CSAE Conference 2008
Economic Development in Africa
16th - 18th March 2008, St Catherine's College,Oxford

South Africa's real exchange rate and the commodities boom: a Markov regime switching approach

Bassam Fattouh
Centre for Financial and Management Studies
SOAS, University of London

Kostas Mouratidis
School of Business and Economics
University of Swansea

Laurence Harris*
Centre for Financial and Management Studies
SOAS, University of London

Abstract
In this paper we use a Markov switching model to estimate non-linearities in the long-run relation between the rand exchange rate and fundamentals. In data for the period 1975 to 2007 we use the Johansen Maximum Likelihood method to estimate a cointegrating vector of the real exchange rate, with a South Africa/US interest differential, a measure of South Africa/US inflation differential and the world price of gold. The cointegration suggests that the long run equilibrium real exchange rate is determined by only two variables, central to uncovered interest parity and relative purchasing power parity theories, plus the gold price. The non-linear model identifies two regimes, a regime in which the exchange rate equation yields errors with high variance and one in which the errors have low variance. The dates at which the model switches from one regime to another indicate that structural shifts in the relationship between the real exchange rate and its determinants are associated with major changes in South Africa's international environment. We find evidence consistent with the hypothesis that the regime switch that occurred at the end of 2001 was related to a belief that rising commodity prices were evidence of a commodity super-cycle.

JEL Classification C22, F31,F41,O24,O55

* Corresponding author, Laurence Harris, CeFiMS, SOAS, University of London, Thornhaugh Street, London WC1H 0XG, LH@soas.ac.uk

South Africa's real exchange rate and the commodities boom: a Markov regime switching approach

Bassam Fattouh
Centre for Financial and Management Studies
SOAS, University of London

Kostas Mouratidis
School of Business and Economics
University of Swansea

Laurence Harris*
Centre for Financial and Management Studies
SOAS, University of London

(This draft 29th February 2008)

1 Background and motivation

The South African economy's openness has given the behaviour of the real exchange rate prominence in discussions of monetary (and fiscal) policy. 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 head of the list. The government judged that periods of high volatility have constrained investment and its effects have been compounded by periods when the SA rand has been overvalued with dutch-disease type effects (Republic of South Africa, 2006).

* Corresponding author, Laurence Harris, CeFiMS, SOAS, University of London, Thornhaugh Street, London WC1H 0XG, LH@soas.ac.uk

Such concern over the real value of the currency in the middle of this decade had arisen as, following a long real depreciation from 1983 to 2001, South Africa's nominal and real exchange rate appreciated strongly from 2002 to 2006, its commodity exports benefited from a global commodities boom, and the current account deficit increased. These developments fuelled long-standing concerns over the effect of real exchange rates on manufacturing industries' cost competitiveness (Golub 2000), concerns which were consistent with survey evidence on the effect of real exchange rates' levels and volatility on South African industrial sectors (FEASability 2006).

South Africa's developed financial sector enables it to adopt and implement monetary policy regimes, including exchange rate regimes, that are comparable with developed countries'. After having targeted a monetary aggregate since 1986, the SA Reserve Bank has, since 2000, had a formal inflation targeting policy framework with no explicit exchange rate target (although it enters the market to adjust official reserves). But, since the real effective exchange rate may be systematically affected by monetary policy and other policies, concern over the real-economy effects of the real value of the rand and its volatility has led to suggestions for modification of inflation targeting (Frankel, 2007a), and it motivates an enquiry into what determines the real effective exchange rate of SA.

McDonald and Ricci (2004) estimate a long run equilibrium relation between the rand's real effective exchange rate and fundamentals including real interest rate differentials, commodity price changes, productivity differentials, openness, fiscal balance, and net foreign assets. In a data set for the years 1970 to 2002 McDonald and Ricci, using Johansen cointegration and VECM estimates, find that around two thirds of the depreciation from the early 1990s to 2002 can be explained by depreciation in the long run equilibrium rate (the cointegrating vector rate). Frankel (2007b) finds that a high proportion of the variation in the real exchange rate from 1984 to 2007 is accounted for in a regression on the lagged exchange rate, an index of commodity export prices, and a real interest rate differential; the dominant effect on the log current real exchange rate is that of itself lagged one period[1].

Estimating an exchange rate function using long data sets generally increases the importance of recognising the possibility of structural breaks, especially in the case of South Africa which, over the past three decades, has experienced strong shifts domestically and internationally in the institutional and political environment of economic transactions. These, which may shift any relation that exists between exchange rates, interest differentials and other variables include South Africa's foreign debt default (moratorium) of 1984, the relaxation of capital controls from 1995, and the adoption of inflation targeting in 2002. A standard technique for accounting for structural shifts in linear regressions is to impose a dated dummy variable, as does Frankel (2007b).

Alternatively, a non-linear Markov-switching model enables us to identify structural breaks generated within the time series and to obtain estimates with more robust

[1] MacDonald and Ricci (2005), Frankel (2007b), and the present study which estimate an exchange market equilibrium equation. Earlier research by Aron, Elbadawi, and Kahn (1997) estimates a macroeconomic internal-external equilibrium model of the South African exchange rate

properties than linear OLS. Recent research using non-linear models, including Markov-switching models, has demonstrated their relative strength for identifying systematic relations between fundamentals and exchange rates that are consistent with theoretical priors (Sarno and Taylor, 2002a, 2002b;. MacDonald, Menkhoff and Frommel 2005).

In this paper we use a Markov switching model to estimate non-linearities in the long-run relation between the rand exchange rate and fundamentals. Using data for the period 1975 to 2007 we use the Johansen Maximum Likelihood method to estimate a cointegrating vector of the real exchange rate, with South Africa/US interest differential, a measure of South Africa/US inflation differential and the world price of gold. The cointegration suggests that the long run equilibrium real exchange rate is determined by only two variables, central to uncovered interest parity and relative purchasing power parity theories, plus the gold price. Although South Africa's exports include a diverse basket of minerals, we find that the gold price proxy captures the effect of world commodity prices that other studies have found to have a systematic effect on the exchange rates of commodity exporting countries (Cashin, Cespedes, and Sahay, 2003).

Following the two-step error correction method of Krolzig (1996) we use the equilibrium correction error from the cointegration to estimate a Markov regime switching error correction model (MRS-ECM). The non-linear model identifies two regimes, a regime in which the exchange rate equation yields errors with high variance and one in which the errors have low variance. The dates at which the model switches from one regime to another indicate structural shifts in the relationship between the real exchange rate and its determinants are associated with major changes in South Africa's international environment. We also compare out of sample forecasts of the MRS-ECM with those of a linear error correction model.

The overall results of our estimates suggest that South Africa's real exchange rate can be explained by a relationship with two variables related to uncovered interest parity and relative purchasing power principles, plus a proxy for world mineral prices. The relationship is a long term equilibrium relation that, in a Markov switching model, takes account of structural breaks associated with changes in South Africa's policy environment. The results differ from those of McDonald and Ricci, and Frankel because they yield a long term equilibrium relationship which is consistent with structural breaks, is not dependent upon a dominant effect of the lagged exchange rate, and is related to simple economic models with the addition of a mineral price alone as a country-specific determinant (in contrast to previous studies which use exchange rate equations with a larger number of explanatory variables).

2 Estimation method

In what follows, we employ the two-step Markov regime-switching Error Correction Model (MRS-ECM) suggested by Krolzig (1996) outlined here. In the first-step (Section 4.1), we use Johansen cointegration analysis to estimate long-run relationship among the variables of vector X_t. In the second step (Section 4.2) , using the equilibrium correction error obtained from the first step, we estimate a MRS-ECM.

2.1 Cointegration model

A general-to-specific approach is adopted in this section to model both the long-run and short-run structure of vector X_t in four steps. First we use the Johansen Maximum Likelihood approach to estimate and identify cointegrating relationships among the variables of vector X_t . More concretely, we can write as a vector autoregressive process of order p (i.e., VAR(p)):

$$X_t = A_0 + \sum_{i=1}^{p} A_i X_{t-i} + u_t \qquad (1)$$

$$\Delta X_t = A_0 + \Pi X_{t-1} + \sum_{i=1}^{p} \Pi_i X_{t-i} + u_t \qquad (2)$$

$$\Delta X_t = A_0 + \alpha\beta' X_{t-1} + \sum_{i=1}^{p} \Pi_i X_{t-i} + u_t \qquad (3)$$

$$u_t \text{ is iid} \sim N(0,\Sigma)$$

Where β is the $(n\times r)$ matrix of the cointegrating vector and α denotes the $(n\times r)$ matrix of speed of adjustment to last period equilibrium error.

In the second step, we estimate the vector equilibrium correction models presented by (3), where the identified matrix of cointegrating vectors β' is explicitly taken into account:

$$\Delta X_t = \hat{A}_0 + \hat{\alpha}\left(\sum_{i=1}^{r} \hat{\beta}_i' X_{t-1}\right) + \sum_{i=1}^{p} \hat{\Pi}_i X_{t-i} + u_t \qquad (4)$$

At this stage, we re-estimate (4) excluding any insignificant regressors. The resulting parsimonious vector equilibrium correction model (PVECM) is a reduced form model and consequently there are simultaneity effects among the endogenous variables including in X_t. Having estimated the PVECM, we test for exogeneity by testing that the null α_i is not significantly different from zero (i.e., H_0: $\alpha_i = 0$). If the null is true then the variable X_i is exogenous with respect to all cointegrating vectors.

In the third step, we estimate (4) conditional on exogenous variables.

$$\Delta X_{1,t} = \hat{A}_0 + \Delta X_{2,t} + \hat{\alpha}_1\left(\sum_{i=1}^{r} \hat{\beta}_i' X_{t-1}\right) + \sum_{i=1}^{p} \hat{\Pi}_i X_{t-i} + u_{1,t} \qquad (5)$$

$$u_{1,t} \text{ is iid} \sim N(0,\Sigma_1)$$

where $\alpha=[\alpha_1, \mathbf{0}]'$, and X_2 is the vector of exogenous variables. In the fourth step, we model any simultaneous effects in (5). If any of the off diagonal elements of Σ_1 is close to zero we can apply OLS to estimate each equation of (5) separately.

2.2 Markov Regime Switching ECM principles

To estimate the Markov Regime Switching Error Correction Model (MRS-ECM) we estimate the MRS-VAR model as an extension of the basic VAR model. Here we adopt the most general specification of MRS-VAR model by allowing all parameters to be conditioned on the unobserved state variable S_t.

$$X_t = A(S_t)_0 + \sum_{i=1}^{p} A(S_t)_i X_{t-i} + u(S_t)_t \qquad (6)$$

Where S_t follows an ergotic Markov chain with a finite number of states $S_t \in \Omega_t = \{1,2,..,N\}$ which is defined by the transition probability matrix $P = (p_{i,j})_{i,j \in \Omega}$ where $p_{i,j} = P(S_{t+1} = j \mid S_t = i)$.

Conditioned on the estimated cointegrated vector the VAR equation (6) can be extended to the MRS-VECM:

$$\Delta X_t = \hat{A}(S_t)_0 + \hat{\alpha}(S_t)\left(\sum_{i=1}^{r} \hat{\beta}_i' X_{t-1}\right) + \sum_{i=1}^{p} \hat{\Pi}(S_t)_i X_{t-i} + u(S_t) \quad (7)$$

We estimate a univariate version of (7) where X_t includes only the real exchange rate[2].

3. Data

We employ the following variables in our empirical analysis: the logarithm of the real effective exchange rate (LRER); the nominal interest rate differential (NIRF) measured by the difference between South African three-month treasury bill rate and US three-month treasury bill rate; the inflation differential (INFDIF) measured by the difference in year-on-year CPI-based inflation rates between South Africa and the US; and the logarithm of the price of gold in US$ per ounce (LPGOLD). The data set consists of monthly data from January 1975 to April 2007 yielding a sample of 388 observations. The data were all obtained from the *Thomson Datastream* database.

Figure 1a: SA real effective exchange rate (Jan 1975-April 2007, Index: 2000=100)

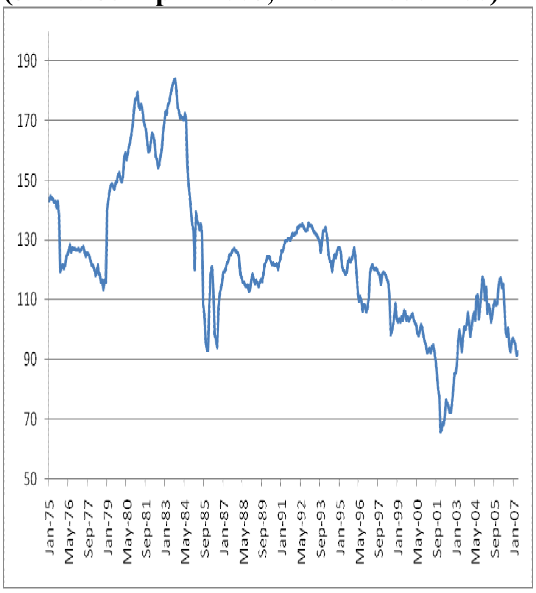
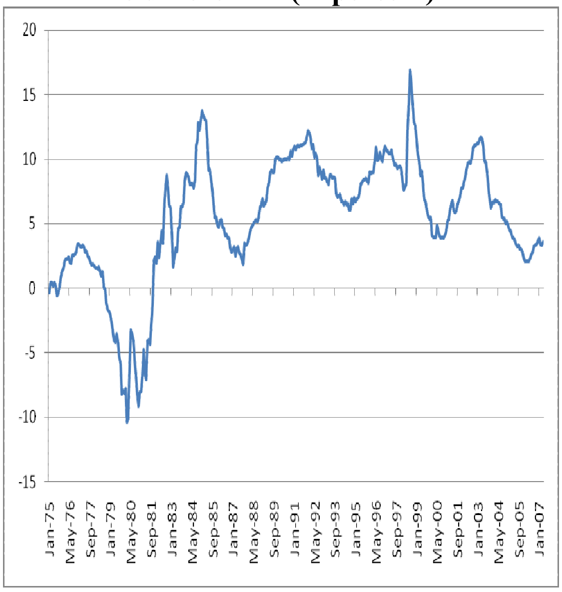


Figure 1b: SA/US nominal interest rate differential (in percent)



[2] This is so because in the last step of cointegration analysis we model any simultaneous interaction among the variables of vector X_t.

Figure 1c: SA/US Inflation rate Differential (in percent)

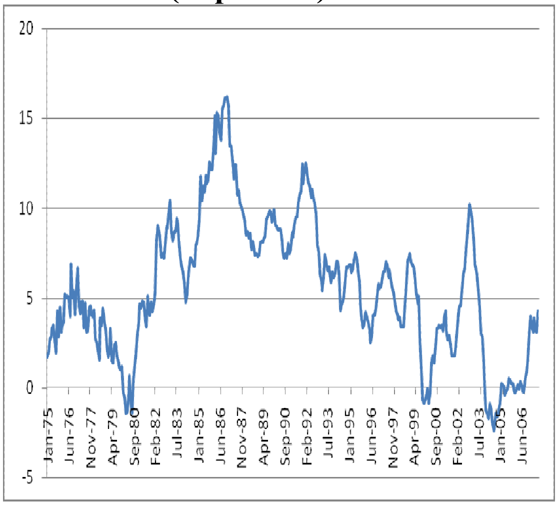
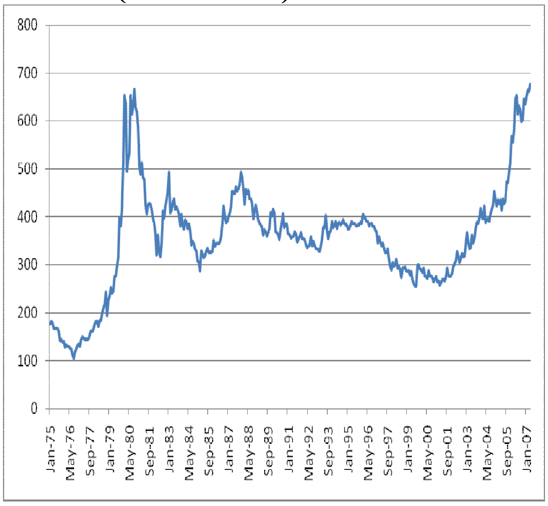


Figure 1d: Price of gold (US$/Ounce)



The variables employed in our empirical analysis are plotted in Figures 1a-1d. Table 1 provides summary statistics for each of these variables. Some interesting patterns can be observed from these descriptive statistics. The first concerns the real effective exchange rate which sharply depreciated from 1995, when South Africa ended its dual rate system, until the end of 2001 when a strong appreciation began. The second concerns the steep rise in the world price of gold since late 2001, which continued until the end of the sample. The third concerns the high variability in nominal interest rate and inflation rate differentials. The summary statistics presented in Table 1 show a mean for interest rate differential of 5.49% with standard error of 4.88%. Similarly, the mean inflation differential is 5.58% with standard error of 3.92 %.

Table 1: Summary Statistics

LREXR		**NIRDIF**	
Observations	388	Observations	388
Sample Mean	4.796901	Sample Mean	5.497371
Standard Error	0.19568	Standard Error	4.880363
Skewness	-0.18835	Skewness	-0.863831
Kurtosis (excess)	0.457251	Kurtosis (excess)	0.804055
Jarque-Bera	5.674197	Jarque-Bera	58.7063
INDIF		**LPGOLD**	
Observations	388	Observations	388
Sample Mean	5.585749	Sample Mean	5.81918
Standard Error	3.922057	Standard Error	0.361542
Skewness	0.298021	Skewness	-1.019524
Kurtosis (excess)	-0.201768	Kurtosis (excess)	1.284431
Jarque-Bera	6.401625	Jarque-Bera	93.887605

4 Empirical Results

4.1 Long run equilibrium exchange rate function

The results of applying Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) to the level of series indicate that the null hypothesis of a unit root cannot be rejected for the level of the series. Applying the DF and ADF test on the first differenced series however indicates that the null hypothesis can be rejected at conventional levels. Hence we conclude that all the variables are integrated of degree one and thus proceed with cointegration analysis.

Table 2: Unit Root Tests

LREXR	**ADF(2)**	**NIRDIF**	**ADF(1)**	**INFDIF**	**ADF(0)**	**LPGOLD**	**ADF(0)**
	-2.067		-2.205		-2.138		-1.388
DLREXR	**ADF(0)**	**DNIRDIF**	**ADF(1)**	**DINFDIF**	**ADF(0)**	**DLPGOLD**	**ADF(0)**
	-15.72		-14.864		-18.861		-20.006

Notes: The DF/ADF and test statistics, based upon an optimal lag selection through the BIC criterion in parenthesis (similar results are obtained using the AIC criterion). The critical values for the ADF from MacKinnon (1996) are –3.46, -2.87, -2.57 for the 1%, 5%, 10% level of significance, respectively.

The VAR length has been set equal to three. The choice of three lags has been made to ensure that the residuals from the estimated VAR model are not auto-correlated. The LM test for autocorrelation (Anderson, 2003) suggests that we can not reject the null hypothesis of auto-correlation at conventional levels. We next test for the rank of the VAR model using the trace statistics. As can be seen from Table 3, the trace statistic suggests the existence of one cointegrating vector. Hence a cointegration rank of one is imposed on the VAR. We next test zero for restrictions on the parameters of the cointegrating vector based on the likelihood ratio test procedure (Johansen, 1996). The statistics in Table 3 show that none of the variables should be excluded from the long run cointegrating vector. We next test for the weak exogeneity of each of the variables in the system (Johansen, 1992, 1996). This involves testing whether the speed of the adjustment coefficient (α) is significantly different from zero. Generally, if α which is related to variable x_1 is zero, then when estimating the parameters of the model there is no loss of information of not including Δx_1. Tests for weak exogeneity show that both NIRF and LPGOLD are weakly exogenous to the system. We impose these weak exogenous restrictions and re-estimate the model. The last section of Table 3 shows the resulting cointegrating vector normalised with respect to the real effective exchange rate. This is a valid normalization since the real effective exchange rate is endogenous to the VAR.

Table 3- The Long Run Equilibrium Exchange Rate Equation (Johansen Procedure)

(*Sample period: Jan 1975 - April 2007)*

Variables entered

LRER: the logarithm of the real effective exchange rate
NIRF: the nominal interest rate differential
INFDIF: the inflation rate differential
LPGOLD: the logarithm of the price of gold

Lag Length of VAR = 3

Test for residual correlation:

LM(1):	$\chi^2(16)= 20.290$	*p-value=0.207*

Test for rank of VAR

Ho: Rank = r	Trace Statistics	*p-value*
r=0	67.257	0.015
r =1	32.574	0.427

Test of exclusion	LR test	
LRER	$\chi^2(1)= 16.426$	*p-value=0.000*
NIRF	$\chi^2(1)= 4.976$	*p-value=0.026*
INFDIF	$\chi^2(1)= 11.167$	*p-value=0.001*
LPGOLD	$\chi^2(1)= 16.138$	*p-value=0.000*
Trend	$\chi^2(1)= 15.518$	*p-value=0.000*

Test of weak exogeneity	LR test	
LRER	$\chi^2(1)= 9.543$	*p-value=0.002*
NIRF	$\chi^2(1)= 0.199$	*p-value=0.660*
INFDIF	$\chi^2(1)= 14.934$	*p-value=0.000*
LPGOLD	$\chi^2(1)= 0.017$	*p-value=0.895*

The cointegrating vector normalized with respect to LRER

LRER = 0.44 ×LPGOLD + 0.022×NIRDIF -0.033×INFDIF -0.003×TREND

The long run relationship between the real effective exchange rate and the proposed variables suggest the following: an increase in the price of gold is associated with an appreciation of the real effective exchange rate; an increase in nominal interest rate differential is associated with an appreciation of the real effective exchange rate; and finally an increase in the inflation differential is associated with a depreciation of the real effective exchange rate.

Estimates of the PVECM are presented in Appendix Table 2. We estimate (5) where X_2=[NIRDIF, LPGOLD]'. We include one lag for changes in each variable.[3] Appendix Table 3 shows results from the conditional parsimonious VECM. The off-

[3] Initially, we included 2 lags but the F-test accepts the null that the second lag is not significantly different from zero.

diagonal element of the residual correlation matrix is close to zero. Thus, there is no evidence of simultaneity and we can model every endogenous variable ignoring the distribution of the other.

As suggested in Section 1 we should expect to find structural breaks or non linearities in long term economic relations. Specification tests show that there is a problem of non-normality and heteroscadasticity, which implies that there might be structural breaks and/or non-linearities that a linear VAR fails to take into account (Canova 2007). It suggests the potential value of a nonlinear Error Correction Model as a superior alternative. The model we use is the MRS ECM which has the advantage of flexibility. As Hamilton (1994) argues, the MRS model allows the researcher to specify a probability model consistent with a broad range of outcomes and choose parameters within the class of this model on the basis of the data alone.

4.2 Estimation of MRS-ECM

We estimate a univariate version of (7)[4] and focus on the equation of the real exchange rate:

$$\Delta LRER_t = b_0(S_t) + b_1(S)ECE_{t-1} + b_2(S_t)\Delta LRER_{t-1} + b_3(S_t)\Delta NIRDIF_{t-1} +$$
$$+ b_4(S_t)\Delta INFDID_{t-1} + b_5(S_t)\Delta LPGOLD_{t-1} + b_6(S_t)\Delta NIRDIF_t + b_7(S_t)\Delta LPGOLD_{t-1} + u(S_t) \quad (8)$$

Variables have the same meanings as previously; *ECE* is the equilibrium correction error constructed from the cointegrating vector (Table 3). We proceed with the identification of regimes. In line with Gray (1996) we identify as regime 1 the regime with low volatility of errors and as regime 2 the regime with high volatility of errors. We consider that in regime 1 agents attach a high probability to the regime continuing, and in regime 2 attach a low probability to it continuing; below we present estimates of the transition probabilities for each regime (Table 4).

In Table 4, the coefficient σ_1 and σ_2 denotes the variance of the error term in regime 1 and regime 2. As can be seen the variance of the error term in regime 2 is more than four time higher than that of regime 1. The transition probabilities indicate that the high volatility regime is less persistent than the low volatility regime. More concretely the expected duration of the high and low volatility regimes are 22.7 and 41.6 months respectively.[5] Table 4 shows that the equilibrium correction error is not only significant in both regimes but its impact is higher in the high volatility regime (the coefficient on the lagged error correction term is higher in regime 2 than in regime 1).

[4] Since the PVECM shows that there is no evidence of simultaneity among the endogenous variables of the system we estimate only the single equation model.

[5] The expected duration of regime j is given by: $E(D=j) = \sum_{i=1}^{\infty} i p_{jj}^{i-1}(1-p_{jj}) = \frac{1}{1-p_{jj}}$.

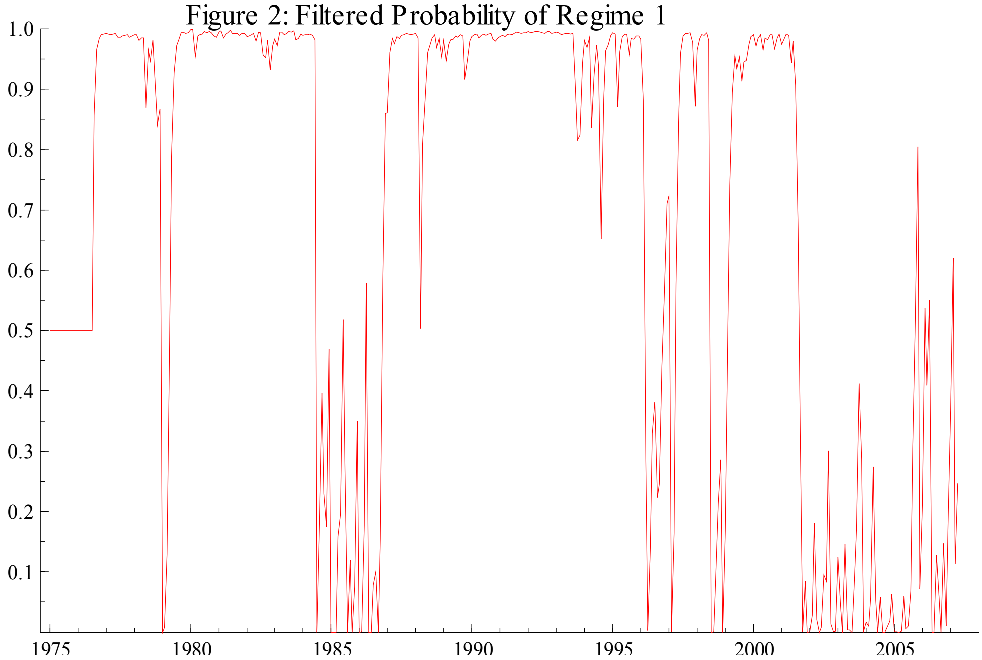
Table 4: Markov Regime Switching Applied to Parsimonious VECM

Regime 1		***Regime 2***	
σ_1	0.012 (0.0009)	σ_2	0.050 (0.004)
P_{12}	0.028 (0.013)	P_{12}	0.057 (0.024)
Intercept	0.059 (0.021)	Intercept	0.410 (0.134)
ECE_{t-1}	-0.020 (0.007)	ECE_{t-1}	-0.145 (0.046)
$DLRER_{t-1}$	0.256 (0.126)	$DLRER_{t-1}$	0.218 (0.121)
$DINFDIF_{t-1}$	-0.000 (0.001)	$DINFDIF_{t-1}$	-0.003 (0.005)
$DLPGOLD_t$	-0.002 (0.030)	$DLPGOLD_t$	0.224 (0.009)
$DLPGOLD_{t-1}$	-0.011 (0.016)	$DLPGOLD_{t-1}$	0.233 (0.108)
$DNIRDIF_t$	-0.003 (0.001)	$DNIRDIF_t$	-0.006 (0.006)
$DNIRDIF_{t-1}$	-0.001 (0.011)	$DNIRDIF_{t-1}$	-0.002 (0.004)

Notes:
Robust standard error in parentheses
P_{12} is probability of moving from regime 1 (low volatility regime) to regime 2 (high volatility regime)
P_{21} is probability of meeting from regime 2 (high volatility regime) to regime 1 (low volatility regime)
ECM is the error correction term.

Figure 2 shows the filtered probability of being in regime 1. The notable feature of this figure is the existence of two substantial periods in which the exchange rate function is in a regime (regime 2) which has low probability of persistence attached to it and hence is distinctly different from most years in the long run sample. The first is 1984-1987 and the second is 2002-2007. The switches between regime 1 and those two periods represent structural shifts in the exchange rate function. What change of circumstances in South Africa's economic environment can account for such regime switches in the Markov series?

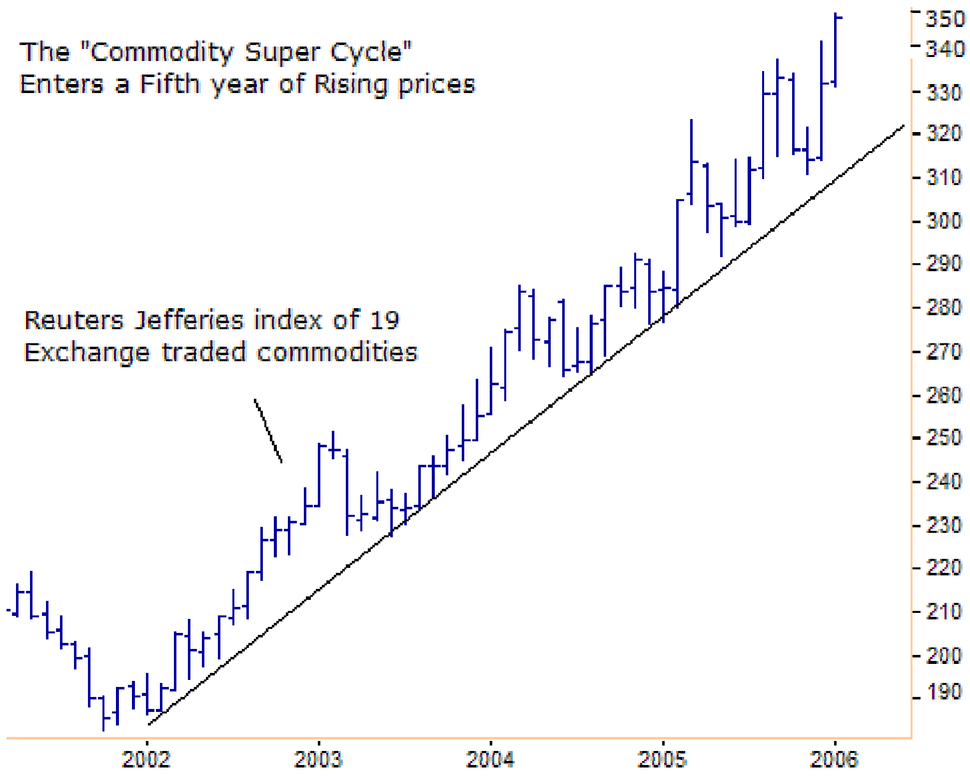
Figure 2: Filtered Probability of Regime 1



The first structural shift covers the years late 1984 to late 1986, immediately around the introduction in 1985 of South Africa's 'debt standstill' covering 60 per cent of South Africa's external debt (a default or forced rescheduling) and introduction of non-resident capital controls through a dual exchange rate system (the financial rand). (Harris, 1986). The dual rate system continued until 1995 but in the initial years the uncertainty and constraints this created appear to have characterised an unstable exchange rate regime.

The second structural shift occurred at the end of 2001 and the new regime continued to the end of our sample period. One potential explanation for the switch from regime 1 to regime 2 from the end of 2001, is the adoption of inflation targeting as a new monetary policy, which disturbed the relationship between the exchange rate, the interest rate differential and the inflation differential. However, the plausibility of that explanation is diminished by the fact that formal inflation targeting was adopted earlier, in 2000. And an informal and incomplete inflation target had been in effect since 1998.

A better explanation is the existence of belief in a commodity super cycle including gold and other metals which are major export earners for South Africa. The price of gold started a sustained and large rise at the end of 2001 and continued through 2007, concomitant with a generalised commodity boom including metals, as illustrated by Reuters Jefferies index:



The rise in the gold price as such is not enough to generate a regime shift, for the coefficients on the variables in the exchange rate equation measure its impact if economic relationships are unchanged. However, this boom generated a global perspective, widely reflected in South African discussions, that it was of a different type from those experienced in recent history. The perception was that instead of being short lived it was a commodity super cycle, driven by fundamental major shifts in the global economy (industrialisation of China especially), and destined to maintain its upward path for perhaps three decades (Heap, 2005; Cuddington and Jerrett, 2007). The change in perceptions is likely to have shifted behavioural relations so that the exchange rate's relation to its determinants came to be located in a new regime and consequently errors from the cointegrating relationship exhibited greater volatility.

We hypothesise that the greater the deviation of the gold price from its pre 2001 trend – which in our sample is correlated with time elapsed from the start of the commodities boom – the greater is agents' confidence that a new international environment exists and will persist as a commodity super cycle. We test that hypothesis by estimating the relationship between the transition probabilities previously calculated and the gold price. Specifically, we are interested in testing whether the gold price affects the probability of moving from one regime to another. To implement this, we follow Diebold and et al (1999) and endogenise the transitional probabilities by making them dependent on the price of gold. As suggested by Filardo (1994) and Diebold et al (1999), the variances and the parameters of each state, the transitional probabilities and their determinants are jointly estimated using the maximum likelihood method. The results are reported in table 5.

Table 5: Estimates of Time Varying Transition Probabilities

Regime 1		*Regime 2*	
σ_1	0.012	σ_2	0.051
Intercept	0.060 (0.018)	Intercept	0.367 (0.111)
ECM_{t-1}	-0.020 (0.006)	ECM_{t-1}	-0.129 (0.038)
$DLRER_{t-1}$	0.262 (0.085)	$DLRER_{t-1}$	0.268 (0.099)
$DINFDIF_{t-1}$	0.0001 (0.001)	$DINFDIF_{t-1}$	-0.004 (0.005)
$DNIRDIF_t$	-0.003 (0.000)	$DNIRDIF_t$	-0.010 (0.006)
$DNIRDIF_{t-1}$	-0.001 (0.001)	$DNIRDIF_{t-1}$	-0.002 (0.004)
LPGOLD		***LPGOLD***	
Intercept	3.100 (6.285)	Intercept	12.404 (6.285)
Slope Coefficient[1]	-1.167 (0.731)	Slope Coefficient[2]	-2.668 (1.133)

Notes:

Robust standard errors in parentheses.

1. Slope coefficient measures the effect of LPGOLD on the probability of moving from regime 1 (low volatility regime) to regime 2 (the high volatility regime).
2. Slope coefficient measures the effect of LPGOLD on the probability from moving from regime 2 (high volatility regime) to regime 1 (the low volatility regime.

Table 5 suggests that increase in the price of gold decreases the probability of moving from regime 2 (the high volatility regime) to regime 1 (the low volatility regime). On the other hand, the price of gold does not impact the probability of moving from regime 1 to regime 2. This evidence is consistent with our proposed explanation that an innovation after the end of 2001, the perception of a super commodity cycle, constitutes a shift of regime inducing parameter shifts in agents' behaviour with respect to the exchange rate, interest and inflation differential and gold price. Additionally it suggests that, as the super commodity cycle belief has become more entrenched, the probability of moving from regime 2 to regime 1 has been on the decline (the expected persistence of the new regime has grown).

4.3 Out of sample forecasts

As we have shown, the Markov Switching Regime Error Correction Model enables us to achieve reliable estimates of the real exchange rate equation in the presence of non-linearities, to identify structural shifts, and calculate regime transition probabilities. We also tested for the model's power by evaluating out of sample forecasting results, and obtain inconclusive results; both linear and non-linear (MRS) error correction models yield poor forecast errors and the MRS`model performs no better than the linear model.

We obtain out of sample forecasts in monthly data, computing recursively from January 1997 to April 2007, and employ Root Mean Square Forecast Error (RMSFE) tests. We use the RMSFE on the assumption that forecasters have a quadratic loss function which generates an optimal forecast of the predicted variable x at forecast horizon h , given by its conditional mean $E(x_{t+h}|\Omega_t)$. This point forecasting approach is required if the conditions for a density forecast are not satisfied. Using the tests derived by Diebold et. al. (1998) and Berkowitz (2001), we find that for both linear and MRS ECM models the density forecast conditions are not satisfied. To test the validity of assuming a quadratic loss function, rather than an asymmetric loss function as has been suggested for macroeconomic policy makers[6] , we test for biasedness.

Using the RMSFE to select an optimal forecast. we find that the linear ECM has slightly lower RMSFE than the MRS-ECM for the whole out of sample period, Figure 1 (or 4???). We test whether the differences between the two models' forecasts are statistically significant using the Diebold and Mariano (1995) (DM) test for the statistical significance of the difference, the results of which accept the null hypothesis that the forecasts of the linear and Markov Switching Regime models are not statistically different[7]. This corresponds with established literature showing that linear models are more robust to structural breaks that occur in out of sample forecasting (Diebold and Nason 1990; Clement and Hendry 2000).

The DM test results are shown in Figure 5 the ρ value, measured on the vertical axis, is below the critical value of 0.5 and thus indicating a significant difference only from 2003 to 2005.[8].

The use of RMSFE to select an optimal forecast would be invalid if the assumption that forecasters do not have an asymmetric loss function were false. An implicit test for an asymmetric loss function is to test whether forecasts are biased. Forecasts are unbiased when a sequence of forecasts is not over or underpredicted, in other words, the mathematical expectation E_t of forecast errors should be zero:

$$E\left(x_{t+h} - x_{t+h|t}\right) = 0$$

Due to the fact that we have only one observation at each point of time, in order to test for unbiasedness, we should examine whether the sample mean of the forecast errors is statistically different from zero:

---

[6] Nobay and Peel (2003) and Capistran (2006) show that central banks might have asymmetric loss over forecast errors rather than MSE loss. Ruge-Murcia (2000; 2003) finds empirical evidence that European Central Bank, Bank of England, Bank of Canada and Sweden have asymmetric loss function

[7] Diebold and Mariano (1995) proposed a test for the null that two forecasts are statistically different: $DM = \dfrac{\bar{d}}{\sqrt{\hat{V}(\bar{d})}}$ where $\bar{d} = \dfrac{1}{T}\sum_{i=1}^{T}\hat{d}_i$, $\hat{d}_i = \hat{u}_{1i}^2 - \hat{u}_{2i}^2$, $\hat{u}_{1i}^2$ and $\hat{u}_{2i}^2$ are forecast errors from model 1 and 2, T is the number of out-of-sample forecast errors and $\hat{V}(\bar{d})$ is the variance of $\bar{d}$. Harvey et.al., (1997) propose a small-sample correction of DM test where now $\sqrt{\hat{V}(\bar{d})} = T^{-1}(\hat{\gamma}_0 + \sum_{i=1}^{h-1}\hat{\gamma}_i)$, $\hat{\gamma}_i, i = 1,2,..h-1$ are the estimated autocovariancies of $\{\hat{d}_i\}_{i=1}^{T}$.

[8] However, even for this period the null is rejected marginally. More concretely, on average the ρ value for this period is 0.04.

$$\left\{e_{t+h|t}\right\}_{t=0}^{T}=\left\{x_{t+h}-x_{t+h|t}\right\}_{t=0}^{T}$$

A satisfactory test of unbiasedness for $h=1$ is via a test of $\mu=0$ in the following regression:

$$e_{t+1}=x_{t+1}-x_{t+1|t}=\mu+\varepsilon_{t+1} \qquad (9)$$

The results of the test for unbiasedness are presented in Figures 7, showing that both the linear and non-linear ECM are unbiased (recursive estimates of the p-values of H0: $\mu=0$ are above the threshold of 0.05).

Unbiasedness tests are reinforced by the rationality test. We assume that forecasters have an unknown asymmetric loss function and we test for rationality by using the quantile test suggested by Patton and Timmermann (2007).[9] Figure 8 shows the recursively estimated t-statistic of the null $\beta_1=0$. Evidence that the t-statistics are below the critical value of 1.96 implies that both models use efficiently the information set which is available at the time that the forecast is made. Thus, both the rationality and unbiasedness tests imply that the assumption of forecasters having a quadratic loss function is not rejected and our use of the RMSFE criterion is valid.

5 Conclusion

Our study of the South African rand, uses cointegration with a Markov Switching Regime Error Correction Mechanism. We have found that a long run equilibrium real effective exchange rate for South Africa can be explained by two variables whose role is accounted for by standard economic theories, a nominal interest rate differential and an inflation differential, augmented only by a country specific variable, the price of gold. The finding that the rand's real value is explained by a few fundamentals contrasts with existing literature which includes a larger range of explanatory variables, each requiring specific rationalisations.

We would not have expected to find that an exchange rate equation fits a data set spanning over three decades, during which South Africa and the international economy have experienced numerous major changes, without taking account of structural breaks. A priori it is plausible to expect that several would have occurred, perhaps to the extent of making it impossible to identify any stable relationships. The

[9] Patton and Timmermann (2006) established a test of forecast optimality when the type of loss function is unknown. They argue that when the dynamics in the mean and variance of the predicted variables are unknown, forecast optimality can be tested using the following regression:

$$I_{t+h,t}=\beta_0+\beta_1 Z_t+u_{t+h}$$

where $I_{t+h,t}=1(y_{t+h}<\hat{y}_{t+h})$. 1(Y) is an indicator function that equals to one if Y is true and otherwise it is zero. A test of the null hypothesis, $\beta_1=0$, shows that the indicator function is uncorrelated with the vector Z_t of information variables that were known at the time that the forecast was made.

major shocks to the system experienced by South Africa between 1975 and 2007 included anti-apartheid riots, the growth of a powerful trade union movement, increased doubts in the 1980s over the viability of the apartheid system; a foreign capital 'sudden stop', default, and the introduction of a dual exchange rate; the creation of a democratic state; changes in corporate governance and large transfers of equity (black economic empowerment policies) ; trade and currency liberalization; and shifts in the objectives and tools of monetary policy and exchange rate policy. Externally, global changes in manufacturing and mining were reflected in major declines in cost competitiveness for important South African sectors, while global financial market liberalization, innovation in financial institutions and markets, and technological innovations changed the conditions under which flows between the rand and other currencies occurred.

Trying to account for structural breaks related to those innovations by using dummy variables[10] would involve an arbitrary a priori selection of shift dates and potentially too many to be meaningful. Instead, we use a Markov Regime Switching Error Correction Mechanism model to identify significant structural breaks endogenously. With that non-linear approach we obtain the unexpected result that South Africa's exchange rate relationship experienced only two periods where the regime changed: late 1984 to late 1986, and end 2001 to 2007. The first was initiated by the withdrawal of US investment funds and then US banks from rand assets and the 'debt standstill' (default) accompanied by new exchange controls. The second is consistent with the notion that the rise in global commodity prices from the end of 2001 generated a belief that the global economy's structure (and South Africa's place in it) had changed and was experiencing the start of a long lasting commodity super cycle, and that belief changed the relationship between the real exchange rate and its long run determinants.

[10] As does Frankel 2007b.

Figure 3: Actual and Fitted values of Linear and MRS ECM

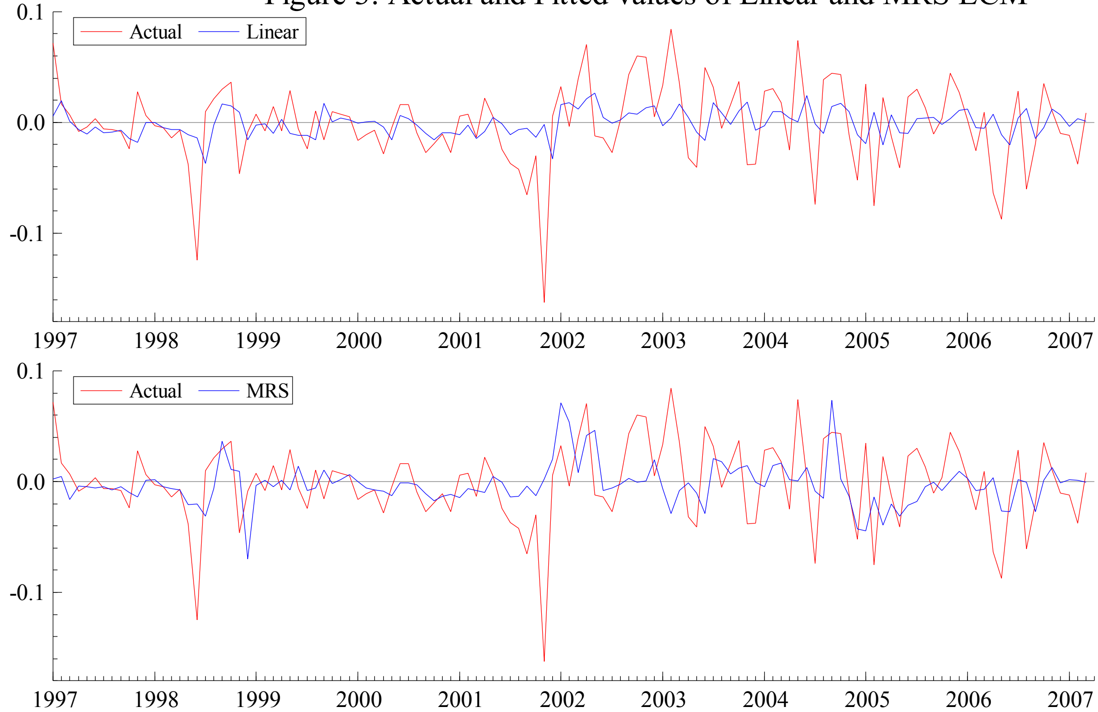
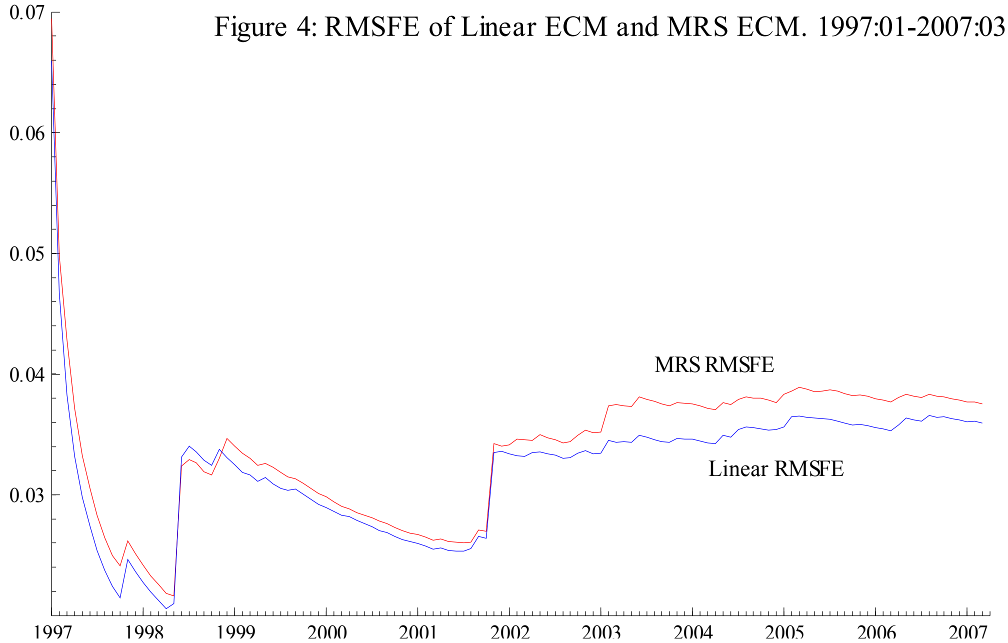


Figure 4: RMSFE of Linear ECM and MRS ECM. 1997:01-2007:03



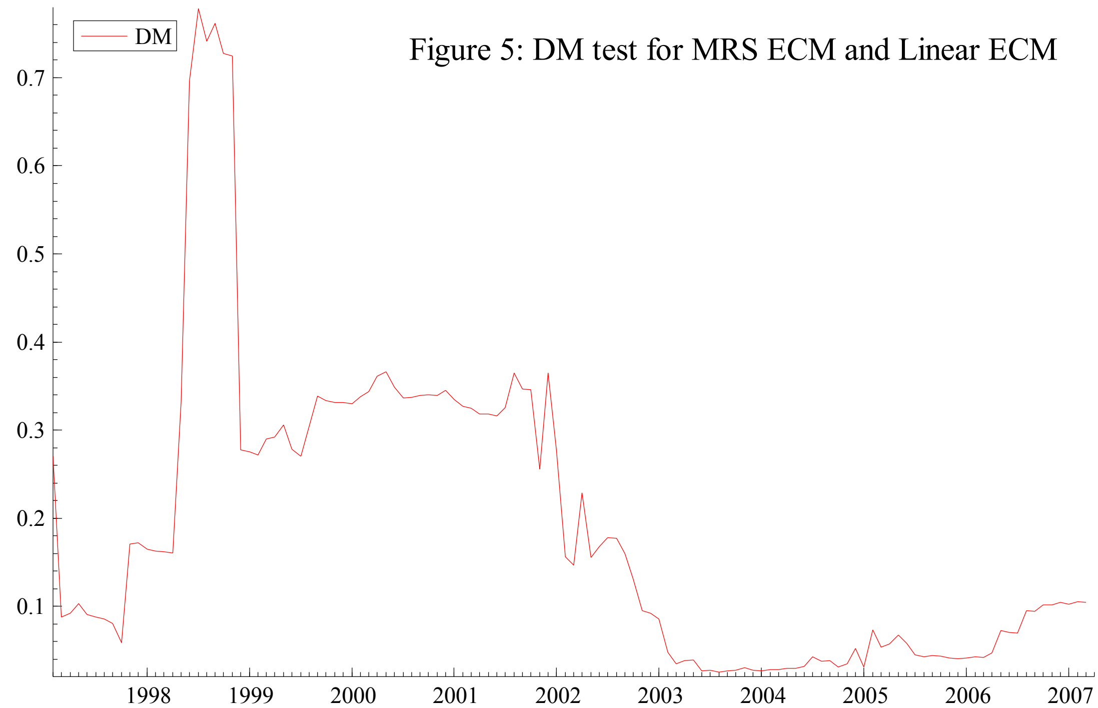


Figure 5: DM test for MRS ECM and Linear ECM

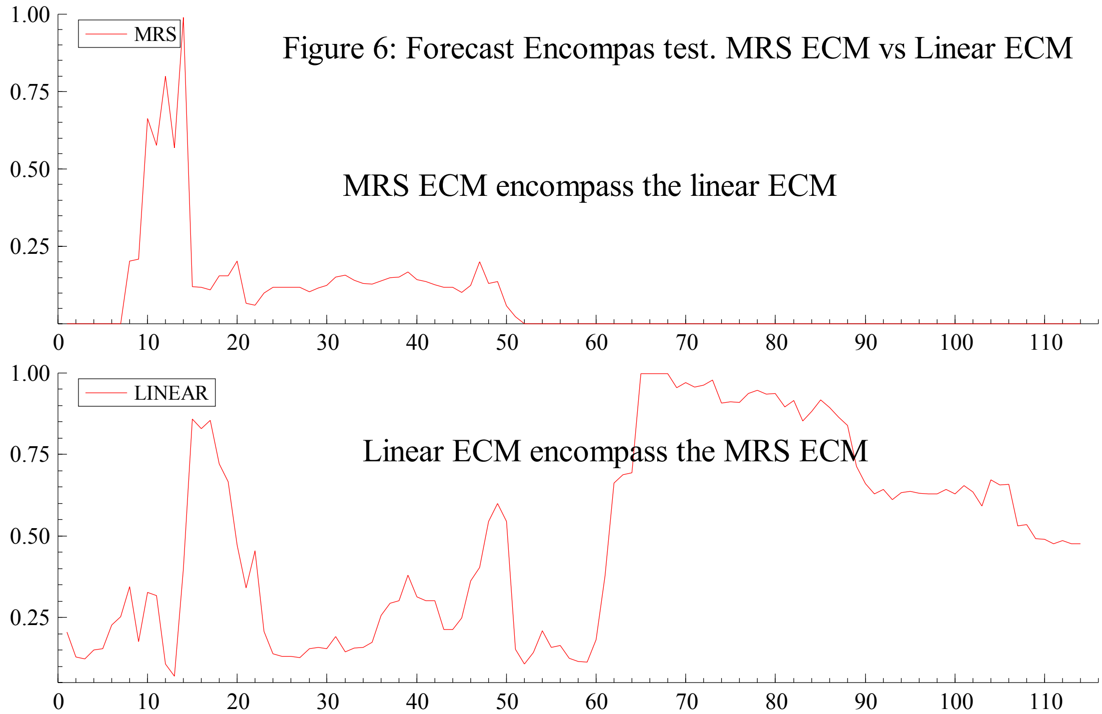


Figure 6: Forecast Encompas test. MRS ECM vs Linear ECM

Figure 7: P-values if Unbiasedness Test

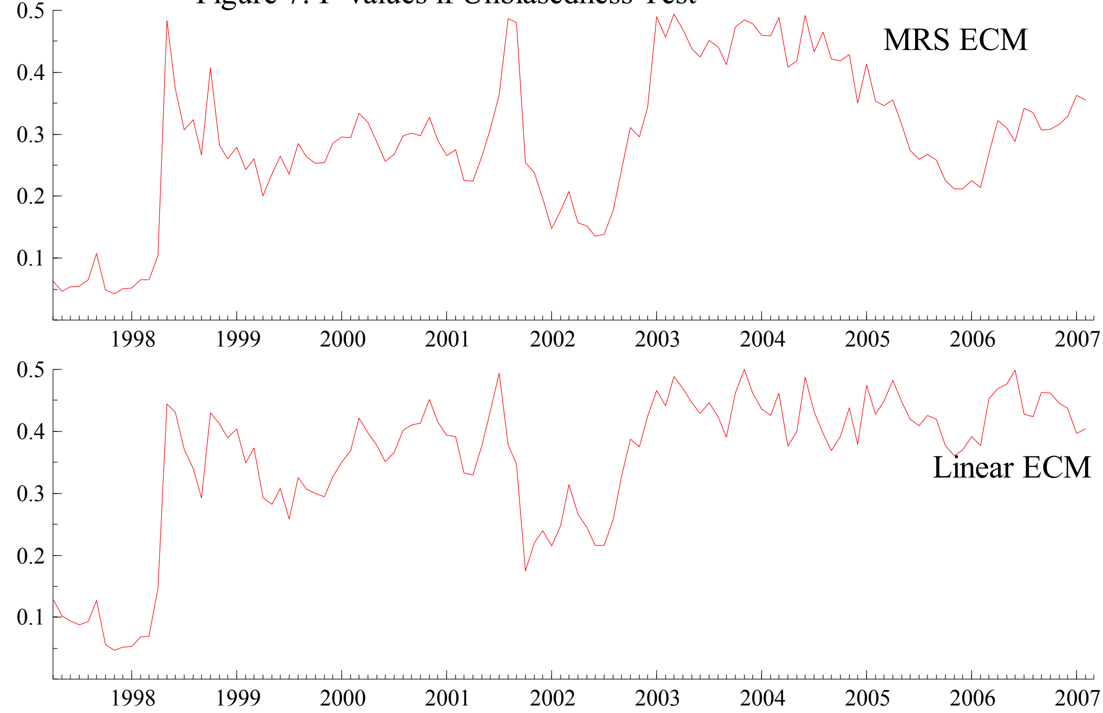
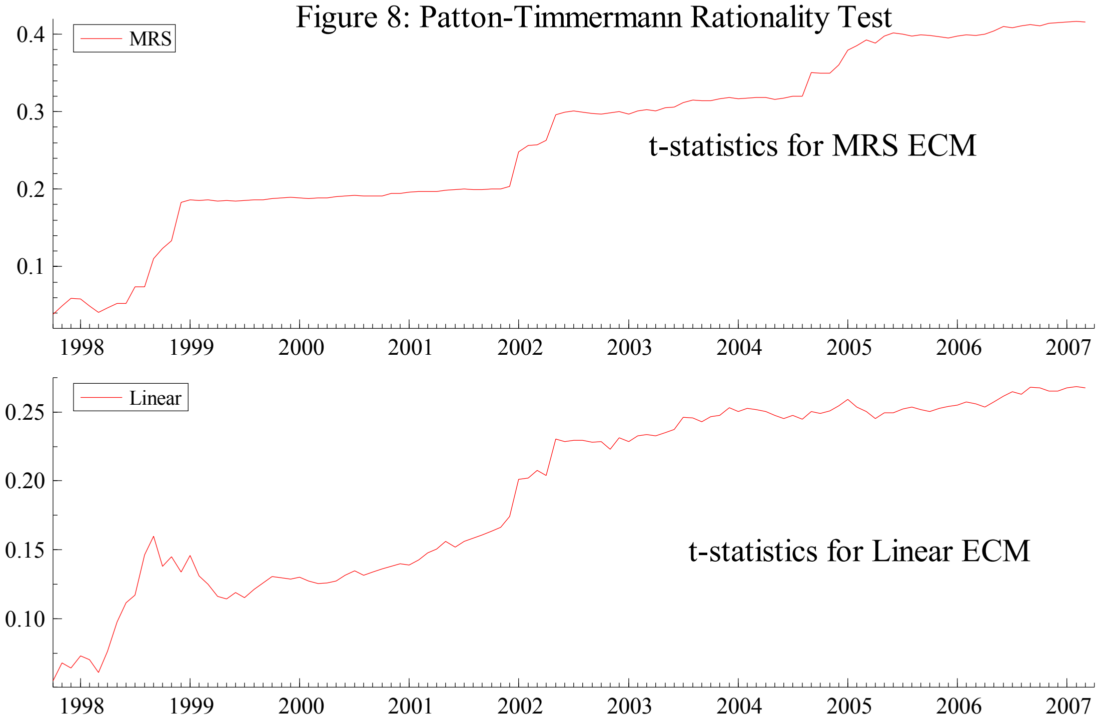


Figure 8: Patton-Timmermann Rationality Test



Appendix

Table 2: PVECM

	DLRER		DLPGOLD		DNIRDIF		DINFDIF	
Variable	Coefficient	t-prob	Coefficient	t-prob	Coefficient	t-prob	Coefficient	t-prob
DLrer_1	0.243	0.000	0.145	0.095	-0.198	0.867	-2.638	0.031
DLrer_2	0.058	0.273	-0.147	0.094	-0.465	0.695	2.250	0.069
DLpgold_1	0.029	0.359	-0.019	0.726	-3.058	0.000	-1.302	0.084
DLpgold_2	-0.060	0.069	-0.024	0.654	-2.990	0.000	-0.415	0.588
Dnirdif_1	-0.005	0.017	0.003	0.365	0.216	0.000	0.145	0.007
Dnirdif_2	0.005	0.038	-0.006	0.092	0.044	0.394	0.025	0.642
Dinfdif_1	-0.002	0.426	0.001	0.807	-0.011	0.814	0.046	0.362
Dinfdif_2	0.000	0.939	-0.003	0.325	-0.012	0.798	0.069	0.158
CI_1	-0.059	0.000	-0.004	0.869	0.201	0.576	-1.913	0.000
Constant	0.165	0.000	0.016	0.829	-0.536	0.596	5.377	0.000

Table 3:Conditional PVECM 1

URF	DLRER		DINFDIF	
Variable	Coefficient	t-prob	Coefficient	t-prob
Dlrer_1	0.228	0.000	-2.227	0.062
Dinfdif_1	-0.001	0.634	0.041	0.404
ECM_1	-0.051	0.001	-1.751	0.000
DLpgold	0.032	0.301	-1.469	0.042
DLpgold_1	0.028	0.389	-1.299	0.087
Dnirdif	-0.004	0.057	-0.041	0.434
Dnirdif_1	-0.002	0.286	0.171	0.001
Constant	0.143	0.001	4.923	0.000

References

Anderson, T.W. (2003), An Introduction to Multivariate Statistical Analysis (3rd edition). John Wiley: New York.

Aron, J, I Elbadawi, and B Kahn (1997) Determinants of the real Exchange Rate in South Africa, CSAE Working Paper WPS/97-16

Berkowitz, J., (2001), Testing density forecasts, with applications to risk management, *Journal of Business and Economic Statistics*, Vol. 19, pp. 465-474.

Canova, F., (2007), Methods for Applied Macroeconomic Research., Princeton, University Press.

Capistran, C. C., (2006). Bias in Federal Reserve Inflation Forecast: Is the Federal Reserve Irrational or just Cautious? working paper, UCSD.

Cashin P. Céspedes L. F. and Sahay R. (2003) Commodity Currencies and the Real Exchange Rate, Central Bank of Chile Working Papers No 236

Clements, M., and D. Hendry, (2000), Economic forecasting in the face of structural breaks, in Holly, S. and Weale, M. (eds.) Econometric Modelling: Techniques and Applications, Cambridge University Press. (2000). 3-37.

Cuddington, J., and D. Jerrett, (2007) Super Cycles in Metals Prices? unpublished seminar paper, Department of Economics, University of Colorado, 24 October.

Diebold FX., and Nason JA (1990), Nonparametric exchange rate prediction, *Journal of International Economics*, Vol. 29, pp. 315-332.

Diebold, F. X., T.A. Gunther, and A. S. Tay, (1998), Evaluating Density Forecast, *International Economic Review*, Vol. 39, 863-883.

Diebold, F., and R. S. Mariano, (1995), Comparing Predictive Accuracy, Journal *of Business and Economic Statistics*, Vol. 13, pp. 253-263.

Diebold, F., Lee J-H, and Weinbach G.C (1999) Regime switching with time-varying transition probabilities, in Business cycles: durations, dynamics and forecasting, Princeton University Press.

FEASability pty (2006) Exchange rates, growth and unemployment: Progress Report, HSRC, SA unpublished

Filardo, A. (1994) Business-cycle phases and their transitional dynamics, Journal of Business and Economic Statistics, 12, 3, 299-308.

Frankel, J (2007a), Making Inflation Targeting Appropriately Flexible, unpublished presentation at SA National Treasury, Pretoria 16th January 2007

Frankel, J (2007b) On the Rand: Determinants of the South African Exchange Rate South African Journal of Economics Vol. 75:3, 425-441

Golub, S.S. (2000) South Africa's International Cost Competitiveness, SA department of Finance, unpublished

Hamilton J., D., and Lin., G., (2006), Stock Market Volatility and the Business Cycle, *Journal of Applied Econometrics*, Vol. 11, No. 5, pp. 573-593.

Hamilton, J., D., (1994), Time Series Analysis, Princeton University Press.

Harris L (1986) South Africa's External Debt Crisis, *Third World Quarterly,* vol.8, no.3, July

Heap, A. (2005). *China - The Engine of a Commodities Super Cycle*. New York City: Citigroup Smith Barney.

Johansen, S. (1992), "Testing for weak Exogeneity and the Order of Cointegration in the UK Money Demand Data", *Journal of Policy Modelling*, Vol.14, 313-34

Johansen, S., (1992), Determination of Cointegration Rank in the Presence of a Linear Trend, *Oxford Bulletin of Economics and Statistics* Vol. 54, pp. 383-397.

Johansen, S. (1996), Likelihood-Based Inference in Cointegrated Vector Autoregressive Models, 2nd edition. Advanced Texts in Econometrics, Oxford University Press: Oxford.

MacDonald, R., and Ricci, L. (2004). 'Estimation of the equilibrium real exchange rate for South Africa', IMF Working Paper No WP/03/44 and *South African Journal of Economics*, vol. 72(2), (June), pp. 282-304

MacDonald, R., Menkhoff, L. and Frömmel, M. (2005). 'Do fundamentals matter for the D-Mark/Euro Dollar? A regime switching approach', *Global Finance Journal*, vol. 15(3), (February), pp. 321-335

Nobay, R.A., and D.A. Peel, (2003), Optimal Discretionary Monetary Policy in a Model of Asymmetric Central Bank Preferences. *The Economic Journal*, Vol. 113, pp. 657-665.

Patton, A., and Timmermann, A. (2007), Testing Forecast Optimality Under Unknown Loss. *Journal of American Statistical Association*, Vol. 102(480), pp. 1172-1184.

Republic of South Africa (2006), Background Document:Accelerated and Shared Growth-South Africa (ASGISA), SA Deputy President Phumzile Mlambo-Ngcuka,6 February 2006.

Ruge-Murcia, F.J., (2000), Uncovering Financial Market's Beliefs about Inflation Targets, Journal of Applied Econometrics, Vol. 15, pp. 483-512.

Ruge-Murcia, F.J., (2003), Inflation Targeting Under Asymmetric Preferences. Journal of money Credit and Banking, Vol. 35, pp. 763-785.

Sarno, L and MP Taylor (2002b) The economics of exchange rates, Cambridge: Cambridge University Press.

Sarno, L and MP Taylor, (2002a) Purchasing Power Parity and the Real Exchange Rate, IMF Staff Papers,49,1;