NOMINAL EXCHANGE RATES IN KENYA. ARE SHOCKS TRANSITORY OR PERMANENT? AN EMPIRICAL INVESTIGATION BASED ON FRACTIONAL INTEGRATION

ABSTRACT

This article deals with the analysis of several nominal exchange rates in Kenya, examining, by means of fractional integration, if shocks in the series are transitory or permanent. The results support the view that most of the exchange rates are nonstationary and non-mean-reverting. This result proves to be critical in the case of the Kenya shilling US dollar rate which is found to be non-mean-reverting and so provides important insights into the 2011 exchange rate crisis in Kenya in which the Central Bank of Kenya initially delayed before taking strong policy actions. Only when it took such actions did the Kenya Shilling dollar rate revert to more normal levels in late 2011. Evidence of mean reversion is obtained for the cases of Canada and India, and to a lesser extent for Sweden and Uganda. Thus, shocks affecting the rates against these currencies are expected to be transitory, disappearing in the long run. Policy implications are derived at the end of the article.

Keywords: Nominal exchange rates; mean reversion; fractional integration
JEL Classification: C22.

1. Introduction

This paper deals with the analysis of nominal exchange rates in Kenya. We focus on the issue of persistence, examining the degree of dependence of the series from a fractional viewpoint. This is important in the sense that by allowing for a fractional degree of integration (i.e. using I(d) models with d as a real value) we permit a richer degree of flexibility in the dynamic specification of the series, not achieved when using the classical approaches based on integer degrees of differentiation, i.e. d = 0 in the case of stationary series, and d = 1 in the case of unit roots.

An analysis of the degree of persistence in the nominal exchange rates is essential since it reflects the stability of the exchange rate of the country in question. Furthermore, this kind of information is important for policy makers in the event of an exogenous shock, when different policy measures have to be adopted depending on the nature of the shock, which is clearly related to the degree of persistence of the series.

The research objective of this study is to extend the current empirical research represented by Fang et al. (1994), Crato and Ray (2000) and Wang (2004) among others by introducing fractional integration with which to analyze the degree of dependence in the nominal exchange rates in Kenya. This is important in that it can tell us which currencies are more persistent in the Kenyan foreign exchange rates market. In the case of a high level of persistence, in the event of exogenous shocks, strong policy measures will be required to recover their original long term projections. On the other hand, with currencies which are less persistent in the exchange market, there will be no need for strong measures in the case of shocks since the series will return by themselves to their original trends. In other words, the key issue becomes the analysis of mean reversion in the exchange rates, and, using fractional integration, the differencing parameter d will be crucial to determine if mean reversion occurs if $d < 1$. Moreover, the lower the parameter d is, the faster the process of convergence will be to its long run mean.

The question of establishing the process of convergence to the long run mean is critical for the Kenyan economy because the Kenya shilling has, since the beginning of 2011, experienced considerable depreciation against key international currencies, especially the US dollar, against which it had depreciated by 33% by September 2011 compared to January 2011. By September 2011 it was considered as the second most volatile currency in the world in the year so far. The question of the type of policy intervention required to stabilize the currency therefore becomes very critical and has indeed been very controversial. The Central Bank of Kenya has tried several measures to stabilize the shilling since the beginning of the year including increasing the Central Bank Rate in August 2011, first by 400 basis points to stand at 11% and then shortly afterwards in October 2011 by 550 basis points to 16.5%, the highest increase in a single sitting since the benchmark rate was introduced in June 2006. Eventually, the rate rose to 18% by December 2011. The Monetary Policy Committee also increased the cash reserve ratio twice from an initial figure of 4.5% so that it eventually stood at 5.25% on December 1st 2011 with a view to mopping up 7.5 billion Kenya shillings from banks. The shilling has recently regained some ground to trade at about 83 Kenya shillings against the dollar from an all time low of 107 owing to the strong policy measures taken by the Central Bank of Kenya. The question of mean reversion of the nominal exchange rate therefore becomes very critical in the current policy context since it gives an indication of how robust the policy intervention needs to be, based on mean reversion of the Kenya shilling against key international currencies.

A key external shock experienced earlier in 2011 was the substantial increase in international oil prices associated with the Arab Spring. In Kenya imports of oil make up about 15% of total imports and so substantial price increases often put downward pressure on the exchange rate. The Central Bank of Kenya had also steadily been boosting its foreign exchange reserves by buying foreign currencies on the domestic market between November and April 2011 putting further downward pressure on the Kenya shilling. The Euro zone crisis also contributed to an appreciating dollar as some investors in euro-denominated assets shifted to US dollars.

The Kenyan economy is classified by the International Monetary Fund as operating an independent float between 1992 and 1997 and a managed float since 1998. On balance, exchange rate evidence in Kenya suggests that, during periods of relative tranquility in foreign exchange markets prior to 2011, the Central Bank of Kenya has generally been able to smooth out exchange rate volatility with relatively modest fluctuations in its intervention policy (O'Connell et al., 2011). In 2011, however, Central bank monetary policy tools seemed, for a long time, to have limited effectiveness against the depreciation of the Kenya shilling against major international currencies, especially the US dollar. This raised the importance of the issue of mean reversion in the nominal exchange rate.

The structure of the paper is as follows. Section 2 presents a literature review on empirical issues dealing with exchange rates dynamics. Section 3 describes the methodology used in the paper. Section 4 is devoted to the data and the empirical results, while Section 5 contains some concluding comments.

2. Literature review

There have been a number of empirical studies investigating the dynamics of short and the long run nominal and real exchange rates. Anthony and MacDonald (1999) analysed the prediction of the target zone in the exchange rate of seven currencies in the Exchange

Rate Mechanism (ERM). They used standard univariate unit root tests and found some evidence of stationarity, and therefore, mean reversion. Mean reversion in real exchange rates was also tested by Bleaney et al. (1999), using data from countries with episodes of high inflation. Their results showed that a unit root model is more appropriate than models with fixed rates of mean reversion. Sollis et al. (2002) proposed new tests to analyze mean reversion, based on smooth transition autoregressive (STAR) models. The methods introduced included a test that forced mean reversion to be symmetric about the integrated process central case, and another that permits asymmetry. In comparison to the usual unit root tests, the authors found stronger evidence against the unit root null hypothesis in a number of real exchange rates against the US dollar and the Deutsche mark. In the same line, using a sample of G7 countries, Campa and Wolf (1997) found strong evidence of mean reversion, and showed that the deviations of real exchange rates and trade from trend are uncorrelated.

The above literature focuses on integer degrees of differentiation for the series of interest. In the context of long memory and fractional integration, Booth et al. (1982) was the first to apply long range dependence techniques to exchange rates. He employed R/S techniques to daily rates for the British pound, French franc and Deutsche mark, and found long term persistence during the flexible exchange rate period (1973-1979) but anti-persistence during the fixed exchange rate period (1965-1971). Later, Cheung (1993) also found evidence of long memory behaviour in foreign exchange markets during the managed floating regime. On the other hand, Baum et al. (1999) estimated fractional ARIMA (ARFIMA) models for real exchange rates in the post-Bretton Woods era and found almost no evidence to support long run PPP. Additional papers on exchange rate dynamics using fractional integration are Fang, Lai and Lai (1994), Crato and Ray (2000) and Wang (2004).

All the above literature focuses mainly on developed countries. There is far less literature on exchange rates dynamics in developing countries and it often focuses on integer degrees of differentiation. Studies in developing countries have also generally focused more on real exchange rate dynamics rather than nominal exchange rates. Recent exchange rate literature in developing countries has focused on several themes. The issue of flexible versus fixed exchange rates continues to dominate the literature on exchange rates in developing countries. Hoffmann (2007) for example investigates the hypothesis that in a small open economy flexible exchange rates act to mitigate the effects of external shocks more effectively than fixed exchange rate regimes using a sample of forty two developing countries. The paper assesses whether the response of real GDP, the trade balance and the real exchange rates to world output and world real interest rates differs across exchange rate regimes. The paper demonstrates that there are significant differences in the variability of macroeconomic aggregates under fixed and flexible regimes. The literature on exchange rate crises especially in Latin America is also quite prevalent. Such literature often considers the theme or overvaluation of currencies in fixed exchange rate systems that contribute to exchange rate crises. Kildegaard (2006), for example, examines the role of structural factors in the Mexican real exchange experience, especially the crisis of December 1994. The cointegrating equation indicates a severe undervaluation during the 1980s and only modest overvaluation in the period immediately preceding the devaluation in December 1994. A theme that also emerges in the exchange rate literature of developing countries is the impact of exchange rate volatility on macroeconomic variables. Arize, Osang and Slottje (2000), for example, empirically investigate the impact of real exchange rate volatility on export flows of thirteen developing countries over the quarterly period 1973-1996. The issue of volatility has in turn found recent expression in the empirical analysis of mean reversion of exchange rates in developing countries.

On the other hand, mean reversion of the exchange rates and its implications for policy has only recently begun to receive attention in the literature. Arize (2011) analyses mean reversion in real exchange rates in developing countries based on data from 1980 to 2009. His results reveal strong evidence in favour of mean reversion in both linear and non-linear mean reversion with non-linear mean reversion dominating beyond the 2005 sample period. Anouro, Liew and Elike (2006) find, via formal linearity tests, that all thirteen selected African countries exchange rates are non-linearly behaved. They further reveal, through non-linear stationarity testing, that eleven out of the thirteen African exchange rates examined are non-linear stationary. Their results indicate that bilateral nominal exchange rates are mean reverting towards the PPP equilibrium positions. Literature on exchange rates in Kenya does not specifically consider the issue of mean reversion and none of the existing literature adopts a fractional integration approach. It focuses mainly on using cointegration approaches to investigate the relationships between exchange rates and other influencing variables. Ndung'u (2000) analyses the relationship between the real exchange differential on the one hand and the implications they have on portfolio capital flows on the other. His results indicate that the nominal exchange rate deviates from the perceived long run equilibrium by the purchasing power parity relationship and that these deviations are governed by the interest rate differential. Were et al. (2001) analyse Kenyan exchange rate movements in a liberalized environment. Using an error correction model, the paper adopts a general empirical specification of the exchange rate equation encompassing the interest rate and price differential as well as the current account balance and net external inflows in the 1990s. Kiptui and Kipyegon (2008) adopt an error correction model to capture the long run and short run dynamics of the impact of external shocks on real exchange rates in Kenya including the terms of trade, net foreign exchange flows and openness based on monthly data from 1996 to 2001. Their results show that, to a large extent, external shocks influence the real exchange rate as demonstrated by the significance of terms of trade and openness in long run and short run estimations. They also find that real exchange rates in Kenya are integrated of order one (i.e. I(1). O'Connell et al. (2011) analyse the challenge of exchange rate management in Kenya in the context of increasing capital mobility. They argue that Kenya adopted an open capital account with a view to widening portfolio opportunities for domestic residents, attracting foreign capital inflows and enhancing financial development. Their analysis shows that the Central Bank of Kenya can achieve considerable short-term smoothing of the exchange rate during normal times but this proves more challenging in the presence of large shocks. There is therefore a need to investigate specifically the question of whether shocks to nominal exchange rates in Kenya are transitory or permanent and to consider the policy implications, which is precisely the subject of the present paper.

3. Methodology

As mentioned earlier, in this study we focus on fractional integration. For this purpose we need, first, to define an integrated of order 0 or I(0) process. We say that a covariance stationary sequence $\{x_t, t = 0, \pm 1, ...\}$ is I(0) if the infinite sum of the autocovariances is finite. Alternatively, it can be defined in the frequency domain as a process with a spectral density function that is positive and bounded at all frequencies. Standard examples of I(0) processes are the white noise case, and allowing weak autocorrelation, the type of stationary AutoRegressive Moving Average (ARMA) models.

Having said this, we say that a process $\{y_t, t = 0, \pm 1, ...\}$ is integrated of order d, and denoted as $y_t \approx I(d)$, if it can be represented as:

$$(1 - L)^d y_t = x_t, \qquad t = 1, 2, ..., \tag{1}$$

with $y_t = 0$, $t \leq 0$, and where x_t is I(0). Note that the polynomial on the left-hand-side of equation (1) can be expanded as

$$(1 - L)^d = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j = 1 - d L + \frac{d(d-1)}{2} L^2 - \dots, \tag{2}$$

implying that the higher d is, the higher the level of dependence is between the observations.

Processes with d > 0 in (1) display the property of "*long memory*", characterized because the autocorrelations decay hyperbolically slowly and the spectral density function is unbounded at the origin. The origin of these processes dates back to the 1960s, when Granger (1966) and Adelman (1965) pointed out that most aggregate economic time series have a typical shape where the spectral density increases dramatically as the frequency approaches zero. However, differencing the data frequently leads to overdifferencing at the zero frequency. Fifteen years later, Robinson (1978) and Granger (1980) showed that aggregation could be a source of fractional integration. Since then, fractional processes of form as in (1) have been widely employed to describe the dynamics of many economic and financial series (see, e.g. Diebold and Rudebusch, 1989; Sowell, 1992; Baillie, 1996; Gil-Alana and Robinson, 1997).

It is well documented that the findings on persistence and long memory can vary substantially depending on the method used. Thus, a robustness check that uses different approaches is crucial. Our analysis is initially based on a parametric approach developed by Robinson (1994). It allows us to test the null hypothesis:

$$H_o : d = d_o, \tag{3}$$

in a model given by:

$$y_t = \beta' z_t + x_t; \quad (1 - L)^d x_t = u_t, \quad t = 1, 2, ..., \tag{4}$$

for any real value d_o, where z_t in (4) is a (kx1) vector of deterministic regressors that may include, for example, an intercept or an intercept with a linear time trend. This method has several advantages compared with other approaches: first, it allows us to test any real value d_o, encompassing thus stationary ($d_o < 0.5$) and nonstationary ($d_o \geq 0.5$) hypotheses; second, the limit distribution is standard normal and this limiting behavior holds independently of the type of deterministic terms included in the model and of the way of modeling the I(0) error term u_t; this method is also the most efficient one in the Pitman sense against local departures from the null. Moreover, it does not require Gaussianity and it is supposed to be robust against conditional heteroscedastic errors.[1] On the other hand, two estimation procedures based on the Whittle function in the frequency domain, one parametric (Dahlhaus, 1989) and the other semiparametric (Robinson, 1995) will also be implemented in the paper.

4. Data and empirical results

The data examined in this work correspond to the Kenyan nominal exchange rates against the US dollar, the Canadian dollar, the Danish kroner, the Indian rupee, the Japanese yen, the South African rand, the Swedish kroner, the Swiss franc, the Tanzanian shilling, the Ugandan shilling and the UK Sterling pound. The time period examined is in all cases 1996Q1 – 2009Q2 except for the Tanzanian shilling and Ugandan shilling where the starting date is 1996Q3. The source of data is the Central Bank of Kenya.

[1] Empirical applications using this approach can be found in Gil-Alana and Robinson (1997), Gil-Alana (2000), Gil-Alana and Henry (2003), etc.

We start this section by estimating the value of the fractional differencing parameter d in the following model,

$$y_t = \alpha + \beta t + x_t; \quad (1 - L)^d x_t = u_t, \quad t = 1, 2, \ldots, \qquad (5)$$

where y_t is the time series we observe (in our case the nominal exchange rates); α and β are the coefficients corresponding to the intercept and the linear time trend; d is the (possibly fractional) differencing parameter and u_t is an I(0) process that we initially assume is white noise though later we also allow for weak autocorrelation.

Note that this is a quite general specification, and permitting d to be a real value we also consider other standard approaches as particular cases of interest. Thus, if d = 0, we have the classical "trend stationary" representation, while d = 1 implies unit roots or the I(1) processes advocated by many authors.[2] Note, however, that we allow for the possibility of fractional values of d.

We consider the three standard cases examined in the literature, that is, the case of no regressors (α = β = 0 a priori in (5)), an intercept (α unknown and β = 0 a priori), and an intercept with a linear time trend (α and β unknown in equation (5)).

[Insert Table 1 about here]

Table 1 reports the Whittle estimates of d along with their corresponding 95% confidence bands of the non-rejection values of d using Robinson (1994) for the three cases specified above of no regressors, an intercept, and an intercept with a linear trend. Several features are noticeable in the table: first, all the estimated values of d are strictly above 0.5, implying thus nonstationary behavior. The time trend is found to be statistically significant in five cases: Canada, Denmark, Switzerland, Tanzania and Uganda, while in the remaining cases (India, Japan, South Africa, Sweden, UK and USA) only an intercept seems to be required. If we focus on the estimated values of d we

[2] The I(1) approach has been employed in many empirical papers following the seminal work by Nelson and Plosser (1982).

observe only two cases where d is found to be significantly smaller than 1 and thus showing mean reversion behaviour: India (d = 0.652) and Canada (d = 0.536).[3] For the remaining cases, the unit root null (i.e. d = 1) cannot be rejected; in seven of these cases (Uganda, Switzerland, UK, Denmark, Tanzania, Sweden and South Africa) the estimated value of d is found to be smaller than 1 while in another two cases (Japan and USA) the estimated d's are above 1.

[Insert Table 2 about here]

Table 2 displays the estimated coefficients for each series for the selected models depending on the specification for the deterministic terms. The series are ordered according to their degree of dependence, noticing that the highest values correspond to the USA and Japan, while the lowest ones (those in fact displaying mean reversion) to India and Canada.

[Insert Table 3 about here]

According to these results, in the event of an exogenous shock in the exchange rates, its effect will be permanent in the cases of the rates against the US dollar, the Japanese yen, the South African rand, the Swedish kroner, the Tanzanian shilling, the Danish kroner, the UK sterling pound, the Swish franc and the Ugandan shilling. On the contrary, for the Indian rupee and the Canadian dollar, the effects will be transitory disappearing in the long run. Thus stronger policy measures are required in the former cases than in the latter ones.

The results presented so far may be biased due to the lack of autocorrelation for the error term u_t. In what follows, we assume that u_t may be weakly autocorrelated. However, instead of using the standard approach based on ARMA specifications, we use a non-parametric method due to Bloomfield (1973). This model is implicitly determined

[3] Note that for these two countries the confidence intervals exclude the case of d = 1.

by its spectral density function and produces autocorrelations decaying exponentially as in the ARMA case. Moreover, this method is very convenient in the context of the tests of Robinson (1994) employed in this work.[4] The results using this approach are presented in Table 4.

[Insert Table 4 about here]

The first noticeable feature in Table 4 is that the time trend is now required in all except one series (India). The estimated values of d are all smaller than 1 and mean reversion (i.e., values of d significatively smaller than 1) are found in four series (India, Canada, Sweden and Uganda).

[Insert Tables 5 and 6 about here]

Table 5 displays the coefficients of the selected models. For seven of the series (Tanzania, U.S.A., U.K., Switzerland, Denmark, South Africa and Japan) the estimated values of d are between 0 and 1, though the I(1) hypothesis cannot be rejected. For the remaining four series (India, Canada, Sweden and Uganda) the values are close to 0, and the I(0) hypothesis cannot be rejected.

As a conclusion we can summarize the results presented so far by saying that India and Canada present the lowest degrees of persistence and strong evidence of mean reversion. That means that shocks disappear in these two series by themselves in the long run and therefore, there is no need for strong policy measures to recover their original trends. Some evidence of mean reversion (in particular if the disturbances are autocorrelated) is also obtained in the currencies corresponding to Sweden and Uganda. For the remaining cases the series are clearly non-mean-reverting.

Given the disparity observed in the results in some of the series (e.g. Sweden and Uganda) depending on the specification for the error term, in what follows we employ a

[4] See Gil-Alana (2004) for the suitability of the Bloomfield (1973) model in the context of Robinson (1994).

semiparametric method, where no functional form is required for the I(0) term u_t. In particular, we use a "local" Whittle estimator, initially developed by Robinson (1995) and further developed later by Shimotsu and Phillips (2005) and Abadir et al. (2007). It uses a band of frequencies that degenerates to zero. The estimator is implicitly defined by:

$$\hat{d} = \arg\min_d \left(\log \overline{C(d)} - 2\, d\, \frac{1}{m} \sum_{s=1}^{m} \log \lambda_s \right), \tag{6}$$

$$\overline{C(d)} = \frac{1}{m} \sum_{s=1}^{m} I(\lambda_s)\, \lambda_s^{2d}, \qquad \lambda_s = \frac{2\,\pi\, s}{T}, \qquad \frac{1}{m} + \frac{m}{T} \to 0,$$

where m is a bandwidth parameter, $I(\lambda_s)$ is the periodogram of the raw time series given by:

$$I(\lambda_s) = \frac{1}{2\pi T} \left| \sum_{t=1}^{T} x_t\, e^{i\,\lambda_s t} \right|^2,$$

and d ∈ (-0.5, 0.5). Under finiteness of the fourth moment and other mild conditions, Robinson (1995) proved that:

$$\sqrt{m}\,(\hat{d} - d_o) \to_{dtb} N(0, 1/4) \quad as\ T \to \infty,$$

where " $\to_{dtb}$ " stands for convergence in distribution, and d_o is the true value of d. This estimator is robust to a certain degree of conditional heteroskedasticity (Robinson and Henry, 1999) and is more efficient than other semi-parametric competitors. Abadir et al. (2007) extended this approach by using an extended Fourier transform in the computation of the periodogram, implying then that no prior differentiation is required when estimating the parameter d in nonstationary contexts. Using this approach the results are presented in Table 7 for a selected group of bandwidth numbers.[5]

[Insert Table 7 about here]

We see that the results are completely in line with those reported earlier in the paper with the parametric case. Thus, evidence of mean reversion (i.e., $d < 1$) is clearly obtained in the cases of the Kenyan exchange rate against the Canadian dollar and the Indian rupee, and this happens for all examined bandwidth numbers. For the remaining cases, though we observe some estimates which are smaller than 1, the unit root cannot be rejected, or if it is rejected, it is in favour of higher orders of integration clearly implying lack of mean reversion and thus permanency of the shocks.

5. Concluding comments

According to the aforementioned results based on data from 1996 to 2009, in the event of an exogenous shock in the Kenyan exchange rates, the effect will be permanent in the cases of the rates directed against the US dollar, the Japanese yen, the South African rand, the Swedish kroner, the Tanzanian shilling, the Danish kroner, the UK sterling pound, the Swish franc and the Ugandan shilling. On the contrary, for the Indian rupee and the Canadian dollar, the effects of the shocks will be transitory, disappearing in the long run. This implies that stronger policy measures are required in the former cases than in the latter ones.

These findings are very informative for the Kenyan policy response in the light of the current exchange rate crisis faced by Kenya in 2011. The largest depreciation of the Kenya shilling during most of 2011 was against the US dollar which according to our

[5] The choice of the bandwidth clearly deals with the classical trade-off bias/asymptotic variance: the asymptotic variance is decreasing with m while the bias is growing with m. Some authors use an interval of values of m (Lobato and Savin, 1998).

results shows that in the event of an exogenous shock such as that experienced earlier in the year, there would be a permanent effect. This implies that more drastic actions are required by the Kenyan monetary authorities in relation to the US dollar Kenya shilling nominal exchange rate. Such actions should include a coordinated response of fiscal and monetary policy for greater effectiveness in addressing response to the exchange rate shock. So far the response has focused mainly on monetary policy tools using changes in the Central Bank Rate and the cash reserve ratio. It is vital that changes also occur in fiscal policy especially in terms of substantial reductions in government expenditure and increased taxation on non-essential imported goods. Kenya also has substantial trade with many of the countries where the effects of shocks on nominal exchange rates have been found to be permanent, for example, Uganda, Tanzania, the United States, Japan and South Africa. Strong policy measures should therefore be taken where substantial adverse movements occur in nominal exchange rates with respect to these currencies arising from external shocks. On the other hand, Kenya does not have very substantial trade with India and Canada where the effects have been shown to be transitory.

References

Abadir, K.M., W. Distaso, and L. Giraitis (2007) Nonstationarity-extended local Whittle estimation, Journal of Econometrics 141, 1353-1384.

Adelman, I. (1965). Long cycles: Fact or artifacts. American Economic Review, 55, 444-463.

Anthony, M., MacDonald, R., 1999, The width of the band and exchange rate mean-reversion: some further ERM-based results, Journal of International Money and Finance 18, 411-428.

Anouro E, Liew V.K and Elike U., 2006, "Non-Linear Exchange Rate Behaviour: Are African Currencies Exceptional?" International Research Journal of Finance and Economics, Issue 1, 2006.

Arize A.C. (2011) "Mean Reversion in Real Exchange Rates in Seventy LDCs: An Empirical Investigation, International Journal- of Economics and Finance, Vol. 3, No. 5, October 2011.

Arize A.C, T. Osang and D. J. Slottje (2000) "Exchange Rate Volatility and Foreign Trade: Evidence from Thirteen LDCs," Journal of Business and Economic Statistics, Vol. 18, No. 1, Jan, pp. 10-17.

Baillie, R.T. (1996). Long memory processes and fractional integration in econometrics. Journal of Econometrics, 73, 5-59.

Baum, C.F., Barkoulas, J., Caglayan, M., 1999. Persistence in the international inflation rates. Southern Economic Journal 65, 900-913.

Bleaney, M.F, Leybourne, S.J., Mizen, P. 1999. Mean Reversion of Real Exchange Rates in High-inflation countries. Southern Economic Journal 65, 839.

Bloomfield, P. (1973). An exponential model in the spectrum of a scalar time series. Biometrika, 60, 217-226.

Booth, G.G., Kaen, F.R.,Koveos, P.E., 1982. R/S analysis of foreign exchange markets under two international monetary regimes. Journal of Monetary Economics 10, 407-415.

Campa, J.M, Wolf, H. C. 1997. Is Real Exchange Rate Mean reversion Caused by Arbitrage?, NBER Working Paper No. 6162.

Cheung, Y.W., 1993. Long memory in foreign exchange rates. Journal of Business and Economic Statistics 11, 93-101.

Crato, N. and Ray, B.K. 2000. Memory in returns and volatilities of future's contracts. Journal of Futures Markets 20, 525-543.

Dahlhaus, R. (1989). Efficient parameter estimation for self-similar process. Annals of Statistics, 17, 1749-1766.

Diebold, F.X. & Rudebusch, G.D. (1989). Long memory and persistence in aggregate output. Journal of Monetary Economics, 24, 189-209.

Fang, H., Lai, K.S., Lai, M. 1994. Fractal structure in currency futures price dynamics. Journal of Futures Markets 14, 169-181.

Gil-Alana, L.A., 2000, Mean reversion in the real exchange rates, Economics Letters 69, 285-288.

Gil-Alana, L.A., 2004. The use of the Bloomfield model as an approximation to ARMA process in the context of fractional integration, Mathematical and Computer Modelling 39, 29-426.

Gil-Alana, L.A. and B. Henry, 2003, Fractional integration and the dynamics of UK unemployment, Oxford Bulletin of Economics and Statistics 65, 221-240.

Gil-Alana, L.A. and P.M. Robinson (1997). Testing of unit roots and other nonstationary hypotheses in macroeconomic time series. Journal of Econometrics, 80, 241-268.

Granger, C.W.J. (1966). The typical spectral shape of an economic variable. Econometrica, 37, 150-161.

Granger, C.W.J. (1980). Long memory relationships and the aggregation of dynamic models. Journal of Econometrics, 14, 227-238.

Hoffmann M (2007), "Fixed and Flexible Exchange Rates: Evidence from Developing Countries, Economica 74, 295, 425-449.

Kildegaard A. (2006) "Fundamentals of Real Exchange Rate Determination: What Role in the Peso Crisis," Estudos Economicos 21, 1 (41), 3-22.

Kiptui M. and L. Kipyegon (2008) External Shocks and Real Exchange Rate Movements in Kenya, Paper Presented at the 13th Annual Conference on Econometric Modelling, University of Pretoria, 9-11 July 2008.

Lobato, I.N. and N.E. Savin (1998) Real and spurious long memory properties of stock market data, Journal of Business and Economic Statistics 16, 3, 261-268.

Nelson, C.R. and C.I. Plosser, 1982, Trends and random walks in macroeconomic time series. Some evidence and implications, Journal of Monetary Economics 10, 2, 139-162.

Ndung'u N (2000) "The Exchange Rate and Interest Rate Differential in Kenya: A Monetary and Fiscal Policy Dilemma" KIPPRA Discussion Paper No. 1, January 2000.

O'Connell et. al (2011), "Capital Mobility, Monetary Policy, and Exchange Rate Management in Kenya" in C.S. Adam, P. Collier and N. Ndungu (eds) *Kenya: Policies for Prosperity* (Oxford: Oxford University Press).

Robinson, P.M. (1978). Statistical inference for a random coefficient autoregressive model. Scandinavian Journal of Statistics, 5, 163-168.

Robinson, P.M. (1994). Efficient tests of nonstationary hypotheses. Journal of the American Statistical Association, 89, 1420-1437.

Robinson, P.M., 1995. Gaussian semi-parametric estimation of long range dependence, Annals of Statistics 23, 1630-1661.

Robinson, P.M. and M. Henry (1999) Long and short memory conditional heteroskedasticity in estimating the memory in levels. Econometric Theory 15, 299-336.

Shimotsu, K. and P.C.B. Phillips, 2005. Exact local Whittle estimation of fractional integration. Annals of Statistics 33, 1890-1933.

Sollis, R., Leybourne, S. J., Newbold, P., 2002. Tests for Symmetric and Asymmetric Nonlinear Mean Reversion in Real Exchange Rates, Journal of Money, Credit, and Banking, 34, pp. 686-700.

Sowell, F. (1992). Modelling long run behaviour with the fractional ARIMA model. Journal of Monetary Economics, 29, 277-302.

Were M, Geda A, Karingi S.N. and Ndung'u N.S., 2001 "Kenya's Exchange Rate Movement in a Liberalised Environment: An Empirical Analysis" KIPPRA Discussion Paper No. 10, August 2001.

Wang, C., 2004. Futures trading activity and predictable foreign exchange movements. Journal of Banking and Finance 28, 1023-1041.

Table 1: Estimates of d and 95% confidence bands using white noise disturbances

	No regressors	An intercept	A linear time trend
CANADA	0.907 (0.682, 1.118)	0.662 (0.572, 0.842)	**0.536** **(0.317, 0.883)**
DENMARK	0.888 (0,695, 1.153)	0.867 (0.428, 1.083)	**0.849** **(0.665, 1.086)**
INDIA	0.902 (0.722, 1.167)	**0.652** **(0.428, 0.956)**	0.655 (0.433, 0.956)
JAPAN	0.930 (0.766, 1.156)	**1.013** **(0.796, 1.308)**	1.012 (0.797, 1.309)
SOUTH AFRICA	0.917 (0.725, 1.200)	**0.981** **(0.770, 1.260)**	0.982 (0.787, 1.258)
SWEDEN	0.894 (0.661, 1.235)	**0.945** **(0.734, 1.290)**	0.943 (0.704, 1.290)
SWITZERLAND	0.879 (0.688, 1.134)	0.823 (0.682, 1.030)	**0.799** **(0.611, 1.031)**
TANZANIA	0.863 (0.619, 1.191)	0.880 (0.767. 1-076)	**0.861** **(0.704, 1.078)**
UGANDA	0.982 (0.770, 1.287)	0.735 (0.519, 1.255)	**0.768** **(0.410, 1.249)**
U.K.	0.960 (0.743, 1.281)	**0.830** **(0.668, 1.108)**	0.846 (0.685, 1.107)
U.S.A.	0.939 (0.786, 1.155)	**1.028** **(0.840, 1.314)**	1.026 (0.852, 1.311)

In parenthesis the 95% confidence band for the values of d.
In bold the significant model with respect to the deterministic terms.

Table 2: Estimates of coefficients of the selected models

	d (Conf. Band)	Intercept (t-value)	Time trend (t-value)
U.S.A.	1.028 (0.840, 1.314)	58.399 (20.188)	-----
JAPAN	1.013 (0.796, 1.308)	55.220 (12.316)	-----
SOUTH AFRICA	0.981 (0.770, 1.260)	14.846 (17.196)	-----
SWEDEN	0.945 (0.734, 1.290)	8.656 (17.185)	-----
TANZANIA	0.861 (0.704, 1.078)	10.311 (15.184)	0.133 (2.271)
DENMARK	0.849 (0.665, 1.086)	9.969 (17.718)	0.084 (1.843)
U.K.	0.830 (0.668, 1.108)	90.772 (16.056)	-----
SWITZERLAND	0.799 (0.611, 1.031)	46.909 (15.509)	0.431 (2.059)
UGANDA	0.768 (0.410, 1.249)	18.803 (17.927)	0.154 (2.272)
INDIA	**0.652** **(0.428, 0.956)**	**1.658** **(28.562)**	**-----**
CANADA	**0.536** **(0.317, 0.883)**	**40.705** **(19.457)**	**0.495** **(6.457)**

In bold the cases of significant evidence of mean reversion.

Table 3: Summary based on Tables 1 and 2

Series	I(0) stationarity	I(d, d < 1) Mean Reversion	I(1) Unit Roots
CANADA		X	
DENMARK			X
INDIA		X	
JAPAN			X
SOUTH AFRICA			X
SWEDEN			X
SWITZERLAND			X
TANZANIA			X
UGANDA			X
U.K.			X
U.S.A.			X

X indicates non-rejection at the 5% level

Table 4: Estimates of d and 95% confidence bands based on Bloomfield disturbances

	No regressors	An intercept	A linear time trend
CANADA	0.440 (0.153, 1.058)	0.629 (0.495, 0.819)	**0.111** **(-1.135, 0.750)**
DENMARK	0.692 (0.213, 1.186)	0.789 (0.522, 1.335)	**0.701** **(0.049, 1.328)**
INDIA	0.689 (-0.005, 1.139)	**0.289** **(-0.091, 0.957)**	0.299 (-0.113, 0.957)
JAPAN	0.859 (0.024, 1.359)	0.443 (0.081, 1.224)	**0.488** **(0.025, 1.216)**
SOUTH AFRICA	0.671 (0.136, 1.180)	0.561 (0.011, 1.222)	**0.671** **(0.157, 1.220)**
SWEDEN	0.148 (0.074, 0.920)	0.550 (0.332, 0.931)	**0.070** **(-0.438, 0.910)**
SWITZERLAND	0.761 (0.194, 1.260)	0.771 (0.502, 1.402)	**0.707** **(0.061, 1.410)**
TANZANIA	0.431 (0.278, 1.024)	0.933 (0.734, 1.298)	**0.869** **(0.181, 1.318)**
UGANDA	0.598 (0.055, 1.241)	0.430 (0.249, 0.644)	**-0.110** **(-0.603, 0.570)**
U.K.	0.442 (0.114, 1.074)	0.731 (0.541, 1.135)	**0.716** **(0.423, 1.108)**
U.S.A.	0.939 (0.569, 1.360)	0.748 (0.370, 1.211)	**0.799** **(0.539, 1.179)**

In parenthesis the 95% confidence band for the values of d.
In bold the significant model with respect to the deterministic terms.

Table 5: Estimates of coefficients of the selected models in Table 4

	d (Confidence Band)	Intercept (t-value)	TIme trend (t-value)
TANZANIA	0.869 (0.181, 1.318)	10.315 (15.165)	0.132 (2.202)
U.S.A.	0.799 (0.539, 1.179)	57.989 (20.679)	0.348 (1.796)
U.K.	0.716 (0.423, 1.108)	89.528 (16.283)	0.744 (2.484)
SWITZERLAND	0.707 (0.061, 1.410)	45.895 (15.790)	0.427 (2.762)
DENMARK	0.701 (0.049, 1.328)	9.687 (18.257)	0.083 (3.012)
SOUTH AFRICA	0.671 (0.157, 1.220)	14.127 (17.982)	-0.093 (-2.456)
JAPAN	0.488 (0.025, 1.216)	51.834 (15.529)	0.411 (3.545)
INDIA	**0.289 (-0.091, 0.957)**	1.626 (65.497)	-----
CANADA	**0.111 (-1.135, 0.750)**	39.531 (43.643)	0.522 (18.777)
SWEDEN	**0.070 (-0.438, 0.910)**	7.553 (46.894)	0.054 (10.799)
UGANDA	**-0.110 (-0.603, 0.570)**	19.002 (90.742)	0.147 (20.654)

In bold the cases of significant evidence of mean reversion.

Table 6: Summary based on Tables 4 and 5

Series	I(0) stationarity	I(d, d < 1) Mean Reversion	I(1) Unit Roots
CANADA	X	X	
DENMARK			X
INDIA	X	X	
JAPAN			X
SOUTH AFRICA			X
SWEDEN	X	X	
SWITZERLAND			X
TANZANIA			X
UGANDA	X	X	
U.K.			X
U.S.A.			X

X indicates non-rejection at the 5% level

Table 7: Estimates of d based on the "local" Whittle semiparametric method

Series / m	5	7	8	10	15	20
CANADA	**0.416***	**0.465***	**0.501***	**0.503***	**0.477***	**0.620***
DENMARK	0.636	0.797	0.834	1.027	0.938	0.876
INDIA	**0.477***	**0.480***	**0.503***	**0.544***	**0.756***	**0.730***
JAPAN	0.646	0.742	0.887	1.156	1.019	1.006
SOUTH AFRICA	1.077	1.012	0.798	1.022	0.985	0.997
SWEDEN	0.916	0.998	0.785	0.750	0.868	0.913
SWITZERLAND	0.708	0.698	0.717	0.863	0.917	0.908
TANZANIA	1.424	1.240	1.061	1.077	0.903	0.941
UGANDA	0.918	1.124	1.058	0.952	0.948	0.973
U.K.	0.838	0.890	0.830	0.868	0.901	0.890
U.S.A.	1.409	0.905	0.870	1.008	0.984	1.097
95% lower C.band	0.632	0.689	0.709	0.739	0.787	0.816
95% upper C. band	1.367	1.310	1.290	1.260	1.212	1.184

* and in bold: Evidence of mean reversion at the 5% level.