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Nonlinearities in real exchange rate determination: Do African exchange rates follow a random walk?

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Abstract

In this paper we aim at modelling the long run behaviour of the Real Effective Exchange Rates (REER) for a pool of African countries. Not much attention has been paid to this group of countries, in particular, to the existence of nonlinearities in the long run path of such a variable. Controlling for two sources of nonlinearites, i.e. asymmetric adjustment to equilibrium and nonlinear deterministic trends allows us to gain some insight about the behaviour of the African REER. We find that these sources of nonlinearites help us to explain the apparent unit root behaviour found applying linear unit root tests for most of the countries.

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1 Introduction

Real exchange rate modelling has become a very popular topic within international economics during the last decades. Its understanding has several implications not only from then theoretical, but also from the applied point of view. From the theoretical point of view, many macroeconomic models assume a constant equilibrium real exchange rate. Moreover, the real exchange rate can be used as a means to asses the overvaluation or undervaluation of currencies and, therefore, it is a tool for exchange rate policy making. In addition, real exchange rate misalignment can be understood as a measure of economic integration (in the real markets) among countries (Wei and Parsley, 1995). Finally, as Faria and León-Ledesma (2003, 2005) point out, the real exchange rate has an impact on long run relative growth rates and unemployment. This adds another feature to exchange rate modelling: it can be used as a means of promoting growth, in particular in developing countries.

During the last decades, several authors have tried to contribute to the literature on the issue of whether monetary models of exchange determination are able to explain observed exchange rates movements. However, the results of the empirical applications have not been that successful (see Taylor, 2006, for a recent literature review).

Recent contributions to this literature claim that the failure of the former empirical literature lies on the fact of not taking into account the possibility of nonlinearities in the long run path of the real exchange rates. There are several reasons for assuming a nonlinear behaviour in this variable. First, as Dumas (1992), Taylor and Peel (2000), Taylor et al. (2001), Kilian and Taylor (2003) point out, the existence of trade barriers, transport costs and exchange rate intervention, may prevent the economic agents from getting a profit from arbitrage. This implies that the real exchange rate might behave as a unit root process within a threshold, but as a stationary variable outside of the threshold, yielding to an asymmetric speed of adjustment towards the equilibrium, i.e. the exchange rate tends to revert faster to the equilibrium the more misaligned it becomes

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after a shock.

Furthermore, Taylor (2004) and Reitz and Taylor (2008) evidence non-linearities in real and nominal exchange rate movements through the effect of official exchange rate intervention operations or announcements -'oral intervention'-, since the intervention is more likely to occur and be effective the greater the degree of misalignment is. The coordinating role of the authorities' intervention or the 'coordination channel' is then displayed, i.e. given their coordinating influence on informed traders, they raise the confidence of the latter and lead to the stabilisation of the exchange rates (see also Taylor, 1995, and Sarno and Taylor, 2001).

In addition, nonlinearities may affect the real exchange rate in the form of structural changes. Devaluations/revaluations of the nominal exchange rates, as well as other exogenous events that can affect the exchange rates with permanent effects, might generate a broken time trend.

These sources of nonlinearities might affect the power of the unit root tests (see Kapetanios et al., 2003, and Bierens, 1997 among others), and, therefore, traditional (linear) unit root tests may be biased towards the non-rejection of the unit root hypothesis. In economic terms this implies that these unit root tests may incorrectly conclude that there is no evidence of stationary behaviour in the real exchange rate, rejecting, thus, the Purchasing Power Parity (PPP) hypothesis.

In this paper we aim at contributing on the analysis of the empirical fulfilment of PPP in a group of African countries, in order to gain some insight about the long run behaviour of their real exchange rates. As Kargbo (2006) pin points, African countries have performed a number of economic reforms (see section 3), mainly focussed on exchange rates adjustment in order to improve their competitiveness and economic growth, based on the assumption that PPP holds. Hence, the empirical analysis of PPP becomes the corner stone for the success of these policies and neglecting these reforms and its effects

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on the real exchange rate, may yield to wrong conclusions.

In order to test for the order of integration of the African real exchange rates, we apply three groups of unit root tests. The first are the Ng and Perron (2001) (linear) unit root tests that are modified version of existing unit root tests in order to obtain tests with better properties in terms of size and power. Second, we account for the possibility of asymmetric speed of mean reversion by means of the Kapetanios et al. (2003) unit root test. For those stationary real exchange rates, we also estimate an exponential smooth transition autoregressive (ESTAR) model so as to understand the equilibrium values of the variable. Finally, for those countries where we cannot reject the null hypothesis of unit root, we apply the Bierens (1997) unit root tests, that takes the possibility of nonlinear trend stationary under the alternative.

The remainder of this paper is organized as follows. Section 2 summarizes the PPP theory and the recent contributions on the PPP analysis for African countries, applying nonlinear techniques. Section 3 describes the data. In section 4, we describe the methodology applied in this empirical research and the results. Finally, the last section summarizes the main contributions.

2 The PPP hypothesis

The PPP hypothesis, in its strict form (or so-called Absolute PPP) contends that the nominal exchange rate[1] (e_t) between two countries should be equal to the price ratio between the countries. This means that the real exchange rate, defined as

$$q_t = \frac{e_t p_t}{p_t^*} \tag{1}$$

should be equal to one, where p_t^* and p_t are respectively the foreign and domestic price indices. However, rejection of the absolute version of PPP does not necessarily imply

[1]Units of foreign currency for a unit of domestic currency.

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that prices in common currency are only explained by idiosyncratic factors; prices may still react with proportional instead of identical deterministic components. In this case it is said that is the Relative version of PPP the one that holds, i.e. the real exchange rate is equal to a constant different from one.

It is well known that PPP only holds in the long run, which implies that the real exchange rate has to be a I(0) process, thus, testing for unit roots in real exchange rates became a popular way to test for PPP empirically (see Meese and Rogoff, 1988; Mark, 1990; Ardeni and Lubian, 1991; Huizinga, 1987; and Chowdhury and Sdogati, 1993, among the more relevant contributions), although the results are ambiguous.

However, as mentioned in the previous section, the traditional unit root tests might not be able to distinguish between a unit root and a stationary process when the autoregressive parameter is near unity and data generating process follows a nonlinear path. For these reasons, another strand of the literature apply univariate techniques that takes into account asymmetric adjustment towards the equilibrium (Obstfeld and Taylor, 1997; Michael et al., 1997; Taylor and Peel, 2000; Taylor et al., 2001; and Baum et al., 2001), and structural changes and nonlinear trends (Papell, 2002; Sollis, 2005; Camarero, et al., 2006, 2008; Cushman, 2008; Cuestas, 2008; and Christopoulos and León-Ledesma, 2007, among others), finding more favourable results towards the stationarity of real exchange rates.

Although the evidence is pretty abundant for industrialized countries, the empirical literature for African countries is quite scarce. As Kargbo (2006) summarizes, there is a number of authors that have applied multivariate techniques, within the linear framework, finding fairly mixed results for this group of countries. Nevertheless, more recently some other authors (Anoruo et al., 2006; Bahmani-Oskooee and Gelan, 2006) have started to apply unit root tests controlling for nonlinearities, when testing for the order of integration of the real exchange rates for African countries. Whereas Anoruo et al. (2006) applies the Kapetanios et al. (2003) unit root test to the real exchange rate vs. the US dollar of

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a sample of 13 African countries, finding evidence of nonlinear mean reversion for only 4 of them, Bahmani-Oskooee and Gelan (2006) apply the same unit root test for a pool of 21 Real Effective Exchange Rates (REER), obtaining evidence of asymmetric mean reversion for 8 of them.

Although in both papers the authors consider the case of a linear trend, it is not considered the case of a nonlinear trend, as a proxy of structural changes. Hence, in this paper we aim at complementing Bahmani-Oskooee and Gelan (2006) first using a Modified Information Criterion (Ng and Perron, 2001) to obtain the lag length in the auxiliary regression of the Kapetanios et al. (2003) nonlinear unit root test, estimating a nonlinear ESTAR model for the exchange rates and, finally, allowing for the possibility of nonlinear trend stationary real exchange rates under the alternative hypothesis.

3 The Data

The data used for this empirical application are REER for a pool of African countries, i.e. Burkina Faso, Burundi, Cameroon, Ivory Coast, Egypt, Ethiopia, Gabon, Ghana, Kenya, Madagascar, Mauritius, Morocco, Niger, Nigeria, Rwanda, Senegal, Seychelles, Sierra Leone, South Africa, Tanzania and Togo. The data for REER have been obtained from Bahmani-Oskooee and Gelan (2007). These authors compute the nominal and effective exchange rates for several African countries as a weighted average of indexes of real bilateral exchange rates of all trading partners, i.e.

$$REER_j = \sum_{i=1}^{20} \omega_{ij} \left[\frac{(p_j e_{ij}/p_i)_t}{(p_j e_{ij}/p_i)_{2003}} \times 100 \right] \tag{2}$$

where ω_{ij} is the imports share of partner i with country j[2]; e_{ij} is the nominal bilateral exchange rate defined as number of units of country i's per unit of j's currency; p_i is country i price level; and p_j is country j price level.

[2] For major trading partners see Bahmani-Oskooee and Gelan (2007) Table 1.

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We have displayed the series of REER in figures 1 - 4. From these graphs we can highlight several features; first, for some of the countries there is clear downward pattern, implying a general depreciation of the currencies in real terms. Secondly, it is pretty obvious that these countries' REER have suffered for a number of structural changes. For instance, the sharp fall in the REER of Burkina Faso, Ivory Coast, Niger, Senegal and Togo in 1994 was caused by a devaluation of their currencies by the Central Bank of West African countries against the French Frank. Likewise, the Central Bank of Central African States devaluated the currencies of Cameroon and Gabon at the end of 1993 that were pegged with the French Frank. For the case of Egypt, in 1990 the president Mubarak devaluated the currency against the US dollar. In the case of Ghana, there is a significant appreciation of the national currency during 1982-1984. This was due to an economic and political turmoil that raised inflation and the national currency was overvalued. The currency returned to its normal levels after the devaluation of 1984 when an economic recovery programme was set.

All these interventions and structural changes might have generated a nonlinear behaviour in the long run paths of these countries' REER.

4 Unit root testing and nonlinear modelling

4.1 Unit root testing

In order to test for the order of integration of the REER in this group of countries, in this section we apply two different types of unit root tests.

The first are the Ng and Perron (2001) unit root tests. Following these authors, traditional unit root tests might suffer from two main problems. First, they might suffer from power problems when the autoregressive parameter is close to 1 and, second, when the errors of a Moving Average process are close to -1, it is necessary a high lag length in order to avoid size problems. However, the Akaike Information Criterion (AIC) and

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Bayesian Information Criterion (BIC) tend to select a low order of the lag length. In order to overcome these issues, Ng and Perron (2001) propose a Modified Information Criterion (MIC) that controls for the sample size. Additionally, the authors propose a method to avoid the power problem associated to the aforementioned traditional unit root tests. Combining these two approaches, Ng and Perron (2001) obtain the following unit root tests: MZ_α and MZ_t that are the modified versions of the Phillips (1987) and Phillips and Perron (1988) Z_α and Z_t tests; the MSB that is related to the Bhargava (1986) R_1 test; and, finally, the MP_T test that is a modified version of the Elliot et al. (1996) Point Optimal Test.

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Moroever, the presence of the aforementioned nonlinearities in the REER has implications for the power of the technique and, therefore, traditional (linear) unit root tests tend to accept a false unit root null hypothesis (Kapetanios et al., 2003, among others). Thus, nonlinearities can be present in REER in the form of different behaviour of the variable depending on its values, i.e. the variable behaves as a nonstationary process when it is within a band, but is stationary when it is outside of the band. As stated by Dumas (1994) and Michael et al. (1997), among others, it is sensible to assume that the shift between regimes is smooth rather than sudden, due to time aggregation and individuals' behaviour.

In order to account for this source of nonlinearity we apply the Kapetanios, Shin and Snell (2003) (KSS) unit root tests. KSS propose a unit root test that takes into account the possibility of smooth transitions between regimes. Thus, the null hypothesis of unit root is tested against the alternative of globally stationary nonlinear process, i.e.

$$q_t = \beta q_{t-1} + \phi q_{t-1}(1 - exp\{-\theta q_{t-1}^2\}) + \epsilon_t \tag{3}$$

where $\epsilon_t \sim iid(0, \sigma^2)$. Note that KSS assume that the transition function is an exponential smooth transition autoregressive (ESTAR) one. A ESTAR function is appropriate to model REER movements, since this type of function assume that the shocks have a symmetric effect over the variable, regardless the sign of the shock (Taylor and Peel, 2000). It is a common practice to reparameterize equation (3) as

$$\Delta q_t = \alpha q_{t-1} + \gamma q_{t-1}(1 - exp\{-\theta q_{t-1}^2\}) + \epsilon_t. \tag{4}$$

in order to test for unit roots. KSS impose $\alpha = 0$, which is equivalent to saying that the variable is a I(1) process in the central regime. In order to test the null hypothesis $H_0 : \theta = 0$ against $H_1 : \theta > 0$ outside of the threshold[3], Kapetanios et al. (2003) propose a Taylor approximation of the ESTAR model since, in practice, the coefficient γ cannot be identified under H_0. Thus, under the null, the model becomes

$$\Delta q_t = \delta q_{t-1}^3 + \eta_t \tag{5}$$

where η_t is an error term. Now, it is possible to apply a t-statistic to test whether q_t is a I(1) process, $H_0 : \delta = 0$, or is a I(0) process, $H_1 : \delta < 0$.

The results of these unit root tests are reported in table 6. The first feature is that the results of applying both types of tests are pretty similar, i.e. for Ethiopia, Gabon, Ghana, Sierra Leone and Tanzania it is not possible to reject the null hypothesis of unit root at the 5% significance level. In addition, with the KSS test we find that the REER of Burundi and Egypt is also stationary. Finally, with the Ng-Perron test, it is possible to reject the null hypothesis, but only at the 10% significance level. In summary, we find that the REER of these countries is stationary with drift in 8 up to 21 countries[4].

For these 8 countries we estimate nonlinear models in order to gain some insight about their long run behaviour (see next section).

[3]Note that the process is globally stationary provided that $-2 < \phi < 0$.

[4]Note that our results are similar to those obtained by Bhamani-Oskooee and Gelan (2006), although we have used the MAIC in order to select the lag length of the auxiliary ADF regression.

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4.2 Estimated models

4.2.1 Foundations

Smooth transition (ST) models are members of the family of state-dependent models; these are a local linearization of the general nonlinear specification. The data generating process is a linear one that switches between a certain number of regimes according to some rule; the regime is characterized as a continuous function of a predetermined variable, so that interactions between variables are permitted, as well as intermediate states between the extreme regimes.

This paper focuses on STs over other usual specifications because, among other reasons: their flexibility enables the description of a wide range of nonlinear behaviours; there exists a modelling cycle; once the state is given, the model is locally linear and easy to interpret; and standard nonlinear estimation techniques can be used. See Granger and Teräsvirta (1993), Teräsvirta (1994, 1998) and van Dijk et al. (2002) for a deeper insight on STs.

The smooth transition autoregression (STAR) is the basic univariate version of the ST model; all predetermined variables are lags of the dependent variable and regimes are endogenously determined. Suppose y_t a stationary, ergodic process, the STAR model of order p is given by

$$y_t = \pi_0 + \sum_{i=1}^{p} \pi_i y_{t-i} + F(y_{t-d}) \left[\theta_0 + \sum_{i=1}^{p} \theta_i y_{t-i} \right] + u_t \tag{6}$$

where $F(y_{t-d})$ is a transition function customarily bounded between 0 and 1 that makes the STAR coefficients vary between π_i and $\pi_i + \theta_i$ $(i = 1, ..., p)$, respectively; d is the transition lag; and u_t is an error process, $u_t \sim Niid(0, \sigma^2)$. The transition variable, y_{t-d}, and the associated value of $F(y_{t-d})$ determine the regime at each t. In its basic version, the regime-switching STAR model considers two extreme regimes, corresponding to $F = 0$ and $F = 1$; the transition from one regime to the other is smooth over time,

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involving that parameters in (6) gradually change with the state variable.

The STAR model links two linear components by means of $F(y_{t-d})$, so that the formulation for this function comes out a key issue when investigating nonlinearities. Two specifications are the most commonly used. First, the logistic function, which has the form:

$$F(y_{t-d}) = \frac{1}{1 + exp\left[-\gamma(y_{t-d} - c)\right]} \tag{7}$$

and the resulting model is the Logistic STAR or LSTAR. This function usually represents the odd case in the literature, meaning that $F(-\infty) = 0$ and $F(\infty) = 1$. The slope parameter γ ($\gamma > 0$) determines the smoothness of the transition; the higher it is, the more rapid the change from one extreme regime to the other. The location parameter c indicates the threshold between the two regimes; here, $F(c) = 0.5$, so that regimes are associated with low and high values of y_{t-d} relative to c.

Second, the exponential function

$$F(y_{t-d}) = 1 - exp\left[-\gamma(y_{t-d} - c)^2\right] \tag{8}$$

is the habitual one in the case of an even transition and provides the Exponential STAR or ESTAR model. This specification implies $F(c) = 0$ and $F(\pm\infty) = 1$ for some finite c, defining the inner and the outer regime, respectively.

LSTAR and ESTAR models describe quite different types of behaviour. In the logistic model the extreme regimes are associated with y_{t-d} values far above or below c, where dynamics may be different; the ESTAR model suggests rather similar dynamics in the extreme regimes, related to low and high y_{t-d} absolute values, while it can be different in the transition period.

As it was anticipated in 4.1, the exponential specification comes out the most suitable one for describing the evolution of real exchange rates; the appeal of this function lies

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in allowing a symmetric adjustment of the variable for deviations below and above the equilibrium value. The location parameter would reflect the equilibrium real exchange rate and the dynamics of the variable would change according to the proximity/distance to the equilibrium state; in the last case, there would not be differences between highly overvaluated or highly undervaluated exchange rates.

Going now into the modelling cycle for STAR models, it has traditionally been based on reproducing Box and Jenkins (1970) iterative methodology; search for specification, estimation and evaluation of the model. There is a well-established STAR modelling procedure in the literature (see Granger and Teräsvirta, 1993; Teräsvirta, 1994). Nonetheless, the most recent investigations do not follow this strategy in a strict manner. Öcal and Osborn (2000), van Dijk et al. (2002) and Sensier et al. (2002), among others, argue that it is possible to develop valid nonlinear formulations that improve the fit of the linear ones by means of an extensive search of STAR models; any possible inadequacy of the models will be unveiled at the validation stage.

We define several combinations of p and d, trying for different values of γ and using a value close to the sample mean of the transition variable for c; the lag order p ranges from 1 to 8 (eighth-order dynamics seem to be general enough for quarterly data) and the transition lag d varies from 1 to p. STAR models are estimated by nonlinear least squares; following the recommendations of Teräsvirta (1994), the argument of the exponential transition function is scaled by dividing it by the variance of the dependent variable. Where parameter convergence is reached, models are subject to further refinement. First, nonsignificant coefficients are excluded to conserve degrees of freedom; then, we simplify this first set of estimations through cross-parameter restrictions in order to increase efficiency. We take 1.6 as the limit t-value for these coefficients, since the variables are stationary.

The properties of the finally proposed models are evaluated using several misspecification tests. We consider the test of no autoregressive conditional heteroskedasticity with

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four lags (ARCH) and the test of business cycle heteroskedasticity (BCH) posed by Öcal and Osborn (2000). Particular attention is also paid to significant estimated coefficients, the features of the transition functions and the results of the following diagnostic statistics: the residual standard error (s), the adjusted determination coefficient ($\bar{R}^2$) and the variance ratio of the residuals from the nonlinear model and the best linear specification (s^2/s_L^2).

4.2.2 Empirical results

As a starting point, we determine the linear models that would describe the behaviour of the exchange rates in the eight countries where we find that the REER is stationary around a drift. An ordinary least squares estimation is carried out, where the proper number of lags is selected using the Akaike Information Criterion. These models are not reported for space reasons but they are available from authors upon request.

The next step involves estimation and specification of the STAR models for all countries following the aforementioned extensive search strategy. A large number of STAR models are generated, although parameter convergence is not attained in some of them[5]. The final selected models are reported in table 3, along with several descriptive statistics and diagnostic tests. These specifications reflect how high absolute values for the lagged exchange rate generate nonlinear effects on the variable at present time. Linearity tests against the estimated ESTAR models prove evidence of nonlinearities for all countries; table 4 displays the p-values of the F tests (significance level, 0.05).

The dependence of the exchange rates on their own recent history is moderate for some countries. The transition from one regime to the other is rather fast in the majority of the countries; in particular, Burundi and Sierra Leone models behave like threshold ones, as they display sharp switches. The values of the location parameter are reasonably

[5] For Tanzania and Ghana it is not possible to obtain adequate specifications with endogenous transition that explain their respective behaviours; convergence problems in the estimation, as well as not very satisfactory results of the latter, are the main justifications

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close to the sample means of the exchange rates for four out of the six countries; for the two remaining, Gabon and Nigeria, almost all the observations lie to the right of c, making the transition a logistic one in practice (see figure 5). The latter would mean the pretty absence of a devaluation regime in both countries, which is confirmed by figures 2 and 3. The reason for this behaviour may be found in that Gabon and Nigeria are leading oil exporters; this brings about a surplus in their current account balances, causing appreciation in the currencies.

The evaluation stage does not unveil signs of misspecification in the models, so the proposed ESTARs are adequate. According to the variance ratio, the estimated nonlinear models explain 7% to 24% of the residual variance of the best linear autoregression in all six countries.

The evaluation of the estimated nonlinear specifications is completed with the analysis of their local dynamic properties, which will allow us to better characterize the variables. The necessary information is obtained from the roots of the characteristic polynomials associated to the models; in this paper, we compute the roots for the two extreme values of the transition function, $F = 0$ and $F = 1$. To save space, table 5 only displays the dominant root, i.e., the root with the highest modulus that is determining the long-run behaviour of the series within each regime.

Table 5 reveals, on the one hand, that the estimated ESTAR models are stable in the inner and in the outer regime for three out of the six countries, i.e. Ethiopia, Gabon and Nigeria, especially in Gabon, which might have to do with being the only country with fixed (strict) exchange rates. On the other hand, Burundi, Egypt and Sierra Leone have globally stationary although locally unstable models. This second group of countries shows explosive roots in the inner regime, involving that exchange rates pass this regime rapidly on their way up or down; the outer regime is a stable one, so that once the variable undergoes an overvaluation or an undervaluation, it will tend to remain there unless an exogenous shock occurred.

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The asymmetry between the extreme regimes is particularly severe in Egypt and Sierra Leone. Both countries are characterized by the following pattern of two opposite forces; on the one hand, they have a well-recognized and high-added value main export product (oil and derivatives in Egypt, diamonds in Sierra Leone) but, on the other hand, they are great importers (Egypt is a major food importers on worldwide scale and Sierra Leone highly demands hydrocarbons and food). The dynamics of these two opposite forces are expected to determine the evolution of the exchange rates in both countries, as they are not fixed (rigid) ones.

Lastly, in Burundi the transition between the extreme regimes is nearly abrupt and, moreover, the unstable inner regime contains few observations. As a result, the currency almost always suffers from a greater or lesser overvaluation/undervaluation in this agricultural economy whose foreign trade depends on imports and on a vulnerable exporting balance caused by food products with volatile prices in international markets.

5 Allowing for deterministic trends

In this section we relax the assumption that the REER is stationary around a drift under the alternative hypothesis and we allow for the possibility of nonlinear deterministic trends. It is well known within the literature that misspecification of the deterministic components might bias the results of the unit root tests towards the acceptance of the null hypothesis of unit root (Perron and Phillips, 1987; West, 1988; Bierens, 1997). As it was mentioned earlier, nonlinearites can be present in the form of structural changes and, therefore, these nonlinearites affect the deterministic components. Many authors have applied unit root tests with structural changes in order to test for PPP, finding in general more favourable results towards PPP empirical fulfillment. However, a broken time trend can be understood as a particular case for a nonlinear trend. Thus, even unit root tests with structural changes may suffer from power problems (Bierens, 1997).

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Bierens (1997) proposes some tests which take into account the possibility of stationarity around a nonlinear deterministic trend under the alternative hypothesis. This author generalizes the ADF auxiliary regression to incorporate Chebishev polynomials to approximate the nonlinear deterministic trend. Following Bierens (1997), the reason for using Chebishev polynomials instead of regular time polynomials (Park and Choi, 1988; Ouliaris et al., 1989) to approximate the nonlinear deterministic trend is that the former create less power distortions. Thus, the auxiliary regression becomes

$$\Delta q_t = \alpha q_{t-1} + \sum_{j=1}^{p} \phi_j \Delta q_{t-j} + \theta^T P_{t,n}^{(m)} + \varepsilon_t \tag{9}$$

where $P_{t,n}^{(m)}$ are the Chebishev polynomials and m is the order of the polynomials, such that $P_{0,t}$ through $P_{m,t}$, where $P_{0,t}$ equals 1, $P_{1,t}$ is equivalent to a linear trend, and $P_{2,t}$ through $P_{m,t}$ are cosine functions. In order to test for the order of integration of the variables, Bierens (1997) proposes several tests. In this paper we apply the t-test over the coefficient α, $\hat{t}(m)$. Moreover, this can be tested by means of the $\hat{A}(m) = \frac{n\hat{\alpha}}{|1-\sum_{i=1}^{p}\hat{\phi}_i|}$ test. The last one is an F-test, $\hat{F}(m)$, for the joint hypothesis that $\hat{\alpha}$ and the last m components of the parameter vector θ in model (9) are zero under the null. The conclusions from these tests are different under the alternative hypothesis depending upon the side of the rejection. Whereas with the $\hat{A}(m)$ and t test right side rejection implies stationarity around a nonlinear trend, left side rejection does not allow us to distinguish between mean stationarity, linear trend stationary and nonlinear trend stationarity. With the F test it is not possible to distinguish between those alternatives (see table 2).

A number of authors have applied Bierens' (1997) approach in order to test for the order of integration of the RER (see Sollis, 2005; Assaf, 2006; Cushman, 2008; Camarero et al., 2008; Cuestas, 2008; and Cuestas and Regis, 2008) finding that in general it is possible to reject the null hypothesis once nonlinear deterministic trends are accounted for.

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In table 6 we present the results of the Bierens (1997) tests. Following this author, these tests suffer from important size distortions (Bierens, 1997), therefore the critical values have been obtained by Monte Carlo experiment based on 5,000 replications of a Gaussian $AR(p)$ process for Δq_t. The parameters and error variances are equal to the estimated $AR(p)$ null model, where the order p of the ADF auxiliary regression has been obtained by the AIC and the initial values have been taken from the actual series. Note that we have only performed the Bierens' tests for those countries we could not reject the null hypothesis with the Ng-Perron and KSS tests. The results show that in 8 of the remaining 13 countries, we can reject the null hypothesis of unit root. Additionally, it is worth highlighting that in all cases, except for Senegal, we obtain left side rejection which means that the series could be stationary around a drift, linear or nonlinear trend. For Senegal the results point to nonlinear trend stationarity process. This is not surprising since the order to the Chebishev polynomial necessary to reject the null hypothesis is pretty high[6]. Figures 6 - 7 display the nonlinear trends adjustment for these 8 countries. Since a higher order of m implies a higher degree of nonlinearity, for some countries such as Senegal of Seychelles it is necessary to use a high order for m in order to approximate structural changes.

This implies that for these 8 countries the strict hypothesis of PPP needs to be relaxed in order to account for structural changes, finding thus, a non-constant or time varying equilibrium real exchange rate.

In parallel to this analysis of nonlinearities in real exchange rate adjustment, Lothian and Taylor (2000) and Lothian and Taylor (2008) evidence the so-called Harrod-Balassa-Samuelson effect for certain real exchange rates, that is, the dependence of the equilibrium real exchange rate on productivity differentials; the first empirical work considers nonlinear time trends prox-

[6]As Bierens (1997) claims, there is not a unique way to choose the order of m. Since we know the ADF test tends to over-accept the null hypothesis, we have selected the order of m that yields more evidence against the alternative hypothesis.

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ying for this effect. These findings justify our interest in allowing for real shocks in the African countries and, as a future research, we intend to consider the effect of real shocks to the equilibrium real exchange in this group of countries.

6 Conclusions

Aimed at contributing to the literature on PPP in Africa, in this paper we have applied several unit root tests that account for different forms of nonlinearities, i.e. asymmetric speed of adjustment and nonlinear deterministic trends. In the former, stationarity is found in 8 up to 13 countries and ESTAR models can adequately capture the behaviour of the real exchange rates for some of these countries; in fact, modelling the nonlinearity in the data explains 7% to 24% of the residual variance of the best linear autoregression. In most of the countries, the transition from the equilibrium location to the overvaluation/undervaluation stage takes place quite fast, which strengthens the nonlinearity hypotheses; in addition, the threshold value between both regimes is generally in the neighbourhood of the variable's sample mean. Furthermore, half of the countries evidence explosive roots in the inner regime, involving real exchange rates pass rapidly this stage to reach the outer regime. The intrinsic features of each economy may underlie this behaviour.

Further, the evidence in favour of the alternative hypothesis of stationarity increases when allowing for nonlinear deterministic trends, approximated by means of Chebishev polynomials. This implies that the real exchange rate tends to revert to a time dependent equilibrium value. This feature is important since structural changes might drive us to wrong conclusions when testing for PPP, in particular for developing countries.

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Table 1: Ng-Perron and KSS unit root test results

Country	MZ_a	MZ_t	MSB	MP_T	$\hat{t}_{NL}$	$\hat{t}_{NLD}$
Burkina Faso	0.786	0.773	0.983	64.943	-1.893	-0.907
Burundi	0.871	1.276	1.464	138.026	-4.122**	-3.542**
Cameroon	-6.521	-1.771	0.271	3.878	-0.892	-1.914
Ivory Coast	-6.297	-1.756	0.278	3.950	-0.839	-2.281
Egypt	-6.632	-1.655	0.249	4.259	-2.230*	-3.949**
Ethiopia	-5.797*	-1.591	0.274*	4.570	-2.257**	-4.571**
Gabon	-9.905**	-2.161**	0.218**	2.727**	-1.531	-2.926**
Ghana	-16.190**	-2.843**	0.175**	1.518**	-8.131**	-8.687**
Kenya	-2.051	-0.808	0.394	10.135	-0.550	-2.559
Madagascar	-2.549	-0.897	0.351	8.628	-1.246	-2.384
Mauritius	-1.670	-0.759	0.454	12.279	-1.238	-2.469
Morocco	0.736	0.649	0.881	53.002	-1.889	-1.043
Niger	-0.764	-0.379	0.496	16.409	-1.311	-1.925
Nigeria	-7.337*	-1.892*	0.257*	3.425*	-1.645	- 1.628
Rwanda	-2.510	-1.090	0.434	9.610	-0.980	-1.056
Senegal	-3.511	-1.167	0.332	6.971	-1.013	-1.524
Seychelles	-1.321	-0.729	0.552	16.293	-0.513	-1.916
Sierra Leone	-8.399**	-2.038**	0.242*	2.958**	-2.359**	-3.216**
South Africa	-2.738	-1.050	0.383	8.558	-1.229	-2.316
Tanzania	-7.849*	-1.931*	0.246*	3.3115	-3.670**	-5.665**
Togo	-0.375	-0.210	0.560	20.684	-1.314	-1.600

Note: The order of lag to compute the tests has been chosen using the modified AIC (MAIC) suggested by Ng and Perron (2001). The Ng-Perron tests include an intercept, whereas the KSS test has been applied to the raw data, $\hat{t}_{NL}$ say, and to the demeaned data, $\hat{t}_{NLD}$ say. The symbols * and ** mean rejection of the null hypothesis of unit root at the 10% and 5% respectively. The critical values for the Ng-Perron tests have been taken from Ng and Perron (2001), whereas those for the KSS have been obtained by Monte Carlo simulations with 50,000 replications:

	MZ_a	MZ_t	MSB	MP_T	$\hat{t}_{NL}$	$\hat{t}_{NLD}$
5%	-8.100	-1.980	0.233	3.170	-2.196	-2.925
10%	-5.700	-1.620	0.275	4.450	-1.906	-2.633

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Table 2: Alternative hypotheses

Test	Left-side rejection	Right-side rejection
$\hat{t}(m)$	MS, LTS or NLTS	NLTS
$\hat{A}(m)$	MS, LTS or NLTS	NLTS
$\hat{F}(m)$	-	MS, LTS or NLTS

Note: MS= mean stationarity, LTS= linear trend stationarity, NLTS= nonlinear trend stationarity.

Table 3: Estimated ESTAR models for exchange rates

BURUNDI

$$e_t = \underset{(873.88)}{2369.69} + \underset{(0.26)}{1.80}e_{t-1} - \underset{(0.13)}{0.39}e_{t-2} + \underset{(0.08)}{0.19}e_{t-3} - \underset{(2.13)}{5.84}e_{t-6} + \underset{(0.29)}{0.45}e_{t-7} + \left(\underset{(873.85)}{-2367.55} - \underset{(0.25)}{0.62}e_{t-1} + \underset{(2.13)}{5.84}e_{t-6} - \underset{(0.29)}{0.45}e_{t-7}\right) \times \left[1 - exp\left\{\underset{(68.97)}{-160.11} \times 2.12\left(e_{t-6} - \underset{(2.89)}{489.26}\right)^2\right\}\right] + a_t$$

s=18.97; $\bar{R}^2 = 0.99$; $s^2/s_L^2 = 0.84$; ARCH=0.54 (0.71); BCH=0.02 (0.88)

EGYPT

$$e_t = \underset{(31.08)}{-26.32} + \underset{(0.10)}{1.24}e_{t-1} + \underset{(0.20)}{0.73}e_{t-4} - \underset{(0.18)}{0.87}e_{t-5} + \left(\underset{(30.49)}{46.71} - \underset{(0.15)}{0.58}e_{t-1} + \underset{(0.12)}{0.27}e_{t-2} - \underset{(0.21)}{0.79}e_{t-4} + \underset{(0.18)}{0.87}e_{t-5}\right) \times \left[1 - exp\left\{\underset{(0.95)}{-1.56} \times 0.0003\left(e_{t-3} - \underset{(9.95)}{254.04}\right)^2\right\}\right] + a_t$$

s=17.62; $\bar{R}^2 = 0.91$; $s^2/s_L^2 = 0.76$; ARCH=0.72 (0.58); BCH=0.005 (0.94)

ETHIOPIA

$$e_t = \underset{(74.93)}{34.63} + \underset{(0.25)}{1.56}e_{t-1} - \underset{(0.46)}{1.09}e_{t-2} + \underset{(0.53)}{0.92}e_{t-3} - \underset{(0.48)}{0.53}e_{t-4} + \left(\underset{(77.34)}{-11.93} - \underset{(0.26)}{0.89}e_{t-1} + \underset{(0.46)}{1.09}e_{t-2} - \underset{(0.53)}{0.92}e_{t-3} + \underset{(0.49)}{0.64}e_{t-4}\right) \times \left[1 - exp\left\{\underset{(0.57)}{-1.11} \times 0.0003\left(e_{t-4} - \underset{(15.64)}{220.26}\right)^2\right\}\right] + a_t$$

s=16.73; $\bar{R}^2 = 0.91$; $s^2/s_L^2 = 0.79$; ARCH=0.89 (0.47); BCH=0.97 (0.33)

GABON

$$e_t = \underset{(62.82)}{114.78} - \underset{(0.81)}{0.40}e_{t-1} + \left(\underset{(59.64)}{-79.66} + \underset{(0.71)}{1.67}e_{t-1} - \underset{(0.17)}{0.54}e_{t-2}\right) \times \left[1 - exp\left\{\underset{(0.37)}{-0.42} \times 0.0022\left(e_{t-1} - \underset{(14.96)}{62.12}\right)^2\right\}\right] + a_t$$

s=7.64; $\bar{R}^2 = 0.86$; $s^2/s_L^2 = 0.93$; ARCH=0.06 (0.99); BCH=1.62 (0.20)

NIGERIA

$$e_t = \underset{(22.28)}{7.33} + \underset{(0.23)}{0.93}e_{t-1} + \left(\underset{(21.34)}{6.07} + \underset{(0.25)}{0.96}e_{t-1} - \underset{(0.14)}{0.95}e_{t-2}\right) \times \left[1 - exp\left\{\underset{(4.46)}{-2.67} \times 0.00005\left(e_{t-1} - \underset{(39.65)}{85.29}\right)^2\right\}\right] + a_t$$

s=27.68; $\bar{R}^2 = 0.96$; $s^2/s_L^2 = 0.81$; ARCH=0.01 (0.99); BCH=2.01 (0.16)

SIERRA LEONE

$$e_t = \underset{(76.34)}{-172.52} + \underset{(0.06)}{0.76}e_{t-1} + \underset{(0.40)}{1.06}e_{t-2} - \underset{(0.52)}{0.90}e_{t-3} + \underset{(0.40)}{0.54}e_{t-4} + \underset{(0.34)}{0.83}e_{t-7} + \left(\underset{(76.80)}{200.70} - \underset{(0.40)}{1.06}e_{t-2} + \underset{(0.52)}{0.90}e_{t-3} - \underset{(0.42)}{0.71}e_{t-4} + \underset{(0.12)}{0.33}e_{t-5} + \underset{(0.11)}{0.30}e_{t-6} - \underset{(0.36)}{1.29}e_{t-7}\right) \times \left[1 - exp\left\{\underset{(7.28)}{-19.07} \times 0.0004\left(e_{t-6} - \underset{(1.51)}{134.43}\right)^2\right\}\right] + a_t$$

s=25.43; $\bar{R}^2 = 0.73$; $s^2/s_L^2 = 0.76$; ARCH=2.20 (0.07); BCH=1.38 (0.24)

Notes: e_t denotes the real exchange rate. Values under regression coefficients are standard errors of the estimates; s is the residual standard error; $\bar{R}^2$ the adjusted determination coefficient; s^2/s_L^2 is the variance ratio of the residuals from the nonlinear model and the best linear AR selected with AIC; ARCH is the statistic of no ARCH based on four lags; BCH is a business cycle heteroskedasticity test. Numbers in parentheses after values of ARCH and BCH are p-values.

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Table 4: Linearity tests against estimated ESTAR models

Country (lag order)	p-value
Burundi (p=7)	0.00003
Egypt (p=5)	0.000002
Ethiopia (p=4)	0.00001
Gabon (p=2)	0.0363
Nigeria (p=2)	0.00001
Sierra Leone (p=7)	0.0001

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Table 5: Local dynamics: dominant roots in each regime

Country	Regime (value of F)	Root	Modulus
Burundi	Inner (F=0)	$1.5796 \pm 0.4820i$	1.65
	Outer (F=1)	0.9745	0.97
Egypt	Inner (F=0)	1.2636	1.26
	Outer (F=1)	0.8919	0.89
Ethiopia	Inner (F=0)	$0.8889 \pm 0.2481i$	0.92
	Outer (F=1)	0.8570	0.86
Gabon	Inner (F=0)	-0.3996	0.40
	Outer (F=1)	$0.6352 \pm 0.3674i$	0.73
Nigeria	Inner (F=0)	0.9352	0.93
	Outer (F=1)	$0.9470 \pm 0.2354i$	0.97
Sierra Leone	Inner (F=0)	1.3833	1.38
	Outer (F=1)	$0.4383 \pm 0.8342i$	0.94

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Table 6: Bierens (1997) unit root tests

Country	m	p	t(m)	A(m)	F(m)
Burkina Faso	2	0	0.0166	0.0130	0.9736
Cameroon	-	-	*I(1)*	*I(1)*	*I(1)*
Ivory Coast	-	-	*I(1)*	*I(1)*	*I(1)*
Kenya	4	0	0.0162	0.0138	0.9610
Madagascar	-	-	*I(1)*	*I(1)*	*I(1)*
Mauritius	9	2	0.0376	0.0354	0.9234
Morocco	5	1	0.0010	0.0004	0.9994
Niger	2	0	0.0094	0.0204	0.9904
Rwanda	-	-	*I(1)*	*I(1)*	*I(1)*
Senegal	20	0	0.9614	0.9558	*I(1)*
Seychelles	17	0	0.0328	0.0608	0.9000
South Africa	-	-	*I(1)*	*I(1)*	*I(1)*
Togo	2	0	0.0050	0.0026	0.9926

Note: The values on the table are p-values.

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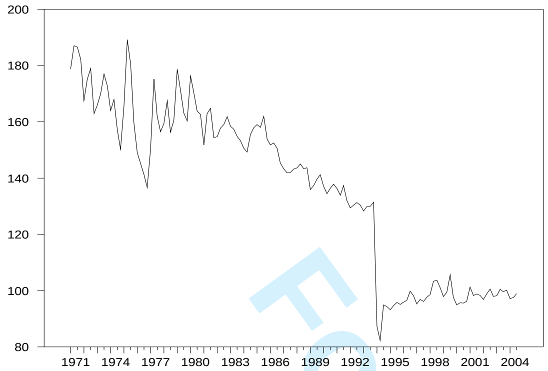
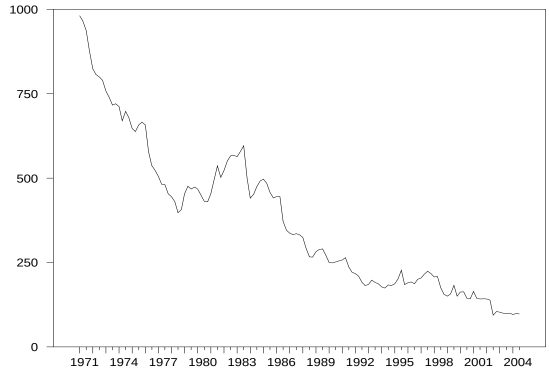
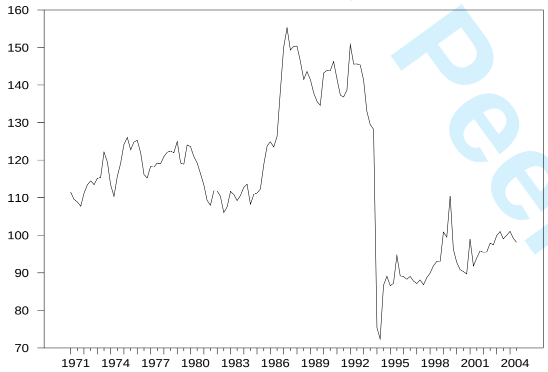
Figure 1: Real exchange rates



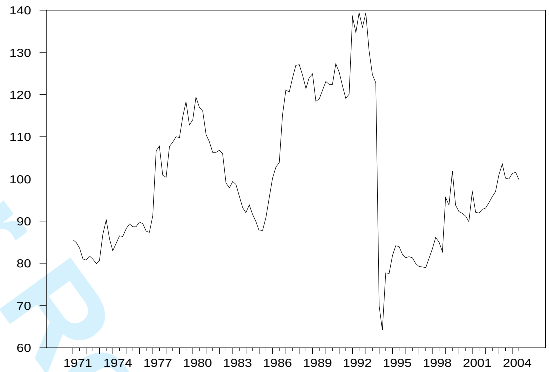
(a) Burkina Faso



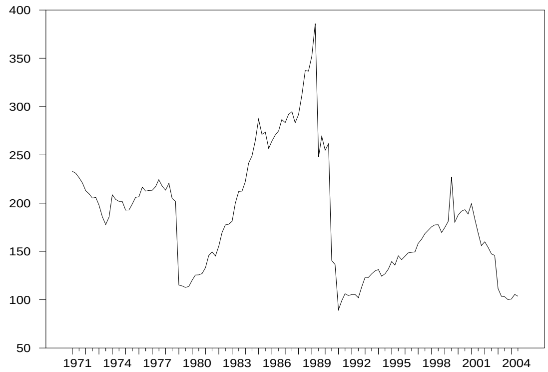
(b) Burundi



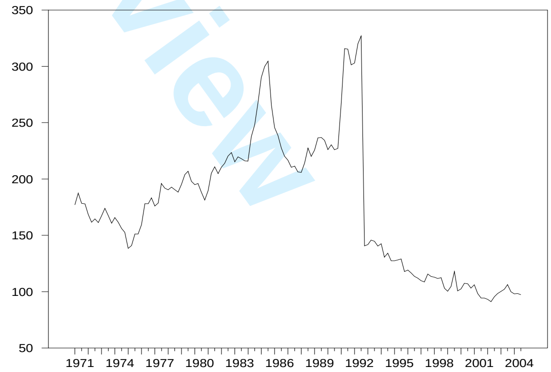
(c) Cameroon



(d) Cote d'Ivoire



(e) Egypt



(f) Ethiopia

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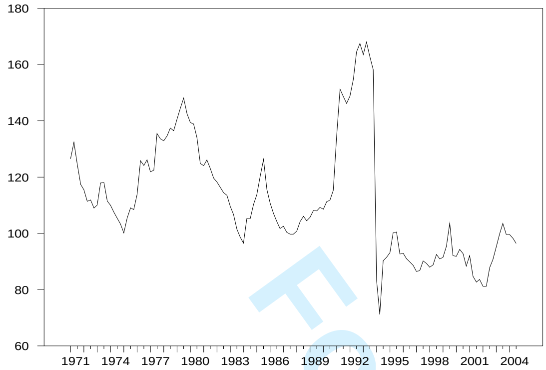
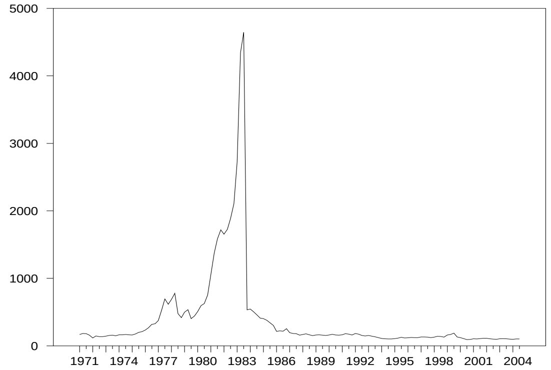
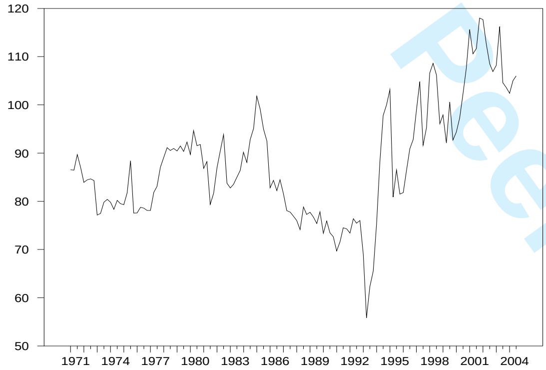
Figure 2: Real exchange rates (cont.)



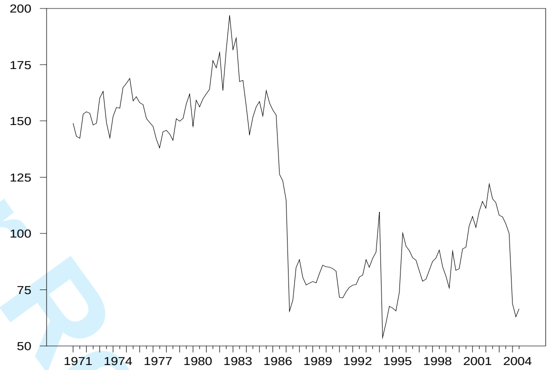
(a) Gabon



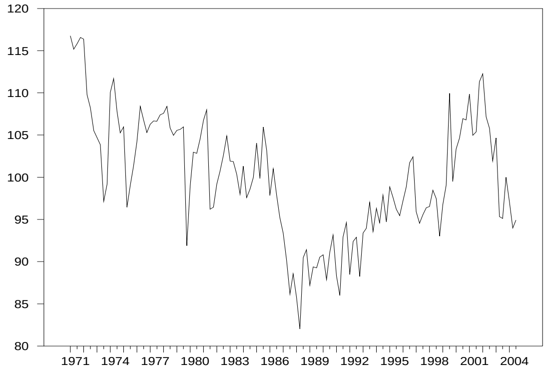
(b) Ghana



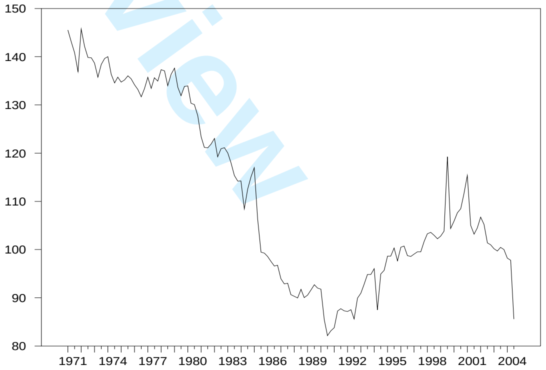
(c) Kenya



(d) Madagascar



(e) Mauritius



(f) Morocco

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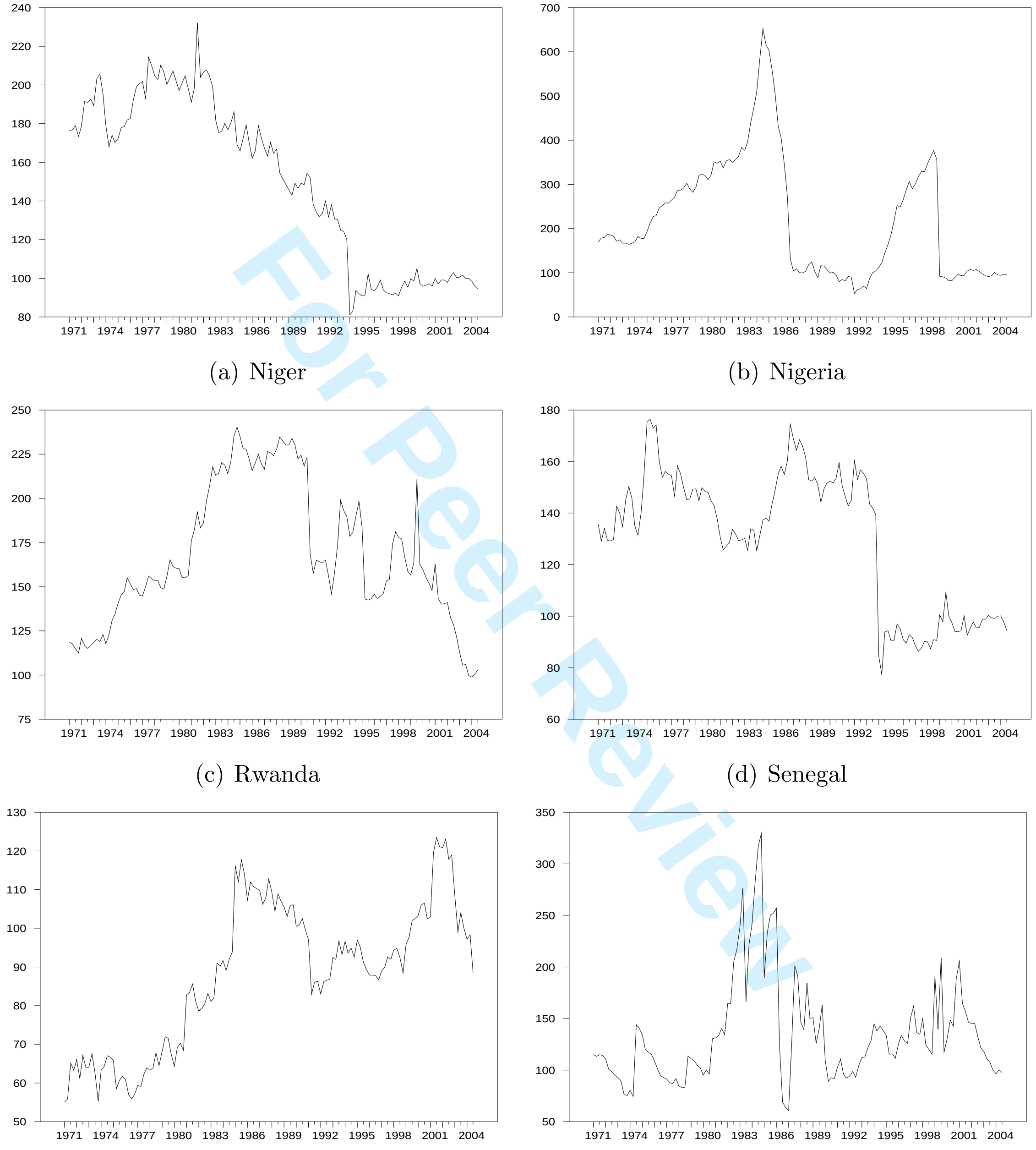
Figure 3: Real exchange rates (cont.)



(a) Niger (b) Nigeria

(c) Rwanda (d) Senegal

(e) Seychelles (f) Sierra Leone

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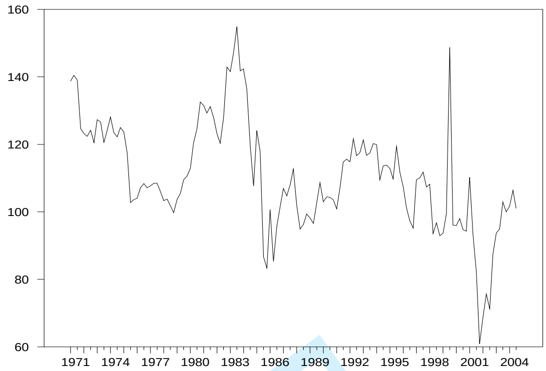
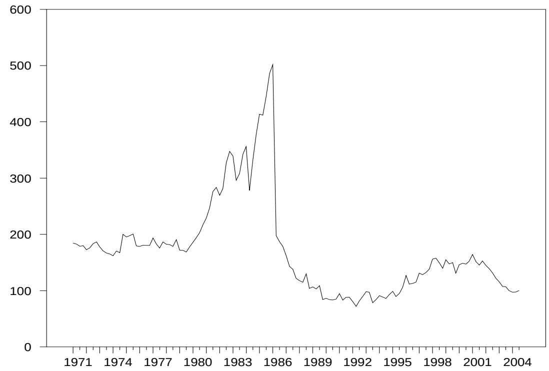
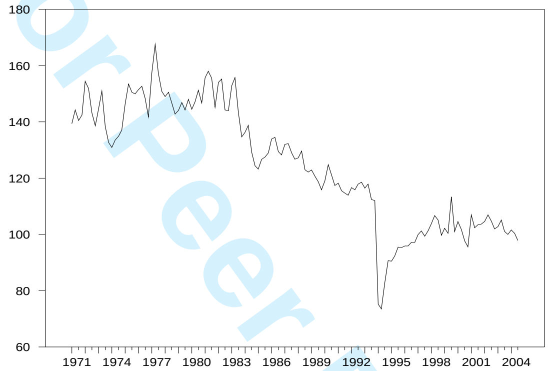
Figure 4: Real exchange rates (cont.)



(a) South Africa



(b) Tanzania



(c) Togo

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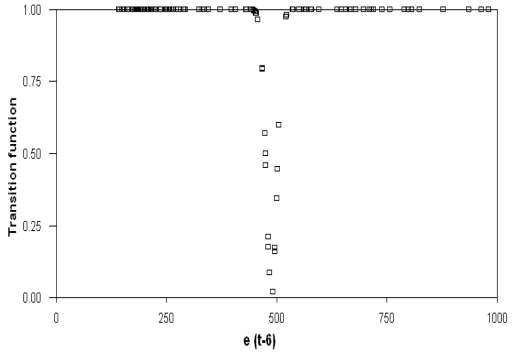
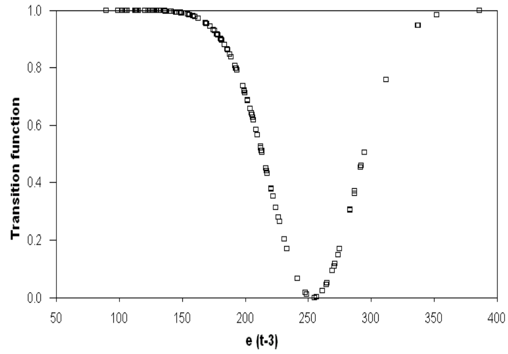
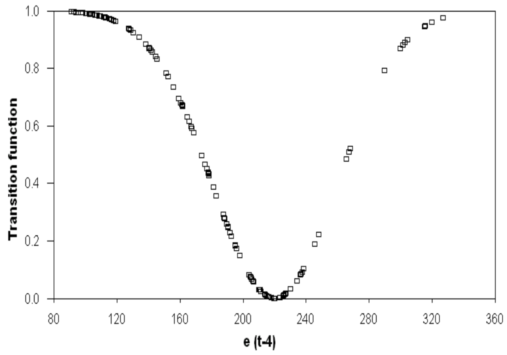
Figure 5: Estimated transition functions



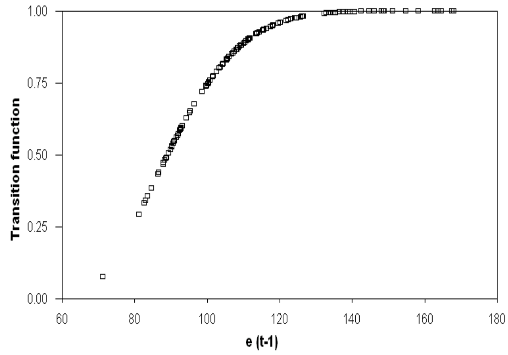
(a) Burundi



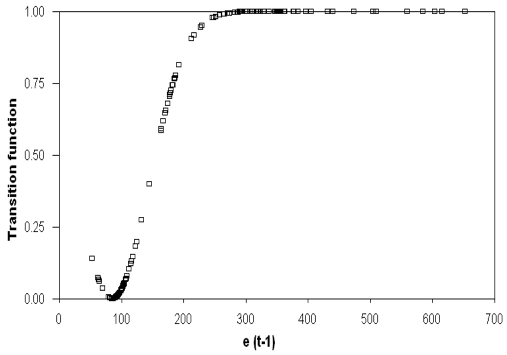
(b) Egypt



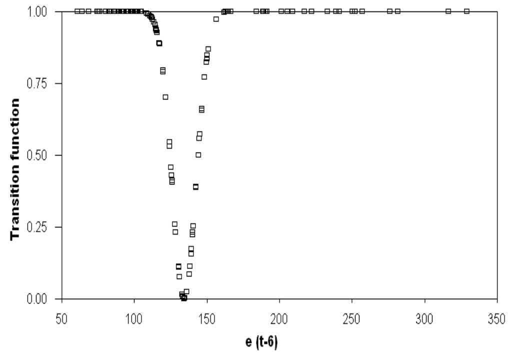
(c) Ethiopia



(d) Gabon

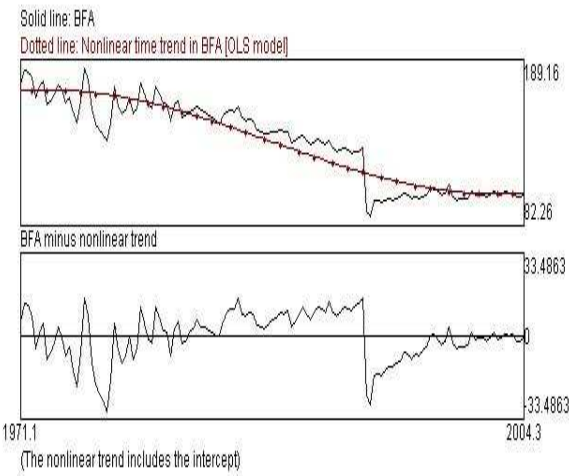


(e) Nigeria

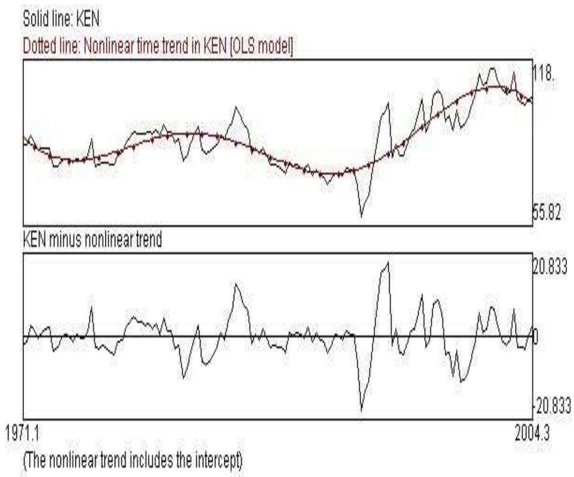


(f) Sierra Leone

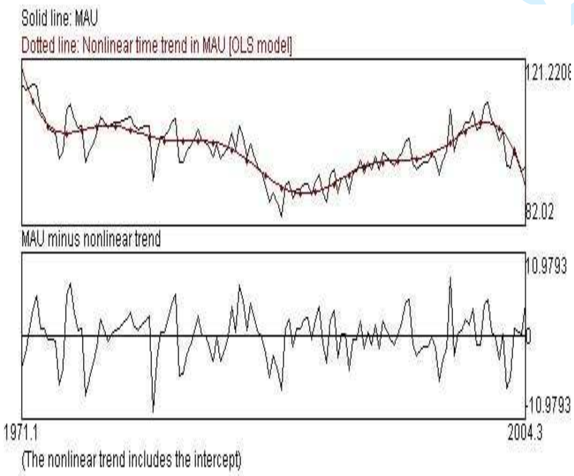
Figure 6: Nonlinear trends



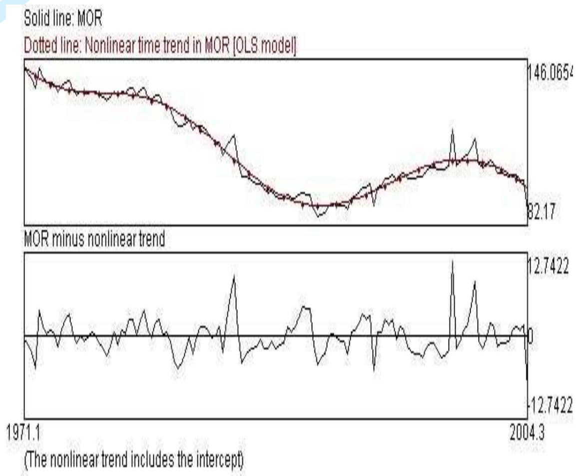
(a) Burkina Faso



(b) Kenya

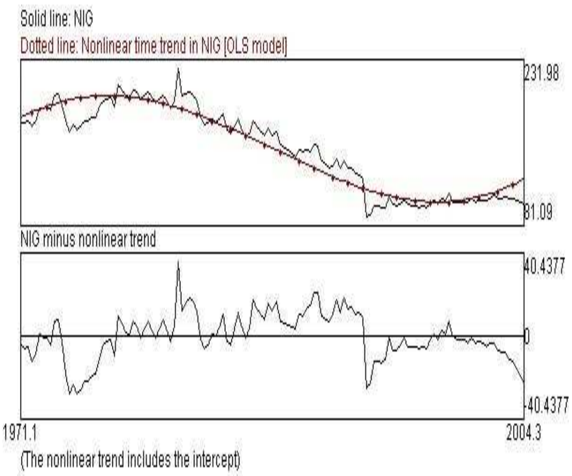


(c) Mauritius

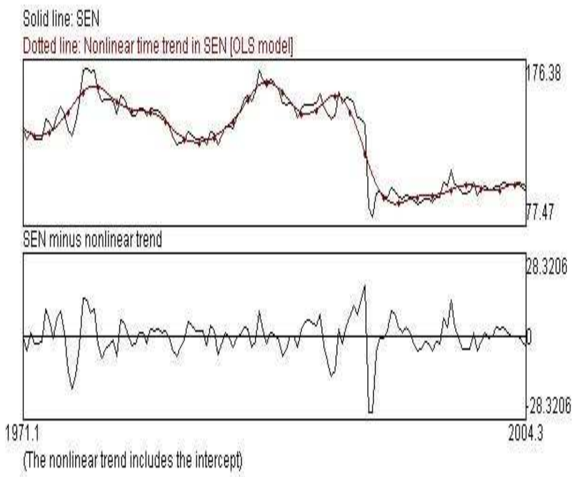


(d) Morocco

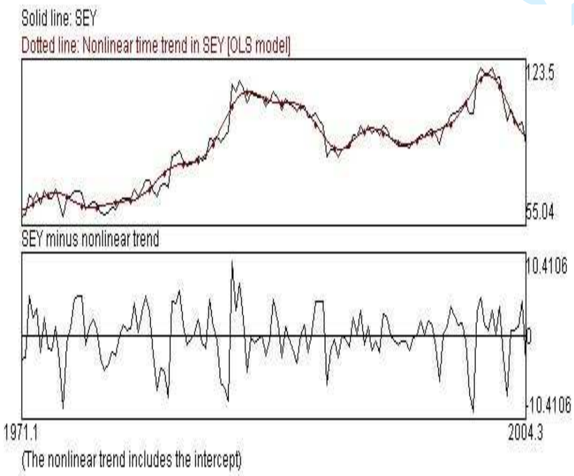
Figure 7: Nonlinear trends (cont.)



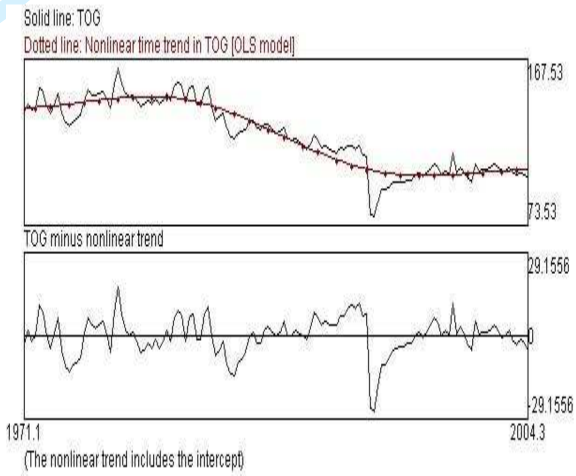
(a) Niger



(b) Senegal



(c) Seychelles



(d) Togo