academicJournals

Vol. 5(4), pp. 57-63, August 2013
DOI: 10.5897/JASD2012.0218
ISSN 2141-2189 ©2013 Academic Journals
http://www.academicjournlas.org/JASD

Journal of African Studies and Development

Full Length Research Paper

Testing the PPP hypothesis in the Sub-saharan countries

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Accepted 1 July, 2013

This paper examines the Purchasing Power Parity (PPP) hypothesis in a number of Sub-Saharan countries by testing the order of integration in the log of their real exchange rates vis-à-vis the US dollar. I(d) estimation techniques based on both asymptotic and finite sample results are used. The test results led to the rejection of PPP in all cases: orders of integration below 1 are found in fourteen countries, but the unit root null cannot be rejected.

Key words: Purchasing Power Parity (PPP), fractional integration.

INTRODUCTION

This paper examines the Purchasing Power Parity (PPP) hypothesis in a number of Sub-Saharan countries using a time series approach. PPP is a crucial issue in international economics. It is assumed to hold continuously in flexible-price models of the exchange rate, whilst it is a long-run property in sticky-price ones. In the new open economy models it is a condition for market completeness (Chortareas and Kapetanios, 2009).

Several studies have analysed it empirically. Some of them have tested for cointegration between nominal exchange rates and prices (Kim, 1990; McNown and Wallace, 1994; Serletis and Goras, 2004; Gouveia and Rodrigues, 2004; etc.). Others have applied unit root tests to real exchange rate (these are the so-called "stage-two" tests (Froot and Rogoff, 1995), although these have turned out to be unable to distinguish between random-walk behaviour and very slow mean-reversion to the long-run equilibrium level (Frankel, 1986; Lothian and Taylor, 1997), as 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 (Caporale et al., 2003). Panel methods have recently been used to increase the power of PPP tests (Chortareas and Kapetanos, 2009).

The present study makes a twofold contribution. First, it adopts a more general framework than the standard unit root tests to investigate the presence of mean-reverting behaviour in the real exchange rate. Specifically, it uses fractional integration or I(d) techniques allowing the degree of integration d to be any real number, therefore introducing a higher degree of flexibility in the dynamic specification of the stochastic processes followed by the variables of interest. Second, it focuses on a span of data from 1970 to 2012 for a large set of 44 Sub-Saharan countries whose exchange rates to our knowledge have not been previously analysed using advanced time series methods. Earlier studies of this type have normally focused on the developed countries and analysed some of the major currencies (Booth et al., 1982; Cheung,1993;

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JEL classification: C22, F31.

Baum et al., 1999). By contrast, the only previous empirical study on Sub-Saharan Africa is due to Olayungbo (2011), but it considers a smaller subset of 16 countries over a relatively short sample period and carries out standard unit root tests whose low power has already been mentioned as well as panel unit root tests, the limitations of which have also been highlighted and extensively discussed in many studies. Evidence on PPP in the Sub-Saharan countries is particularly interesting in view of the current discussion on creating an African Union that would eventually have its own currency and central bank, as its feasibility would also depend on the degree of conformity to PPP.

The layout of this paper is as follows. Section 2 outlines the econometric approach. Section 3 describes the data and presents the empirical results. Section 4 offers some concluding remarks.

METHODOLOGY

We consider the following model,

$$y_t = \alpha + \beta t + x_t, \qquad t = 1, 2, ..., \tag{1}$$

where y_t is the observed time series, α and β are the coefficients on the deterministic terms (an intercept and a linear trend), and x_t is assumed to be I(d) and defined as,

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

with $x_t = 0$ for $t \leq 0$, and where L is the lag operator ($Lx_t = x_{t-1}$), d can take any real value and u_t is assumed to be I(0). The fractional differencing parameter,d, may be equal to 0, a fraction between 0 and 1, 1, or even above 1. When it is not an integer, the process is said to be fractionally integrated. In this context, the parameter d plays a crucial role for the degree of persistence of the series. If d = 0 in (2), $x_t = u_t$, and the series is I(0). If d belongs to the interval (0, 0.5) the series is still covariance stationary but the autocorrelations take a longer time to disappear than in the I(0) case. If d is in the interval [0.5, 1) the series is no longer stationary; however, it is still mean-reverting in the sense that shocks affecting it disappear in the long run. Finally, if d ≥ 1 the series is nonstationary and non-mean-reverting. Thus, by allowing d to be any real value, we introduce more flexibility in the dynamic specification of the series than in the classical I(0) and I(1) representations. These processes (with non-integer d) were first considered by Granger (1980), Granger and Joyeux (1980) and Hosking (1981) and since then have been widely employed to describe the behaviour of many economic time series.

In the empirical analysis we test the null hypothesis:

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

in a model given by the equations (1) and (2) where d_o in (3) can be any real value. Thus, under the null hypothesis (3) the model becomes:

$$y_t = \alpha + \beta t + x_t; \qquad (1-L)^{d_o} x_t = u_t; \qquad t = 1, 2, ... \tag{4}$$

This is a very general specification that includes many cases of interest. Thus, for example, if we cannot reject the null with $d_o = 0$, we are in the classical trend-stationary representation with or without weak (ARMA) autocorrelation in u_t.[1] On the other hand, if we cannot reject the null with $d_o = 1$, the unit root model advocated by many authors is given support. Moreover, we can also consider cases where d_o can be a real value between 0 and 1, or even above 1. As mentioned before, the estimation of d_o is crucial to determine the degree of persistence: the higher is the degree of integration, the higher is the level of dependence across the observations, and if $d_o < 1$ the series will be mean- reverting with shocks disappearing in the long run.

To test the null hypothesis (3), we employ a parametric approach developed by Robinson (1994). This is a general testing procedure based on the Lagrange Multiplier (LM) principle that uses the Whittle function in the frequency domain. The test statistic is given by:

$$\hat{r} = \frac{T^{1/2}}{\hat{\sigma}^2} \frac{\hat{a}}{\hat{A}^{1/2}},$$

where T is the sample size, and

$$\hat{a} = \frac{-2\pi}{T} \sum_{j}^{T-1} \psi(\lambda_j) g_u(\lambda_j; \hat{\tau})^{-1} I(\lambda_j);$$

$$\hat{\sigma}^2 = \sigma^2(\hat{\tau}) = \frac{2\pi}{T} \sum_{j=1}^{T-1} g_u(\lambda_j; \hat{\tau})^{-1} I(\lambda_j),$$

$$\hat{A} = \frac{2}{T}\left(\sum_{j}^{T-1} \psi(\lambda_j)\psi(\lambda_j)' - \sum_{j}^{T-1} \psi(\lambda_j)\hat{\varepsilon}(\lambda_j)' \left(\sum_{j}^{T-1} \hat{\varepsilon}(\lambda_j)\hat{\varepsilon}(\lambda_j)' \right)^{-1} \sum_{j}^{T-1} \hat{\varepsilon}(\lambda_j)\psi(\lambda_j)' \right)$$

$$\psi(\lambda_j) = \log\left| 2\sin\frac{\lambda_j}{2} \right|; \qquad \hat{\varepsilon}(\lambda_j) = \frac{\partial}{\partial \tau} \log g_u(\lambda_j; \hat{\tau});$$

with $\lambda_j = 2\pi j/T$, $I(\lambda_j)$ is the periodogram of $\hat{u}_t = (1-L)^{d_o} y_t - \hat{\beta}' \bar{z}_t$, with

$$\hat{\beta} = \left(\sum_{t=1}^{T} \bar{z}_t \bar{z}_t' \right)^{-1} \sum_{t=1}^{T} \bar{z}_t (1-L)^{d_{1o}} y_t;$$

$\bar{z}_t = (1-L)^{d_o} z_t, \quad z_t = (1, t)^T$; and $\hat{\tau} = \arg\min_{\tau \in T^*} \sigma^2(\tau)$, with T^* being a suitable subset of the R^q Euclidean space. Finally, the function g_u above is a known function coming from the spectral density of u_t:

$$f(\lambda) = \frac{\sigma^2}{2\pi} g_u(\lambda; \tau), \qquad -\pi < \lambda \leq \pi.$$

Note that these tests are purely parametric and, therefore, they require specific modelling assumptions about the short-memory specification of u_t. Thus, if u_t is white noise, $g_u \equiv 1$, and if u_t is an AR process of the form $\phi(L)u_t = \varepsilon_t$, $g_u = |\phi(e^{i\lambda})|^{-2}$, with $\sigma^2 = V(\varepsilon_t)$, so that the AR coefficients are a function of τ.

Robinson (1994) showed that, under certain very mild regularity conditions, the LM-based statistic $(\hat{r})$

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[1]Note that u_t is I(0) and therefore could incorporate stationary and invertible ARMA sequences.

$$\hat{r} \rightarrow_{dtb} N(0,1) \quad as \quad T \rightarrow \infty,$$

where " $\rightarrow_{dtb}$ " stands for convergence in distribution, and this limit behaviour holds independently of the regressors used in (1) and the specific model for the I(0) disturbances u_t. The functional form of this procedure can be found in any of the numerous empirical applications based on these tests.

As in other standard large-sample testing situations, Wald and LR test statistics against fractional alternatives will have the same null and limit theory as the LM test of Robinson (1994). In fact, Lobato and Velasco (2007) essentially employed such a Wald testing procedure, and though this and other recent methods such as the one developed by Demetrescu et al. (2008) have been shown to be robust with respect to even unconditional heteroscedasticity (Kew and Harris, 2009), they require a consistent estimate of d, and therefore the LM test of Robinson (1994) seems computationally more attractive.

DATA AND EMPIRICAL RESULTS

We use data on real exchange rates, in logged form, for forty-four Sub-Saharan countries, for the time period 1970 – 2012 (with 2005 as the base year), obtained from the Economic Research Service, US Department of Agriculture (http://www.ers.usda.gov).[2]

We consider the model given by the equations (1) and (2), testing H_o (3) for values of d_0 from 0 to 2 with 0.001 increments, that is, d_o = 0, 0.001, 0.002, …, 1.999 and 2. We report in Table 1 the estimates of d based on the Whittle function in the frequency domain (Dahlhaus, 1989) along with the 95% confidence interval of non-rejection values of d using Robinson's (1994) tests, under the assumption that the error term u_t in (4) is a white noise process. Weakly (ARMA) autocorrelated errors were also considered and led to very similar results.

Table 1 displays the results for the three standard cases usually analysed in the literature, that is, with no regressors in the undifferenced regression model in (4) ($\alpha = \beta = 0$ a priori); with an intercept (α unknown and $\beta = 0$ a priori); and with an intercept and a linear time trend (α and β unknown); statistically significant deterministic terms are in bold. It appears that the time trend is only required for four series, namely those for the real exchange rates of Gambia, Guinea Bissau, Malawi and Sudan. In all the remaining cases, an intercept is sufficient to describe the deterministic part. Focusing now on the estimated orders of integration of the series (for the selected models), we see that for fourteen countries the value of d is smaller than 1 - these are Malawi, Guinea Bissau, Liberia, Swaziland, Sudan, Gambia, Madagascar, Comoros, Angola, Togo, Botswana, Senegal, Ivory Coast and Central Africa. However, in all these cases, the confidence intervals for the values of d imply that the unit root null hypothesis (d = 1) cannot be rejected. For the remaining countries, the estimated d is above 1, and the unit root null is rejected in favour of d > 1 in the cases of Djibouti, Sierra Leone, Mauritania, Cape Verde, Eritrea, Uganda, Sao Tome, Tanzania and Ghana.

Table 2 summarizes the results in terms of the degree of persistence. The countries are divided in three groups according to the statistical significance of the estimated values of d: mean reversion (d < 1); unit roots (with d < 1 and with d > 1), and explosive behaviour (d > 1). In other words we distinguish between the following cases of: a) statistical evidence against the unit root and in favour of mean reversion (d < 1); b) the unit root cannot be rejected (d = 1); and c) statistical evidence against the unit root and in favour of d > 1. In case b) (unit root) we also distinguish between estimates which are smaller and greater than 1.

The results indicate that there is no single country where mean reversion is statistically significant, implying that PPP does not hold anywhere. However, although the unit root null hypothesis cannot be rejected in 35 countries, in 14 of them the estimated value of d is below 1 implying that PPP might hold in the very long run. Another group of nine countries displays explosive behaviour. Overall, the evidence does not support PPP, consistently with the findings of Olayungbo (2011), who reports that it holds only in Ghana and Uganda; in fact even for these two countries PPP is rejected according to our results, since they are found to belong to the group with the highest degree of persistence.

We also examined the finite sample behaviour of sized-corrected versions of Robinson (1994) tests by means of Monte Carlo simulations, and compared the results with those based on the asymptotic critical values. Note that in the original paper by Robinson (1994) he stressed large sample theory and suggested approximate critical values. Thus, we calculated the empirical size of the test statistic $\hat{r}$ for a sample size T = 42 as in our case, based on 10,000 replications, for the three cases of no regressors, an intercept, and an intercept with a linear time trend. In all cases, we assume u_t is Gaussian white noise process with zero mean and variance 1, generated by the routines GASDEV and RAN3 of Press et al. (1986).

Table 3 displays for each country both the asymptotic and the finite sample 95% confidence intervals for the fractional differencing parameter d. We notice that in all cases the intervals are shifted to the right, implying higher degrees of integration, and therefore, even less evidence of PPP for the Sub-Saharan countries. These results are presented in Table 4. There are four countries (Mozambique, Seychelles, Burundi and Zambia) where the unit root cannot be rejected in Table 2, and is rejected in favour of d < 1 in Table 4.

Conclusion

This paper applies long-range dependence or fractional integration techniques to test for PPP in a set of 44 Sub-

[2]Monthly data are available only for 17 countries, and therefore carrying out the analysis at a monthly frequency would considerably reduce the sample of countries; moreover, the seasonal component of these series could generate a potential bias. For these reasons we have decided to use the annual data set.

Table 1. Estimates of d and 95% confidence intervals.

Country	No. of regressors	Intercept	A linear time trend
ANGOLA	0.93 (0.74, 1.23)	0.95 (0.68, 1.36)	0.95 (0.64, 1.36)
BURKINA FASO	0.86 (0.64, 1.17)	1.04 (0.83, 1.35)	1.04 (0.81, 1.35)
BENIN	0.89 (0.69, 1.18)	1.13 (0.92, 1.45)	1.13 (0.92, 1.45)
BOTSWANA	0.74 (0.40, 1.10)	0.96 (0.73, 1.34)	0.96 (0.63, 1.34)
BURUNDI	0.88 (0.66, 1.18)	1.23 (0.98, 1.63)	1.23 (0.98, 1.63)
CAPE VERDE	0.88 (0.67, 1.19)	1.30(1.08, 1.66)	1.30(1.078 1.67)
CAMEROON	0.87 (0.66, 1.17)	1.05 (0.77, 1.43)	1.05 (0.79, 1.42)
CENTRAL AF.	0.85 (0.63, 1.15)	0.99 (0.80, 1.29)	0.99 (0.77, 1.29)
CHAD	0.86 (0.64, 1.16)	1.03 (0.82, 1.35)	1.03 (0.80, 1.35)
COMOROS	0.86(0.643 1.16)	0.95 (0.76, 1.25)	0.95 (0.74, 1.25)
CONGO REP.	0.87 (0.67, 1.17)	1.10 (0.62, 1.51)	1.10 (0.78, 1.49)
DJIBOUTI	0.90 (0.70, 1.19)	1.22 (1.03, 1.56)	1.22 (1.03, 1.56)
EQ. GUINEA	0.85 (0.63, 1.17)	1.08 (0.92, 1.31)	1.08 (0.92, 1.31)
ERITREA	0.86 (0.63, 1.21)	1.31 (1.10, 1.64)	1.30 (1.09, 1.65)
ETHIOPIA	0.72 (0.49, 1.07)	1.10 (0.90, 1.42)	1.10 (0.89, 1.42)
GABON	0.86 (0.64, 1.17)	1.11 (0.90, 1.41)	1.11 (0.91, 1.39)
GAMBIA	0.84 (0.59, 1.15)	0.89 (0.72, 1.17)	0.87 (0.60, 1.17)
GHANA	1.38 (1.11, 1.87)	1.45 (1.15, 2.00)	1.45 (1.15, 2.00)
GUINEA B.	0.86 (0.63, 1.16)	0.84 (0.71, 1.07)	0.83 (0.68, 1.07)
GUINEA	0.74 (0.41, 1.12)	1.01 (0.83, 1.29)	1.00 (0.80, 1.29)
IVORY COAST	0.88 (0.67, 1.18)	0.99 (0.70, 1.37)	0.99 (0.72, 1.37)
KENYA	0.89 (0.68, 1.20)	1.06 (0.90, 1.30)	1.07 (0.90, 1.31)
LESOTHO	0.73 (0.17, 1.13)	1.00 (0.71, 1.46)	1.00 (0.68, 1.46)
LIBERIA	0.80 (0.64, 1.08)	0.84 (0.69, 1.10)	0.82 (0.63, 1.10)
MADAGASCAR	0.90 (0.70, 1.19)	0.93 (0.77, 1.19)	0.93 (0.76, 1.19)
MALAWI	0.85 (0.62, 1.16)	0.82 (0.68, 1.17)	0.74 (0.37, 1.17)
MAURITANIA	0.84 (0.60, 1.16)	1.29 (1.10, 1.56)	1.29 (1.10, 1.54)
MAURITIUS	0.87 (0.65, 1.19)	1.05 (0.81, 1.42)	1.05 (0.78, 1.42)
MOZAMBIQUE	0.91 (0.69, 1.27)	1.18 (0.97, 1.56)	1.19 (0.96, 1.55)
NAMIBIA	0.78 (0.46, 1.15)	1.13 (0.84, 1.58)	1.13 (0.84, 1.58)
NIGER	0.86 (0.63, 1.18)	1.07 (0.90, 1.34)	1.07 (0.89, 1.34)
NIGERIA	0.84 (0.58, 1.19)	1.12 (0.82, 1.54)	1.12 (0.81, 1.54)
REUNION	0.85 (0.63, 1.16)	1.00 (0.82, 1.27)	1.00 (0.81, 1.26)
SIERRA LEONE	0.84 (0.60, 1.15)	1.26 (1.04, 1.63)	1.26 (1.04, 1.60)
SOUTH AFRICA	0.14 (0.08, 0.68)	1.23 (0.52, 2.21)	1.22 (0.30, 2.19)
SAO TOME	0.85 (0.62, 1.17)	1.37 (1.14, 1.75)	1.36 (1.14, 1.74)
SENEGAL	0.86 (0.64, 1.17)	0.98 (0.78, 1.28)	0.98 (0.76, 1.28)
SEYCHELLES	0.84 (0.64, 1.14)	1.22 (0.95, 1.62)	1.20 (0.96, 1.54)
SUDAN	0.91 (0.74, 1.19)	0.87 (0.71, 1.13)	0.86 (0.67, 1.14)
SWAZILAND	0.85 (0.57, 1.22)	0.85 (0.58, 1.32)	0.85 (0.54, 1.32)
TANZANIA	0.89 (0.67, 1.19)	1.42 (1.19, 1.79)	1.42 (1.19, 1.79)
TOGO	0.88 (0.66, 1.19)	0.95 (0.74, 1.27)	0.95 (0.73, 1.27)
UGANDA	0.78 (0.47, 1.14)	1.35 (1.11, 1.70)	1.34 (1.16, 1.68)
ZAMBIA	0.880 (0.669, 1.178)	1.304 (0.995, 1.811)	1.306 (0.998, 1.810)

Saharan countries. The advantage of this approach is its generality and flexibility in comparison to standard time series methods restricting the degree of integration to integer values. Previous evidence (Olayungbo, 2011) was only available for a smaller subset of countries and a short sample period and was based on low-power unit root tests as well as panel tests whose drawbacks are also well known.

Table 2. Summary based on the asymptotic results.

Mean reversion (d < 1)	Unit root (d = 1)		Explosive behaviour (d > 1)
	d < 1	d > 1	
xxx	Malawi (0.744)	Reunion (1.002)	Djibouti (1.228)
	Guinea Bis. (0.831)	Lesotho (1.009)	Sierra Leone (1.265)
	Liberia (0.845)	Guinea (1.011)	Mauritania (1.293)
	Swaziland (0.859)	Chad (1.035)	Cape Verde (1.308)
	Sudan (0.861)	Burkina Faso (1.041)	Eritrea (1.314)
	Gambia (0.870)	Mauritius (1.051)	Uganda (1.358)
	Madagascar (0.937)	Cameroon (1.053)	Sao Tome (1.371)
	Comoros (0.957)	Kenya (1.068)	Tanzania (1.428)
	Angola (0.959)	Niger (1.078)	Ghana (1.459)
	Togo (0.959)	Eq. Guinea (1.085)	
	Botswana (0.967)	Ethiopia (1.102)	
	Senegal (0.986)	Congo Rep. (1.106)	
	Ivory Coast (0.996)	Gabon (1.115)	
	Centr. Africa (0.997)	Nigeria (1.122)	
		Namibia (1.131)	
		Benin (1.138)	
		Mozambique (1.189)	
		Seychelles (1.223)	
		Burundi (1.233)	
		South Africa (1.239)	
		Zambia (1.304)	

Table 3. Asymptotic and finite sample confidence intervals for the values of d.

Country	Asymptotic	Finite samples
ANGOLA	(0.68, 1.36)	(0.75, 1.50)
BURKINA FASO	(0.83, 1.35)	(0.88, 1.47)
BENIN	(0.92, 1.45)	(0.98, 1.57)
BOTSWANA	(0.73, 1.34)	(0.79, 1.48)
BURUNDI	(0.98, 1.63)	(1.05, 1.79)
CAPE VERDE	(1.08, 1.66)	(1.14, 1.79)
CAMEROON	(0.77, 1.43)	(0.85, 1.56)
CENTRAL AF.	(0.80, 1.29)	(0.85, 1.40)
CHAD	(0.82, 1.3 5)	(0.87, 1.46)
COMOROS	(0.76, 1.25)	(0.81, 1.35)
CONGO REP.	(0.62, 1.51)	(0.83, 1.64)
DJIBOUTI	(1.03, 1.56)	(1.08, 1.70)
EQ. GUINEA	(0.92, 1.31)	(0.96, 1.39)
ERITREA	(1.10, 1.64)	(1.15, 1.78)
ETHIOPIA	(0.90, 1.42)	(0.95, 1.56)
GABON	(0.90, 1.41)	(0.96, 1.52)
GAMBIA	(0.60, 1.17)	(0.68, 1.27)
GHANA	(1.15, 2.00)	(1.22, 2.22)
GUINEA B.	(0.68, 1.07)	(0.74, 1.18)
GUINEA	(0.83, 1.29)	(0.87, 1.39)
IVORY COAST	(0.70, 1.37)	(0.79, 1.50)
KENYA	(0.90, 1.30)	(0.94, 1.38)
LESOTHO	(0.71, 1.46)	(0.78, 1.63)

Table 3. Contd.

LIBERIA	(0.69, 1.10)	(0.73, 1.20)
MADAGASCAR	(0.77, 1.19)	(0.81, 1.29)
MALAWI	(0.37, 1.17)	(0.48, 1.33)
MAURITANIA	(1.10, 1.56)	(1.15, 1.66)
MAURITIUS	(0.81, 1.42)	(0.86, 1.56)
MOZAMBIQUE	(0.97, 1.56)	(1.02, 1.70)
NAMIBIA	(0.84, 1.58)	(0.92, 1.75)
NIGER	(0.90, 1.34)	(0.94, 1.44)
NIGERIA	(0.82, 1.54)	(0.90, 1.68)
REUNION	(0.82, 1.27)	(0.87, 1.36)
SIERRA LEONE	(1.02, 1.63)	(1.09, 1.77)
SOUTH AFRICA	(0.52, 2.21)	(0.63, 2.44)
SAO TOME	(1.14, 1.75)	(1.20, 1.91)
SENEGAL	(0.78, 1.28)	(0.83, 1.38)
SEYCHELLES	(0.95, 1.62)	(1.02, 1.75)
SUDAN	(0.67, 1.14)	(0.75, 1.24)
SWAZILAND	(0.58, 1.32)	(0.65, 1.51)
TANZANIA	(1.19, 1.79)	(1.25, 1.93)
TOGO	(0.74, 1.27)	(0.80, 1.39)
UGANDA	(1.11, 1.70)	(1.18, 1.82)
ZAMBIA	(0.99, 1.81)	(1.07, 1.99)

On the whole, our results suggest that PPP does not hold in this group of countries. This is in contrast with the

Table 4. Summary based on the finite sample results.

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xxx	Malawi (0.744)	Reunion (1.002)	Mozambique (1.189)
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	Madagascar (0.937)	Cameroon (1.053)	Zambia (1.304)
	Comoros (0.957)	Kenya (1.068)	Cape Verde (1.308)
	Angola (0.959)	Niger (1.078)	Eritrea (1.314)
	Togo (0.959)	Eq. Guinea (1.085)	Uganda (1.358)
	Botswana (0.967)	Ethiopia (1.102)	Sao Tome (1.371)
	Senegal (0.986)	Congo Rep. (1.106)	Tanzania (1.428)
	Ivory Coast (0.996)	Gabon (1.115)	Ghana (1.459)
	Centr. Africa (0.997)	Nigeria (1.122)	
		Namibia (1.131)	
		Benin (1.138)	
		South Africa (1.239)	

available evidence for developed countries based on long-memory models. For instance, using a similar version of Robinson's (1994) tests to the one adopted here, Gil-Alana (2000) found mean reversion in the US real exchange rates vis-à-vis five major currencies with weakly autocorrelated disturbances. Similar conclusions were reached applying fractional integration and cointegration techniques by Masih and Masih (2004) for the Australian dollar real exchange rate vis-à-vis seven major OECD trading partners. Finally, Yoon (2009) applied the Exact Local Whittle estimators of Shimotsu and Phillips (2005) to estimate the long memory parameters of the real exchange rates for more than 100 years in 16 developed countries and concluded again that PPP holds in most of these countries.

The implications of our results for policy makers and for the Sub-Saharan countries are as follows. First, they suggest that the degree of conformity to PPP is much less in these countries compared to the developed ones, and, as already pointed out by Olayungbo (2011), this should be taken into account when considering the proposed African Union and the creation of a common currency, since the absence of PPP relationships between its prospective members raises some doubts about its feasibility or at least long-run sustainability. Second, it is well known that there is a negative relationship between any misalignment of the real exchange rate (RER) and economic performance (economic growth, imports, exports, saving and investment (Ghura and Grennes, 1993), and therefore the lack of PPP is a reason for concern about growth in these countries and calls for exchange rate management policies. Such policies appear to be crucial in this group of countries also because RER misalignment has a negative effect on export performance (Sekkat and Varoudakis, 2000). Overall, our analysis highlights the fact that managing the exchange rate effectively is one of the key challenges in the Sub-Saharan countries and one of the most important issues that should be addressed in that region given the adverse impact on the economy of RER misalignments.

ACKNOWLEDGMENTS

*Luis A. Gil-Alana gratefully acknowledges financial support from the Ministerio de Ciencia y Tecnología ECO2011-2014 ECON Y FINANZAS, Spain) and from a Jeronimo de Ayanz project of the Government of Navarra. Comments from the Editor and two anonymous referees are also gratefully acknowledged.

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