Long-Run Purchasing Power Parity in Eastern and Southern African Countries: Evidence from panel data stationary tests with multiple structural breaks

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Abstract. The panel data stationary test with multiple structural breaks developed by Carrion-i-Silvestre, Del Barrio-Castro and Lopez-Bazo (*Econometrics Journal*, 2005) is used to study the Purchasing Power Parity in a set of 17 Eastern and Southern African economies for the period 1970-2007. This latter is considered both in its traditional view, *i.e.* the reversion of the real exchange rate to a constant mean, and in the Balassa-Samelson view, *i.e.* the reversion of the real exchange rate to a constant trend. The results indicate that taking into account the presence of structural breaks changes dramatically the conclusion about long-run Purchasing Power Parity. Indeed, in contrast with standard stationary tests, our study shows that Purchasing Power Parity holds in the long-run for our panel when the Balassa-Samuelson version is considered.

Keywords: Panel data, PPP, real exchange rate, structural breaks, Africa

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1. Introduction

Testing for long-run Purchasing Power Parity (PPP) remains a major problem in international economics for many reasons (Sarno and Taylor, 2002). Firstly, the validity of PPP questions a large part of open economy macroeconomic theory that assumes PPP as an equilibrium benchmark. Secondly, even if the PPP hypothesis may not be regarded as an explicit exchange rate theory, it may still serve to provide fundamentals determinants which can be used to assess the appropriate level of exchange rate and then possible situations of misalignment. Thirdly, since the development of the Optimal Currency Area theory, we know that PPP implies price convergence in a given zone (Drine and Rault, 2006).

Traditionally, the PPP hypothesis considers a proportional relation between the nominal exchange rate and the relative price ratio, which implies that the real exchange rate is constant over time. So, the most common way to test for PPP consists in investigating unit roots in real exchange rate. If the unit root can be rejected in favour of level stationarity, then deviations from parity are temporary and PPP is said to hold in the long run. Here, PPP is defined in the spirit of Cassel (1925) as reversion of the real exchange rate to a constant mean (Hereafter, C-PPP). However, Balassa (1964) and Samuelson (1964) drew attention to the fact that divergent international productivity level could lead to permanent deviations from PPP. Assuming that PPP holds for traded goods, their argument is based on the fact that productivity differentials between countries determine the domestic relative prices of nontraded goods, leading in the long-run to trend deviations from PPP. Thus, PPP in the spirit of Balassa-Samuelson is defined as reversion of the real exchange rate to a constant trend (hereafter, BS-PPP) (Papell and Prodan, 2006; Darné and Hoarau, 2007).

Empirical research on PPP has therefore focused on the credibility of the unit root finding and on why deviations from PPP exist. And, despite a decade of multiple applications of the unit root tests in analysing PPP, we are still unable to draw homogenous conclusions (Taylor and Taylor, 2004; Taylor, 2006). This lack of consensus has been attributed to the low power of the tests used. As a result, the recent literature, mainly focused on industrial and large emerging economies, has moved on in two new directions. While some researchers have turned to panel unit root tests (Bai and Ng, 2004; Pedroni, 2004; Alba and Papell, 2007), others have opted for nonlinear unit root tests (Chortareas and Kapetanios, 2004; Paya and Peel, 2004; Bahmani-Oskooee *et al.*, 2007). However, few studies have been conducted using data from small developing countries and, in particular, from Africa. Besides, robust researches on PPP are very important for African countries in the extent that these ones have implemented, since the early 1990s, modifications in their exchange rate policies which were based on the assumption of PPP validity (Kargbo, 2003).

Recently, some studies have examined in detail whether or not there is empirical support for long-run PPP in African countries (Krichene, 1998; Odokekun, 2000; Holmes, 2001; Nagayasu, 2002; Kargbo, 2003; Hassanain, 2004, Akinboade and Makina, 2006; Bahmani-Oskooee and Gelan, 2006; Chang *et al.*, 2006). But, once again, a strong consensus could not be reached even if more results converge towards PPP validity. As mentioned by Kargbo (2006), the mixed results on PPP reflect the data and econometric techniques used in addition to the time period under investigation. Unfortunately, except for Akinboade and Makina (2006)[2], all these studies did not consider the presence of structural breaks in data.

And, since the work of Perron (1989), we know that ignoring structural changes could lead to the erroneous acceptance of non-stationarity. However, in a pure conceptual sense, putting forward structural changes in real exchange rate series implies that PPP does not hold because the mean and/or the time trend are not constant even if stationarity is found. Dornbusch and Vogelsang (1991) argue that a “qualified” version of PPP can still be claimed in the presence of one-time shift in the mean level of the real exchange rate that is determined exogenously. They interpret their findings as supporting the Balassa-Samuelson model. Hegwood and Papell (1998) formalize and generalize this idea, allowing for multiple structural breaks that are determined endogenously. Then, PPP can be tested in a lighter version, called by the author as “quasi” PPP, that is the stationarity of the real exchange rate but around a mean which is subject to occasional structural changes, *i.e.* there is reversion to a changing mean. This concept of PPP may also focus on trend stationarity, *i.e.* a reversion to a broken trend (Papell and Prodan, 2006). Besides, in what follows, when we talk about long-run PPP in the framework of structural breaks, we refer to this “quasi” version of Hegwood and Papell (1998)[3].

In this regard, the aim of this paper is to continue the investigations about long-run PPP for a panel of African countries, and more precisely for 17 Eastern and Southern African economies, covering the period 1970-2007. Note that all these countries belong to the same regional arrangement, the so-called COmmon Market of Eastern and Southern Africa

[2] The authors used unit root tests with structural breaks but within a univariate framework for South Africa.
[3] Papell and Prodan (2006) show that in the presence of multiple structural changes, one still can find “strict PPP” (the Cassel alternative of a constant mean or the Balassa-Samuelson alternative of a constant trend) if the changes are offsetting. To this end, they developed unit root tests that restrict the coefficients on the dummy variables that depict the breaks to produce a constant mean or trend in the long-run.

(COMESA)[4]. Indeed, we implement, with standard univariate (Kwiatkowski *et al.*, 1992) and panel data (Hadri, 2000) stationary procedures, the new panel data stationary test of Carrion-i-Silvestre, Del Barrio-Castro and Lopez-Bazo (hereafter, CDL) (2005) which allows for the structural changes to shift the mean and/or the trend of the individual time series. Ultimately, our study result in a very interesting finding given that the PPP hypothesis is confirmed for these economies when the Balassa-Samuelson version is considered.

The remainder of the paper is organized as follows. Section 2 briefly presents the data. Section 3 gives the results obtained from implementing standard stationary tests. Section 4 develops the new test CDL and displays the main findings. Finally, section 5 concludes.

2. The data

The empirical study employs the annually bilateral real exchange rate against the US dollar for 17 selected Eastern and Southern African countries (Egypt, Libya, Comoros, Djibouti, Eritrea, Ethiopia, Kenya, Madagascar, Malawi, Mauritius, Seychelles, Sudan, Swaziland, Tanzania, Uganda, Zambia, Zimbabwe) over the period 1970-2007[5] (646 observations). The real exchange rate (e_t) is given (in logarithms) as $e_t = s_t + p_t^{us} - p_t$ where s_t is the nominal (official) exchange rate defined in local currency units per US dollar, and p_t and p_t^* are the domestic and US price levels, respectively. The Consumer Price Index (CPI) is used in our work. The source of the data is the ERS-USDA database[6] and a summary of the statistics is given in Table 1.

The distributional properties of time series reveal wide variability and significant deviation from normality of several real exchange rate series during the past three decades. The coefficient of variation ranges from 11.7% in Djibouti to 288.2% in Zimbabwe. The Table 1 also provides preliminary evidence about the mean reverting behaviour of real exchange rates and PPP in Eastern and Southern Africa.

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[4] Four other economies also belong to the COMESA: Angola, Burundi, Democratic Republic of Congo and Rwanda. However, we did not have reliable data on these countries. Tanzania does not belong to COMESA at date but is trying to negotiate its future reintegration.

[5] Figures for 2007 are calculated as an average on the first three months of 2007.

[6] This database is available at http://www.ers.usda.gov/data/macroeconomics/.

Table 1: Summary statistics on real exchange rates in Eastern and Southern Africa, 1970-2007

	Mean	Max.	Min.	Std. Dev.	Skewness	Kurtosis	JB	Prob.	CV
Comoros	56.963	104.118	24.952	23.931	0.367	1.980	2.498	0.287	42.011
Djibouti	105.473	129.671	80.852	12.333	0.279	2.429	1.007	0.604	11.693
Egypt	98.855	164.177	46.293	31.485	0.328	2.227	1.627	0.443	31.849
Erithrea	86.463	164.258	46.032	26.489	0.820	4.107	6.197**	0.045	30.636
Ethiopia	57.166	116.723	27.048	29.633	0.641	1.804	4.866*	0.088	51.837
Kenya	70.346	149.215	36.997	25.804	0.662	3.361	2.981	0.225	36.681
Libya	128.957	363.507	62.446	97.309	1.814	4.464	24.246***	0.000	75.458
Madagascar	70.681	104.234	36.671	23.010	-0.163	1.449	3.976	0.137	32.555
Malawi	60.894	113.313	36.741	22.781	0.695	2.160	4.178	0.124	37.412
Mauritus	84.197	108.272	58.781	13.887	-0.165	2.086	1.496	0.473	16.494
Uganda	60.505	116.356	3.841	35.874	-0.053	1.728	2.578	0.276	59.292
Seychelles	118.310	231.950	85.081	33.038	1.773	5.821	32.503***	0.000	27.925
Sudan	52.446	100.416	10.464	30.084	0.375	1.535	4.291	0.117	57.363
Swaziland	76.230	133.690	43.660	21.254	0.500	2.884	1.607	0.448	27.882
Tanzania	83.252	147.375	23.644	40.895	-0.124	1.621	3.108	0.211	49.122
Zambia	35.475	100.000	0.427	41.301	0.502	1.416	5.566*	0.062	116.422
Zimbabwe	397.441	5218.781	36.193	1145.369	3.603	14.666	297.725***	0.000	288.186

Notes: The Jarque-Bera [JB] test is a test for normality of the series (Jarque-Bera, 1980). (***), (**) and (*) indicate the reject of the hypothesis of a normal distribution at the 1%, 5% and 10% level, respectively. The Coefficient of Variation [CV] is given in percentage.

3. The standard stationary tests

3.1. The univariate KPSS approach

To provide a benchmark for our panel data results, we first performed a standard univariate stationary test, namely the Kwiatkowski *et al.* (KPSS) (1992) test using both Bartlett and Quadratic Spectral kernel, on real exchange rate of each country in the panel. In this regard, we consider two specifications that is to say *(i)* a model without a deterministic trend but with an intercept (model 1) and *(ii)* a model with an intercept and a deterministic trend (model 2). The first model allows for verify the long-run C-PPP hypothesis although the second one permits to test the BS-PPP alternative. However, in all cases the stationarity means a reverting behaviour. The results are displayed in Table 2. On the whole, the results are mixed but the KPSS test seems to invalidate the PPP hypothesis in most cases. More precisely, three main findings can be put forward.

Table 2: The univariate KPSS test for PPP in Eastern and Southern Africa, 1970-2007

	KPSS			
	Intercept (model 1)		Trend & Intercept (model 2)	
	Bartlett	Quadratic spectral	Bartlett	Quadratic Spectral
Comoros	0.354(18.2)*	0.336(23.3)*	**0.133(9.61)***	**0.199(10.9)***
Djibouti	**0.218(19.7)**	**0.521(25.6)****	0.220(21)**	0.664(27.7)**
Egypt	0.357(14.1)*	0.312(17.2)	**0.151(9.11)****	**0.154(10.2)****
Eritrea	0.216(14)	0.248(17)	**0.144(14)***	**0.215(18.5)****
Ethiopia	0.583(43.3)**	1.486(66)**	**0.112(10.8)**	**0.121(12.5)***
Kenya	0.332(22.1)	0.429(29.5)*	**0.133(14.4)***	**0.214(17.6)****
Libya	0.656(49)**	2.490(76.4)**	**0.397(30.7)****	**2.533(43.7)****
Madagascar	0.425(31.2)*	0.783(44.5)**	**0.087(8.91)**	**0.083(9.97)**
Malawi	0.352(21)*	0.390(27.6)*	**0.083(5.94)**	**0.077(6.23)**
Mauritus	0.354(23.3)*	0.460(31.3)*	**0.075(7.99)**	**0.103(8.77)**
Uganda	0.344(23.8)	0.490(32.1)**	**0.096(9.82)**	**0.117(11.2)**
Seychelles	0.423(10.7)*	0.337(12.4)	**0.170(7.27)****	**0.147(7.86)****
Sudan	0.329(12.1)	0.273(14.3)	**0.107(7.02)**	**0.099(7.55)**
Swaziland	0.332(14.4)	0.284(17.6)	**0.111(8.69)**	**0.117(9.67)**
Tanzania	0.623(46.4)**	2.129(71.6)**	**0.159(15)****	**0.323(18.5)****
Zambia	1.010(75.8)***	5.451(129)**	**0.418(25)****	**7.033(34.1)****
Zimbabwe	0.373(17.8)*	0.347(22.7)*	**0.516(38.2)****	**5.529(56.8)****
Critical Values 5%	0.463	0.463	0.146	0.146
Critical Values 10%	0.347	0.347	0.119	0.119

Notes: Figures in bold correspond to the good specification. (**) and (*) indicate the reject of the null hypothesis of stationarity at the 5% and 10% level, respectively. The long run variance is estimated using both the Bartlett and Quadratic Spectral kernel with automatic spectral window bandwidth selection (figures in parentheses) as in Andrews (1991).

Firstly, the null hypothesis of stationarity around a constant mean is strongly rejected in 4 (Ethiopia, Libya, Tanzania, Zambia) with the Bartlett kernel, and in 7 (Djibouti, Ethiopia, Libya, Madagascar, Uganda, Tanzania, Zambia) with the Quadratic Spectral kernel, out of 17 cases at the 5% significance level. Moreover, if the 10% level is considered, 7 other countries (Comoros, Egypt, Madagascar, Malawi, Mauritius, Seychelles, Zimbabwe) with the Bartlett kernel and 5 others (Comoros, Kenya, Malawi, Mauritius, Zimbabwe) with the Quadratic Spectral kernel are also non stationary. All in all, the C-PPP is validated significantly only for 3 countries (Eritrea, Sudan, Swaziland). Then, C-PPP holds just for 17.6% of the sample.

Secondly, the null hypothesis of stationarity around a constant trend is strongly rejected in 7 (Djibouti, Egypt, Libya, Seychelles, Tanzania, Zambia, Zimbabwe) with the Bartlett kernel, and in 9 (Djibouti, Egypt, Eritrea, Kenya, Libya, Seychelles, Tanzania, Zambia, Zimbabwe) with the Quadratic Spectral kernel, out of 17 cases at the 5% significance level. Considering the 10% level, we have to add 3 other countries (Comoros, Eritrea, Kenya) with the Bartlett kernel and 2 others (Comoros, Ethiopia) with the Quadratic Spectral kernel. So, only 6 countries (Madagascar, Malawi, Mauritius, Uganda, Sudan, Swaziland) are stationary for all conditions considered. Finally, BS-PPP holds for 35.3% of the sample.

Thirdly, except for Djibouti, all the series of the panel have a significant individual time trend component. Thus, even when the stationarity is found, only the Balassa-Samuelson approach of long-run PPP can be at best supported.

3.2. The standard panel stationary test

However, as usually noted, the main drawback of the individual KPSS test is its low power, and especially in short samples, which can result in rejecting too easily the null hypothesis of stationarity. So, the standard panel stationary test of Hadri (2000) (we use both the homogeneous and the heterogeneous tests) is applied on the set of Eastern and Southern African bilateral real exchange rates.

This one is similar to the KPSS test, and has a null hypothesis of no unit root in any of the series in the panel. Like the KPSS test, the Hadri test is based on the residuals from the individual Ordinary Least Squares (OLS) regressions of $y_{i,t}$ on an intercept (model 1), or on an intercept and a time trend (model 2). So the test is built on the general specification:

(1) $$y_{i,t} = \alpha_i + \beta_i t + \varepsilon_{i,t}$$

Then, given the residuals $\hat{\varepsilon}$ from the individual regressions, the LM statistics can be formed:

(2) $$LM = N^{-1}\sum_{i=1}^{N}\left(\hat{\sigma}_{\varepsilon}^{-2}T^{-2}\sum_{t=1}^{T}\hat{S}_{i,t}^{2}\right)$$ for the homogeneous case

(3) $$LM = N^{-1}\sum_{i=1}^{N}\left(\hat{\sigma}_{\varepsilon,i}^{-2}T^{-2}\sum_{t=1}^{T}\hat{S}_{i,t}^{2}\right)$$ for the heterogeneous case

where $\hat{S}_{i,t} = \sum_{j=1}^{t}\hat{\varepsilon}_{i,j}$ is the partial sum of the residuals, with $\hat{\sigma}_{\varepsilon,i}^{2}$ being a consistent estimator of the long-run variance ($\sigma_{\varepsilon,i}^{2}$) of $\varepsilon_{i,t}$ under the null hypothesis[7], and $\hat{\sigma}_{\varepsilon}^{2} = N^{-1}\sum_{i=1}^{N}\hat{\sigma}_{\varepsilon,i}^{2}$. Lastly, Hadri shows that under mild assumptions it is possible to obtain a Z-statistic like:

(4) $$Z = \frac{\sqrt{N}(LM - \xi)}{\zeta} \xrightarrow{d} N(0,1)$$

where $\xi = \frac{1}{6}$ and $\zeta = \frac{1}{45}$, if the model only includes intercepts, and $\xi = \frac{1}{15}$ and $\zeta = \frac{11}{6300}$, otherwise[8].

Table 3: The standard panel stationary tests for Eastern and Southern Africa, 1970-2007

	Model			
Hadri tests	**Individual intercept (model 1)**		**Individual trend & intercept (model 2)**	
	Bartlett (Prob.)	**Quadratic Spectral (Prob.)**	**Bartlett (Prob.)**	**Quadratic Spectral (Prob.)**
Z-stat	14.254 (0.000)**	14.188 (0.000)**	9.049 (0.000)**	8.613 (0.000)**
Heteroscedastic Consistent Z-stat	9.951 (0.000)**	9.754 (0.000)**	5.885 (0.000)**	5.030 (0.000)**

Notes: Figures in () give the probability of rejecting the null hypothesis of stationarity in the whole panel. (**) and (*) indicate the reject of the null hypothesis of panel stationarity at the 5% and 10% level, respectively. The long-run variance is estimated using both the Bartlett and the Quadratic Spectral kernel with automatic spectral window bandwidth selection as in Newey-West (1994).

[7] A consistent estimator of the long-run variance can be obtained using one of the HAC estimators in the market. In this work, we have implemented the estimator of Newey and West (1994).

[8] Where $\xrightarrow{d}$ denotes weak convergence in distribution.

The results for the two specifications are given in Table 3. On the whole, the standard panel tests corroborate the univariate KPSS test in the extent that the null hypothesis of stationarity is strongly rejected whatever the Hadri version retained and the kernel method used. Thus, in accordance with the panel data stationary tests, we can conclude that PPP does not hold for the Eastern and the Southern African countries. Moreover, long-run PPP is rejected for the two models, indicating that both C-PPP and BS-PPP are not valid.

4. The panel stationary test with multiple structural breaks

4.1. The CDL stationary test

The specific time period covered by the variables, on the one hand, and the information shown in the graphs (see Figure 1[9]), on the other hand, indicate that there might be some structural breaks affecting the time series. Nevertheless, no considering the presence of structural breaks in the series can deteriorate the results obtained from standard procedures. Indeed, as pointed out in Huang *et al.* (1997), the presence of structural breaks affects the limit distribution of the univariate KPSS so that they should be taken into account when testing the stationarity of the panel.

To address the multiple structural breaks problem, this study employs a new panel testing procedure based on Carrion-i-Silvestre *et al.* (2005). Following the test of Hadri (2000), this new test still considers the null hypothesis of stationarity for all cross-sections but the influence of structural breaks is taken into account in a very convenient way. Indeed, the procedure is general enough to allow the following characteristics: *(i)* the structural breaks can have different effects on each individual time series, *(ii)* they can be located at different dates and *(iii)* individuals can have different number of structural breaks.

Consider the following regressions which encompass $i = 1,...,N$ individuals and $t = 1,...,T$ time periods:

(5) $$y_{i,t} = \alpha_{i,t} + \beta_i t + \varepsilon_{i,t}$$

and

[9] The graphs are displayed in appendix to save space.

(6) $$\alpha_{i,t} = \sum_{k=1}^{m_i} \theta_{i,k} D(T_{b,k}^i)_t + \sum_{k=1}^{m_i} \gamma_{i,k} DU_{i,k,t} + \alpha_{i,t-1} + \upsilon_{i,t}$$

where $\upsilon_{i,t}$ ~i.i.d.$(0, \sigma_{\upsilon,i}^2)$ and $\alpha_{i,0} = \alpha_i$, a constant. The dummy variables $D(T_{b,k}^i)_t$ and $DU_{i,k,t}$ are defined as $D(T_{b,k}^i)_t = 1$ for $t = T_{b,k}^i + 1$ and 0 elsewhere, and $DU_{i,k,t} = 1$ for $t > T_{b,k}^i$ and 0 elsewhere, with $T_{b,k}^i$ giving the k th date of the break for the i th individual, $k = 1,...,m_i$, $m_i \geq 1$. Moreover, note that the stochastic processes {$\varepsilon_{i,t}$} and {$\upsilon_{i,t}$} are taken to be mutually independent across the two dimensions of the panel data set. So, if we state the condition $\sigma_{\upsilon,i}^2 = 0$ for all $i = 1,...N$, *i.e.* the null hypothesis of a stationary panel, substituting (6) in (5) results in:

(7) $$y_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{i,k,t} + \beta_i t + \sum_{k=1}^{m_i} \gamma_{i,t} DT_{i,k,t}^* + \varepsilon_{i,t}$$

with the dummy variable $DT_{i,k,t}^* = t - T_{b,k}^i$ for $t > T_{b,k}^i$ and 0 elsewhere, $k = 1,...,m_i$, $m_i \geq 1$.

Then, the test of the null hypothesis of a stationary panel follows the proposal of Hadri (2000), who designed a test statistic that is simply the average of the univariate stationary test in KPSS. The general expression for the test statistic is:

(8) $$LM(\lambda) = N^{-1} \sum_{i=1}^{N} \left(\hat{\omega}^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}_{i,t}^2 \right)$$ for the homogeneous case

(9) $$LM(\lambda) = N^{-1} \sum_{i=1}^{N} \left(\hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}_{i,t}^2 \right)$$ for the heterogeneous case

where $\hat{S}_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ denotes the partial sum process that is obtained using the estimated OLS residuals of (7), with $\hat{\omega}_i^2$ being a consistent estimate of the long-run variance[10] of $\varepsilon_{i,t}$,

[10] Carrion-i-Silvestre *et al.* (2005) suggest that the long-run variance estimate should be obtained as in Sul *et al.* (2003) with the specification of an AR(p) process for the prewithening and using both the quadratic and the Bartlett spectral window with automatic bandwidth selection as indicated in Andrews (1991), Andrews and

$\omega_i^2 = \lim_{T\to\infty} T^{-1}E(S_{i,T}^2), i = 1,...,N$, and $\hat{\omega}^2 = N^{-1}\sum_{i=1}^{N}\hat{\omega}_i^2$. Note that λ is used in (8) and (9) to denote the dependence of the test on the dates of break. For each individual i, it is defined as the vector $\lambda_i = (\lambda_{i,1},...,\lambda_{i,m_i})' = (T_{b,1}^i/T,...,T_{b,m_i}^i/T)$ which indicates the relative positions of the dates of the breaks on the entire time period T. Finally, by defining $\overline{\xi} = N^{-1}\sum_{i=1}^{N}\xi_i$ and $\overline{\zeta}^2 = N^{-1}\sum_{i=1}^{N}\zeta_i^2$, with ξ_i and ζ_i^2 the individual mean and variance of $\hat{\omega}_i^{-2}T^{-2}\sum_{t=1}^{T}\hat{S}_{i,t}^2$, respectively, Carrion-i-Silvestre *et al.* (2005) show that under mild assumptions the test statistic for the null hypothesis of a stationary panel with multiple shifts is :

$$(10) \qquad Z(\lambda) = \frac{\sqrt{N}(LM(\lambda) - \overline{\xi})}{\overline{\zeta}} \xrightarrow{d} N(0,1)$$

In addition, note that Carrion-i-Silvestre *et al.* (2005) demonstrate that this test gives good performance in finite samples by Monte Carlo simulations.

4.2. The determination of the breaks

At this stage, the break fraction vector has been considered as given. However, this latter is usually unknown and must therefore be estimated. Consequently, computing the test statistic requires to detect the breaks in each one of the individual time series as a preliminary step. In this regard, we apply the procedure implemented by Bai and Perron (1998) which computes the global minimization of the sum of squared residuals. Then, the estimate of the dates of the breaks results from the argument that minimizes the sequence of individual sum of squared residuals $S(T_{b,1}^i,...,T_{b,m_i}^i)$ computed from (7):

$$(\hat{T}_{b,1}^i,...,\hat{T}_{b,m_i}^i) = \underset{(T_{b,1}^i,...,T_{b,m_i}^i)}{\arg\min}\; S(T_{b,1}^i,...,T_{b,m_i}^i)$$

Monahan (1992) and Sul. *et al.* (2003). The order of the autoregressive process used in the prewithening stage is fixed using the BIC information criterion with a maximum of 4 lags.

Once the dates for all possible $m_i \leq m^{\max}$, $i = 1,...,N$, have been estimated, a crucial step is to find the exact number of structural changes, *i.e.* the optimal m_i. Several procedures for selecting the dimension of a model are available (Jouini and Boutahar, 2005): *(i)* those testing whether a fixed value of m_i corresponds to the real structural changes number (the application of pseudo F-type test statistics) and *(ii)* those where an *a priori* value of m_i is not needed (the use of information criteria). However, Bai and Perron (2001) compare the procedures and conclude that the second ones present better performance than the first ones. Thus, following this finding, we chose to select the optimal number of breaks according to the Bayesian Information Criterion (BIC). The results are displayed in Table 4.

Table 4: The estimates of the break dates for Eastern and Southern African countries, 1970-2007

	Model			
	Individual intercept (model 1)		**Individual trend & intercept (model 2)**	
	Number	**Break dates**	**Number**	**Break dates**
Comoros	0	..	2	1993; 1996
Djibouti	2	1995; 2001	4	1994; 1997; 2000; 2003
Egypt	0	..	3	1977; 1996; 1999
Erithrea	0	..	5	1972; 1976; 1989; 1992; 1995
Ethiopia	2	1994; 2000	3	1993; 1996; 2000
Kenya	1	1975	0	..
Libya	2	1993; 2002	4	1991; 1994; 1998; 2002
Madagascar	2	1993; 1999	4	1992; 1995; 1999; 2003
Malawi	0	..	3	1990; 1994; 1997
Mauritus	2	1992; 2002	4	1991; 1994; 1998; 2002
Uganda	2	1994; 2000	4	1975; 1993; 1996; 1999
Seychelles	0	..	3	1974; 1993; 1996
Sudan	0	..	4	1990; 1993; 1997; 2001
Swaziland	2	1992; 2001	3	1983; 1991; 1995
Tanzania	0	..	3	1993; 1996; 2000
Zambia	2	1993; 1999	4	1992; 1995; 1999; 2003
Zimbabwe	0	..	2	1994; 1997

Notes: The number of break points has been estimated using the BIC information criteria allowing for a maximum of $m^{\max} = 6$ structural breaks. The trimming is taken on the interval [0.1T, 0.9T].

We can remark than most of the break dates occurs from the beginning of the 1990s. Actually, this finding is not surprising. Indeed, for much of the past two decades, African countries have been implementing structural and macroeconomic adjustment programmes designed to improve the external competitiveness and economic growth of these economies (Kargbo, 2004). In particular, exchange rate policy reforms were the focal point of the adjustment programmes. So, since the end of the 1980s and the early 1990s, there were several waves of depreciation of domestic currencies in order to stop the extensive overvaluation of exchanges rates in Africa during the 1970s and the 1980s.

4.3. The empirical results

So, it is evident that structural changes are present in the data. Then, the conclusions derived from the Hadri tests are biased and we have to implement the CDL tests. Moreover, some preliminary remarks must be stated. Firstly, the CDL tests are done with the optimal number of break dates given by the BIC information criteria[11]. Secondly, these tests suppose that individuals are cross-section independent. However, this assumption is rarely found in practice, especially in a free trade area as COMESA where the shocks overpass the borders of the economies. Moreover, cross-sectional dependence is expected in panels on real exchange rates if a common currency such as the US dollar is used as a base. Unfortunately, this problem can have negative consequences on the statistical properties of panel unit root tests, raising the significance levels of tests with a nominal size of 5% to as much as 50% (O'Connell, 1998)[12]. Thus, following Maddala and Wu (1999), we have computed the bootstrap distribution in order to take into account for cross-sectional dependence of the statistics[13].

Finally, Table 5 shows that the allowance of structural breaks changes the findings obtained from the standard stationary tests concerning the specification with a deterministic trend (model 2). Indeed, the null hypothesis of panel stationarity cannot be rejected at the 10% level of significance for both the Bartlett and Quadratic spectral kernel regardless of the assumption concerning the heterogeneity in the long-run variance estimate. However, the null is rejected at the 5% level for the homogeneous case. Nevertheless, as mentioned above, we have to compare the statistics with the critical values of the bootstrap distribution. This allows

[11] The LWZ criteria corroborate the results given by the BIC criteria.
[12] The implication is that studies not accounting for cross-sectional dependence are likely to falsely reject a unit root.
[13] The details of the bootstrap method can be found in Maddala and Wu (1999) with 2,000 replications.

us to find that we cannot reject the null hypothesis of stationarity at the 10% and at the 5% levels of significance, whatever the assumption about the kernel specification and the nature of the long-run variance. Therefore, taking into account the presence of structural changes leads to the acceptance of the long-run BS-PPP. More precisely, we put forward a special type of the BS-PPP which we call the “quasi Balassa-Samuelson PPP”, that is to say the reversion in the real exchange rate around a broken or a changing time trend.

On the contrary, the allowance of structural breaks does not change the evidence shown by the standard stationary tests concerning the model without a time trend (model 1) significantly, since the null hypothesis of panel stationarity still can be rejected for almost all conditions. Thus, the long-run C-PPP does not hold. Nevertheless, when we look at the critical values of the bootstrap distribution, the heterogeneous statistic supports the C-PPP. Finally, concerning the hypothesis of reversion to a changing mean, the evidence is mixed.

Table 5: The panel stationary tests with multiple structural breaks for Eastern and Southern Africa, 1970-2007

	Model			
	Individual intercept (model 1)		**Individual trend & intercept (model 2)**	
	Bartlett (Prob.)	**Boostrap critical values 10% (5%)**	**Bartlett (Prob.)**	**Boostrap critical values 10% (5%)**
Z(λ)-stat	4.792 (0.000)**	3.191 (4.569)**	1.643 (0.050)*	4.299 (5.285)
Heteroscedastic Consistent Z(λ)-stat	2.019 (0.022)**	3.317 (4.401)	0.154 (0.439)	3.933 (4.647)
	Model			
	Individual intercept (model 1)		**Individual trend & intercept (model 2)**	
	Quadratic spectral (Prob.)	**Boostrap critical values 10% (5%)**	**Quadratic spectral (Prob.)**	**Boostrap critical values 10% (5%)**
Z(λ)-stat	4.797 (0.000)**	3.435 (4.719)**	1.648 (0.050)*	4.054 (5.368)
Heteroscedastic Consistent Z(λ)-stat	2.091 (0.018)**	3.359 (4.808)	0.240 (0.405)	3.750 (4.664)

Notes: (**) and (*) indicate the reject of the null hypothesis of panel stationarity at the 5% and 10% level, respectively. The long-run variance is estimated using both the Bartlett and the Quadratic Spectral kernel with automatic spectral window bandwidth selection as in Andrews (1991), Andrews and Monahan (1992) and Sul *et al.* (2003). The bootstrap distribution is based on 2,000 replications.

5. Conclusion

African countries, in general, and the members of COMESA, in particular, have been in the process of broadening and deepening their level of regional and international integration in trade, financial markets and fiscal affairs. Exchange rate policy has been the centrepiece of macroeconomic adjustment programmes in Africa during much of the past two decades. This policy, designed to improve the external competitiveness of African economies, is based on the assumption that PPP holds in Africa (Kargbo, 2003). However, the empirical studies lead to mixed results about the evidence of PPP hypothesis.

So, this article implemented, with standard stationary methods, the new stationary test of Carrion-i-Silvestre *et al.* (2005) to determine whether or not there is support for long-run PPP in a panel of 17 Eastern and Southern African countries over the period 1970-2007. This procedure has the decisive advantage to allow for the presence of multiple structural breaks both in the mean and/or in the trend of the individual time series.

Finally, the results indicated that taking into account the presence of structural breaks changes dramatically the conclusion about long-run PPP. Indeed, in contrast with standard stationary tests, we showed that PPP holds in the long-run for our panel when the Balassa-Samuelson version is retained, *i.e.* the real exchange rate is trend-stationary but around a broken trend. In other words, in this paper, we validated the “quasi” PPP alternative, as in hegwood and Papell (1998), but with a deterministic trend.

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APPENDIX

Figure 1: The bilateral real exchange rates for Eastern and Southern Africa, 1970-2007