African Development Review
Revue africaine de développement

African Development Review, Vol. 27, No. 2, 2015, 161–170

Testing PPP for the South African Rand/US Dollar Real Exchange Rate at Different Data Frequencies*

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***Abstract*:** This paper tests the PPP hypothesis for the South African rand/US dollar real exchange rate using a fractional integration framework. The results suggest that the real exchange rate of the South African rand with respect to the US dollar is a highly dependent variable with an order of integration very close to 1. This finding is not affected by the data frequency considered (daily, weekly or monthly). Also, there appears to be a single break in December 2001 (possibly corresponding to a change in the monetary policy framework), with the unit root null being rejected in favour of $d > 1$ for the periods before the break, but not afterwards. Thus, our results strongly reject the PPP hypothesis for the South African rand/US dollar rate across data frequencies, since shocks are found to affect the exchange rate forever.

1. Introduction

Purchasing power parity (PPP) is a central tenet in international economics. It is assumed to hold continuously in flexible-price models of the exchange rate, whilst in sticky-price ones it is a long-run property, temporary deviations from the long-run equilibrium being possible. In the new open economy models it is a condition for market completeness (Chortareas and Kapetanios, 2009). Establishing whether PPP holds is also crucial in order to assess the effects of a devaluation.

Empirical studies have used different methods to examine the validity of PPP. Some of them have tested for cointegration between nominal exchange rates and prices (Kim, 1990; McNown and Wallace, 1989 1994; Serletis and Goras, 2004; Gouveia and Rodrigues, 2004). Others have applied unit root tests to real exchange rates (these are the so-called 'stage-two' tests — see Froot and Rogoff, 1995). However, such tests have been found to be unable to distinguish between random-walk behaviour and very slow mean-reversion to the long-run equilibrium level (see, e.g., Frankel, 1986; Lothian and Taylor, 1997): in small samples they have very low power against alternatives such as trend-stationary models (DeJong *et al*., 1992), structural breaks (Campbell and Perron, 1991), regime-switching (Nelson *et al*., 2001), or fractionally integration (Diebold and Rudebusch, 1991; Hassler and Wolters, 1994; Lee and Schmidt, 1996).

Moreover, at times they exhibit erratic behaviour, suggesting the presence of endemic instability (see Caporale *et al*., 2003 and Caporale and Hanck, 2009 the latter finding that this also characterizes cointegration tests, and therefore is not due to arbitrarily imposed symmetry/proportionality restrictions); however, adjusting the residuals for non-normality and heteroscedasticity using a wild bootstrap method attenuates this type of behaviour considerably (see Caporale and Gregoriou, 2009), and the latter disappears almost completely if panel tests are performed, the evidence for PPP becoming much stronger (see Caporale and Hanck, 2010).

The aforementioned time series studies all restrict themselves to the cases of stationary I(0) and non-stationary I(1) processes. The more recent literature has stressed the importance of considering the possibility of non-integer values for the degree of integration. In this case, PPP is satisfied if the fractional differencing parameter d is strictly smaller than 1, although the higher d is, the longer it takes for the adjustment to the long-run equilibrium to be completed. Alternatively, panel methods have been used to increase the power of tests for PPP (see, e.g., Chortareas and Kapetanios, 2009 and some of the references therein).

*The second-named author gratefully acknowledges financial support from the Ministerio de Ciencia y Tecnología (ECO2008-2011 (03035/28196) ECON Y FINANZAS, Spain). Comments from the Editor and two anonymous referees are gratefully acknowledged.
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In the present paper we test the PPP hypothesis in South Africa by using a fractional integration framework which allows for long memory and also for a much richer dynamic specification than the classical methods. Earlier studies of this type have normally focused on the developed countries and analysed some of the major currencies. For instance, Booth *et al.* (1982) found positive memory ($d > 0$) during the flexible exchange rate period (1973–79) but a negative one ($d < 0$, i.e., anti-persistence) during the fixed exchange rate period (1965–71) for the British pound, French franc and Deutsche mark; Cheung (1993) also found evidence of long memory behaviour during the managed floating regime. On the other hand, Baum *et al.* (1999) estimated ARFIMA models and found no evidence of long-run PPP in the post-Bretton Woods era (see also Fang *et al.*, 1994; Crato and Ray, 2000; and Wang, 2004). Caporale and Gil-Alana (2010) also provide some evidence for the Latin American countries.

There are only a few papers testing the PPP hypothesis for African countries. Nagayasu (2002) examined the long-run PPP hypothesis in 17 African countries using panel cointegration tests. His results support a weak form of this hypothesis. Mkenda (2001) carried out a similar study for a sample of 20 African countries and found that PPP holds when using import-based multilateral indices and bilateral indices, but is rejected when using trade-weighted multilateral indices. More recently, Olayungbo (2011) examined a subset of 16 sub-Saharan countries over a relatively short sample period; he performed standard unit root tests as well as panel unit root tests, and therefore his study has the same limitation as previous ones of only allowing for integer degrees of differentiation. He found little evidence of the PPP hypothesis in the countries examined. Other papers also investigating African economic growth and development include Ghirmay (2004) Adjasi and Biekpe (2006) Deverajan and Kasekende (2011), Ngepah (2012) and Walle (2014).

The contribution of the present study is threefold. First, it conducts the analysis for the exchange rate of the South African rand vis-à-vis the US dollar using a fractional integration approach to test the PPP hypothesis. This is particularly appropriate because it allows for a much higher degree of flexibility than the standard methods based on the I(0) versus I(1) dichotomy; moreover, it enables us to measure the speed of adjustment following shocks as well as to determine whether departures from the null imply mean reversion ($d < 1$) or explosive behaviour ($d > 1$). Second, it offers new evidence on a key African economy, and therefore adds to the limited existing literature on PPP in the African continent. Third, unlike most existing studies, it tests for PPP at different data frequencies (daily, weekly and monthly), and in this way also sheds light on whether PPP results are robust to the frequency examined.

The outline of the paper is as follows. Section 2 describes the data. Section 3 presents the empirical results, and Section 4 offers some concluding remarks.

2. Data

The series used for the analysis is the real exchange rate for the South African rand vis-à-vis the US dollar, for the time period 2 January 1990–31 December 2008, at the daily, weekly and monthly frequency; the data are obtained from the 'Statistics South Africa' (http://www.statssa.gov.za), and the Federal Reserve Bank of St. Louis database. Thus, we focus on the post-apartheid period.

Figure 1: Real exchange rate (South Africa rand/US dollar)

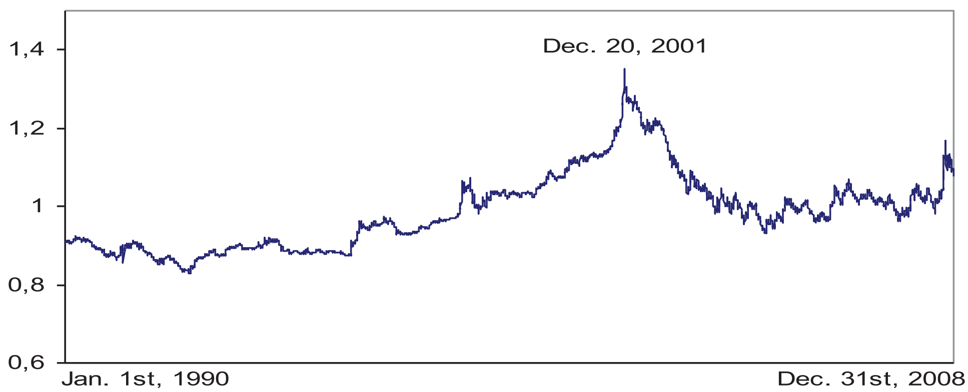
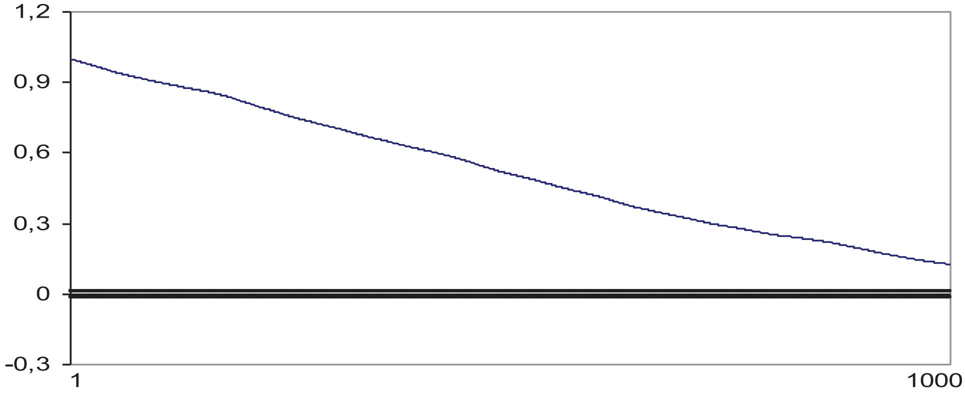


Figure 2: Correlogram of the real exchange rate



Note: The thick lines refer to the 95% confidence band for the null hypothesis of no autocorrelation.

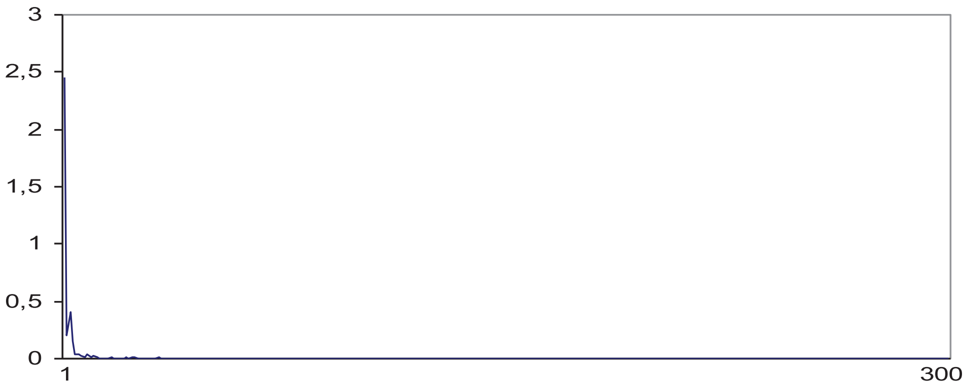
Figure 1 plots the series at the daily frequency. At first sight, it appears to be stationary, but to exhibit some degree of dependence. However, the correlogram and the periodogram, plotted in Figures 2 and 3 respectively, both clearly indicate that the series is non-stationary: the former displays values decaying very slowly to zero and the latter has its highest value at the zero frequency.

Figures 4–6 show similar plots for the first differenced data. These exhibit higher values towards the end of the sample, which may be consistent with conditional heteroscedastic models. In this paper, however, we focus on the degree of dependence and use a procedure that is robust to heteroscedastic errors. The correlogram and the periodogram of the differenced data indicate that the series may be stationary or I(0).

3. Empirical Results

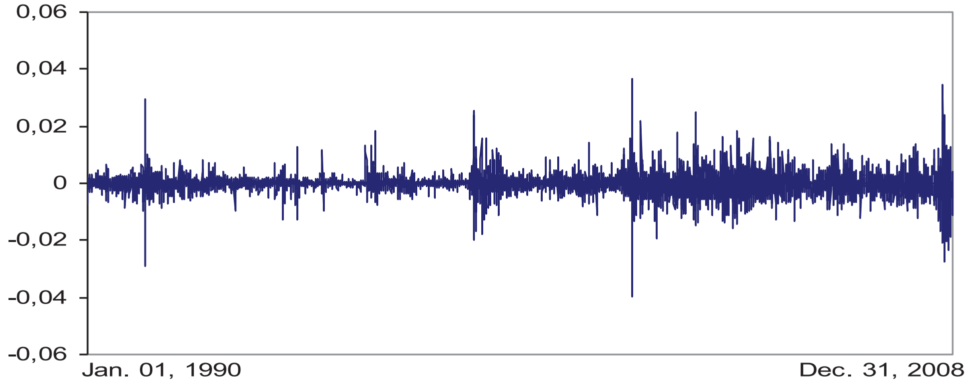
As a first step we carry out standard unit root tests, specifically ADF (Dickey and Fuller, 1979), PP (Phillips and Perron, 1988), and KPSS (Kwiatkowski *et al*., 1992) tests to determine whether the series is non-stationary I(1) or stationary I(0). The results (not reported for reasons of space) strongly support the presence of a unit root. However, they should be taken with caution, as these methods have extremely low power if the alternatives are of a fractional form.

Figure 3: Periodogram of the real exchange rate (South Africa rand/US dollar)



Note: The horizontal axis refers to the discrete Fourier frequencies $\lambda_j = 2\pi j/T$.

Figure 4: First differences of the real exchange rates



Next we consider the possibility of fractional integration and examine first a model of the form:

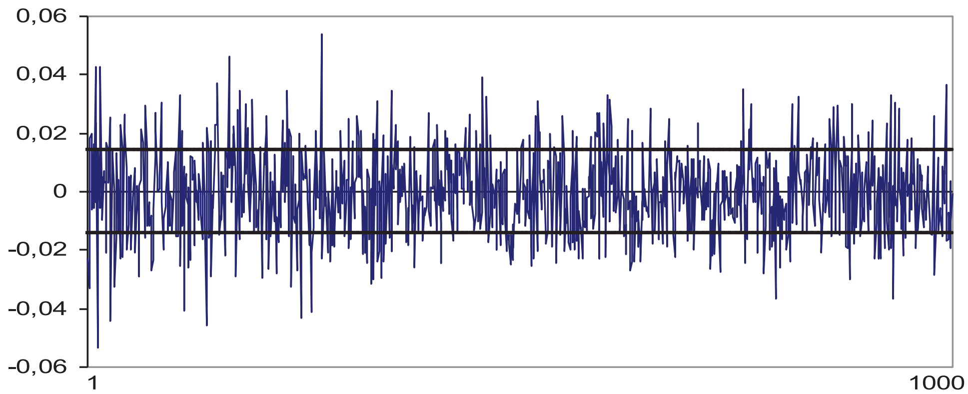
$$y_t = \alpha + \beta t + x_t; (1 - L)^d x_t = u_t, t = 1, 2, ..., \quad (1)$$

where y_t is the time series observed, α and β are the coefficients on the deterministic terms (an intercept and a linear time trend respectively), and x_t is assumed to be I(d), where d can be any real number. Different assumptions will be made about the error term u_t in (1).

We estimate d in (1) using a Whittle function in the frequency domain (Dahlhaus, 1989). In addition, we also employ a testing procedure developed by Robinson (1994) that is very general since it allows testing any real value d in (1) (i.e., including non-stationarity, $d \geq 0.5$) with a standard normal asymptotic distribution. This limit distribution holds independently of the modelling assumptions about the I(0) error term.[1]

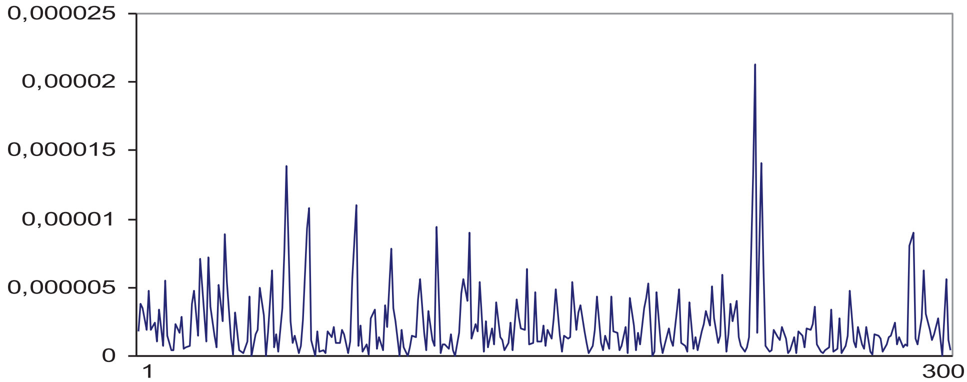
First we assume that the disturbances (u_t in Equation 1) are white noise. We report the estimates of d along with the 95 per cent confidence band of the non-rejection values using Robinson's (1994) parametric approach. The results are displayed in the first row in Table 1. It can be seen that if regressors are not included (i.e., $\alpha = \beta = 0$ in Equation 1) the estimated value of d is 0.990, and the unit root null cannot be rejected at the 5 per cent level.[2] However, when including an intercept or an intercept with a linear time trend, the estimated value of d is around 0.978 and the unit root null is rejected in favour of slow mean reversion.[3]

Figure 5: Correlogram of the first differenced data



Note: The thick lines refer to the 95% confidence band for the null hypothesis of no autocorrelation.

Figure 6: Periodogram of the first differenced data



Note: The horizontal axis refers to the discrete Fourier frequencies $\lambda_j = 2\pi j/T$.

Next, we allow for weak autocorrelation in the error term and assume that u_t in (1) is an AR(1) process. Higher AR orders were also considered obtaining very similar results. In this case, if regressors are not included the estimated value of d is significantly above 1; however, with an intercept and/or a linear trend the estimated d is below unity and the unit root null cannot be rejected in these two cases.

The results presented so far seem to indicate that the exchange rate of the South African rand with respect to the US dollar is a highly dependent variable with an order of integration very close to 1. Next we check if the above result holds for different data frequencies. The results on a weekly basis are reported in Table 2, while those based on monthly data are given in Table 3.

Starting with the weekly case, it can be seen that in the case of white noise residuals the unit root null cannot be rejected, even though the estimate of d is smaller than 1 in the case of no regressors and above 1 with an intercept and/or a trend. If u_t is autocorrelated the results are similar to the daily case, and provide evidence of $d > 1$ for the case of no regressors and I(1) with deterministic terms. Finally, when using monthly data (see Table 3) the same conclusions hold, that is, the unit root cannot be rejected for the case of uncorrelated errors, d is above 1 with AR(1) u_t with no regressors, and the I(1) hypothesis cannot be rejected when including an intercept and/or a linear trend. Therefore, at least for the data analysed here, the results seem to be robust across data frequencies.

On the basis of LR tests and the t-values for the deterministic terms, a model with an intercept and AR(1) disturbances appears to be the most adequate specification for each series. Thus, the orders of integration are 0.980, 1.059 and 0.961 respectively for the daily, weekly and monthly exchange rates, and the unit root null cannot be rejected for any of the three series, suggesting that shocks have permanent effects. This implies that in the event of an exogenous shock decisive policy action must be taken for mean reversion to occur and the equilibrium relationship to be restored.

The potential presence of structural breaks should also be investigated. Note that fractional integration and structural breaks are closely related issues. For example, Bhattacharya *et al*. (1983), Teverovsky and Taqqu (1997), Diebold and Inoue (2001), Granger and Hyung (2004) and Ohanissian *et al*. (2008) among others show that fractional integration may be a spurious phenomenon caused by the existence of breaks in short-memory I(0) contexts. Similarly, Kuan and Hsu (1998), Wright (1998) and Krämer and Sibbertsen (2002) showed that evidence of structural change might be spurious since most commonly employed tests for breaks are biased towards an over-rejection of the null of no change when the process exhibits long memory.

Table 1: Estimates of *d* using daily data

	No regressors	An intercept	A linear trend
White noise	0.990 (0.973, 1.009)	0.978 (0.961, 0.996)	0.978 (0.961, 0.996)
AR (1) errors	1.300 (1.257, 1.344)	0.980 (0.956, 1.006)	0.980 (0.957, 1.006)

Notes: The values in parentheses refer to the 95% confidence band of the non-rejection values of d.

Table 2: Estimates of *d* using weekly data

	No regressors	An intercept	A linear trend
White noise	0.997 (0.959, 1.041)	1.033 (0.995, 1.078)	1.033 (0.995, 1.078)
AR (1) errors	1.319 (1.232, 1.415)	1.059 (0.982, 1.150)	1.059 (0.982, 1.150)

Note: The values in parentheses refer to the 95% confidence band of the non-rejection values of d.

In this paper we employ a recent technique developed by Gil-Alana (2008) that allows for breaks at unknown periods of time with different orders of integration across subsamples. In its simplest form (i.e., with a single break) it takes the following form:

$$y_t = {\beta_1}^T z_t + x_t; (1-L)^{d_1} x_t = u_t, t = 1, ..., T_b \tag{2}$$

and

$$y_t = {\beta_2}^T z_t + x_t; (1-L)^{d_2} x_t = u_t, t = T_b + 1, ..., T, \tag{3}$$

where the β's are the coefficients of the deterministic terms, d_1 and d_2 can take real values, u_t is I(0), and T_b is the unknown break date. This method is based on minimizing the residuals sum of squares in the two subsamples and can be easily extended to the case of two or more breaks (see Gil-Alana, 2008).

The results based on the above approach are displayed in Tables 4 and 5 respectively for the two cases of uncorrelated and AR(1) errors. In all cases the model is specified so as to include an intercept but not a linear trend. The first noticeable feature is the presence of a single break in December 2001 in all three series. This might be interpreted as a consequence of the change in the monetary policy framework which took place the year before, when inflation targeting was introduced. It is also consistent with other studies focusing on inflation and interest rate expectations data and forward interest rate data to show the increased credibility and reasonable predictability of monetary policy since the adoption of inflation targeting in 2000 (see Aron and Muellbauer, 2007). Also, the order of integration decreases slightly after the break.

Starting with the case of white noise disturbances (Table 4), the estimated orders of integration for the pre-break period are 1.003, 1.087 and 1.161 for daily, weekly and monthly exchange rates respectively, the unit root null being rejected in the last two cases in favour of values of d above 1. Following the break the orders are 0.986, 0.964 and 0.972, and the unit root null is never rejected. In the case of AR(1) disturbances (see Table 5) the values are slightly different but the same conclusions hold. Thus, the unit root null is rejected in favour of $d > 1$ for weekly and monthly exchange rates for the periods before the break, but it cannot be rejected for any of the three series for the post-break period.

To summarize, we find a slightly lower degree of integration after the break in 2001, although PPP still does not hold. This is in contrast to the results of other studies on the PPP hypothesis focusing on the South African rand (see Akinboade and Makina, 2006a) and finding some evidence supporting it provided breaks are taken into account. However, it is in line with the findings reported in Caporale and Gil-Alana (2013) for a number of sub-Saharan countries including South Africa and rejecting PPP.[4] As already pointed out in that study and in a related paper by Olayungbo (2011) the lack of conformity to PPP has important implications for a prospective African Union and the creation of a common currency, since it raises the question of its feasibility and long-run sustainability.

Table 3: Estimates of *d* using monthly data

	No regressors	An intercept	A linear trend
White noise	0.991 (0.914, 1.088)	1.035 (0.953, 1.143)	1.035 (0.953, 1.143)
AR (1) errors	1.336 (1.182, 1.507)	0.961 (0.783, 1.149)	0.961 (0.787, 1.149)

Note: The values in parentheses refer to the 95% confidence band of the non-rejection values of d.

Table 4: Estimates of *d* with a single break and white noise disturbances

Data frequency	Break date	d_1	d_2
Daily	20 December 2001	1.003 (0.984, 1.026)	0.986 (0.957, 1.018)
Weekly	14 December 2001	1.087 (1.044, 1.139)	0.964 (0.880, 1.072)
Monthly	December 2001	1.161 (1.026, 1.343)	0.972 (0.859, 1.138)

Note: The values in parentheses refer to the 95% confidence band of the non-rejection values of *d*.

Table 5: Estimates of *d* with a single break and AR (1) disturbances

Data frequency	Break date	d_1	d_2
Daily	20 December 2001	1.021 (0.992, 1.054)	0.987 (0.942, 1.038)
Weekly	14 December 2001	1.153 (1.063, 1.269)	0.859 (0.683, 1.148)
Monthly	December 2001	1.177 (0.963, 1.695)	0.917 (0.689, 1.153)

Note: The values in parentheses refer to the 95% confidence band of the non-rejection values of *d*.

4. Conclusions

This paper has tested the PPP hypothesis for the South African rand/US dollar real exchange rate using a fractional integration framework. The results suggest that this is a highly dependent variable with an order of integration very close to 1. This finding is not affected by the data frequency considered (daily, weekly or monthly). Also, there appears to be a single break in December 2001 (possibly corresponding to a change in the monetary policy framework), with the unit root null being rejected in favour of $d > 1$ for the periods before the break, but not afterwards. Thus, although the degree of dependence is lower after the break, no evidence of mean reversion is found, implying that PPP does not hold since the effects of shocks to the real exchange rate last forever.

Our results are consistent with those from other studies also examining African countries and finding partial or little support for the PPP hypothesis in contrast to developed countries (see, e.g. Nagayasu, 2002; Mkenda, 2001; and Olayungbo, 2011). The weaker evidence for PPP in comparison to the developed countries suggests that an African Union might encounter some difficulties in the absence of appropriate policy actions.

Notes

1. For a review of fractional integration and its application in economics and finance, see Gil-Alana and Hualde (2009). More details on Robinson's (1994) method can be found in any of the numerous empirical applications of this tests (e.g. Gil-Alana and Robinson, 1997; Gil-Alana, 2000; Gil-Alana and Henry, 2003).
2. Note that the 95% confidence interval is now (0.973, 1.009) including the case of a unit root (i.e., $d = 1$).
3. In these cases, the interval is (0.961, 0.996), therefore excluding the unit root case.
4. Unlike the present one, that study did not consider different data frequencies and did not model breaks, therefore the evidence presented here is much more comprehensive and robust.

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