REAL EXCHANGE RATE SHOCKS, ASYMMETRIC ADJUSTMENT AND LONG-RUN EQUILIBRIUM IN LESS DEVELOPED COUNTRIES

by

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Abstract

This paper investigates the possibility that long-run relative purchasing power parity is dependent upon the nature of real exchange shocks that are experienced. While existing studies involving developed and less developed countries often find against purchasing power parity having employed linear tests of non-stationarity or non-cointegration, we employ a new cointegration test, recently advocated by Enders and Siklos and Enders and Dibooglu, that tests for an asymmetric adjustment towards parity with respect to positive and negative real exchange rate shocks. Using a sample of ten African economies with data taken from the post-Bretton Woods floating exchange rate era, long-run purchasing power parity holds in eight of these cases if an explicit distinction is made between positive and negative shocks. Across the sample, we find variation in the type of asymmetry experienced and the roles played price and nominal exchange rate adjustment.

JEL Codes: C5, F3, F4.

1. Introduction

Policy makers in less developed countries (LDCs) have a potential interest in purchasing power parity (PPP) for a number of reasons. First, PPP becomes a prediction model for exchange rates and a criterion for judging over- and undervaluation of currencies. This may be particularly relevant for small open economies and those experiencing large inflation differentials between domestic and foreign inflation rates. Second, many exchange rate theories employ some notion of PPP in their construction. Thus the quality of policy advice, insofar as it is based on these theories, may depend on the validity of PPP. However, the evidence on PPP for developed countries and LDCs has generally provided ambiguous results without a conclusive answer, for example Balassa (1964) and Hakkio (1984) find in favor of PPP while Dornbusch (1980) and Frenkel (1981) find no evidence in favor of PPP. However, Frenkel (1978) suggests that PPP holds during periods of high inflation. More specific evidence concerning LDCs is also mixed [see, *inter alia*, McNown and Wallace (1989), Liu (1992), Bahmani-Oskooee (1993), Mahdavi and Zhou (1994), Holmes (2001), Nagayasu

(2002)]. In terms of methodology, the recent studies of long-run PPP in LDCs have utilised augmented Dickey Fuller (ADF) tests for unit roots in real exchange rates and cointegration between various measures of domestic and foreign prices and nominal exchange rates. Typically, the null of non-stationarity or non-cointegration is accepted by the data and therefore purchasing power parity is rejected. The conclusions drawn for LDCs in these studies are primarily based on *linear* tests for stationarity or cointegration.[1] However, evidence of asymmetries in key economic variables has been established for developed economies in recent years. For example, Ramsey and Rothman (1996) and Verbrugge (1998) identify asymmetries in inflation and attribute them to downward price rigidities. Cover (1992), Rhee and Rich (1995), Karras (1996) and Madsen and Yang (1998) offer more general empirical evidence that corroborates the implications of price adjustment models where prices are primarily sticky in a downward direction. Studies by van Dijk and Franses (1997), Enders and Granger (1998), Enders and Siklos (1999) find evidence in asymmetries in nominal interest rates while Coakley and Fuertes (2002) consider real interest rates. Enders and Dibooglu (2001) identify asymmetries in OECD real exchange adjustment towards PPP. Since the vast majority of the existing work on PPP in LDCs is based on *linear* tests for mean-reversion in real exchange rates, no allowance is made for the type of shock (positive or negative) that is experienced.

This paper offers an important contribution to the PPP debate by considering whether long-run PPP holds for LDCs against a background of asymmetric responses to positive and negative real exchange rate shocks. When considering the long-run adjustment towards PPP, one can point to the presence of transactions costs that inhibit international goods arbitrage, product markets that are characterised by extensive government intervention through

[1] In an attempt to overcome low test power, the studies by Holmes (2001) and Nagayasu (2002) employ (linear) panel data unit root and cointegration tests. Whereas these studies identify PPP for various LDC panels as a whole, there is little evidence to suggest that anything more than a few countries are responsible for rejecting the nulls of non-stationarity or non-cointegration.

taxation, price setting and public trading monopolies [Collier and Gunning (1999)] and the reluctance of central banks to facilitate depreciation of the nominal exchange rate in a regime of managed floating [Hossain and Chowdhury (1998)]. These factors may be relevant to how real exchange rate adjustment responds to positive and negative shocks.[2] Using quarterly data for the period 1973Q2 to 2002Q1, ten African countries are analysed using a new test for asymmetric adjustment towards PPP advocated by Enders and Siklos (2001) and Enders and Dibooglu (2001). Our analysis indicates that PPP is confirmed where the speed of adjustment towards equilibrium depends on the nature of the shock that the economy experiences.

The paper is organised as follows. The following section describes the nature of asymmetries in adjustment towards PPP and the methodological approach. This approach enables us to formally test for cointegration distinguishing between the roles played by positive and negative real exchange rate shocks. The third section describes the data. The fourth section discusses the results and the final section concludes.

2. Cointegration, Asymmetric Adjustment and PPP

Let the natural logarithm of the domestic price index (p) be and determined in the long-run as

$$p_t = \beta_0 + \beta_1 p_t^* + \beta_2 e_t + \mu_t \qquad (1)$$

where p^* is the natural logarithms of the foreign price index, e is the nominal exchange rate measured as the domestic price of foreign currency and μ is the error term. For a number of reasons the use of this estimating equation is appropriate for testing *relative* rather than *absolute* long-run PPP. First, empirical studies of PPP typically employ data on price *indices* rather than price *levels*. Price indices contain base periods where the nominal exchange rate

[2] In addition to this, the presence of asymmetries in real exchange rates can be attributed to the heterogeneity of participants in the foreign exchange market in terms of agents' expectations formation or investors' objectives (see Sarantis (1999) and references contained therein).

can equal the price ratio by construction. Therefore, a test for a unit root in μ_t is in fact a test for a unit root in the change from base period. This is the usual test of long-run relative PPP in the literature arguing that the percentage change in the nominal spot exchange rate should equal the inflation differential between country i and the base country.[3] Second, the actual exchange rate may deviate from its parity value on account of imperfections in published prices indices (for example, in reality the price indices of different countries do not reflect the same basket of goods). Third, deviations from absolute PPP may occur on account of transport costs, tariffs and differential speeds of adjustment in the goods and foreign exchange markets. Assessing relative PPP allows for any constant of proportion based on these factors that drives a wedge between p_t and $e_t + p_t^*$. For these reasons, we do not impose the strict conditions $\beta_0 = 0$, $\beta_1 = 1$ and $\beta_2 = 1$ but *a priori* we would expect $\beta_1 > 0$ and $\beta_2 > 0$. The usual test for linear adjustment towards PPP is based assessing the unit root properties of μ_t through the estimation of

$$\Delta\mu_t = \rho\mu_{t-1} + \sum_{i=1}^{k} \psi_i \Delta\mu_{t-i} + \varepsilon_t \tag{2}$$

where ε_t is a white noise residual. However, the adjustment towards PPP may in fact be asymmetric and can be justified on a number of grounds. For a given country, suppose there is positive shock to PPP with $\mu_t > 0$ in equation (1) that represents a decline in competitiveness. Adjustment back towards PPP can be achieved through a rise in *e* (nominal depreciation), an increase in p^*, and/or a decrease in *p*. Conversely, a negative shock in equation (1) can be represented by $\mu_t < 0$ or an increase in competitiveness. This time, adjustment will require a decrease in *e* (nominal appreciation), a decrease in p^*, and/or an increase in *p*. Agents in the economy may have different attitudes towards increases and

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3 See Crownover, Pippenger and Steigerwald (1996) for an elaboration on this point.

decreases in e and p. For example, some central banks may view devaluations of the nominal exchange rate as being inflationary with no long-run beneficial effects on the real exchange rate.[4] There may be political stigma attached to devaluation [see, for example, Aghevli *et al.* (1991)]. On the other hand, an overvalued nominal exchange rate may be seen as detrimental to the trade balance. In the case of price adjustment, there is the possibility of price controls or simply the relative downward rigidity with respect to p even in the absence of price controls. For these reasons, it is possible that the speed of adjustment back towards PPP depend on the sign of the shock to μ. When testing for cointegration and long-run PPP in equation (1), it therefore seems appropriate to explicitly distinguish between positive and negative real exchange rate shocks accordingly,

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \varepsilon_t \tag{3}$$

where the stationarity of μ is achieved where $-2 < \rho_1, \rho_2 < 0$ and I_t is the Heaviside indicator with the following function

$$I_t = \begin{cases} 1 & if \ \mu_{t-1} \geq 0 \\ 0 & if \ \mu_{t-1} < 0. \end{cases} \tag{4}$$

On the basis of the above arguments, we might expect $\rho_1 \neq \rho_2$. In the case of goods market arbitrage, transactions costs may straddle the equilibrium value of PPP so that as one moves further away from central parity, arbitrage becomes feasible (see, *inter alia*, Obstfeld and Taylor (1997). We may therefore define a threshold (τ) below which adjustment towards PPP will not occur. Making the appropriate adjustment to the Heaviside indicator means that

$$I_t = \begin{cases} 1 & if \ \mu_{t-1} \geq \tau \\ 0 & if \ \mu_{t-1} < \tau. \end{cases} \tag{5}$$

[4] Few studies have identified a long-run relationship between the nominal and real exchange rate of LDCs [see, *inter alia*, Connolly and Taylor (1976, 1979), Donovan (1981), Bautista (1981), Morgan and Davis (1982), Edwards (1988, 1994)]. However, recent evidence on LDCs from Bahmani-Oskooee and Miteza (2002) argues that devaluations of the nominal exchange rate do actually result in devaluations of the real exchange rate.

Finally, it might be necessary to include lags of $\Delta\mu_t$ to ensure that ε_t is white noise.

$$\Delta\mu_t = I_t \rho_1 \mu_{t-1} + (1 - I_t)\rho_2 \mu_{t-1} + \sum_{i=1}^{k} \psi_i \Delta\mu_{t-i} + \varepsilon_t \qquad (6)$$

Enders and Dibooglu (2001) and Enders and Siklos (2001) also advocate testing for asymmetric cointegration using the momentum threshold autoregressive (MTAR) model. This requires the amendment of equation (5) so that

$$I_t = \begin{cases} 1 & if\ \Delta\mu_{t-1} \geq \tau \\ 0 & if\ \Delta\mu_{t-1} < \tau. \end{cases} \qquad (7)$$

It is argued that this model is appropriate if μ exhibits more momentum in one direction than another. Rather than adopt a methodology based on the application of smooth transition autoregressive (STAR) modelling to real exchange rate behaviour (see, for example, Terasvirta (1994) and Sarantis (1999)), this study investigates adjustment towards long-run PPP through the employment of a threshold cointegration approach. For the purposes of testing for asymmetries in a long-run cointegrating relationship, two tests are proposed. The first is the *F*-test for cointegration in equation (6) where the null is specified as $\rho_1 = \rho_2 = 0$ and the test statistic is defined as Φ_μ**.** This test follows a non-standard distribution and we employ the critical values reported in Enders and Dibooglu (2001), Table 1. The second test focuses on the possibility of asymmetries within the cointegrating relationship and is on $\rho_1 = \rho_2$. This is a standard *F*-test. The estimation of equation (6) requires the consistent estimation of τ. This is achieved through a grid search procedure for τ where the boundaries are defined between the largest and smallest values of μ_t. The chosen value for τ is the one that minimises the residual sum of squares during the OLS estimation of (6).

3. Data and Results

The empirical analysis involves the following ten African countries: Botswana, Burundi, Cameroon, Egypt, Ivory Coast, Kenya, Morocco, Nigeria, South Africa and Tanzania. All price and exchange rate data are taken from the *International Financial Statistics* database. The price series are based on the consumer price index (line 64) and the nominal exchange rate data are end of period spot rates with respect to the US dollar (domestic price of the U.S. dollar). Quarterly data for the period 1973Q2-2002Q2 provide a sample size of 117 observations on each series for each country where the use of quarterly data is dictated by data availability across this large sample.[5] The start of 1973 is consistent with Bahmani-Oskooee (1993), Mahdavi and Zhou (1994) and Holmes (2001) in their investigations of PPP in LDCs and can be regarded as the start of modern "floating rate" period with respect to the US dollar.

Table 1 reports the application of the Engle-Granger procedure to equations (1) and (2). For each country, the lag length k was selected using the Akaike Information Criteria (AIC). We are unable to reject the null of non-cointegration at the 10% significance level in all cases. As argued above, the general absence of long-run PPP in these initial tests might be attributable to the employment of linear tests for mean reversion whereas there are in fact asymmetries in any adjustment towards long-run PPP with respect to positive and negative shocks. Moreover, these tests for symmetric cointegration have low test power against a background of asymmetric adjustment. To accommodate this, we set the indicator function I_t as shown in equations (5) for the TAR model and (7) for the MTAR model and estimate equation (6). Given the presence of measurement errors and/or adjustment costs, there is no reason to presume that the threshold is identically equal to zero. To estimate the threshold, we use the grid search method described above which provides a consistent estimate of the threshold. Table 2 reports the results. Using the AIC model selection criterion, the MTAR model is favored in the cases of Botswana, Burundi, Ivory Coast, and Morocco. A major difference from the results previously reported in Table 1 is that the case for cointegration and long-run PPP is substantially strengthened when asymmetries are allowed for because in eight out of ten cases, asymmetric cointegration is now confirmed. The Φ_{μ} statistic indicates that the null $\rho_1 = \rho_2 = 0$ is rejected at the 5% significance level or better and each of these cases and this is supported by the rejection of the null $\rho_1 = \rho_2$. In all cases except Cameroon, South Africa and Tanzania there is evidence that $|\rho_1| > |\rho_2|$ implying that the speed of adjustment towards long-run relative PPP is faster in the case of a positive shock with respect to μ. The value of the estimated threshold τ, expressed as a percentage of the mean dependent variable, assumes a range of values. For most countries it is less than 3% though Nigeria and Tanzania are characterised by larger thresholds suggesting that adjustment towards PPP will only occur with respect to large shocks.

The final part of our investigation is to estimate asymmetric error-correction models that are based on positive and negative shocks to equilibrium previously employed in the earlier results. The purpose of this procedure is to examine how the burden of adjustment towards long-run equilibrium is shared between the adjustment of domestic prices and the nominal exchange rate. Specifically, we estimate the following system of asymmetric error correction models for each country,

$$\Delta p_t = A_{11}(L)\Delta p_{t-1} + A_{12}(L)\Delta p^*_{t-1} + A_{13}(L)\Delta e_{t-1} + \rho_{11} z^+_{t-1} + \rho_{12} z^-_{t-1} \tag{8}$$

$$\Delta e_t = A_{21}(L)\Delta p_{t-1} + A_{22}(L)\Delta p^*_{t-1} + A_{23}(L)\Delta e_{t-1} + \rho_{21} z^+_{t-1} + \rho_{22} z^-_{t-1} \tag{9}$$

$$\Delta p^*_t = A_{31}(L)\Delta p_{t-1} + A_{32}(L)\Delta p^*_{t-1} + A_{33}(L)\Delta e_{t-1} + \rho_{31} z^+_{t-1} + \rho_{32} z^-_{t-1} \tag{10}$$

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[5] Data limitations prevent the use of wholesale price or effective exchange rate indices across all these countries.

where $z_t^+ = I_t \mu_t$, $z_t^- = (1 - I_t)\mu_t$, μ_t is the residual from the long-run equation (1), I_t is the indicator defined in the previous section, and $A_{ij}(L)$ is a polynomial in the lag operator L. The choice of the appropriate lag length is based on the multivariate AIC. The choice of non-zero threshold follows the same procedure outlined earlier. Table 3 reports the error-correction equations. In the cases of the Cameroon, Morocco, Nigeria, South Africa and Tanzania, ρ_{12} is significantly different from zero at the 5% significance level or better. This is consistent with the upward adjustment of p in order to help facilitate the movement towards weak PPP in the event of a negative shock to μ in equation (1) described above. In each of these countries, we also find that ρ_{11} is insignificantly different from zero which is consistent with the downward rigidity of domestic prices that inhibits any move towards long-run equilibrium following a positive shock to μ. In the cases of the Egypt, Ivory Coast and Kenya, a different set of circumstances applies. The Ivory Coast and Kenya are characterised by ρ_{11} rather than ρ_{12} which is negative and significant. For these countries, domestic prices exhibit upward rather than downward price rigidity in any adjustment towards long-run PPP. In the case of Egypt, the emphasis is on nominal exchange rate adjustment because we find that both ρ_{11} and ρ_{12} are insignificantly different from zero and so domestic prices play no significant role in any adjustment. The results reported in Table 3 highlight the more general role played by nominal exchange rate adjustment. We find that ρ_{21} is positive and significant in the cases of Egypt, Ivory Coast, Kenya, Nigeria and Tanzania where a positive shock to μ and decline in competitiveness will result in a depreciation of the nominal exchange rate. In the cases of Egypt and South Africa, ρ_{22} is positive and significant indicating an appreciation of nominal exchange rate in the event of a negative shock to μ. However, the dynamics of nominal exchange rate adjustment are complex. With $1 < \rho_{21} < 2$ in the cases of Egypt and Tanzania, there is evidence of overshooting on the part of the nominal exchange rate in any adjustment towards PPP equilibrium. The *t* statistics indicate that $\rho_{21} < 0$ in the case of Morocco and $\rho_{22} < 0$ in the cases of the Cameroon, Morocco and Nigeria. Although asymmetric cointegration was confirmed earlier in Table 2, these results point to potentially destabilising behaviour on the part of the nominal exchange rate. Governments may prefer to use the nominal exchange to protect any gains in competitiveness following a negative shock to PPP, or may be wary of allowing a nominal depreciation and any inflationary consequences following a positive shock to PPP. Finally, the estimates for ρ_{31} and ρ_{32} indicate that there is little in the way of p^* adjustment to facilitate long-run PPP. At the 5% significance level, we can only identify $\rho_{32} \neq 0$ in one case. This supports the view that the asymmetric adjustment towards long-run PPP relies on domestic prices and the nominal exchange rate rather than U.S. prices.

4. Summary and Conclusion

The majority of empirical work that addresses whether or not PPP holds in LDCs is based on linear tests for a long-run relationship between domestic prices, foreign prices and the nominal exchange rate. Most of these studies have been generally unsupportive of PPP. However, it can be argued that the sign of any disturbance to PPP is important with respect to any adjustment towards long-run equilibrium. Depending on upward and downward rigidities with respect to domestic prices and the nominal exchange rate, it is reasonable to argue that the speed of adjustment back towards long-run equilibrium will depend on the sign of shock. This paper explores the possibility of asymmetric adjustment towards long-run PPP for a sample of ten African countries over a thirty year study period. In answer to the question of whether or not long-run PPP holds in African countries, we find in favour of asymmetric PPP. Initial tests for linear cointegration are unable to find in favour of PPP. However, using a new test advocated by Enders and others, asymmetric cointegration is confirmed where the speed of adjustment back towards long-run equilibrium will depend on the whether the shock was positive or negative. Estimation of the error correction models indicates that most of the sample are characterised by downward price rigidity where domestic prices only respond to a negative real exchange rate shock. However, there is evidence that adjustment back towards long-run equilibrium is shared between domestic prices and the nominal exchange rate. Avenues for future research might include country-specific investigations into the precise channels through which adjustment occurs following external shocks.

Table 1. Tests for Symmetric Cointegration

Country	β_0	β_1	β_2	ADF
Botswana	-1.892	1.250	0.587	-1.959
Burundi	-4.188	1.163	0.625	-2.359
Cameroon	-2.914	1.384	0.175	-2.921
Egypt	-4.221	1.776	0.459	-2.442
Ivory Coast	-2.801	1.368	0.166	-2.988
Kenya	-3.313	0.929	0.891	-3.019
Morocco	-1.531	1.268	0.105	-3.112
Nigeria	-6.086	1.956	0.602	-1.845
South Africa	-2.829	1.422	0.553	-2.004
Tanzania	-8.293	1.972	0.590	-1.885

Notes. These results are based on the OLS estimation of equation (1) followed by ADF unit root tests defined in equation (2).

Table 2. Tests for Asymmetric Cointegration

Country	Model	Γ	ρ_1	ρ_2	Φ_μ [a]	$F_{\rho_1=\rho_2}$ [b]
Botswana	MTAR	0.000	-0.058	-0.120	3.128	0.707
Burundi	MTAR	1.578	-0.091	-1.762	5.327	4.895**
Cameroon	TAR	1.576	-0.054	-0.238	7.729**	6.486**
Egypt	TAR	4.467	-0.586	-0.043	11.291***	16.479***
Ivory Coast	MTAR	1.903	-0.253	-0.022	10.793***	11.734***
Kenya	TAR	4.508	-0.729	-0.071	17.791***	24.521***
Morocco	MTAR	1.741	0.032	-0.121	8.250**	6.318**
Nigeria	TAR	23.766	-0.488	-0.026	17.771***	28.374***
South Africa	TAR	1.961	0.005	-0.287	10.624***	16.637***
Tanzania	TAR	11.465	-1.150	-0.034	35.219***	64.707***

Notes. These estimates are based on the estimation of equations (5) and (6), or (6) and (7) for the TAR and MTAR models respectively. Γ is the value of the threshold expressed as a percentage of the mean of the dependent variable across the study period.

[a] Entries in this column are the *F*-statistics for the null hypothesis $\rho_1 = \rho_2 = 0$. ***, ** and * denote rejection of the null at the 1, 5 and 10% significance levels respectively. This test follows a non-standard distribution so the test statistics are compared with critical values reported by Enders and Dibooglu (2001): 9.94, 7.53 and 6.35 for a sample size of *n*=100.

[b] Entries in this column are conventional *F*-statistics.

Table 3. Error Correction Modelling with Asymmetric Adjustment

	ρ_{11}	ρ_{12}	ρ_{21}	ρ_{22}	ρ_{31}	ρ_{32}
Cameroon	-0.058	-0.297***	-0.159	-0.748***	0.015*	0.026**
	(-1.381)	(-4.103)	(-1.100)	(-3.023)	(1.966)	(2.024)
Egypt	-0.052	-0.014	1.177***	0.083*	-1.024*	0.005*
	(-0.734)	(-0.953)	(5.027)	(1.679)	(-1.932)	(1.736)
Ivory Coast	-0.257***	-0.031	0.393*	-0.047	0.007	0.015*
	(-3.592)	(-0.752)	(1.705)	(-0.309)	(0.581)	(1.919)
Kenya	-0.106*	-0.033	0.654***	-0.017	-0.011	0.007
	(-1.689)	(-1.490)	(3.905)	(-0.296)	(-0.789)	(1.392)
Morocco	0.087	-0.104***	-0.466**	-0.314**	0.024	0.010
	(1.544)	(-2.633)	(-2.190)	(-2.104)	(1.077)	(0.632)
Nigeria	-0.033	-0.040***	0.674***	-0.077*	0.000	0.001
	(-0.889)	(-3.017)	(5.291)	(-1.711)	(0.067)	(0.499)
South Africa	-0.014	-0.046**	-0.118	0.311**	-0.005	-0.006
	(-1.408)	(-2.447)	(-1.578)	(2.278)	(-1.011)	(-0.623)
Tanzania	0.064	-0.059**	1.790***	-0.038	-0.009	-0.001
	(-0.549)	(-2.056)	(12.122)	(-1.036)	(-0.839)	(-0.207)

Notes. Results for equations (8), (9) and (10) taken from the system of error correction equations presented in the text. These results exclude Botswana and Burundi for whom the null of non-cointegration was accepted using the tests for asymmetric cointegration in Table 2. ***, ** and * denote rejection of the null of a zero coefficient at the 1, 5 and 10% significance levels where the figures in parentheses are t-statistics.

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