor, 1993, World Bank, 1999). We will now discuss the market based hedging strategies for oil exporters although the strategies apply well for importers as well. Broadly there are two main ways in which governments can hedge oil price risk: selling oil price forward and buying insurance against adverse price falls using options or a combination of both.

Oil price risk can be hedged by selling a specific quantity of crude oil forward. This strategy can be achieved through an organised futures exchange such as the New York Mercantile Exchange (NYME) trading forward contracts in the Over the Counter (OTC) market. In Figure 5.8, we illustrate various market based hedging solutions that are traded in the capital markets with the specific example of Nigeria. Assuming the country expected to sell 200 million barrels of oil at a budget price of $38/barrel in 2016, if the government decided to do nothing to hedge the price risk, the outcome is represented by the black bold 45 degree line. The effective price in future could be anything as the spot price cannot be predicted with certainty. Using a futures strategy however, the government can sell 200 million litres of crude oil at $38/barrel (the horizontal line in Figure 5.8) effectively locking in a fixed price for its oil output. The effective price of $38 is achieved by two operations, selling the oil price in the spot market and by the gain/loss from the futures contract.

Figure 5.8: Hedging Scenarios



Source: Author

The downside of using standardised futures contracts is basis risk, that is, the standard futures contracts may not match the exact needs of the hedger and therefore leaving some exposures unhedged. Stated differently, exchange traded instruments may only be weakly correlated with government revenue leaving some of the oil price risk unhedged.[79] To overcome basis risk the hedger can use customised forwards and swaps in the OTC market to achieve the same results.

In a forward contract, the hedger agrees to buy/sell a specific volume of crude oil at a certain future date at a pre-set price. The contract is similar to a futures contract except that it involves physical delivery of the product and it is done directly between the buyer and seller without involving an organised exchange. For this reason, the counterparties in a forward contract assume greater credit risk.

Commodity swap contracts are agreements between counterparties to buy and sell a specific volume of the commodity at a fixed price for many periods in the future. In its "plain vanilla" form, a swap contract involves one party exchanging (swapping) a fixed price for a floating price or vice versa. For example, a producing country could agree with a financial institution to lock in a price for its oil output, say 5 million barrel per half year for the next ten years. If the spot price were to fall (rise) below (above) the agreed price, the financial institution would pay (receive payment) the producer. There are several variations of swaps available in financial markets. Producing countries can use swap markets to establish price floors and price caps. The upshot of using swaps is significant certainty brought to the country's future oil revenues although both parties to the transaction are exposed to significant credit risk. There are several other variations of OTC instruments, including commodity bonds and/or loans and hybrid combinations of the instruments.

[79] Research shows that for crude oil exports, 80% and 70% of oil price risk can be eliminated by using futures contracts for horizons of less than 6 months and horizons longer than 6 months respectively (Daniel, 2001).

From Figure 5.8, a hedger can combine the spot and futures strategy on say a 50-50 ratio. In this case half of the output is sold in the futures market while the other half is sold in the spot market. The strategy still provides downside protection to the oil producer and allows them to benefit from higher oil prices in the future at the cost of bearing part of lower prices.

Using an options strategy, the producer effectively buys insurance against lower future oil prices. To use this strategy a producer would set a minimum price that it would receive in the future. Suppose that the Nigerian Government decides that while its 2016 budget price is set at $38/barrel, oil prices falling below $30/barrel will cause significant challenges. The government can buy a put option (right but no obligation to sell) at $30/barrel. If the 2016 oil prices were to fall below $30, the producer would be compensated by exercising their right to sell at $30/barrel, bringing the effective price to $30/barrel less the premium (the dotted kinked line in Figure 5.8). At prices higher than $30/barrel, the producer does not exercise their put option, instead they benefit from the higher spot price. For this reason option contracts involve upfront payment of the option premium. To offset the cost of the put option, the producer can sell a call option, that is, the obligation to buy at a specific price over a given period, effectively establishing a price "ceiling". In the so called "costless collars" a producer can sell a call option and use the premium to buy a put option. Several other exotic variations exist in financial markets.

In using market based hedging instruments, oil producing countries can benefit from two sources of welfare gains, the "consumption smoothing channel" and the "balance sheet" channel (Borensztein et al., 2013). By bringing certainty to export revenues, oil producing countries effectively bring certainty to expenditure that leads to steadier level of consumption. By using market based instruments, oil producers can save on their foreign exchange reserves which can be made available to service external debts. The upshot of this "balance sheet channel" is that oil exporters can then reduce their default risk, improving its credit worthiness and therefore access to international debt markets. What is more, oil producing governments have some time to adjust in an orderly fashion to long large and prolonged oil price shocks. Using private sector tools to solve macro-economic problems promotes private sector involvement and therefore further development of financial markets in developing countries.[80] Further, there are potential positive global externalities associated with reducing volatility in the oil price. Volatile oil prices create substantial inefficiencies in consuming countries that are likely to find it difficult to sustain fuel efficiency and conservation efforts. Equally, producing countries are likely to face substantial economic and political instability with volatile oil prices (Daniel, 2001). Research suggests that producers will generate higher revenues on average and volatility in oil prices would be significantly reduced (at least in theory) with market based hedging strategies (Devlin and Titman, 2004).

The benefits of market based hedging instruments suggest that international financial institutions such as the IMF have a potentially important role in helping developing countries to implement market based hedging programs. The IMF, for example can help developing countries with education and publicity necessary for them to explore the scope and possibilities of market based hedging programs. Such programs can then be implemented with the technical assistance of agencies with specialist knowledge in these fields such as the World Bank. Moreover, international financial institutions are placed well to help develop instruments to facilitate securitisation of oil proceeds. They can also help to encourage oil companies to work with developing countries (both net importers and exporters) in sharing oil price risk. Such efforts are likely to pay off by increasing the chances of success of the IMF programs in developing oil exporters in Africa.

[80] The establishment of SAFEX, a division of the JSE in South Africa is an example. The AFDB (2011) notes that SAFEX is Africa's most important commodity derivatives market

Finally, small savings and stabilisation funds can be combined with financial instruments to hedge oil price risk. The advantage of combination strategies is that hedgers can then avoid paying premiums and where financial instruments are used, the risk of depleting the fund is reduced in the event of a permanent price shock (Larson and Varangis 1996 and Humphreys 2000).

5.7 Discussion and Conclusions

In this chapter, we show the impact of oil shocks on financial market variables. We discuss the conjectured theoretical links between the oil price and financial market variables and embark on an empirical investigation using monthly nominal data for Kenya and Nigeria. We highlight the significance of oil in these two countries, measured by the share of export revenues and import bill for Nigeria and Kenya respectively. We compare the impact of oil price shocks on financial markets between a net oil importer and a net oil exporter in a structural VAR model.

We confirm, in line with other authors that the oil price structural break of 2008 was significant. We address the oil price shock in two ways. First we introduce a break dummy variable and estimate the full sample for both countries. Second, we split our sample into two, namely the pre-crisis and post-crisis samples on a break date. We employ impact response functions and forecast error variance decompositions in our analysis.

Our results indicate that the "wealth transfer" argument from net importer to net-exporters exists between Kenya and Nigeria in the short run although it is not robust to sample specification. Theory predicts that a positive shock in the oil price would imply "wealth transfer" from producers to consumers of oil (see Golub 1983 and Krugman 1983). If this theory holds, ceteris paribus, one would expect the financial markets to respond positively and negatively to a positive innovation in the international price of oil for a net oil exporter and importer respectively. The response of the exchange rates, while consistent with theoretical predictions for Nigeria, it is less clear for Kenya, especially after the 2008 structural break in the oil price. Equity prices responses are in-line with theory only for SS2 and full sample for Nigeria and SS1 for Kenya. Interest rates appear to move in different directions only in SS1 and appear to move in the same direction in SS2.

This chapter finds that the nexus between financial markets and oil prices is much stronger and statistically significant for an oil exporter (Nigeria) and weaker and statistically insignificant for a net oil importer (Kenya) after July 2008. Prior to the 2008 oil price shock, the results are roughly the opposite of the post oil shock period for both countries. Our results highlight that it is important to account for major structural shifts in modelling the impact of oil prices in developing countries (Le and Chang, 2011). Without accounting for this structural break, one may wrongly conclude that there is, for example, a weak relationship between Nigerian stock market prices and oil prices (e.g. Fowowe, 2013). This result supports the findings of Pershin et al. (2016) and Roberedo (2012) among others.

These results raise the issue of whether the two countries in our study should consider hedging the oil price risk at a macro-economic level. Our results show that Nigeria has a bigger case for active management of the oil price risk after the 2008 oil price shock than Kenya. In addition to Nigeria's heavy dependence on oil for export revenues (over 90%, see Table 5.1), our results show that the pass through of the oil price shock to financial markets is greater and more aggressive for Nigeria than Kenya.

We evaluate the various oil price risk management strategies broadly divided as ex post savings and stabilisation funds and ex ante market based commodity derivatives. We highlight the inherent flaws and limitations of stabilisation funds in Africa and make a case for market based hedging instruments. We argue for the adoption of market based hedging instruments given the promising growth in financial markets of developing African countries in spite of several thorny implementation difficulties. There is a potentially important role for international financial institutions in helping developing countries to develop and implement hedging programs for the significant oil price risk that they are often ill-equipped to bear. There are several positive externalities of using market based hedging instruments, including, involvement of the private sector in managing macro-economic risk, and the positive economic, social welfare and political benefits brought about by reduced oil price volatility. It is potentially appealing too for developing countries to consider a combination of market based instruments and small stabilisation funds to manage the cost of buying options and the risk of depleting their funds in the event of a permanent oil price shock.

The impediments to adoption of market based oil price risk management strategies are unique for each country. Determining the country-specific obstacles and therefore specific recommendations for specific hedging programs in Sub Sharan Africa is a promising area for future research.

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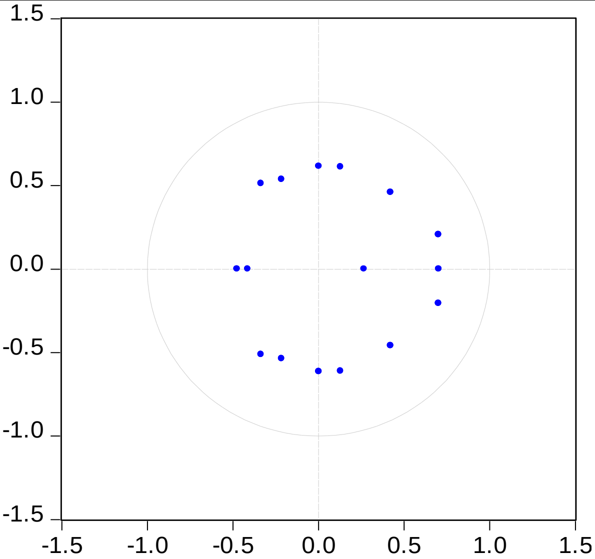
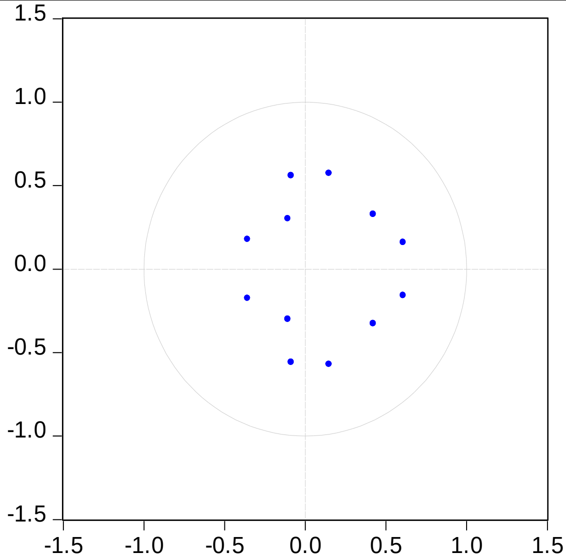
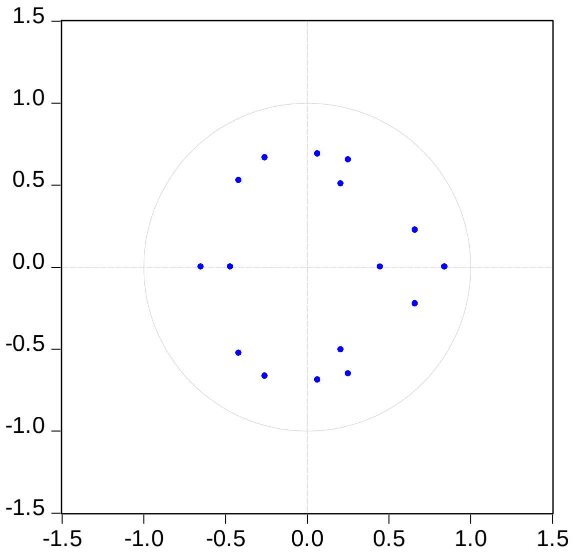
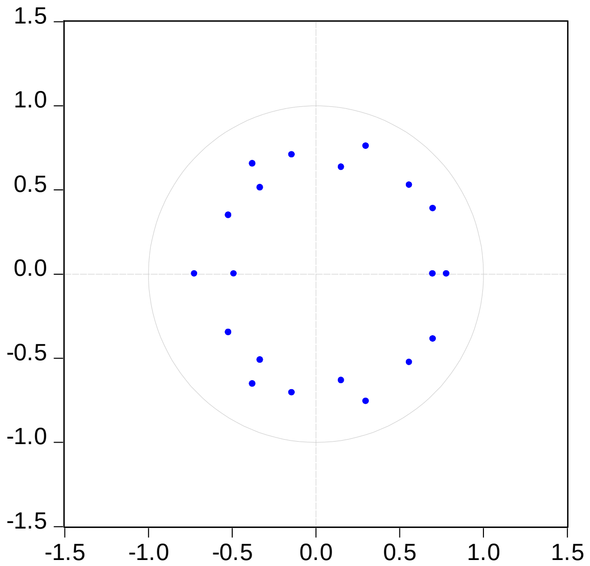
5.9 Appendix

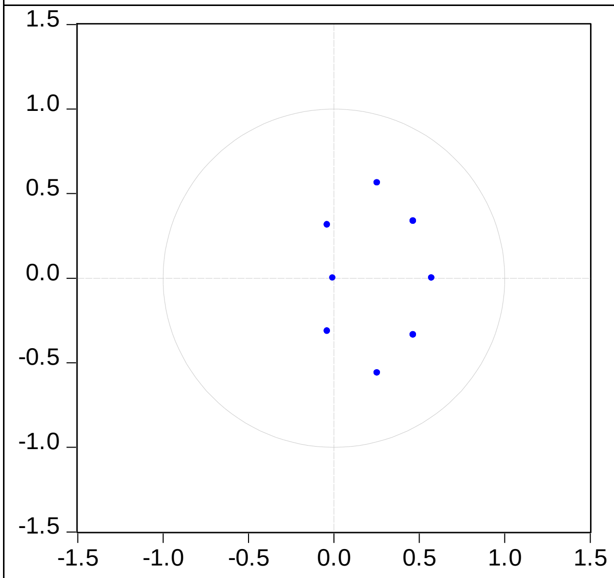
Stability Diagnostics of the underlying VAR

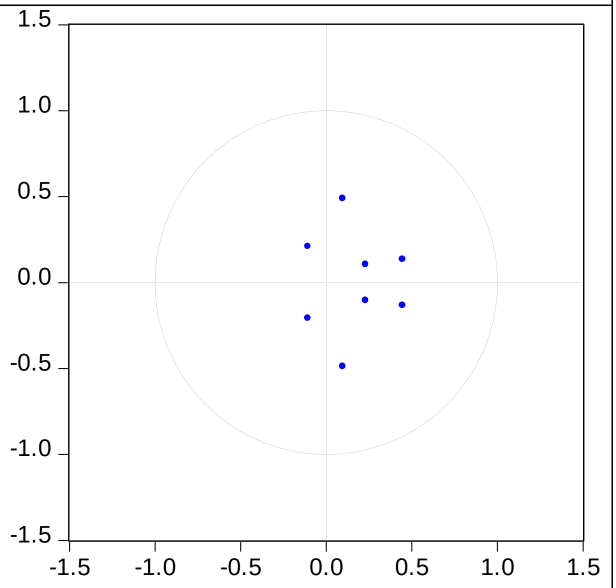
Kenya	**Nigeria**
Full Sample	
	
Sub-sample 1: 1997M09-2008M06	





Sub-sample 2: 2008M07-2015M12





6. Conclusion

Here we summarise the main conclusions of our study:

a) **On the existence of the financial channel** – financial liberalisation in African markets, while laudable, exposes these small commodity exporting economies to the exogenous shocks in international commodity markets. From a country that specialises in the export of a single commodity, to a fairly industrialised country exporting a basket of commodities to a small open economy that imports significant amounts of crude oil, the financial channel appears significant and comparable to other developing economies in Asia and Latin America.

b) **On the predictive content of commodity prices** – commodity prices have robust in-sample and out of sample predictive power, mostly both in the short and long run for most financial market variables. The hypothesis proposed in literature that commodity prices which are transacted in high speed efficient markets should be more information efficient is supported by our findings. In the case of oil, the statistical properties of the oil price such as the structural break of 2008 have equivalent impact on the financial market variables of mainly oil exporters. Movements in financial market variables that are market determined such as the free floating nominal exchange rate in South Africa are not random

c) **On the validity of the "wealth transfer" hypothesis** – results from the shocks in the oil price suggest some validity of this hypothesis within African markets in the short run. Because African markets from the same region have a similar "regional" risk profile, the "wealth transfer" phenomenon within the same markets would be weak in the long run. The hypothesis appears to hold however where bilateral exchange rates with the US dollar are concerned. Weaker commodity prices tend to weaken exchange rates of exporters against the US dollar through the current account channel.

d) **Recommendations for policy and commerce** – overall, for policy, we make a case for market based oil price risk management strategies which can be extended to other commodity classes. The advancements in the financial markets in recent years are promising. The financial market development gains however come at the cost of increased vulnerability of small commodity exporters to exogenous shocks in the commodity markets and their associated negative externalities. It is in the interest of responsible public policy makers to consider market based strategies to protect their economies from the volatility in international commodity markets. Finally, commodity prices can be relied as leading indicators for players in African commodity exporters. There are commodity price-specific drivers of financial market variables that are orthogonal to domestic factors of commodity exporters.