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Multivariate Cointegration Analysis of the PPP and UIP relations between South Africa and the United States: Empirical Evidence

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Declaration

This work has not been previously submitted as a whole, or in part, for the award of any degree. It is my own work. Each significant contribution to, and quotation in, this dissertation from the work, or works, of other people has been attributed, and has been cited and referenced.

Signature: Joel Daniel Mussinga Date: 8 of December of 2007

Abstract

This research assesses whether purchasing power parity (PPP) and uncovered interest parity (UIP) are valid stationary relationships in the long run for South Africa. Quarterly data from January 1972 to June 2006 for South Africa (SA) and the United States (US) is used. The empirical model is a six-dimensional vector autoregressive (VAR) model incorporating oil prices as an exogenous *I(1)* variable. The inclusion of this variable has the advantage of improving estimation efficiency and allows us to investigate the impact of oil price shocks in the variables of interest through impulse responses. The model is further conditioned on the changes in average world gold prices in order to account for fluctuations in South Africa's exchange rate. In addition, the paper also represents an innovative contribution to inflation forecast processes in South Africa by employing probability event forecasts based on stochastic simulation, which account for future uncertainty.

The major findings indicate that the hypothesis of strict PPP is not supported by the data, whilst UIP holds in the long run. Impulse response functions show that oil prices slowly tend to raise inflation, which is followed by a tightening of the monetary policy, as short-term interest rates adjust quickly. Although the PPP hypothesis is rejected, however, the results clearly show that domestic prices, US prices, nominal exchange rate and domestic interest rates form a cointegrated vector for South Africa. Probability event forecasts indicate that inflation forecasts in South Africa are more likely to be in the designated ranges in the presence of incomplete PPP and UIP relations than in a more complete long run relationship, although inflation targeting forecasts are more likely to be met in the complete PPP model.

Dedication

This thesis is dedicated to my parents, Daniel Joel Muzima and Celeste Salomão Nhamunze. Thank you for your unwavering support.

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1. Introduction

In the early years of the post Bretton-Woods era, movements in exchange rates and import prices were expected to be closely associated with movements in national price levels. Moreover, international transmission effects between countries through differences in interest rates were assumed to play an important role in determining exchange rate fluctuations. These relationships, purchasing power parity (PPP) and uncovered interest parity (UIP) respectively, have since attracted the attention of policy makers for various reasons. Firstly, due to the need to understand the impact of exchange rate volatility on domestic inflation, or 'exchange rate pass-through', and secondly due to the consequences that an interest rate differential would imply in terms of a perceived risk premium.

A debate on the validity of long run PPP and UIP has been the focus of interest in a vast number of empirically oriented papers; see, for instance, Kravis and Lipsey (1977), Isard (1977), Richardson (1978), Giovannini (1988), Froot and Thaler (1990). Generally, the hypotheses in support of these fundamental relationships have been empirically rejected.[1] Johansen and Juselius (1992) attribute the failure of these studies to single-equation bias. Furthermore, they point to the importance of considering the interaction between exchange rates, interest rates, and prices in the goods and assets markets, in a simultaneous equation model, as well as the importance of distinguishing between short-run and long run effects as a solution to improving the efficiency of the empirical estimates.

Most recently, Pesaran, et al (2000), and Garratt et al (2003) using a new modeling strategy that provides a practical approach to incorporating long run structural relationships, have found evidence in support of long run UIP, as well as a modified version of the PPP, which accounts for the impact of domestic interest rates on inflation. This work constitutes an extension of the empirical analysis presented in Johansen and Juselius (1992) and Pesaran and Shin (1996).

[1] See, for example, Levich (1985), Dornbusch (1989), and Goldberg and Knetter (1997).

This paper employs a simple framework within which to investigate the long run validity of the PPP and UIP relations between South Africa and the United States. The empirical approach closely follows that of Pesaran et al (2000). A VAR model in levels, under the assumption of cointegration, is thus considered. This model is used to describe the statistical variation of the data without imposing *a priori* restrictions. Instead, this modelling strategy allows us to formulate a set of long run structural hypotheses as suggested by economic theory.

The line of argument presented in this paper is as follows. It begins with a general discussion of the literature surrounding the long run validity of the PPP and UIP in developing countries and in South Africa in particular. This allows us to understand the context within which the empirical study is conducted. It also provides a benchmark with which to compare the results of this study. Thereafter, we discuss the empirical methodology. The vector autoregressive model with Gaussian errors is briefly discussed, as we choose the empirical model. A VAR model with no intercepts and no trends appears to be the most appropriate for modelling PPP and UIP, given the data.[2] The model also includes the effects of oil price shocks into domestic inflation, interest rates and exchange rates. This is done by adding the logarithm of oil prices in the endogenous *I(1)* variables list, and then testing the validity of excluding the level of oil prices from the cointegrating relations. The model is further conditioned on the changes in average world gold prices, which are included as an exogenous *I(0)* variable in order to account for volatility in SA's exchange rate.

The goal of this research is to contribute to the debate regarding the long run validity of two structural relationships, viz. the PPP and UIP hypotheses, in the context of the South African economy. Furthermore, this research represents an innovative contribution to inflation forecasts in South Africa since it employs probability event forecasts based on stochastic simulation in order to account for future uncertainty. This methodology,

[2] The absence of intercepts and linear trends in the model implies that there is no linear growth trend in both PPP and UIP relations. Moreover, the inclusion of a linear trend in the PPP relations is not justified by economic theory (see Pesaran and Shin, 1998; and Garratt et al, 2003). Nevertheless, the inclusion of an intercept in both relationships would reflect the existence of information disparities in the goods markets and risk premia in capital markets. However, cointegrating rank tests did not confirm this hypothesis.

initially developed in Garratt et al (2003b), is used to predict the following events: (i) the probability of inflation falling below a certain level ranging from 1.5% up to 6%; (ii) the probability of the South African Reserve Bank (SARB) keeping inflation target between 3% and 6%. This study also aims to analyse the influence of oil price shocks on the domestic inflation, interest rates and exchange rates. More specifically, the research tests the following hypotheses: (i) that PPP exists, at least in the long run; (ii) that UIP holds in the long run; (iii) whether the logarithm of oil prices does enter directly in the cointegrating vectors; (iv) that Oil prices, US prices, world gold prices and US interest rates are exogenous to the system.

The paper is structured as follows: Section 1 presents the background of the study and outlines the objectives and the relevance of this research. Section 2 presents the research methodology. In this regard, it explains the theoretical background surrounding issues of the PPP and UIP as well as the empirical background from previous studies of PPP and UIP in South Africa, middle income countries, and other developing countries. The methodology of the vector error correction model (VECM) for multivariate cointegration analysis is also discussed in Section 2. Section 3 presents the results of the long run estimates. Different structural relations between prices, exchange rates, and interest rates are tested. Moreover, the section reports on the results of tests of weak exogeneity with respect to the long run parameters. In Section 4, the dynamic stability of the system is analyzed through impulse response functions. This allows us to investigate the impact of oil price shocks in the variables of interest. Section 5 computes probability event forecasts of inflation based on stochastic simulations that account for future uncertainty. Finally, Section 6 concludes the paper.

2. Research Methodology

2.1. Theoretical background

The idea of PPP is based on the presence of goods market arbitrage, and on the assumption that the price of a common basket of goods will be equal in different countries when measured in a common currency. Theorists of PPP argue that information disparities, transportation costs or the effects of tariffs and non-tariff barriers are likely to create considerable deviations from (absolute) PPP in the short run (Patterson, 2000). However, the weaker form of PPP is expected to hold even in the presence of such imperfections, if the size of these influences has a constant mean over time (Garratt et al, 2003). The relative version of PPP can be expressed as:

$$P_{t+1} = E_{t+1} PF_{t+1} \exp(\xi_{ppp,t+1}) \qquad (1)$$

where P_t is the domestic price index, PF_t is the foreign price index, E_t is the nominal exchange rate and the term in brackets captures the deviations from PPP. Here, $\xi_{ppp,t+1}$ is assumed to follow a stationary process[3], capturing short-run variations in information disparities, and the effects of tariffs and non-tariff barriers.

The second arbitrage condition is based on the UIP relationship. This relation represents the equilibrium condition in international capital markets. This states that any differential in interest rates across countries must be offset by expected exchange rate changes to eliminate the scope for arbitrage. However, empirical evidence[4] shows that the presence of transaction costs, risk premia and speculative effects can result in short-run deviations from the UIP. The interest rate parity (IRP) can be defined as follows:

$$(1 + r_t) = (1 + rf_t)\left(1 + \frac{E^e_{t+1} - E_t}{E_t}\right)\exp(\xi_{uip,t+1})$$

$$= (1 + rf_t)\left(\frac{E^e_{t+1}}{E_{t+1}}\right)\left(1 + \frac{\Delta E_{t+1}}{E_t}\right)\exp(\xi_{uip,t+1}) \qquad (2)$$

[3] The basis of this assumption is based on the grounds that a linear combination of domestic prices, foreign prices and the nominal exchange rate represents a stationary long run relationship.
[4] See Levich (1985).

where r_t and rf_t represent domestic and foreign interest rates, respectively, and $\xi_{uip,t+1}$ is the risk premium associated with the effects of bonds and foreign exchange uncertainties on risk averse investors. It is furthermore assumed that this term is stationary[5]. Moreover, assuming that inflation and exchange rate expectations are rational, then we can write:

$$P^e_{t+1} = P_{t+1}\exp(\xi^e_{p,t+1});\ E^e_{t+1} = E_{t+1}\exp(\xi^e_{e,t+1}) \qquad (3)$$

The relations in *(1)* and *(2)* can be written in terms of observables as follows:

$$\ln(P_{t+1}) = \ln(PF_{t+1}) + \ln(E_{t+1}) + \xi_{ppp,t+1} \qquad (4)$$

$$\ln(1+r_t) = \ln(1+rf_t) + \eta_{\Delta e,t+1} + \xi_{uip,t+1} + \eta^e_{e,t+1} \qquad (5)$$

where *(4)* and *(5)* represent the log-linear versions of the PPP and UIP respectively, $\eta_{\Delta e,t+1} = \Delta\ln(E_{t+1})$ is the long run structural disturbance arising from changes in the spot exchange rate, and $\eta^e_{e,t+1}$ is a long run structural disturbance deriving from uncertainties regarding the expected (future) exchange rate.

2.2. Empirical background

The subject of whether PPP holds in South Africa has been investigated in a number of papers. Tsikata (1998) and Subramanian (1998), for example, show that the effective nominal depreciation of the rand during the 1990s is reflected in higher prices of imported goods. However, these results are found to be sensitive to the choice of price aggregates, sample period, and inclusion of structural breaks. In another study by Aron et al (1997), it is found that the real exchange rate in South Africa is non-stationary, which implies that the strict form of PPP does not hold. Nevertheless, fluctuations in the real exchange rate are found to be cointegrated with a set of economic 'fundamental' variables such as trade liberalization, terms of trade, government expenditures, capital flows and official reserves.

In contrast, Jonsson (2001), using a structural vector error-correction model (VECM) to examine money demand relationship and the PPP in South Africa, finds a stable long run relationship among domestic prices, foreign prices and nominal exchange rates.

[5] This term is deemed stationary given that long run disturbances in the capital markets are mean-reverting or stationary.

Furthermore, it is found that shocks to the exchange rate affect domestic prices but have virtually no impact on output.

Calderon and Duncan (2003) using long-span data for Chile found evidence in favour of PPP. Moreover, their results were robust to changes in the domestic price index, to changes in the sample period, and to the econometric technique. In other study for New Zealand, Stephens (2004) uses PPP and UIP to estimate a time-varying equilibrium for the $NZ/$US nominal exchange rate over the period 1992 to 2003. The study found evidence in support of PPP while the data did not support the strictest form of UIP. In contrast, Hatzinikolaou and Polasek (2005) using quarterly data from the post-float period 1984:1-2003:1 for Australia, found evidence in support of PPP and UIP when commodity prices were included in the cointegrating relations.

Studies of the PPP for developing countries have been empirically analysed by imposing homogeneity either on all the coefficients (unit roots test of the real exchange rate), or on some of the coefficients (cointegration tests between the two prices in common currencies), or without imposing any restriction on the coefficients at all. The conclusions obtained vary mainly as a function of the methodology and sample period (see Choudhry, 1999; Nagayasu, 2002; Holmes, 2002; and Achy, 2003).

These studies have not been consistent, but have provided mixed evidence on the validity of the PPP. For example, Bahmani-Oskooee and Goswami (2005) use the Johansen-Juselius (1990) multivariate cointegration framework to test the validity of PPP in the black market exchange rates of emerging economies. They find that, although the variables are cointegrated, domestic price and foreign prices are not weakly exogenous. In addition, they also reject the hypothesis of PPP. Wickremasinghe (2002), using unit roots tests for the real exchange rate in Sri Lanka, overwhelmingly rejects the hypothesis of the long run validity of the PPP.

Pesaran et al (2000), re-examining the work of Johansen and Juselius (1992) for the UK, find evidence in support of a modified version of the PPP that accounts for the impact of domestic interest rates on inflation. Their modelling strategy represents an innovation on empirical grounds because, unlike Johansen and Juselius (1992), they also allow for the presence of exogenous *I(1)* variables. Most importantly, they show that tests of restrictions based on this strategy tend to perform well even in small samples.

Garratt et al (2003), using a new modelling strategy that provides a practical approach to incorporating long run relationships suggested by economic theory to model the UK economy, find evidence in support of both PPP and UIP relationships. In addition, the dynamic properties of their model account for the impact of oil price shocks on domestic inflation, as well as the impact of monetary policy changes on the variables of interest.

2.3.The Estimation Methodology

The basic model consists of six variables of interest, which together form a six-dimensional VECM with Gaussian errors. A *k*-dimensional VAR is employed in this particular estimation using the Johansen multivariate cointegration technique.[6] For empirical purposes, a log-linear approximation of the two equilibrium relationships (PPP and UIP) is employed. These constitute the long run relationships of the model, and assume the following form:

$$p_t - pf_t - e_t = \xi_{ppp,t+1} \qquad (6)$$

$$r_t - rf_t = \xi_{uip,t+1} \qquad (7)$$

where $p_t = \ln(P_t)$, $pf_t = \ln(PF_t)$, $e_t = \ln(E_t)$. The model does not allow for intercepts and linear trends to ensure that the disturbances $\xi_{ppp,t+1}$, and $\xi_{uip,t+1}$ have zero means. These long run reduced form disturbances are related to the long run structural disturbances, the $\eta_i's$:

$$\xi_{ppp,t+1} = \eta_{ppp,t} \qquad (8)$$

$$\xi_{uip,t+1} = \eta_{uip,t+1} + \eta^{e}_{e,t+1} + \eta_{\Delta e,t+1} \qquad (9)$$

[6] See Johansen and Juselius (1990).

According to Garratt et al (2003a:424), the relationships between the long run structural disturbances, $\eta_i's$, and the long run reduced form disturbances, $\xi_i s$, illustrate the difficulties involved in identifying the effects of changes in particular structural disturbances on the dynamic behaviour of the macroeconomy. For example, $\xi_{uip,t+1}$ is composed of the three structural disturbances, $\eta_{uip,t+1}, \eta^e_{e,t+1}, \eta_{\Delta e,t+1}$, which represent the different factors that could be responsible for disequilibria between domestic and foreign interest rates. Without further *a priori* restrictions, the effect of particular structural disturbances $\eta_i's$ cannot be identified, since there are many more long run structural disturbances than there are long run reduced form disturbances.[7]

The two long run relationships of the model, *(6)-(7),* can be written more compactly as:

$$\xi_t = \beta' z_{t-1} \qquad (10)$$

where:

$$z_t = (p_t, pf_t, e_t, r_t, rf_t, o_t)', \text{ and } \xi_t = (\xi_{pppt}, \xi_{uipt}) \qquad (11)$$

and

$$\beta' = \begin{bmatrix} 1 & -1 & -1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & -1 & 0 \end{bmatrix}' \qquad (12)$$

In modelling short-run dynamics, we also include a set of seasonal dummies, corresponding to the first three quarters, *s1, s2, s3*, respectively. The model is also conditioned on changes in world gold prices to account for fluctuations in the nominal exchange rate. For estimation purposes, o_t (logarithm of oil prices) is considered to be a 'long run forcing' variable. Note that foreign prices and interest rates are treated as endogenous variables for pragmatic reasons.[8]

[7] For further details, see Garratt et al (2003a).

[8] Such treatment of foreign variables does not; however, seem to be necessary in the case of small open economies where it is unlikely that changes in domestic variables have a significant impact on the long run evolution of foreign prices or interest rates. Moreover, the endogenous treatment of foreign prices and interest rates would involve loss of efficiency in estimation if they were in fact long run forcing variables.

The treatment of oil prices as 'long run forcing' draws on the new modelling strategy adopted by Pesaran et al (2000) in order to model oil price effects on macroeconomic variables. This approach allows for the possibility of testing the restriction of whether or not oil prices enter into the cointegrating relations. In cases where the restriction is not rejected, then it can be imposed in the underlying long run relationship.

Although oil prices are treated as 'long run forcing', it is nonetheless necessary to specify an oil price equation in order to analyse the short-run dynamics and impulse response functions. Therefore, the following general specification for the evolution of oil prices is adopted:

$$\Delta o_t = \lambda_0 + \sum_{i=0}^{k-1} \lambda_{oi} \Delta z_{t-i} + u_{0t} \qquad (13)$$

where u_{0t} has zero mean and a constant variance, and is uncorrelated with oil price shocks. The specification in *(13)* ensures that oil prices are 'long run forcing' in the sense that lagged changes of the endogenous and exogenous variables of the model are allowed to influence current oil prices, but it also excludes the possibility that error correction terms have any impact on oil price changes.[9]

Assuming that the variables in z_t are *I(1)*, then the modelling strategy consists of estimating a restricted *k-1*-dimensional VAR. According to Johansen and Juselius (1990), given a set of *k*-variables, we may expect to have *r* cointegrating relationships, such that $0 \leq r \leq k\text{-}1$. Under the assumptions discussed above, the analysis will be based on the following conditional vector error correction model (VECM):

$$\Delta z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \prod z_{t-k+1} + \delta_{oi} \Delta go_t + \delta_{oi} \Delta go_{t-1} + \psi_t S_t + \varepsilon_t \qquad (14)$$

where $\epsilon_1 \ldots \epsilon_t$ are i.i.d. $N_p(0, \sum)$, S_t are seasonal dummies, and Δgo_t represents gold price changes. Assuming the following null hypothesis of the existence of *r* cointegrating vectors:

$$H_1(r): \prod = \alpha\beta' \qquad (15)$$

[9] This prediction is supported by the data as it is illustrated by the results in Table 7, in section 3.5.

The hypothesis in *(15)* can be used to test for the number of cointegrating vectors. In this empirical estimation, the anticipated number of cointegrating vectors is two, given the theoretical discussion surrounding long run equilibrium relationships in both goods and asset markets, namely the PPP and UIP conditions. However, since both prices and exchange rates do not enter into the UIP relation and, similarly, as the interest rate differential does not enter directly into the PPP relation, then this will lead to an over-identified structural system:

$$\prod z_{t-k+1} = \alpha\beta' z_{t-k+1} = \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \\ \alpha_{41} & \alpha_{42} \\ \alpha_{51} & \alpha_{52} \\ \alpha_{61} & \alpha_{62} \end{bmatrix} \begin{bmatrix} 1 & \beta_{12} & \beta_{13} & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & \beta_{25} & 0 \end{bmatrix}' \begin{bmatrix} p_t \\ pf_t \\ e_t \\ r_t \\ rf_t \\ o_t \end{bmatrix}_{t-k+1} \quad (16)$$

It is important to note that oil prices are entered in the structural system as an exogenous *I(1)* variable in order to improve estimation efficiency as well as to explicitly account for its impact on the variables of interest. The parameters of the model *(16)* are estimated according to the approach described in Pesaran et al (2000), which comprises the following steps: (i) selecting the order of the underlying VAR model (using model selection criteria such as the Akaike Information Criterion (AIC) or the Schwarz Bayesian Criterion (SBC)); (ii) selecting the deterministic components; (iii) testing for the number of cointegrating relations; and (iv) computing maximum likelihood estimates of the model's parameters subject to exact and over-identifying restrictions and testing their validity; and lastly, (v) computing impulse response functions.

This model also allows us to analyse the dynamic characteristics of the structural system. This is done by means of generalised impulse response functions (IRFs).[10] According to Pesaran and Shin (1998), an impulse response function measures the time profile of the effect of shocks at a given point in time on the future values of the variables in a dynamic

[10] Here we use generalised impulse responses rather than orthogonalised impulse responses since they are not sensitive to the order in which the variables enter into the system. For further details about impulse responses literature see: Lütkepohl and Breitung (1996); Koop et al (1996); Kilian (1997); Pesaran and Shin (1998); and Ivanov and Kilian (2005).

system. An intuitive illustration of how IRFs are computed can be stated as follows: Assume a VAR(1) model of the form: $z_t = \Phi z_{t-1} + \varepsilon_t$; $t=1,...,T$. *(17)*

where z_t is a $kx1$ vector of variables and Φ is a kxk matrix of unknown coefficients. Express, z_t as an infinite vector moving average (VMA) representation:

$$z_t = \varepsilon_t + \Theta_1 \varepsilon_{t-1} + \Theta_2 \varepsilon_{t-2} + ... = \sum_{i=0}^{\infty} \Theta_i \varepsilon_{t-i} \qquad (18)$$

Then the plot of $\frac{\partial z_{t+h}}{\partial \varepsilon_t} = \Theta_{h,1}$ against the horizon $h=0,1,2...$ is called the IRF with respect to the equation of the innovation.

3. Empirical and Estimation Results

3.1. The Data

The quarterly data used in the estimations were obtained from the *International Financial Statistics Online Database*, published by the International Monetary Fund. In order to accommodate the requirements of the econometric model, all variables in the z_t space are expected to be *I(1)*. The basic variables of interest are: $p_t, pf_t, e_t, r_t, rf_t, o_t$. Price variables are measured in terms of the standard CPI index, while nominal exchange rate is defined in terms of units of home currency per unit of foreign. Interest rates are measured by monthly average Treasury bill discount rate and oil prices are represented by average price of crude oil. A detailed description of these variables is given in Table 1. The data correspond to the period 1972q1-2006q2, hence covering the post-Bretton Woods era of a floating exchange rate system. Note that the term 'foreign' is hereafter used as corresponding to the US economy.

Table 1: List of variables and their description in the core model

p_t: natural logarithm of the South Africa's CPI index (2000=100)
pf_t: natural logarithm of the United States' CPI index (2000=100)
e_t: natural logarithm of the nominal exchange rate of R/$
r_t: South Africa's three month Treasury Bill average discount rate
rf_t: United States' three month Treasury Bill average discount rate
o_t: natural logarithm of oil prices, measured as the Average Price of Crude Oil
S1: seasonal dummy, taking the value of 1 in the 1st quarter and 0 otherwise
S2: seasonal dummy, taking the value of 1 in the 2nd quarter and 0 otherwise
S3: seasonal dummy, taking the value of 1 in the 3rd quarter and 0 otherwise
S4: seasonal dummy, taking the value of 1 in the 4th quarter and 0 otherwise

go_t: natural logarithm of gold prices, measured as the Average Price of Gold in London
Source: International Financial Statistics, Online database (IFS, 2006).

The graphs of the series in levels and in first differences are presented in Figures 1 and 2 respectively. For prices and interest rates, the series in first differences display a large degree of fluctuation in both countries. This can be attributed to the impact of the second oil shock (1979) as well as to significant structural changes described in Jonsson (2001:244) and Johansen and Juselius (1992:218) as follows: (i) change in monetary policy regime in SA in 1980q1; (ii) introduction of the Depository Institutions Deregulation and Monetary Control Act in the US in 1980; (iii) introduction of the Depository Institutions Act of 1982 in the US; and (iv) the Reserve Bank's monetary policy regime shift in 1991q1 in SA, during governor Stals' mandate.

It is possible that these events changed some of the parameters[11] of the model, as Lucas (1976) would argue. It is thus necessary to exercise care when interpreting the results. Nevertheless, in this estimation, the impact of the second oil shock is directly accounted for by the inclusion of world oil prices in the system. The effects of other policy interventions are left to be accounted for by the general specification of the short-run dynamics.[12]

The Augmented Dickey-Fuller (ADF) test statistics, computed over the period 1972q1-2006q2, for the levels and first differences of the variables are reported in Table 2. The results suggest that it is reasonable to assume that all variables in the z_t space are *I(1)*. This suggests the appropriateness of the VECM as an estimation methodology. With regard to the variables in the core model, the unit root hypothesis is rejected when the ADF tests are applied to their first differences, but there is no evidence with which to reject the unit root hypothesis when the tests are applied to their levels. Similar results were obtained using Perron tests.

[11] These interventions and shocks to the economy are more likely to have changed the short-run parameters rather then the long run fundamental relationship between the variables in the system.
[12] Explicit efforts to account for the impact of changes of policies in the economy did not improve the quality of the results.

Perron's unit root tests were conducted using the Perron (1994) innovation outlier model which controls for structural breaks. The test consists of testing whether the coefficient on y_{t-1} is statistically different from one in the following regression:

$$y_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \delta DTB_t + \alpha y_{t-1} + \sum_{i=1}^{n} \alpha_i \Delta y_{t-i} + \varepsilon_t \qquad (19)$$

where y_t corresponds to the variables in the core model. $DU_t = 1$; $DT_t = t - T_b$ *if* $t > T_b$ and *0* otherwise. $DTB_t = 1$ *if* $t = T_b + 1$, and *0* otherwise; where T_b denotes the time of structural break. Different structural breaks are imposed: (i) 1991q1 to control for the impact of monetary policy regime shifts in South Africa on changes in the mean value and variance of inflation; (ii) 1972q3 and 1979q1 to control for the impact of the first and second oil crisis on inflation in SA and the US respectively, (iii) 1980q1 to control for the impact of the monetary policy regime change on SA's exchange rates and interest rates, (iv) 1980 and 1982 to account for the introduction of the Depository Institutions Deregulation and Monetary Control Act and its impact on US interest rates.

Looking at the unit root tests of the series in first differences there is, however, some ambiguity regarding the order of integration of price variables. Application of the τ_u, Φ_1[13] and Φ_3[14] tests to Δp_t and Δpf_t resulted in mixed results. For example, the hypothesis that inflation has a unit root is not rejected in at least two of the tests, τ_u, and Φ_1 respectively, but when considering Φ_3, SA's inflation becomes stationary with a time trend while the US's inflation has a unit root. However, the application of the Perron test consistently indicates that Δp_t and Δpf_t are stationary, which implies that prices are *I(1)* variables. Therefore, we proceed to estimate the results under the assumption that all variables of interest are *I(1)* .

[13] Φ_1 Test consists on running the following regression, $\Delta y_t = \mu + (p-1)y_{t-1} + \sum_{j=1}^{p} \alpha_i \Delta y_{t-i} + u_t$ and then testing the joint null hypothesis that: $(p-1)=0$; $\mu = 0$ (series has unit root) against the alternative that $(p-1)<0$; $\mu \neq 0$ (series is stationary).

[14] Φ_3 Test is done by running the regression, $\Delta y_t = \mu + \beta t + (p-1)y_{t-1} + \sum_{j=1}^{p} \alpha_i \Delta y_{t-i} + u_t$ and then testing the joint null hypothesis that: $(p-1)=0$; $\beta = 0$ (series has unit root and no trend) against the alternative that $(p-1)<0$; $\beta \neq 0$ (series is stationary with trend).

Table 2: Augmented Dickey-Fuller Unit Root Tests and Perron Structural Break Tests

Variables	Model	τ	Φ_1	τ_u	μ	τ_τ	β	Φ_3	Perron t-stat	Breaks
				A: Series in levels						
p_t	ADF(8)	-0.002	3.60	-2.11	0.016	1.30	-0.000	3.45	-2.62	Yes
pf_t	ADF(10)	-0.003	5.23	-2.74	0.020	-2.05	0.000	3.09	-4.28	Yes
e_t	ADF(6)	0.004	3.52	-1.16	0.019	-1.79	0.001	2.11	-1.69	Yes
R_t	ADF(1)	-0.007	3.64	-2.83	0.631	-2.71	0.000	3.63	-3.85	Yes
RF_t	ADF(7)	-0.008	1.40	-1.63	0.230	-2.85	-0.004	3.28	-2.46	Yes
o_t	ADF(6)	0.004	7.19	-1.42	0.304	-1.72	0.000	6.86	-2.20	Yes
go_t	ADF(2)	0.002	4.95	-1.97	0.278	-2.23	0.001	4.16	-3.56	Yes
				B: Series in first differences						
Variables	Model	τ	Φ_1	τ_u	μ	τ_τ	β	Φ_3	Perron t-stat	Breaks
Δp_t	ADF(2)	-0.038	4.45	-2.78	0.006	-4.20	0.000	9.24	-6.06	Yes
Δpf_t	ADF(2)	-0.028	2.04	-2.12	0.001	-3.34	0.000	4.60	-5.30	Yes
Δe_t	ADF(5)	-0.639	13.70	-5.05	0.014	-5.12	0.000	13.84	-5.46	Yes
ΔR_t	ADF(0)	-0.571	18.72	-6.90	0.021	-6.95	-0.003	19.30	-7.25	Yes
ΔRF_t	ADF(6)	-0.858	12.80	-4.76	-0.015	-4.76	-0.000	12.75	-5.70	Yes
Δo_t	ADF(5)	-0.715	9.77	-4.70	0.019	-4.71	-0.000	9.89	-5.59	Yes
Δgo_t	ADF(1)	-0.648	20.61	-4.89	0.008	-4.87	-0.000	20.82	-6.50	Yes

Notes: The longest lag in the ADF model is chosen based on Akaike Information Criteria (AIC) in order to ensure no autocorrelation in the errors; e_t is the log of the nominal exchange rate; p_t and pf_t are the logarithms of domestic and foreign prices; R_t and RF_t are domestic and foreign interest rates; o_t is the logarithm of world oil prices; and go_t is the logarithm of world gold prices.
Sample period: 1972Q1 to 2006Q2.

Critical values	τ	τ_u	τ_τ	Φ_1	Φ_3	**Perron's test critical value**	
5%	-1.930	-2.856	-3.408	4.968	6.397	5%	-5.08
10%	-1.604	-2.559	-3.119	3.835	5.433		

3.2. Selection of the Order of the VAR

Table 3 presents the results of the test for selection of the appropriate order of the VAR. The Schwarz Bayesian Criterion (SBC) suggests a VAR of order 1, the Akaike Information Criterion (AIC) selects a VAR of order 4. We select a VAR of order 2 to avoid over-parameterization and the 'vanishing degrees of freedom'[15] problem (see Johnston and DiNardo, 1997; and Kilian, 1997). An inspection of the single equations in the VAR shows that the assumption of normally distributed errors is rejected in the equations for e_t, r_t, rf_t, o_t. This is understandable if we consider the two major hikes in oil prices.

[15] The phenomenon might be described when the number of unknown coefficients can rapidly approach the sample size. For example, a system of 20 variables with 4 lags would require an estimation of at least 80 coefficients in each equation in the VAR.

Table 3: Test Criteria for selecting the order of the VAR

Test Statistics and Choice Criteria for Selecting the Order of the VAR Model Based on 134 observations from 1973Q1 to 2006Q2. Order of VAR = 4						
List of variables included in the unrestricted VAR:						
p_t	pf_t	e_t	r_t	rf_t	o_t	
List of deterministic and/or exogenous variables:						
INPT	S1	S2	S3	Δgo_t	Δgo_{t-1}	
Order		LL	AIC	SBC	LR test	Adjusted LR test
4		1111.2	931.2	670.4	------	------
3		1053.3	909.3	700.7	CHSQ(36)= 115.7 [0.000]	89.8[0.000]
2		1019.1	911.1	754.6	CHSQ(72)= 184.2 [0.000]	142.9[0.000]
1		952.1	880.1	775.6	CHSQ(108)= 318.2 [0.000]	246.9[0.000]
0		-617.6	-653.6	-705.8	CHSQ(144)=3457.6[0.000]	2683.5[0.000]

Notes: AIC=Akaike Information Criterion SBC=Schwarz Bayesian Criterion
The Information Criteria Values (AIC) and (SBC) were computed using MICROFIT 4.1. Note that these values are to be maximized as opposed to the more usual minimization in other software packages.

3.3. Testing for Cointegrating Rank

Selected the order of the VAR, the next stage consists of selecting the appropriate empirical model that fits the data well. Model *(14),* together with the reduced rank hypothesis *(15),* thus constitute the starting point of the empirical analysis. Using a VAR(2) model with no intercepts and linear trends coefficients, and treating the oil price variable, o_t, as a weakly exogenous *I(1)* variable, we computed the Johansen's 'trace' and 'maximal eigenvalue' statistics in Table 4.

Based on the trace statistic and $\lambda_i$max; displayed in Table 4 we uniformly[16] find two cointegrating vectors at 95% critical value. This is consistent with the theoretical discussion regarding the possible number of long run relationships predicted by the theory.[17]

Table 4: Test statistic for cointegrating rank (Model: No Intercepts and no trends; VAR=2)

[16] The complete disagreement between the two procedures for testing the number of cointegrating relations may often result in conflicting conclusions and the decision concerning the choice of r, the number of cointegrating vectors, may be made based on other information, such as economic theory.

[17] The $\lambda_i$max is rejected for the null hypothesis of r=0 and r=1since 43.4 > 36.3, and 37.8 > 29.9, respectively. Using the trace test, we also find that the null hypothesis of r=0 is rejected in favour of r ≥1, given that 115.3 > 83.2. The same conclusion follows for the hypothesis that r≤1 against that of r≥2, since 71.8 > 59.3 (see Table 4).

Null	Alt.	λ_i	λ_i max statistic	95% quantile	Null	Alt.	Trace statistic	95% quantile
r = 0	r =1	0.275	$-T\ln(1-\lambda_1) = 43.4 > 36.3$*		r = 0	r ≥1	$-T\sum_{i=1}^{k}\ln(1-\lambda_i) =$	115.3 > 83.2*
r ≤1	r =2	0.244	$-T\ln(1-\lambda_2) = 37.8 > 29.9$*		r ≤1	r ≥2	$-T\sum_{i=2}^{k}\ln(1-\lambda_i) =$	71.8 > 59.3*
r ≤2	r =3	0.119	$-T\ln(1-\lambda_3) = 17.1 < 23.9$		r ≤2	r ≥3	$-T\sum_{i=3}^{k}\ln(1-\lambda_i) =$	34.0 < 39.8
r ≤3	r =4	0.074	$-T\ln(1-\lambda_4) = 10.4 < 17.7$		r ≤3	r ≥4	$-T\sum_{i=4}^{k}\ln(1-\lambda_i) =$	16.9 < 24.1
r ≤4	r =5	0.043	$-T\ln(1-\lambda_5) = 5.9 < 11.0$		r ≤4	r =5	$-T\sum_{i=5}^{k}\ln(1-\lambda_i) =$	6.5 < 12.4
r ≤5	r =6	0.004	$-T\ln(1-\lambda_6) = 0.5 < 4.2$		r ≤5	r =6	$-T\ln(1-\lambda_6) =$	0.5 < 4.2

*Notes: * indicates rejection of the null hypothesis at 5% level.*
The statistics refer to Johansen's log-likelihood based 'trace' and 'maximal eigenvalue' statistics and are computed using 135 observations for the period 1972q4-2006q2.

3.4. Long run estimates

3.4.1. Long run estimates with oil prices as an exogenous I(1) variable

Given that we have already established the number of long run relationships ($r=2$), we now turn our attention to their estimation and to identifying factors that might be responsible for their possible breakdown. We denote the two cointegrating vectors β^* associated with $z_t^* = (p_t, pf_t, e_t, r_t, rf_t, o_t)'$ by $\beta_1^* = (\beta_{11}, \beta_{12}, \beta_{13}, \beta_{14}, \beta_{15}, \beta_{16})'$ and $\beta_2^* = (\beta_{21}, \beta_{22}, \beta_{23}, \beta_{24}, \beta_{25}, \beta_{26})'$, respectively, with β_1^* viewed as explaining domestic prices and β_2^* explaining domestic interest rates. Exact identification of these vectors requires the imposition of two restrictions per vector. Therefore, the following exact identifying constraints are chosen: $\beta_{11} = 1; \beta_{15} = 0; \beta_{24} = 1; \beta_{21} = 0$, and '*' represents the unrestricted coefficients.

$$H_E : \beta_*' = \begin{bmatrix} 1 & * & * & * & 0 & * \\ 0 & * & * & 1 & * & * \end{bmatrix}'$$

Table 5 presents the results of the long run estimates with oil prices as an exogenous *I(1)* variable. The first two eolumns show the results for the exactly identified model. The first vector (β_1^*) corresponds to the PPP relation, whereas the second (β_2^*) corresponds to the UIP. The first vector is normalized on SA prices, and the second on SA interest rates. All

explanatory variables have the expected signs, and the numerical elasticities are consistent with the theory.[18] However, the coefficient on the US prices is less than unity, which contrasts with the strict form of PPP.

Table 5: Long run estimates with oil prices as an exogenous I(1) variable

Variables	Models							
	Exactly-identified VAR(2)		Over-identified VAR(2) and H_{oil}		PPP, UIP and H_{oil} VAR(2)		Weak form PPP and UIP model and H_{oil}	
	β_1^*	β_2^*	β_1^*	β_2^*	β_1^*	β_2^*	β_1^*	β_2^*
p_t	1.0000	0.0000	1.0000	0.0000	1.0000	0.0000	1.0000	0.0000
pf_t	-0.4513 (0.2151)	-8.6877 (8.1499)	-0.4584 (0.2201)	-9.0011 (8.4807)	-1.0000	0.0000	-0.4142 (0.2381)	0.0000
e_t	-0.8188 (0.4090)	14.7544 (17.3336)	-0.9045 (0.3571)	14.5834 (16.9480)	-1.0000	0.0000	-1.0000	0.0000
R_t	-0.1116 (0.0454)	1.0000	-0.1157 (0.0469)	1.0000	0.0000	1.0000	-0.4482 (0.5612)	1.0000
RF_t	0.0000	-2.4329 (2.6586)	0.0000	-2.5065 (2.7626)	0.0000	-1.0000	0.0000	-1.0000
o_t	-0.0812 (0.1894)	-0.5695 (4.4103)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
LL-value	992.9		992.8		968.3		991.7	
P-value			χ^2 [2]=0.20[0.904]		χ^2 [8]=49.21[0.000][19]		χ^2 [6]= 2.59 [0.148]	

Note: LL is the maximized value of the log-likelihood function. Asymptotic standard errors are given in parentheses.

This model allows us to test some structural relationships between the variables that are of significant importance. First, we consider the hypothesis that the level of oil prices does not enter into the cointegrating relations. This hypothesis is denoted hy H_{oil} with the following over-identifying constraints: $\beta_{11}=1;\beta_{15}=0;\beta_{24}=1;\beta_{21}=0;\beta_{16}=0$ and $\beta_{26}=0$. The corresponding long run estimates are reported in the third and fourth columns of Table 5. The log-likelihood ratio statistic associated with these 2 over-identifying restrictions is a chi-squared test. This test returned the value of 0.20 [0.904], which, thus, does not reject H_{oil}. Accordingly, we now test the PPP and UIP hypotheses given H_{oil}. The underlying over-identifying restrictions for the PPP case

[18] Recall that the strict form of the PPP requires the coefficient on domestic prices to be positive and equal to one, the exchange rate coefficient to be negative and equal to one, and that of foreign price to be negative and equal to unity.

[19] According to Pesaran and Shin (1998), in small samples the chi-squared test may not be valid due to its tendency to over-reject the null hypothesis. Therefore, bootstrapping technique may be required to find the critical values. However, in this paper we consider both models of PPP.

are: $\beta_{11}=1;\beta_{15}=0;\beta_{24}=1;\beta_{21}=0;\beta_{16}=0;\beta_{26}=0;\beta_{12}=-1;\beta_{13}=-1;\beta_{14}=0$, while in terms of the UIP hypothesis and H_{oil}, we have the following restrictions: $\beta_{11}=1;\beta_{15}=0;\beta_{24}=1;\beta_{21}=0;\beta_{16}=0;\beta_{26}=0;\beta_{25}=-1;\beta_{23}=0$; and $\beta_{22}=0$. The results (see Table 11 in appendix) indicate that there is evidence to reject the hypothesis of PPP even at 1% level.[20] In contrast, the hypothesis of UIP jointly with H_{oil} it is not rejected at 1% level.[21]

Next, we test the stability characteristics that are imposed by the theoretical relationships between the variables; that is, the hypothesis of proportionality and symmetry in the PPP and UIP relationships, by imposing the following restrictions: $\beta_{11}=1;\beta_{15}=0;\beta_{24}=1;\beta_{21}=0;\beta_{16}=0;\beta_{26}=0;\beta_{12}=-1;\beta_{13}=-1;\beta_{14}=0,\beta_{22}=0$ $\beta_{23}=0;\beta_{25}=-1$. The results are reported in the fifth and sixth columns of Table 5. The log-likelihood ratio statistic is 49.21[0.000], which is well above the 0.05 critical value of the chi-squared distribution with 8 degrees of freedom. Therefore, we jointly reject the hypothesis of PPP, UIP and H_{oil}.

However, the rejection of the PPP hypothesis is more likely to be related to the impact of domestic interest rates in the domestic price equation and due to the incomplete price 'pass-through' between South Africa and the United States. Therefore, it is possible to propose a modified version of the PPP relation, where the coefficients on US prices and domestic interest rates are unconstrained. The results of this model are reported in the last two columns of Table 5. The log-likelihood ratio statistic for testing the weak form of PPP jointly with UIP and H_{oil} is 2.59 [0.148], which is below the 0.05 critical value of the chi-squared distribution with 6 degrees of freedom. Consequently, the main cause of the breakdown of the PPP in the present application seems to be the existence of a statistically significant and positive relationship between domestic interest rates and

[20] A Log-likelihood statistic value of 30.32[0.000] was obtained. This is well above of the 0.05 chi-squared distribution with 5 degrees of freedom. Thus, the hypothesis of PPP jointly with H_{oil} is strongly rejected.
[21] The Log-likelihood ratio statistic for testing the UIP hypotheses is 14.98[0.010] which is well above the 0.05 critical value of the chi-squared distribution with 5 degrees of freedom, but does not exceed that of the 0.01 critical value.

domestic prices, as well as the incomplete price transmission between South Africa and the United States.

3.4.2. Long run estimates with gold prices as an exogenous I(1) variable

Changes in world gold prices have often been associated with fluctuations in the nominal exchange rate in South Africa. Nell (2000) argues that high gold prices have helped SA's exchange rate to remain fairly stable during the period 1973-1983. Moreover, Aron et al (1997) and MacDonald and Ricci (2004) find a statistically significant long run relationship between gold prices and the real exchange rate in SA. Given such evidence, it seems worthwhile to relax the assumption regarding oil prices and to investigate, instead, the impact of gold priees on the long run relationships. This is done by including gold prices as an exogenous *I(1)* variable, while leaving changes in oil prices as an exogenous *I(0)* variable. The results of this alternative specification are summarised in Table 6.

Table 6: Long run estimates with gold prices as an exogenous I(1) variable

Variables	Models							
	Exactly-identified VAR(2)		Over-identified VAR(2) and H_{gold}		PPP, UIP and H_{gold} VAR(2)		Weak form PPP and UIP model and H_{gold}	
	β_1^*	β_2^*	β_1^*	β_2^*	β_1^*	β_2^*	β_1^*	β_2^*
p_t	1.0000	0.0000	1.0000	0.0000	1.0000	0.0000	1.0000	0.0000
pf_t	0.2692 (0.7444)	7.7292 (12.5149)	-0.3702 (0.2542)	-16.3098 (46.7593)	-1.0000	0.0000	0.0283 (1.7828)	0.0000
e_t	-0.8231 (0.3896)	7.1662 (7.5709)	-0.9689 (0.3976)	30.8859 (99.4031)	-1.0000	0.0000	-1.0000	0.0000
R_t	-0.0924 (0.0673)	1.0000	-0.1451 (0.0721)	1.0000	0.0000	1.0000	-1.1342 (3.3626)	1.0000
RF_t	0.0000	-0.9412 (1.2462)	0.0000	-5.9818 (19.2212)	0.0000	-1.0000	0.0000	-1.0000
go_t	-0.6650 (0.7847)	-10.6823 (11.8563)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
LL-value	1069.7		1067.8		1042.8		1056.5	
P-value			χ^2 [2]=3.74[0.154]		χ^2 [8]=53.83[0.000]		χ^2 [6]=26.40[.000]	

Note: LL is the maximized value of the log-likelihood function. Asymptotic standard errors are given in parentheses.

The assumption that gold prices do not enter in the long run cointegrating relations is not rejected by the data at the conventional significance levels of 5% and 1%. The log-likelihood ratio statistic for testing the joint hypothesis of the over-identified model and H_{gold} is 3.74[0.154], which is below the 0.05 critical value of the chi-squared distribution

with 2 degrees of freedom. Most interesting is the fact that, when gold prices are included, neither PPP nor UIP are supported by the data. This implies that the long run stability conditions of convergence in the goods and asset markets are not met under the gold price specification.

Since the main interest of this research is to investigate the long run validity of the theoretical relationships (PPP and UIP), then we can proceed to examining the dynamic stability estimations only, based on the assumption that oil prices are exogenous *I(1)* due to the convergence properties of this model. The long run cointegrating estimates obtained under this specification have important policy implications. For example, the fact that the strict hypothesis of the PPP is rejected against its weak form points to the need to account for both domestic interest rates and foreign prices when forecasting domestic inflation. Failure to accommodate these effects can result in misrepresentation of the underlying long run dynamics of the variables in the system.

3.5. The Vector Error Correction Model

The estimates of the long run relations and the short-run dynamics of the model are provided in Table 7. The long run relations, which incorporate the joint test of restrictions[22] for the modified version of the PPP and the UIP, are summarised below:

$$\xi_{ppp,t} = p_t - 0.4142pf_t - e_t - 0.4482r_t \qquad (20)$$
$$\qquad\qquad (0.2381) \qquad (0.5612)$$

$$\xi_{uip,t} = r_t - rf_t \qquad (21)$$

The first equation, *(20)*, represents the modified version of the PPP relationship that allows for the impact of domestic interest rates on prices and for the incomplete price transmission in the goods markets between SA and the United States. Rejection of the strict PPP hypothesis in this study contrasts with the findings of Jonsson (2001), who

[22] The chi-squared test of restriction with 6 degrees of freedom returned the value of: χ^2 [6] = 2.59[0.148] which is well below the 0.05 critical value (see Table 5).

found that domestic prices, foreign prices and exchange rate displayed a stable long run relationship. However, the results herein are consistent with the findings of Aron et al (1997), who find that the real exchange rate is cointegrated with a set of 'fundamental' economic variables such as trade liberalization, terms of trade, government expenditures, capital flows and official reserves. In addition, these results are consistent with those in Pesaran et al (2000), who find that domestic prices, foreign prices, exchange rates and domestic interest rates form a cointegrating vector.

The estimates of the error correction coefficients (see Table 7) show that the long run relations make an important contribution in the short-run dynamics of both domestic and foreign prices, in the nominal exchange rate and the domestic interest rate, and that they are statistically significant. Moreover, the speed of adjustment in capital markets is slightly greater than that of the goods market. This is consistent with the prediction that convergence to equilibrium in the goods market tends to be slow due to imperfect information and nominal rigidities. The diagnostic statistics of the single equations in the VAR are generally satisfactory as far as the tests of the residual serial correlation, functional form, and heteroskedasticity are concerned. Figure 3 displays the resulting residuals of the single equations in the VECM.

3.6. Testing for weak exogeneity

This section tests whether $x=(\Delta o_t, \Delta rf_t, \Delta pf_t, \Delta go_t)$ are weakly exogenous for β in the sense that the lagged values of $y=(\Delta p_t, \Delta e_t, \Delta r_t)$ do not improve on the explanation of x obtainable from only the lagged values of x themselves. The test consists of imposing a joint restriction of the form $H_i : \alpha_{il} = 0$, with $i = 1; 2$, in the 'alpha' matrix given by equation *(16)*. The test is χ^2 distributed with four degrees of freedom since it is jointly performed with the hypothesis that the exogenous *I(1)* variable does not enter into the long-run cointegrating relations.

The results were computed by using the PcGive/PcFiml 9.10 version and are reported in Table 8. The null hypothesis of weak exogencity cannot be rejected for US prices since

the $\chi^2[4]=13.5[0.085]$ value is below the 0.05 critical value of the chi-squared distribution with 4 degrees of freedom.

Table 7: Error Correction Specification for the Over-identified Model: 1972Q4 – 2006Q2.

Equation	Δp_t	Δpf_t	Δe_t	Δr_t	Δrf_t	Δo_t
ξ_{1t}	-0.003*	-0.001**	-0.018*	-0.009*	0.149	-0.006
	(0.0009)	(0.0004)	(0.006)	(0.0011)	(0.085)	(0.016)
ξ_{2t}	-0.001*	-0.006*	-0.007*	-0.076***	0.033	-0.002
	(0.0004)	(0.0002)	(0.0003)	(0.048)	(0.037)	(0.006)
Δp_{t-1}	0.491*	0.007	-0.979**	5.835	16.381**	-1.117
	(0.081)	(0.036)	(0.489)	(9.136)	(7.061)	(1.303)
Δpf_{t-1}	0.072	0.581*	-0.356	-10.898	-16.730	1.528
	(0.158)	(0.069)	(0.955)	(17.812)	(13.768)	(2.541)
Δe_{t-1}	0.044*	0.004	0.112	0.279	-1.002	-0.025
	(0.015)	(0.007)	(0.091)	(1.699)	(1.314)	(0.243)
Δr_{t-1}	-0.008	-0.003	-0.005	0.361*	-0.035	-0.010
	(0.007)	(0.003)	(0.004)	(0.083)	(0.064)	(0.012)
Δrf_{t-1}	-0.001	0.002*	0.003	0.113	0.200**	0.017
	(0.001)	(0.004)	(0.006)	(0.113)	(0.087)	(0.016)
Δo_{t-1}	0.003	0.002	-0.042	0.428	0.617	0.204**
	(0.006)	(0.003)	(0.036)	(0.679)	(0.526)	(0.097)
S1	0.005**	0.004*	-0.011	0.506**	0.251	-0.004
	(0.002)	(0.009)	(0.013)	(0.238)	(0.184)	(0.034)
S2	0.008*	0.005*	0.009	0.344	0.141	0.025
	(0.002)	(0.009)	(0.013)	(0.244)	(0.189)	(0.035)
S3	0.006*	0.003*	0.022	0.542	0.306	0.025
	(0.002)	(0.009)	(0.013)	(0.253)	(0.196)	(0.036)
Δgo_t	0.008	0.014	-0.148**	-2.866*	0.713	0.422*
	(0.009)	(0.004)	(0.058)	(1.084)	(0.838)	(0.155)
Δgo_{t-1}	0.004	0.005	-0.060	-1.336	0.879	-0.218
	(0.010)	(0.005)	(0.062)	(1.157)	(0.894)	(0.165)
R^2	0.497	0.772	0.265	0.339	0.189	0.154
σ	0.009	0.004	0.056	1.044	0.807	0.149
$\chi^2_{SC}[4]$	3.828*	5.746*	1.692	1.273	6.016*	1.460
	[0.006]	[0.000]	[0.156]	[0.284]	[0.000]	[0.219]
$\chi^2_{FF}[1]$	4.742**	0.289	1.652	4.482**	0.453	4.253**
	[0.031]	[0.592]	[0.201]	[0.036]	[0.502]	[0.041]
$\chi^2_N[2]$	1.803	1.250	39.329*	664.433*	573.483*	120.348*
	[0.406]	[0.535]	[0.000]	[0.000]	[0.000]	[0.000]
$\chi^2_H[1]$	4.188**	0.124	0.576	0.021	0.001	32.368*
	[0.043]	[0.725]	[0.449]	[0.886]	[0.969]	[0.000]

Notes: The two error correction terms, estimated over the period 1972Q4 2006Q2, are given by

$\xi_{ppp,t} = p_t - 0.4142pf_t - e_t - 0.4482r_t$

(0.2381) (0.5612)

$\xi_{uip,t} = r_t - rf_t$

Standard errors are given in parentheses, "" indicates significance at the 1% level, "**" indicates significance at the 5% level, and "***" indicates significance at the 10% level. The diagnostics are chi-squared statistics for serial correlation (SC), functional form (FF), normality (N) and hetcroskedasticity (H).*

This result is not surprising given the fact that we would not expect a small open economy, like South Africa, to influence the determination of US prices.[23] Similarly, there is evidence to assume that US interest rates are weakly exogenous[24] to the system, since the $\chi^2[4]=7.2$ [0.135] value is now below the 0.05 critical value of the chi-squared distribution with 4 degrees of freedom.

As expected, oil prices are found to be weakly exogenous since the χ^2 [4] = 3.6 [0.501] value is below the 0.05 critical value of the chi-squared distribution with 4 degrees of freedom (see last two columns of Table 8). Similar inferences are obtained in the case of world gold prices (see Table 9). These results suggest that the inclusion of exogenous *I(1)* variables in the model is more likely to improve estimation precision.

Table 8: Tests of weak exogeneity for US prices, interest rates and world oil prices

	Unrestricted α matrix		$H_2 : \alpha_{21} = \alpha_{22} = 0$		$H_5 : \alpha_{51} = \alpha_{52} = 0$		$H_6 : \alpha_{61} = \alpha_{62} = 0$	
Equations	α_{i1}	α_{i2}	α_{i1}	α_{i2}	α_{i1}	α_{i2}	α_{i1}	α_{i2}
Δp_t	-0.0048 [0.0021]	-0.0004 [0.0001]	-0.0028 [0.0025]	0.0001 [0.0000]	-0.0051 [0.0016]	-0.0004 [0.0001]	-0.0044 [0.0021]	-0.0003 [0.0001]
Δpf_t	0.0003 [0.0009]	-0.0002 [0.0001]	0.0000	0.0000	-0.0008 [0.0007]	-0.0003 [0.0001]	0.0004 [0.0008]	-0.0002 [0.0000]
Δe_t	-0.0212 [0.0131]	-0.0019 [0.0006]	0.0035 [0.0161]	0.0007 [0.0002]	-0.0211 [0.0101]	-0.0022 [0.0009]	-0.0192 [0.0132]	-0.0017 [0.0005]
Δr_t	0.8118 [0.2378]	-0.0245 [0.0102]	1.1818 [0.2959]	0.0066 [0.0035]	0.5255 [0.1807]	-0.0603 [0.0162]	0.8476 [0.2403]	-0.0207 [0.0087]
Δrf_t	0.5548 [0.1823]	-0.0182 [0.0078]	0.7848 [0.2231]	0.0039 [0.0026]	0.0000	0.0000	0.5759 [0.1836]	-0.0146 [0.0066]
Δo_t	0.0075 [0.0345]	-0.0182 [0.0015]	-0.0096 [0.0386]	-0.0004 [0.0005]	-0.0105 [0.0264]	-0.0028 [0.0024]	0.0000	0.0000
P-value			$\chi^2(4) = 13.5[0.085]$		$\chi^2(4) = 7.2$ [0.135]		$\chi^2(4) = 3.6$ [0.501]	

Notes: Standard errors are reported in parentheses. The test is chi-squared distributed with four degrees of freedom since it is jointly conducted with the hypothesis that oil prices do not enter in the cointegrating relations. All results are computed by using the PcGive/PcFiml 9.10 version.

[23] Indeed, the link between SA prices and US prices seems to be difficult to explain given the reduced trade flows between these countries (i.e. SA's trade with North America accounted for 15% of total SA's trade, most of which (39%) is concentrated with the European Community; see IMF, 2000).

[24] This can be attributed to the fact that US interest rates are more likely to be determined by developments in the US economy, such as fluctuations in the dollar/euro exchange rate, productivity growth in the US, changes in US inflation, movements in money markets in the US, as well as by the world oil prices rather than by SA's prices, interest rates, and exchange rates.

Table 9: Test of weak exogeneity for world gold prices

	$H_6 : \alpha_{61} = \alpha_{62} = 0$	
Equations	α_{i1}	α_{i2}
Δp_t	-0.0049 [0.0019]	-0.0004 [0.0001]
Δpf_t	0.0005 [0.0007]	-0.0002 [0.0000]
Δe_t	-0.0221 [0.0123]	-0.0021 [0.0006]
Δr_t	0.6951 [0.2295]	-0.0178 [0.0103]
Δrf_t	0.6018 [0.1686]	-0.0098 [0.0075]
Δgo_t	0.0000	0.0000
P-value	$\chi^2(4)=6.6$ [0.1601]	

Notes: Standard errors are reported in parentheses. The test is chi-squared distributed with four degrees of freedom since it is jointly conducted with the hypothesis that oil prices do not enter in the cointegrating relations. All results are computed in PcGive/PcFiml 9.10 version.

4. Impulse Response Analysis

This section analyzes the dynamic impact of shocks in the system. Three major shocks are analysed; oil prices, foreign prices, and foreign interest rates, respectively. The choice of these shocks is driven by the need to understand how South African economy would react to exogenous shocks incorporated into the system. This analysis is carried out by means of impulse response functions. Figure 4 plots the persistence profile of the effect of a system-wide shock on the long run relationships. The results show that, after a one standard error shock to the system, the UIP relation tends to adjust relatively faster than the PPP, though still quite slowly. This is consistent with the perception of a low speed of adjustment in the goods markets because of the prevalence of information disparities, transportation costs and the effects of tariff and non-tariff barriers.

Using the long run relations discussed in sub-section 3.5, estimates of the impulse response functions of the impact of the shocks on the endogenous variables of the model are computed. The analysis is carried out for a unit (one standard error) increase in the respective exogenous variables. Note that, for example, the effects of an oil price shock

are of particular importance since they can provide some insights in terms of the dynamic stability of the macroeconomy in the context of the South African economy.[25]

The impulse response functions with respect to shocks applied to observables, such as the oil price, can be estimated by using the generalized impulse response approach. To compute the impulse response functions, we need an estimate of the oil price equation specified by *(13)*. The resultant oil price equation, estimated over the period 1972q4-2006q2, is given by:[26]

$$\Delta o_t = 0.2536 + 0.2559\,\Delta o_{t-1} - 0.0050\,\Delta rf_{t-1} - 0.0492\,\Delta pf_{t-1} + \xi_{ot} \qquad (22)$$
$$(0.152)\quad (0.084)\qquad (0.005)\qquad (0.032)$$

$$\sqrt{w_{oo}} = 0.1485;\ \chi^2_{SC}[4] = 2.0865\ [0.086];\ \chi^2_N[2] = 237.584\ [0.000]$$

The results show evidence of a positive impact of past changes in oil prices. The hypothesis that the residuals are serially uncorrelated cannot be rejected at the conventional significance levels of 5% and 1%, respectively. Nonetheless, the normality assumption of the residuals is strongly rejected. This can be attributed to the two major increases in world oil prices during the period under analysis. It is important to note that a joint test of significance[27] of the foreign variables (prices, and interest rates) shows that none of these coefficients is significant. These results suggest that oil prices follow a non-stationary process (i.e. a random walk, eventually with drift). Therefore, the following specification of the oil price equation is adopted:

$$\Delta o_t = 0.0179 \quad + \quad 0.2722\,\Delta o_{t-1} + \xi_{ot} \qquad (23)$$
$$(0.0129)\qquad (0.0832)$$

$$\sqrt{w_{oo}} = 0.1488\ ;\ \chi^2_{SC}[4] = 2.3109[0.061];\ \chi^2_N[2] = 283.661[0.000].$$

[25] Studies that attempted to estimate the impact of oil price shocks in the South African economy include Dagut (1978); Kantor and Barr (1986); Van der Merwe and Meijer (1990); Swanepoel (2006); and Wakeford (2006).

[26] This equation excludes domestic variables since we would not expect a small open economy such as South Africa to have any impact on oil prices.

[27] The chi-squared test of restriction with 2 degrees of freedom returned the value of: $\chi^2[2] = 1.2218[0.298]$ which is well below the 0.05 critical value.

4.1. Effects of an Oil Price Shock

Oil price shocks have often been associated with increasing domestic prices and consequent adoption of restrictive monetary policies, which can ultimately result in slow global economic growth (Bernanke, et al, 1997, Kohler, 2006). Figure 5 shows the persistence profiles of the two long run relationships. It can be seen that all the persistence profiles converge toward zero, thereby confirming the cointegrating properties of the long run relations.

Oil constitutes a critical input in many productive processes. Therefore, oil price shocks tend to fuel high costs of production of commodities that have an impact on the CPI inflation (Burbidge and Harrison, 1984; and Mork et al, 1994). Figure 5.1 gives the impulse responses on the levels of all the five endogenous variables. The oil price shock has a permanent effect, reflecting its unit root properties. Initially, the shock raises domestic prices by 0.5% after a year then prices gradually rise before they stabilise close to 1% after 5 years.[28] Similarly, foreign prices also rise, initially by 0.6% after a year before converging to 0.8% after 5 years.[29]

The oil price shock also tends to increase both domestic and foreign interest rates by about 0.2% on impact, eventually in response to rising prices. This outcome is generally in line with the literature on the macroeconomic impacts of oil price shocks, in which tightening monetary policies have often been observed during periods of oil price shocks (see Bohi, 1989, 1991; Bohi and Toman, 1993; and Cochrane, 1994).

Oil price shocks ultimately tend to affect the level of the rand through movements in relative prices and interest rates. For example, the shock initially creates a small appreciation of the nominal exchange rate, since foreign prices tend to rise more than domestic prices, following the shock. This finding is consistent with the PPP definition of

[28] This result parallels to the findings of Kantor and Barr in who estimate that a 10% increase in world oil prices result in a 0.7 percentage point increase in consumer inflation (excluding food prices).
[29] This result is consistent with the findings of EIA (2001) in which sustained high level of oil prices seemed to create adjustment problems for the US economy as a whole, mostly by raising the inflation rate for the Consumer Price Index.

the nominal exchange rate (i.e. $E_t = P_t / PF_t$). However, the process starts to reverse after approximately 6 quarters, as the increasing interest rate differential stimulates capital outflows. As a result, exchange rate starts depreciating[30] and it fully adjusts to its long run equilibrium after 6 years.

4.2.Effects of a Foreign Price Shock

Figure 5.2 displays the persistence profile of the long run relations after a unit shock in the foreign price equation. As in the previous cases, PPP displays negative deviations from its long run equilibrium, while UIP shows positive and small deviations. In the long run, though, the shock tends to raise domestic prices by 0.01% above their base-line values due to an increase in the costs of intermediate goods. Moreover, this small price response can in part be explained by the slow price transmission between South Africa and the United States as well as by the existence of information asymmetries, transportation costs, and the effect of tariffs and non-tariff barriers in the goods markets. For example, US price shock tends to increase domestic prices only by a constant rate of 0.02% even after 25 quarters.

Foreign price shock also sparks monetary policy reactions in both countries as interest rises in order to curb inflationary pressures. For example, the domestic interest rate rises by 0.1% and then by 0.2% before returning to its equilibrium value. Foreign interest rates display a similar behaviour, although their response on impact is much higher (i.e. about 0.3% in the first quarter). Moreover, an increase in US prices initially causes a small appreciation of the nominal exchange rate.[31] However, this process starts to reverse after approximately 2 quarters as increases in foreign prices start to impact on domestic prices, probably through rising costs of intermediate goods. The exchange rate then reaches its

[30] Another explanation for this process is that oil prices have lagged impact on consumer prices. However, since domestic prices are more responsive to changes in oil prices than the US prices (due to rising costs of intermediate inputs), then the nominal exchange will tend to depreciate as domestic prices outpace foreign prices.

[31] This is consistent with the absolute PPP theory. Note that the nominal exchange rate is here defined as $E_t = P_t / PF_t$. In other words, it is measured as the number of units of domestic currency (Rand) per unit of foreign currency (US dollar). Therefore, an increase in the foreign price initially causes an appreciation to the exchange rate

base-line value after 12 quarters. Figure 5.3 depicts the impulse responses functions of the variables.

4.3.Effects of a Foreign Interest Rate Shock

Figure 5.4 plots the persistence profiles of the impact of a unit shock in the foreign interest rate on the long run relations. As in the previous cases, the varying deviations of the cointegrating relations from their long run values are confirmed. Figure 5.5 show that a one unit shock in the foreign interest rates initially raises these rates by 0.75% on impact. This tends to increase the cost of capital in international markets. As a consequence, domestic interest rates also rise (i.e. by about 20 basis points), probably in response to the increasing interest rate differential as predicted by the UIP theory. However, since foreign interest rates rise more than the domestic rates, this tend to create an uncovered interest rate differential. Following the 'flexible price' monetary model of exchange rate determination (Frenkel, 1976; and Frenkel and Mussa, 1985), this will emanate capital outflows, as investors search for profit opportunities in the US market. As a result, nominal exchange rate depreciates by 0.1%, on average (see Figure 5.5(c)).

Finally, foreign interest rate shock sparks an interesting phenomenon in the goods markets: the 'price puzzle'[32]. This is characterized by a temporary increase in prices in response to an interest rate shock. Initially, foreign interest rate shock tends to raise foreign prices but have virtually no impact on domestic prices (by construction). However, domestic prices start to increase after 3 quarters, due to rising costs of capital in the domestic market. Most interestingly, is the fact that in the long run, domestic prices tend to rise much faster than foreign prices, mainly because US tightening monetary policy measures seems to be more effective in curbing inflation than the underlying domestic monetary policy measures. This constitutes another stimulus for depreciating currency. Figures 5.5(a)-(b) and 5.5(d)-(c) depicts the impulse response functions of both domestic and foreign prices, and interest rates, respectively.

[32] Shup (2004) associates this anomaly with the rising costs of capital due to an increasing interest rate. Since an increase in the interest rate raises the replacement costs of capital, if firms are to recover all of their replacement costs, then they eventually raise the output prices.

5. Probability Event Forecasts

This section computes out-of-sample probability event forecasts based on stochastic simulation that account for future uncertainty.[33] Eight different models are estimated, which account for macroeconomic uncertainty arising from: (i) the evolution of world oil prices; (ii) uncertainty associated with the imposition of long run theory restrictions (i.e. exact or over-identified restrictions when there are 2 cointegrating relationships); and (iii) uncertainty regarding the rank of the cointegrating vectors in the context of exact identification (i.e. r = 0; 1; 2). Initially, we estimate all models over the period 1972q4-2004q2 and then out-of-sample forecasts are performed for the last eight quarters (2004q3-2006-q2). The Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC) are used to evaluate out-of-sample forecast performance as suggested by Garratt *et al* (2003a). Two oil price specifications represented by the equations *(22)* and *(23)* are used in the forecasts. In addition, long run estimates of the cointegrating vectors are used to initialize the simulations.[34] Two events are predicted here: first, the probability of inflation falling below a certain level ranging from 1.5% up to 6%; and second, the probability of the South African Reserve Bank (SARB) keeping inflation between 3% and 6%.

Table 10 presents the model selection criteria for event forecasts. The AIC criterion selects the exactly-identified model with two cointegrating vectors and the oil price equation *(23)*, while the SBC criterion selects the over-identified model with two cointegrating vectors associated with the oil price equation *(22)*. The results indicate that the likelihood of inflation forecasts falling below a certain level ranging from 1.5% up to 6% is much higher in a scenario of imperfect PPP and UIP (i.e. AIC model) than in the complete long run relationship case (i.e. SBC model). For example, the probability that inflation forecast will be less than 6% after 10 quarters is estimated at about 17% under AIC, while SBC predicts a level of accuracy of 14%. However, the level of precision in

[33] This new modelling strategy for forecasting is extensively discussed in Garratt, A., Lee, K., Pesaran, Hashem, M., and Shin, Y. (2003b). "Forecasting Uncertainties in Macroeconomic Modeling: An Application to the UK economy".
[34] The simulations are computed using a bootstrap technique with 1000 iterations for each event.

inflation forecasts tends to fall with time which underpins the existence of uncertainty in the behaviour of macroeconomic variables. This is illustrated by Figures 6 and 7.

Another interesting event to analyse is the probability of SARB keeping inflation in the interval ranging from 3% to 6%. The results are depicted in Figure 8. According to the simulations, the forecast of this event is more likely to be accurate in the case of a stable long run relationship (i.e. the SBC model) than in the case of incomplete PPP and UIP (i.e. the AIC model)[35]. This result suggests that the presence of both stationary real exchange rate and interest rate differential seems to be associated with more accurate inflation forecasts. In addition, both models indicate that the probability that the inflation targeting forecast will be correct reaches its peak after approximately 6 quarters (i.e almost one year and half time period). This prediction is slightly below the current two years timeframe adopted by the SARB.

Table 10: Model selection criteria for inflation forecasts

	Model selection criteria		
Models	**AIC**	**SBC**	**HQ**
EX0A	805.7	747.4	782.0
EX1A	822.8	748.1	792.7
EX2A	835.2	748.4	799.9
OV2A	822.4	748.8	791.8
EX0B	806.3	739.4	779.1
EX1B	823.3	740.8	789.8
EX2B	835.7	740.4	797.0
OV2B	823.0	739.1	788.9

Notes: AIC – Akaike Information Criterion; SBC – Schwarz Bayesian Criterion; HQ – Hannan-Quinn Criterion. EX stands for exactly-identified model; OV – over-identified model. 0,1,...2 corresponds to the number of cointegrating relationships. A and B represent the two different oil price equations (22) and (23).

[35] Nevertheless in the long run (i.e. after 4 years) AIC forecasts tend to outperform those of SBC probably due to uncertainty (see Figure 8 in appendix).

6. Conclusions and Evaluations

This paper used a vector error correction model suggested by Johansen and Juselius (1990) and extended by Pesaran et al (2000) to test the structural hypothesis of long run stationarity of the PPP and UIP relationships between South Africa and the United States. Using quarterly data from January 1972 to June 2006, in a six-dimensional system of equations (two prices, two interest rates, exchange rate, and oil prices as a 'long run forcing' variable), the major findings of the paper suggest that: (i) the PPP relation is not stationary in the long run; (ii) UIP is stationary in the long run; (iii) domestic prices, exchange rates, US prices and domestic interest rates form a cointegrating vector in the context of South Africa; (iv) impulse response functions show that oil price shocks tend to raise the level of domestic inflation, which consequently triggers a tightening monetary policy as short-term interest rates quickly adjust; (v) probability event forecasts indicate that inflation forecasts in South Africa are more likely to be in the designated ranges in the presence of incomplete PPP and UIP relations than in a more complete long run relation, although inflation targeting forecasts are more likely to be met in the complete PPP model.

The result regarding the rejection of the PPP hypothesis is somewhat unsurprising given the evidence in the literature, since unity restrictions are often rejected. However, this finding also suggests that deviations from PPP in South Africa only tend to reduce at a relatively slower rate. Consequently, the goods market may take a long time to clear following an exogenous shock in the economy. In contrast, the interest rate differential tends to adjust relatively faster than PPP, though still quite slowly.

Last, but not least important, is the need to improve upon these estimates. Firstly, the results might have been affected by structural changes in the economy. Secondly, the choice of the sample period as well as the methodology may yield different results, especially if different structural breaks are imposed. Thirdly, when the model is extended to include more economic relationships (i.e. money market equilibrium), it may give more comprehensive insights into the behaviour of the South African economy. Therefore, it would be worthwhile to research these factors further.

Appendix

Figure 1: Prices; Exchange Rate; World Oil Prices; World Gold Prices; and Interest rates (levels)

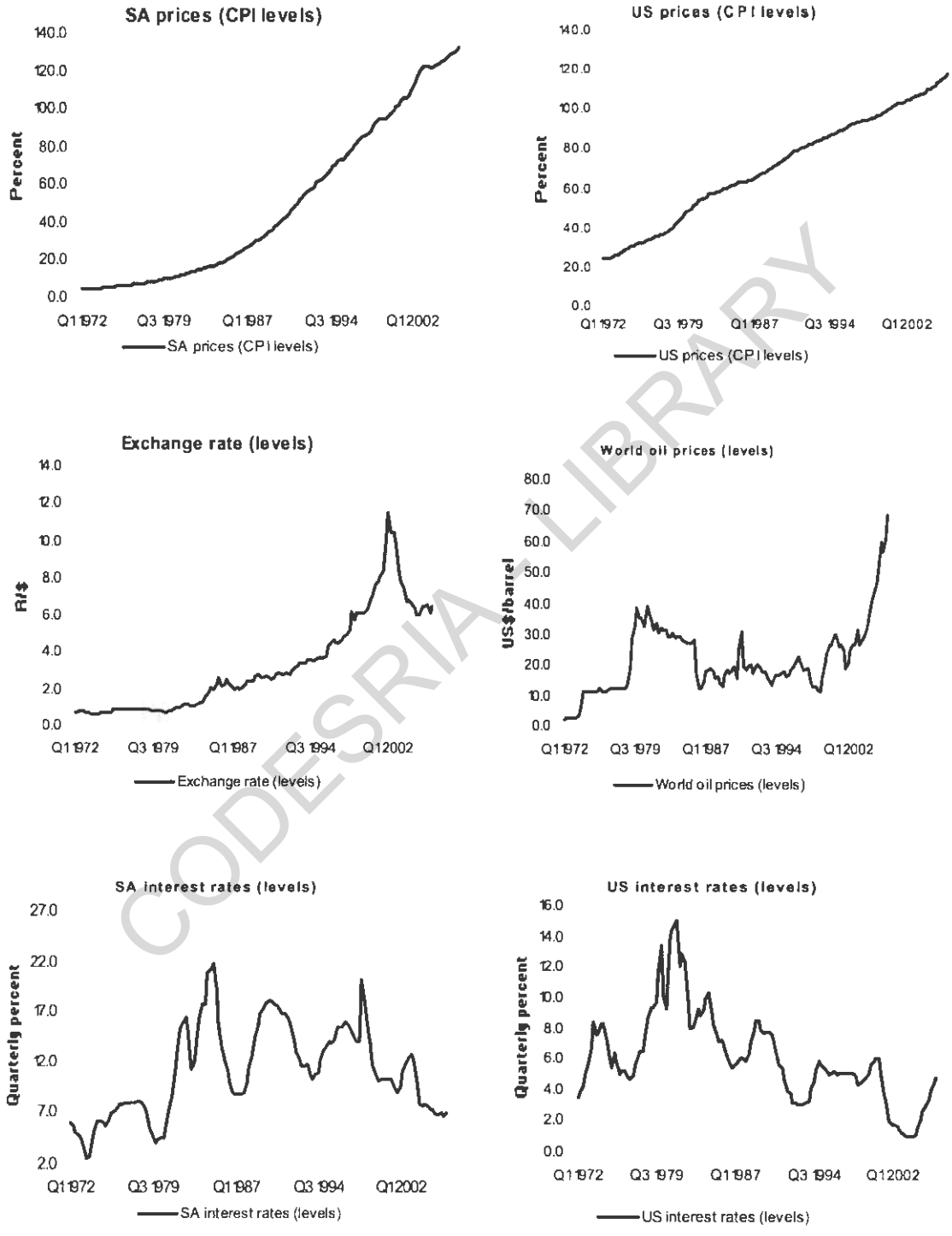
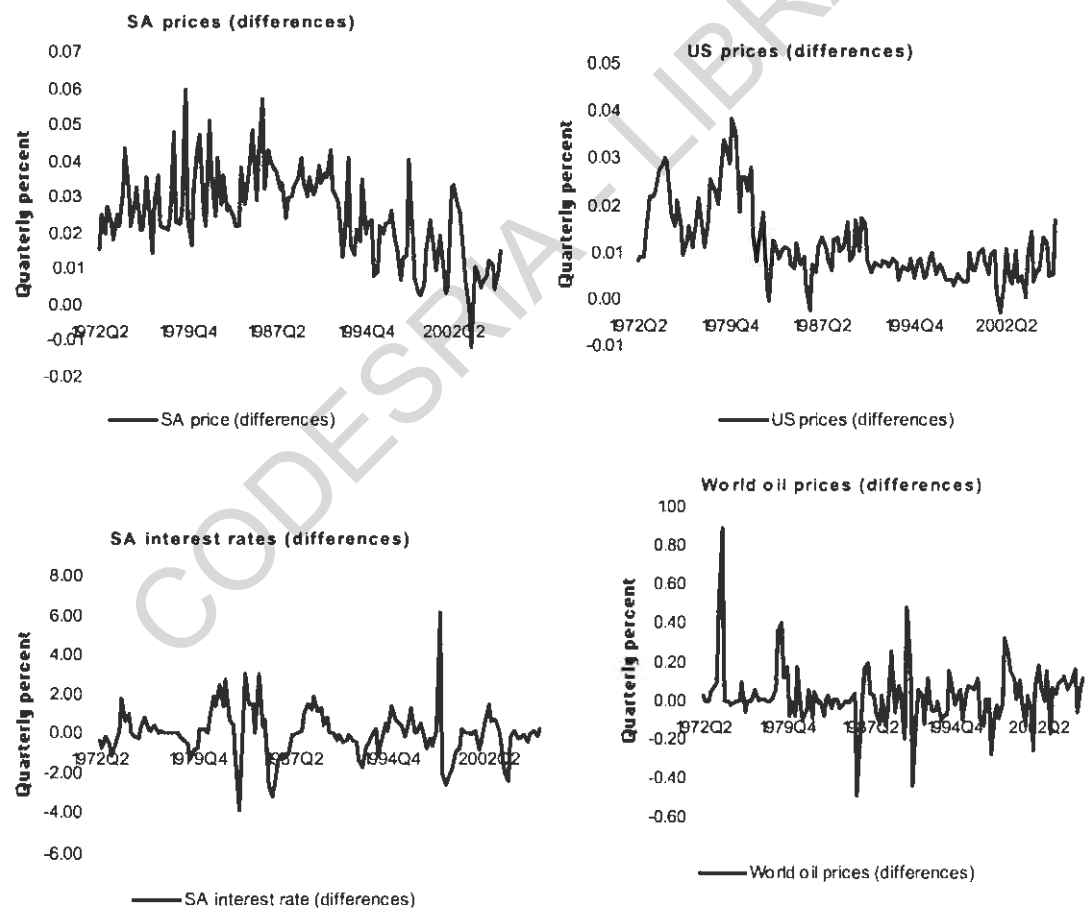


Figure 2: Prices; Exchange Rate; World Oil Prices; World Gold Prices; and Interest rates (differences)



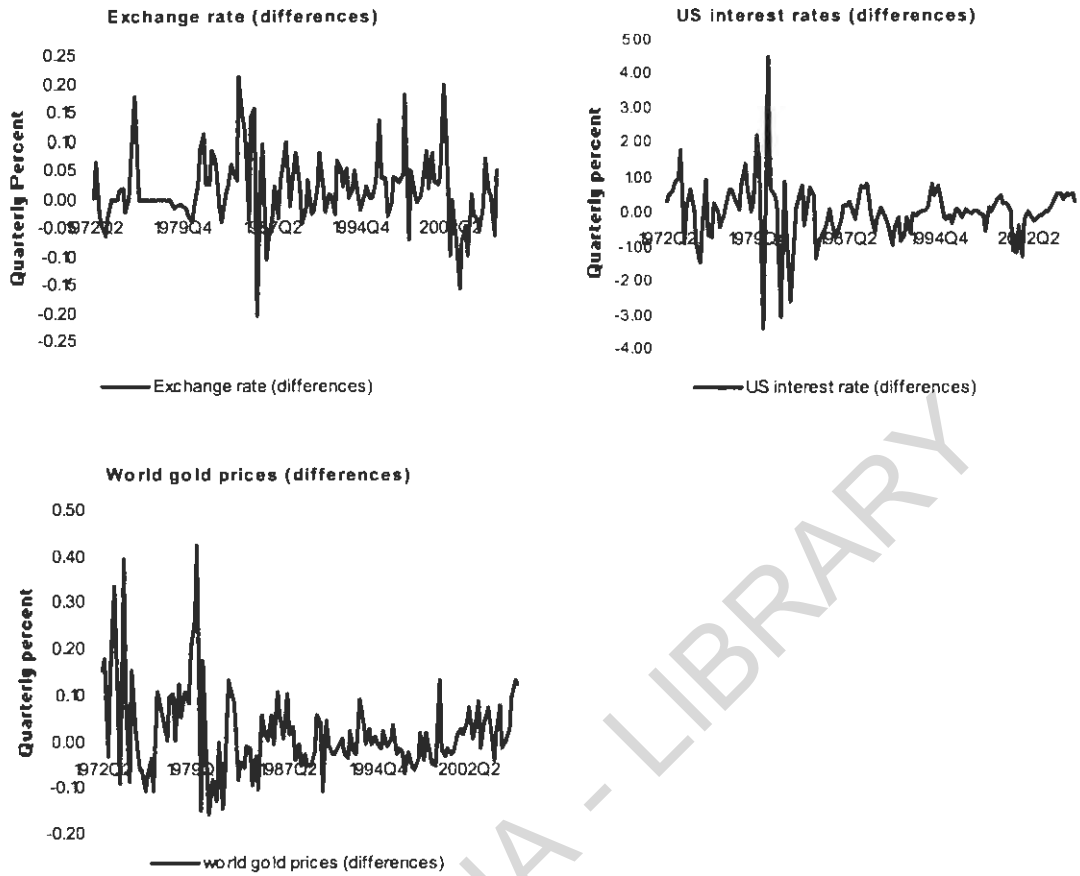
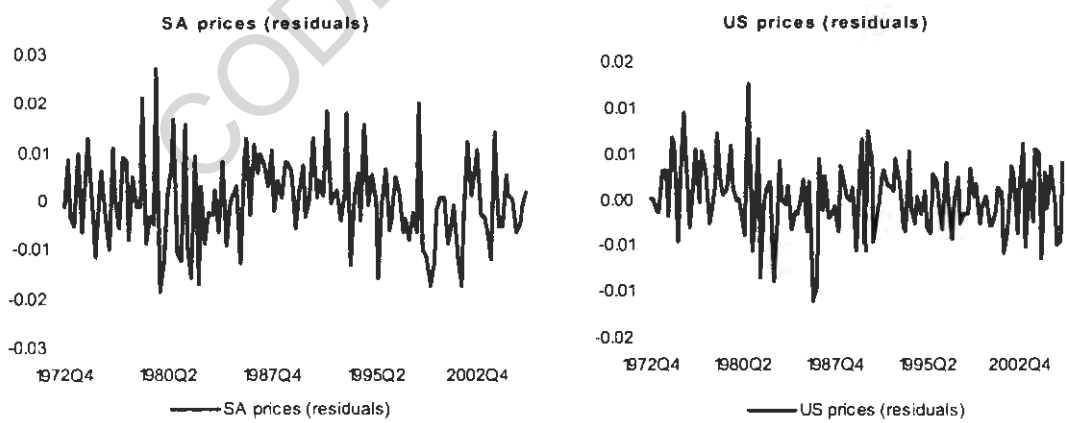


Figure 3: Prices, Exchange Rates, and Interest Rates (residuals from the VAR(2))



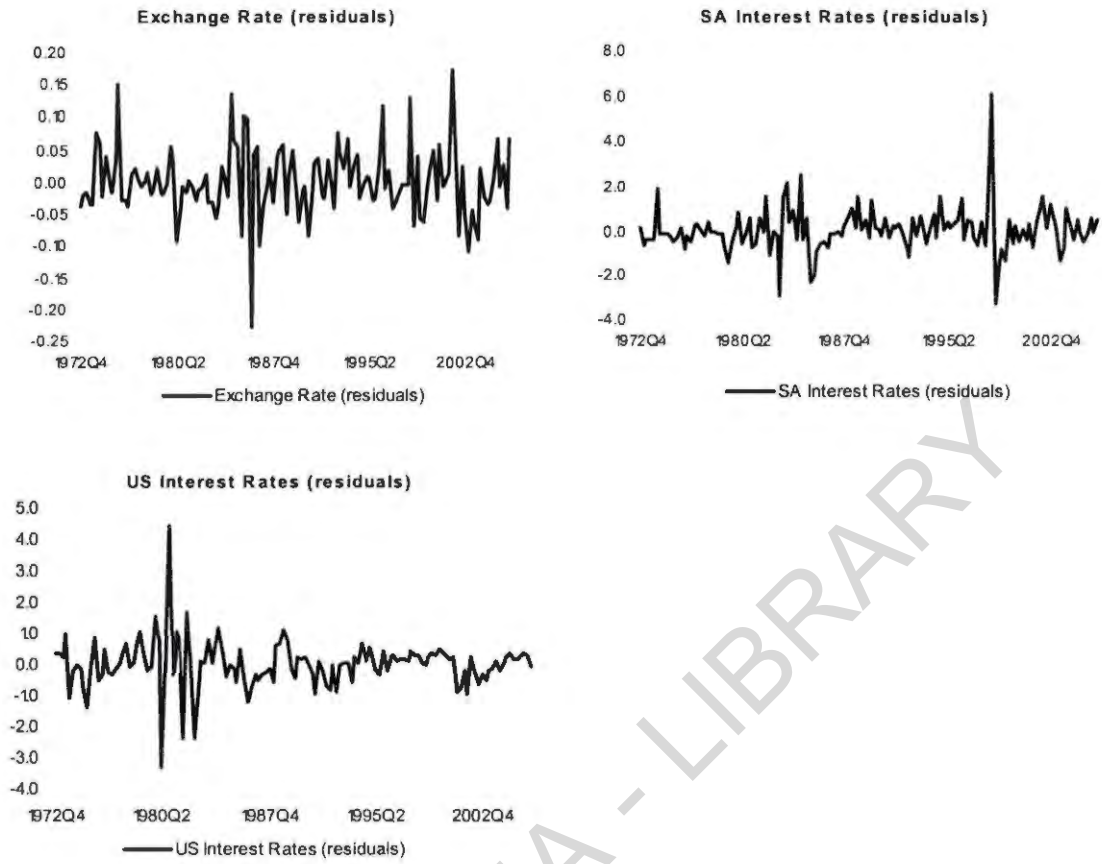
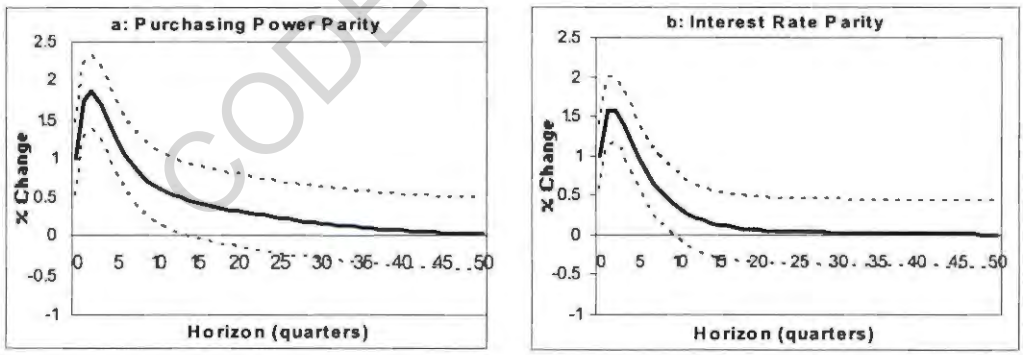


Figure 4: Persistence Profiles of the effect of a system-wide shock to the Long run Relations



Notes: The graphs define the long run relationships as follows; Interest rate parity: r_t - rf_t and the PPP (real exchange rate): $e_t + pf_t - p_t$. The size of the shock is equal to the standard deviation of the selected equation error. The solid and dashed lines plot the point estimates and the ± 2 standard error bands.

Figure 5: Persistence Profiles of the Long run Relations of a Positive Unit Shock to the Oil Price

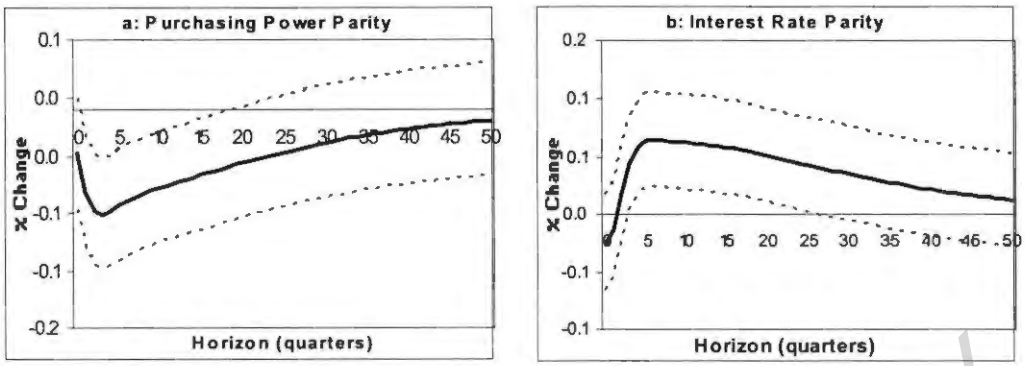
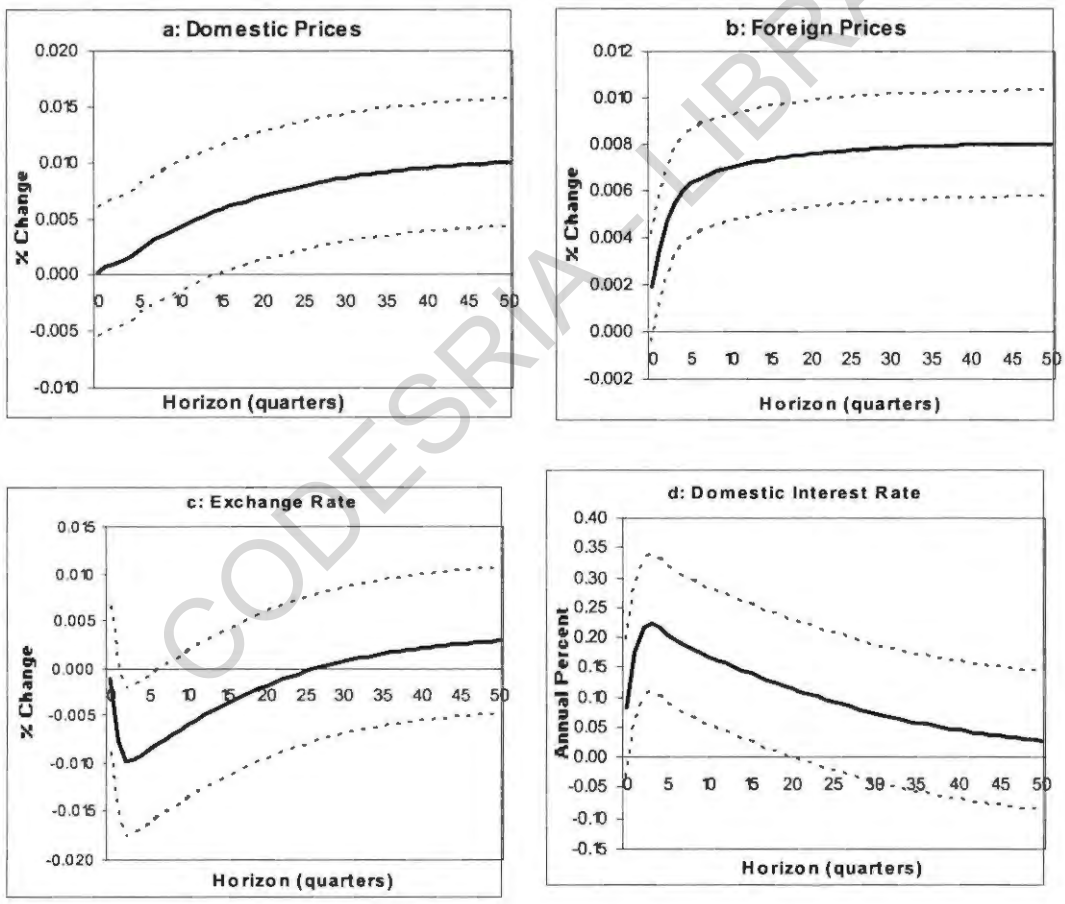


Figure 5.1: Generalized Impulse Responses of a Positive Unit Shock to the Oil Price



Notes: The solid and dashed lines plot the point estimates and the ± 2 standard error bands respectively of the impulse responses.

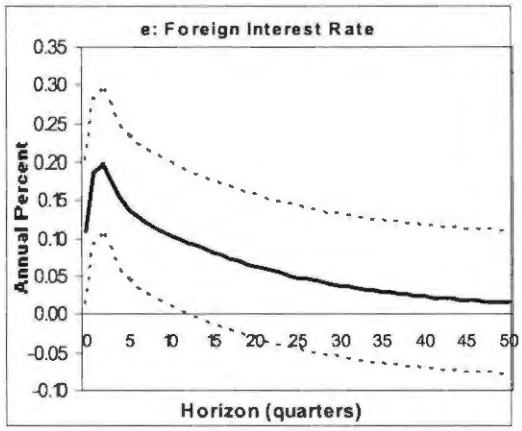


Figure 5. 2: Persistence Profiles of the Long run Relations of a Positive Unit Shock to the Foreign Price Equation

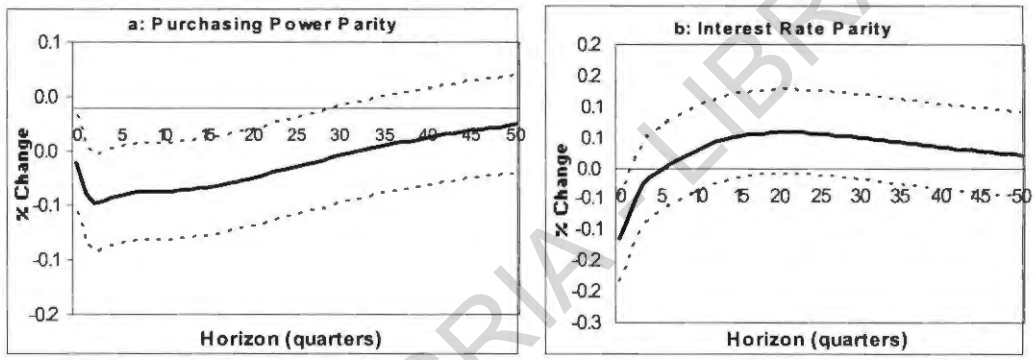
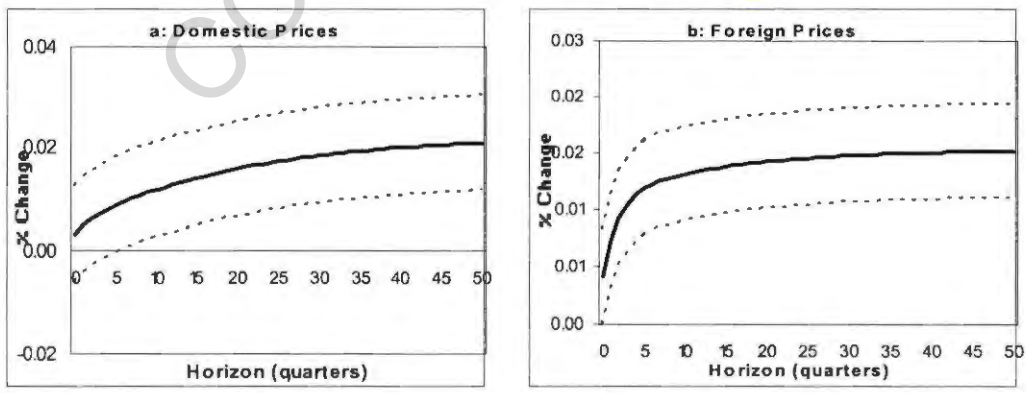


Figure 5.3: Generalized Impulse Responses of a Positive Unit Shock to the Foreign Price Equation



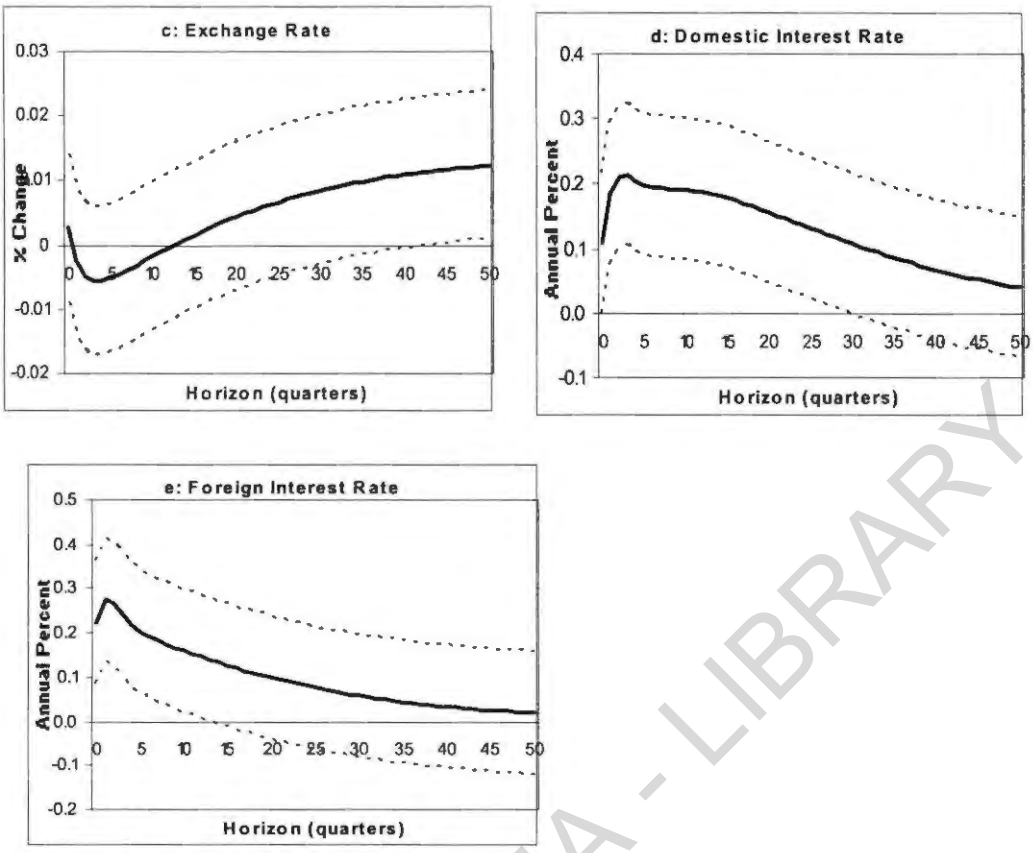


Figure 5.4: Persistence Profiles of the Long run Relations of a Positive Unit Shock to the Foreign Interest Rate Equation

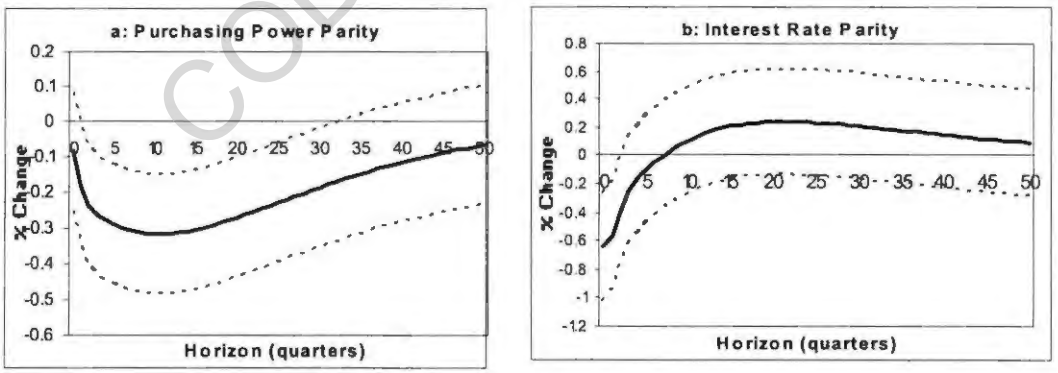


Figure 5.5: Generalized Impulse Responses of a Positive Unit Shock to the Foreign Interest Rate Equation

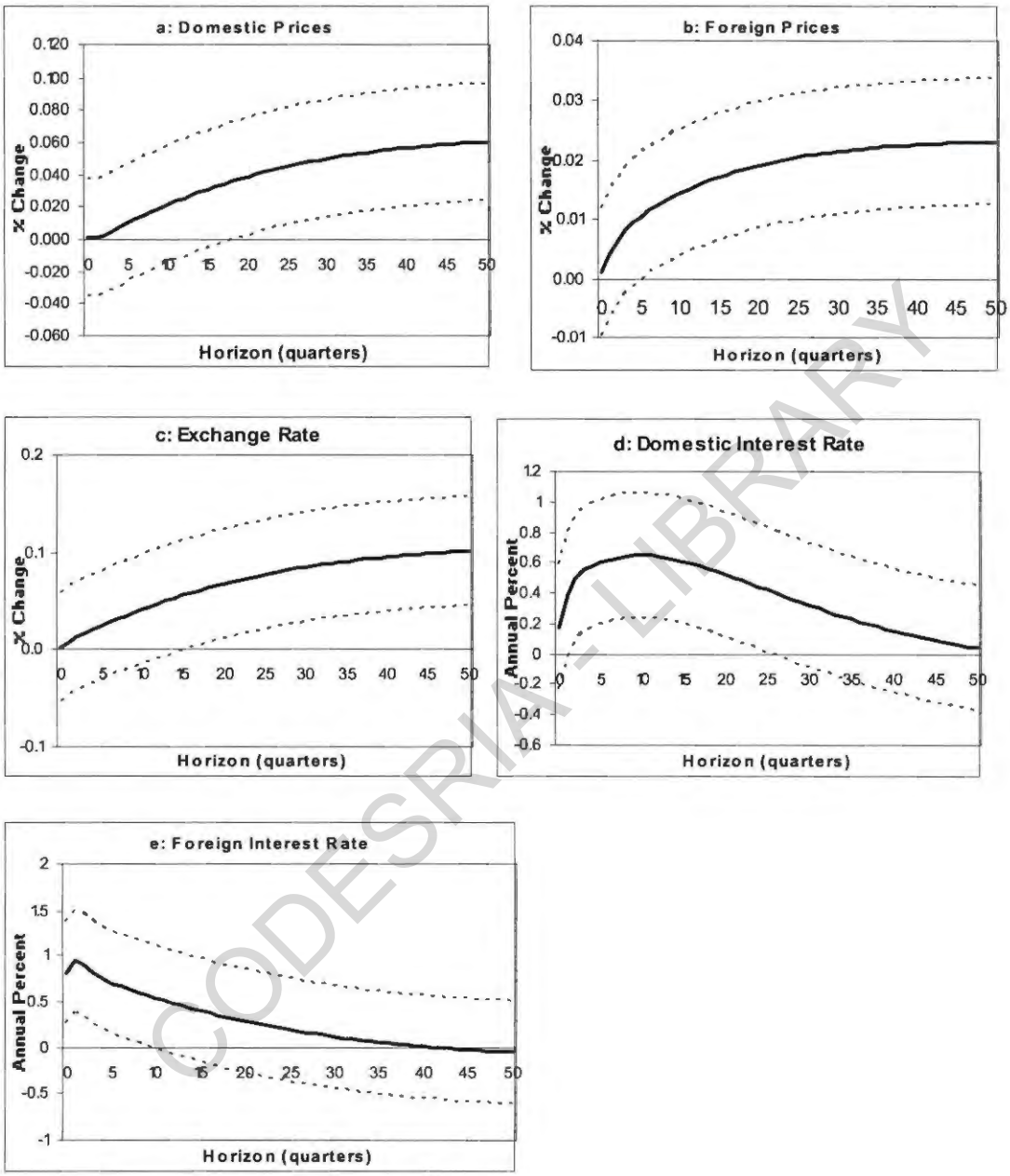


Figure 6: Probability event forecast of inflation falling below 1.5, 2%, 2.5,…,6% (SBC model)

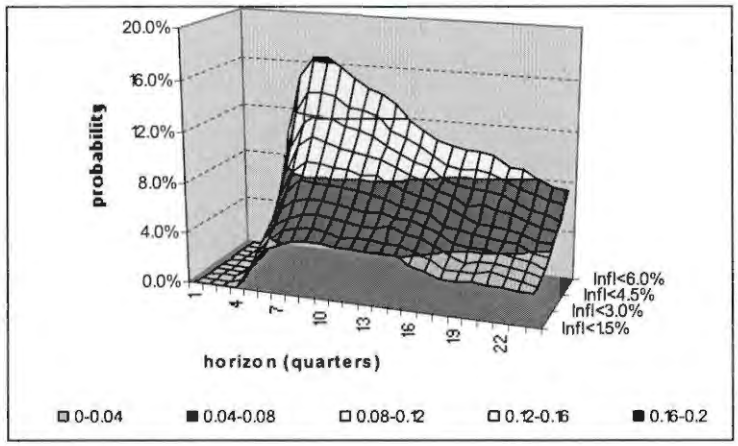


Figure 7: Probability event forecast of inflation falling below 1.5, 2%, 2.5,…,6% (AIC model)

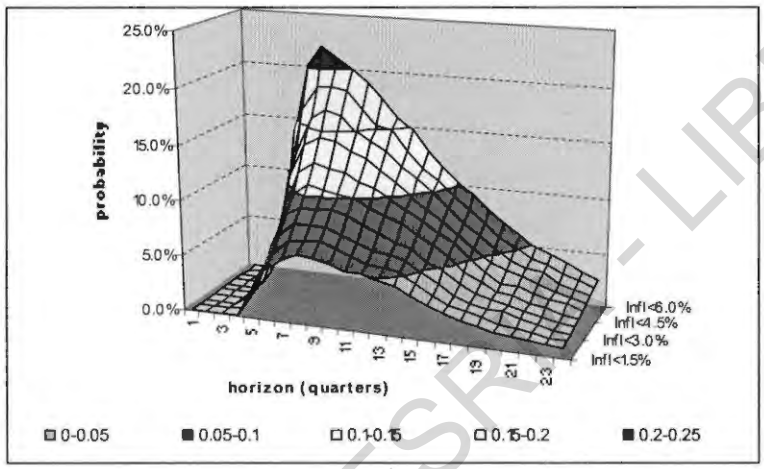
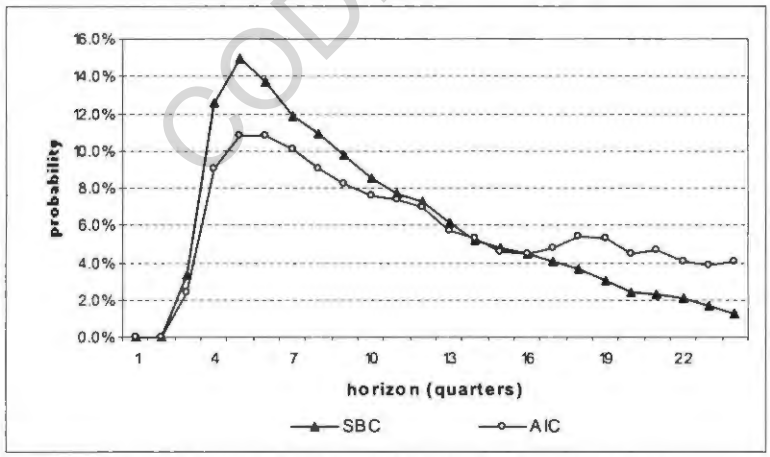


Figure 8: Probability event forecast of SARB keeping inflation between the range 3%<Inflation<6%.



Note: SBC – Schwarz Bayesian Criterion and AIC – Akaike Information Criterion

Table 11*: Tests of PPP and UIP under oil price specification*

	Model			
Variables	Test of PPP and H_{oil} VAR (2)		Test of UIP and H_{oil} VAR (2)	
p_t	1.0000	0.0000	1.0000	0.0000
pf_t	-1.0000	-8.6873 (5.5620)	-1.4520 (1.1668)	0.0000
e_t	-1.0000	13.1835 (13.9378)	0.7151 (2.1517)	0.0000
R_t	0.0000	1.0000	-0.2794 (0.3276)	1.0000
RF_t	0.0000	-2.2462 (2.1794)	0.0000	-1.0000
o_t	0.0000	0.0000	0.0000	0.0000
LL-value	977.7		985.4	
P-value	$\chi^2[5]=30.32[0.000]$		$X^2[5]=14.98[0.010]$	

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