Does the Purchasing Power Parity Hold in Emerging Markets? Evidence from Black Market Exchange Rates

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Abstract

We examine the Purchasing Power Parity (PPP) hypothesis using a unique panel of monthly data on black market exchange rates for twenty emerging market economies over the period 19973M1-1993M12. We apply a large number of recent heterogeneous panel unit root and cointegration tests. Panel unit root tests do not favour mean reversion in the real black market exchange rate. The evidence for non-rejection of the unit root hypothesis remains robust even after allowing for structural breaks. Panel cointegration tests support evidence of cointegration between the nominal exchange rate and relative prices. These results contrast with those obtained from unit root tests. Since we believe that the former may be biased by the imposition of the joint symmetry and proportionality restriction, we test for such a restriction using likelihood ratio tests and find that it is strongly rejected.

Keywords: Black Market Exchange Rates; Purchasing Power Parity; Structural Breaks; Panel Unit Root and Cointegration Tests

JEL Classification: F31; C23; C52

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1. Introduction

The fundamental notion of the Purchasing Power Parity (PPP) hypothesis is that the exchange rate depends on relative prices. Given its importance in international finance, the long-run PPP relationship has been subjected to extensive empirical investigation during the last decade. However, most of that literature has focused on testing for PPP in OECD countries. The consensus amongst researchers seems to be mixed (see, for example, Sarno and Taylor, 2002; O'Connell, 1998; Pappell, 1997; Pedroni, 1997; Lothian, 1997; Frankel and Rose, 1996).

On the other hand, little work has been done for emerging market economies[1]. More importantly, very few papers investigate black market exchange rates behaviour in emerging market economies, which play such a major role in these economies. Phylaktis and Kosimmatis (1994) and Speight and McMillan (1998), who use time series, and Luintel (2000) who employs panel unit root tests, are few examples that consider black market exchange rates, though they cover only a small number of countries.

Black market exchange rates are unofficial rates in the sense that their transactions do not take place in official markets. In most of the countries covered in the present study, black market exchange rates have a long tradition and in many cases have also been supported by governments. In fact, in many cases the volume of transactions in these markets were also much larger than that in the official market.

The data set used in this study includes twenty emerging market economies spaning over the period 1973M1-1993M12. To our knowledge, empirical investigation of

[1] Frankel and Rose (1996) and Lothian (1997) are two exceptions, but these studies use official nominal exchange rates.

the real exchange rate using black market rates of this dimension has not been previously undertaken. Thus this study extends the test of PPP into new directions.

We use a battery of new heterogeneous panel unit root and cointegration tests which have greater power than the time series tests normally used in the literature on PPP. It is interesting to note that empirical evidence for PPP based on panel cointegration tests is very limited. Pedroni (1997) is the exception, though he uses only his own test and OECD data. One important contribution of our paper is that we also apply the recently developed McCoskey and Kao (1998) and Larssson *et al* (2001) panel cointegration tests. In addition, we examine the symmetry and proportionality conditions. Furthermore, we assess the robustness of the evidence from unit roots by testing for structural breaks in the real exchange rate series.

When testing the PPP using emerging markets data, two propositions are normally made. First, it is suggested that real exchange rates in those countries are more volatile than exchange rates in OECD countries. Second, in emerging markets, monetary growth tends to overshadow real factors such that the relative price ratios exhibit excess volatility. The latter may bias evidence in support of PPP. We try to shed some light on the above issues in this paper.

The paper is organised as follows. Section 2 discusses the PPP specification. Sections 3 and 4 outline the panel unti root and cointegration tests used in the study. Section 5 discusses the data on black market exchange rates. This is a unique set of data that has not been previously used in the literature. The empirical results are presented and analysed in Sections 6 and 7. Section 8 concludes.

2. Purchasing Power Parity

Under absolute PPP the nominal exchange rate is proportional to a ratio of domestic to foreign price levels:

$$s_t = \alpha + \beta_0 p_t - \beta_1 p_t^* \tag{1}$$

where s_t is the nominal exchange rate, and p_t, p_t^* are, respectively domestic and foreign prices, all measured in logs.

Equation (1) is known as a trivariate relationship. A bivariate relationship between the nominal exchange rate and the domestic to foreign price ratio is given by:

$$s_t = \alpha + \beta(p_t - p_t^*) + u_t \tag{2}$$

This PPP framework does impose an a-priori restriction on the cointegrating vector. The difference between the PPP framework represented by equation (1) and (2), is that in the latter the symmetry condition on the price coefficients has been imposed.

Another specification of PPP that is commonly used in unit root tests is given by

$$q_t = s_t - p_t + p_t^* \tag{3}$$

Where q_t is the real exchange rate.

The PPP equation (3) requires $\beta = 1$. This also implies $\beta_1 = -\beta_0$, which imposes the joint symmetry/proportionality restriction. Since all unit root tests on the real exchange rate assume implicitly that such a restriction holds, a failure of these tests to find evidence favouring mean reversion in the real exchange rate may be caused by a failure of such a restriction. Various explanations have been oferred for the potential rejection of the symmetry and proportionality conditions. Sarno and Taylor (2002) stress the importance of measurement errors, barriers to trade and other economically unimportant factors, while Froot and Rogoff (1995) suggest the possibility of a common trend in the relative prices of traded and non-traded goods.

3. Testing for a Unit Root in Heterogeneous Panels

In this section we review the new heterogeneous panel unit root tests used in this paper to investigate whether or not the black market real exchange rate has been stationary over the sample period under consideration[2].

Im *et al* (1997) proposed a unit root test for heterogeneous dynamic panels based on the mean-group approach. This test is valid in the presence of heterogeneity across-sectional units and is given by the following equation:

$$t - bar = \frac{\sqrt{N(T)}(t_T - E(t_T))}{\sqrt{VAR(t_T)}} \qquad (4)$$

[2] For a formal description of the tests presented in this and next sections refer to Im *et al* (1997), McCoskey and Kao (1998), Pedroni (1997), and Larsson *et al* (2001).

Where N is the cross sectional dimension, t_T is the average ADF statistic for individual countries, and $E(t_T)$ and $Var(t_T)$ are respectively mean and variance tabulated by Im *et al* (1997). The authors state that the standardized t-bar statistic converges in probability to a standard normal distribution as T, N→∞. Therefore we can compare the t-statistic obtained to the critical values from the lower tail of the normal distribution. We shall be using the demeaned version of the above t-bar test in this study.

While the Im *et al* (1997) t-bar test relaxes the assumption of homogeneity of the root across units, several difficulties still remain. In fact, Im *et al*. assume that T is the same for all the cross-section units and hence the t-bar test requires a balanced panel or complete panel, (i.e. where the individuals are observed over the sample period).

Maddala and Wu (1999) propose another panel unit root test that is valid for unbalanced panels too. Furthermore, by using Monte Carlo simulations, they show that their test is more powerful than the t-bar test. Suppose there are N unit root tests. Let π_i be the observed significance level (π-value) for the ith country. The Π_λ test has a χ^2 distribution with 2N d.f.:

$$\Pi_\lambda = \sum_{i=1}^{N}(-2\log_e \pi_i) \tag{5}$$

Maddala and Wu (1999) suggest obtaining π-values by using bootstrap methods in order to account for cross sectional dependence. However the bootstrap procedure suggested by Maddala and Wu (1999) is extremely time consuming. Furthermore it requires bootstrapping a moving average process. The suggested procedure encounters two practical difficulties. First it is well known that estimation of MA time series models is not as straightforward as the estimation of the AR models. Second it requires the truncation of an infinite sum. Taking into account these drawbacks, Cerrato and Sarantis (2002) suggested a more efficient bootstrap procedure than the one used in Maddala and Wu (1999) and show, by Monte Carlo simulation, that the bootstrap test proposed is free of size distortion. We shall also use this bootstrap test in this paper.

4. Testing for Cointegration in Heterogeneous Panels

McCoskey and Kao (1998) develop a residual-based Lagrange Multiplier test for the null hypothesis of cointegration in panel data. The model they consider allows for varying slopes and intercepts across units:

$$y_{i,t} = \alpha_i + \beta_i x_{it}^{'} + e_{it} \tag{6}$$

where $e_{it} = \theta \sum_{j=1}^{t} u_{ij} + u_{it}$

We test the null hypothesis $H_0 : \theta = 0$ against the alternative $H_0 : \theta \neq 0$. Under the null hypothesis we have $e_{it} = u_{it}$ and the equation above is a system of cointegrated regressors. The test statistic is then given by the following LM statistic:

$$LM = \frac{\frac{1}{N} \sum_{i=1}^{N} \frac{1}{T^2} \sum_{t=1}^{T} S_{it}^{+2}}{s^{+2}} \tag{7}$$

where S_{it}^{+2} is the partial sum of estimated residuals:

$$S_{it}^{+} = \sum_{j=1}^{t} e_{ij}^{*+} \text{ and } s^{+2} = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T} e_{it}^{*+2} \tag{8}$$

The residuals e_{it}^{*} can be estimated using either the dynamic ordinary least squares (DOLS) estimator or the fully modified OLS (FMOLS) estimator, both of which correct for serial correlation and endogeneity of regressors. A comparative study by Kao and Chiang (1999) demonstrate that the DOLS estimator outperforms the FMOLS estimator, so we employ the DOLS method in this study. In cases where there is significant autocorrelation, we use the Stock and Watson (1993) dynamic GLS (DGLS) estimator. McCoskey and Kao (1998) show that the standardised version of the equation is given by:

$$LM^{*} = \frac{[\sqrt{N}(LM - u_v)]}{\sigma_v} \Rightarrow N(0,1) \tag{9}$$

where u_v and σ_v are obtained by Monte Carlo simulation and tabulated by the authors (see McCoskey and Kao, 1998, Table 1).

Pedroni (1999) uses the same heterogeneous model as the one represented by equation (6), but he also assumes individual specific deterministic trends. Furthermore, the null hypothesis in his test is that of no cointegration. Based on this model, he proposes seven panel cointegration statistics. Specifically, four are based on within-dimension approach and three are based on between-dimension approach. In the first group we sum both the numerator and the denominator terms over the N dimension. In the second group we first divide the numerator by the denominator prior to summing over the N dimension separately. Furthermore, Pedroni (1997) shows that the asymptotic distribution of these statistics, under an appropriate standardisation, is a normal distribution, that is:

$$\kappa = \frac{\kappa_{N,T} - u\sqrt{N}}{\sqrt{v}} \Rightarrow N(0,1) \qquad (10)$$

where $\kappa_{N,T}$ is the panel cointegration statistic and u and v are the moments of the Brownian function (i.e. broadly speaking expected mean and variance) that are computed in Pedroni (1999).

A weakness of the tests considered above is that they assume the cointegrating vector to be unique. Such an assumption may be too strong. It constrains researchers to choose a normalisation rule and it is unclear on the basis of what criteria this choice is made. To overcome this problem, system estimation methods have been suggested.

Larsson *et al*. (2001) propose a panel cointegration test analogue of the Johansen maximum likelihood method that allows for multiple cointegrating vectors. Assume that the data generating process for each of the countries is represented by the error correction model,

$$\Delta y_{it} = \Pi_i y_{i,t-1} + \sum_{k=1}^{k_i} \Gamma_{ik} \Delta y_{i,t-k} + \varepsilon_{it} \qquad \text{i=1,…,N} \qquad (11)$$

where Γ_i is of order $p \times p$ (p is the number of variables in each country), y_i is a $p \times 1$ vector of variables and Π_i a $p \times p$ long run matrix. We estimate equation (11) for each individual country using maximum likelihood methods and calculate the trace statistic, LR_i. The panel rank trace statistic, LR_{NT}, can be obtained as the average of the N individual trace statistics, LR_{iT} (H(r) ¦ H(p)). The null and alternative hypotheses are:

$$Ho : rank(\Pi_i) = r_i \leq r \quad \text{for all i = 1,…N}$$

$$H_1 : rank(\Pi_i) = p \qquad \text{for all i = 1,…N}$$

The standardised panel cointegration rank trace test, Y_{LR}, is:

$$Y_{LR}(H(r) \mid H(p)) = \frac{\sqrt{N}(LR_{NT}(H(r) \mid H(p)) - E(Z_k))}{\sqrt{Var(Z_K)}} \Rightarrow N(0,1) \quad (12)$$

where $E(Z_k)$ and $Var(Z_k)$ are the mean and variance of the asymptotic trace statistic. Larsson *et al*. (2001) report the values for the moments of Z_k, and these can be used to calculate the test statistic.

5. An Overview of the Data

We use monthly data on the black market exchange rates for a panel of twenty emerging market countries (Nepal, Pakistan, Philippines, S. Lanka, Thailand, Turkey, Venezouela, Indonesia, Kenya, Korea, Malaysia, Ethiopia, Ghana, Hungary, India, Algeria, Bolivia, Colombia, D. Republic, Egypt) over the period 1973M1-1993M12. The US Dollar is used as numeraire currency. The black market exchange rates are obtained from *Pick`s Currency Yearbook* (various publications), and from the *World Currency Yearbook* (various issues) published by the International Currency Analysis. The consumer price index (CPI) is used as price index. These are the standard sources for black market exchange rate data. Generally, the black market currency is defined as the private dealings of foreign currency bank notes and/or nonblank transfers abroad. We have included only twenty countries in this panel because of the lack of consistent data on the CPI (over the period 1973-1993) for most emerging markets. We have also excluded some countries because the time series for the exchange rate displays exceptionally large jumps due to the re-denomination or large devaluation of the respective domestic currency against the US dollar. The sample ends in 1993 because of the unavailability of data beyond that year.

3.1 Volatility of Exchange Rates and Relative Prices

One of the most important results in the PPP literature is that this parity condition seems to hold pretty well for high inflationary countries while it does not for those countries whose rate of inflation has been relatively low over the sample period under analysis. This is why monetary growth in the former countries is likely to overshadow real factors, and that may bias evidence towards PPP (see Lothian and Taylor, 1996). Since the panel data set used in this study contains data for developing countries, and these countries may have experienced high inflation rates, we calculate the volatility of the relative price and exchange rates. Results are displayed in Table 1.

Comparison of the volatility of the nominal exchange rate with that of the relative price shows that the relative price ratios are less volatile than the nominal exchange rate. That is the monthly absolute rate of change of the nominal exchange rate is always greater than the monthly absolute rate of change of the relative price ratios except Turkey. Following Lothian and Taylor (1996) our result means that the volatility of relative prices in our set of data should not provide a source of bias towards the PPP.

One proposition often presented in the literature is that real exchange rates in developing countries have been more volatile than exchange rates in OECD countries. We compare the results on the real exchange rate ($\triangle q_t$) displayed in Table (1) with those for a panel of twenty OECD countries over the same sample period (see Table 2). On average, the black market real exchange rates seem to be characterised by lower standard deviations than the real exchange rates in OECD countries. In fact the increment between these two data sets is 3.84. However this result is strongly distorted by just one country, that is the UK whose exchange rate has been very volatile (with a the standard deviation of 1.58) over the sample period under consideration. If we drop the UK from our panel the increment falls to 0.18. This result suggests that in terms of volatility of the real exchange rate there is very little difference between these two data sets.

6. Results from Panel Unit Root Tests

We perform standard ADF tests on the real exchange rate of each country in the panel. The number of lags in the ADF specification is chosen using the procedure suggested by Campbell and Perron (1991). The results are displayed in Table 3. On the basis of the individual ADF statistics we are able to reject the unit root null hypothesis in only two countries out of twenty.

The demeaned version of the t-bar test suggested by Im *et al*. (1997) is 2.04, which is considerably larger than its critical value (-1.64). This indicates that the null hypothesis of a unit root in the black market real exchange rate for the full panel of emerging market economies cannot be rejected.

Next we apply the bootstrap panel unit root test proposed by Cerrato and Sarantis (2002). The results are shown in Table 4. The individual π_i probabilities cannot reject the null hypothesis of a unit root at the 5% significance level in any country except for Korea and S. Lanka. The panel unit root test of 42.40 is well below both the 5% and 1% critical values. This result provides strong evidence that the black market real exchange rate in the full panel of emerging markets is an I(1) stochastic process.

Taken together, the above findings indicate a much stronger acceptance of the null hypothesis of a unit root, and hence rejection of the PPP, than the one obtained in Cerrato and Sarantis (2002) for a panel of twenty OECD countries. Furthermore, these results contrast with the ones obtained by Luintel (2000) for black market exchange rates. However that study includes only eight countries in its panel and only five of them are also included in the current investigation. Finally, our findings provide supportive evidence for the "difference in productivity" issue raised by Froot and Rogoff (1995).

6.1 Testing for Structural Breaks

Most of the emerging market economies experienced different exchange rate and policy regimes during our sample period. Could it be that the detection of a unit root in the real exchange rate for individual countries is due to the effects of potential structural breaks in the time series? To investigate this issue we apply the methodologies of Banerjee *et al* (1992) and Zivot and Andrew (1992) which allow us to estimate the break-date endogenously and then to test for the unit root hypothesis conditional on the identified structural break[3]. Consider the following models:

$$y_t = \alpha^A + \theta^A DU_t(\lambda) + \beta^A t + \delta^A y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \qquad (13)$$

$$y_t = \alpha^B + \beta^B t + \eta^B DT_t^*(\lambda) + \delta^B y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \qquad (14)$$

$$y_t = \alpha^C + \theta^C DU_t(\lambda T) + \beta^C t + \eta^C DT^*(\lambda) + \delta^C y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \qquad (15)$$

where λ is the break date; $DU_t = 1$ if $t>\lambda$, and 0 otherwise; $DT_t^* = t - \lambda$ if $t>\lambda$, and 0 otherwise.

Model (13) describes Perron's (1989) crash model that allows one time shift in the mean of the trend of the process. Model (14) specifies a shift in the slope of the trend function, described as the "changing growth" model by Perron (1989). Model (15) allows simultaneously for a shift in both the mean and slope of the trend function (mixed model). The latter model was investigated by Zivot and Andrew (1992), while Banerjee et al (1992) considered only the first two cases.

The true break is assumed to be in the interval $\Psi = [\lambda_0, \lambda_0 + 1, T - \lambda_0]$, where λ_0 is the initial start up sample defined as $\lambda_0 = \gamma_0 T$ and γ_0 the trimming parameter. Equations 5.19-5.21 are estimated for break dates $[\lambda_0, \lambda_0 + 1, \ldots, T-\lambda_0]$ and the sequence of ADF statistics for $H_0 : \delta = 1$ denoted as $t^i_{ADF}(T_b)$ for $T_b = [\lambda_0$ up to $T-\lambda_0]$ with i = A, B

[3] For a similar application of these tests, see Luintel (2000). However, Luintel covered

and C, computed. The minimum sequential ADF statistic as suggested in Banerjee *et* al (1992) and Zivot and Andrew (1992) is given by the statistic that maximises the evidence against the no-break unit root null hypothesis. In this section we select the break point using this methodology.

While these procedures assume that the location of the break point is unknown, they all assume that its specification is known, which is unrealistic. In other words, once the data break has been selected which of the above alternative specifications is to be preferred? Sen (2000) shows that a misspecification of the model under the alternative hypothesis leads to lower power of the test proposed by Banerjee *et al* (1992) and Zivot and Andrews (1992). What he recommends is using the mixed model (15) under the alternative hypothesis, once the break date has been selected.

Table 5 presents the sequential t^i_{ADF} statistics for models (13)-(15). Critical values for these statistics are reported at the bottom of table.. The critical values used for the mean shift model and the trend shift models are taken from Banerjee *et al* (1992, Table 2). The critical value for the mixed model (15) has been taken from Zivot and Andrew (1992, Table 4). The trimming parameter γ_0 is set to 0.15 and the number of lags is set to 4, as suggested in Banerjee *et al* (1992).

Let us consider the mean shift model first. We can reject the null hypothesis of a unit root only in two countries, namely Ghana and Bolivia. In the case of the trend shift model, the unit root null is rejected in four countries (i.e. Ghana, Korea, Nepal, and S. Lanka). Finally, the mixed (mean and trend shift) model also selects four countries (that is Ghana,

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only eight Asian countries.

Indonesia, S. Lanka and Bolivia) for the rejection of the unit root null hypothesis. Ghana and S. Lanka seem to be the most selected countries.

Following Sen's (2000) recommendation, we present estimates of the mixed model (15) for the six countries selected above, with break dates indexed as selected in Table 5. The null hypothesis of a unit root with one-time endogenous structural break corresponds to $\delta=0$. The alternative hypothesis is one-time trend-break stationarity (asymptotic critical values are taken from Zivot and Andrews (1992)). Results are displayed in Table 6. We notice that the null hypothesis is rejected here for four countries, that is Ghana, Indonesia, S. Lanka and Bolivia. We can now assess the significance of the other parameters considering that their t-values are normally distributed (see Perron, 1989). Results for all coefficients and countries are very mixed. The time trend (see coefficient β) is significant only in two out of the four countries, that is, S. Lanka and Bolivia, with the former displaying a significant shift in the mean of the process, and the latter a significant shift in the slope of the trend. For the other countries, the coefficients are often insignificant and display the wrong sign. Therefore, the only countries where there is evidence that the underlying process is a trend-break stationary process are S. Lanka and Bolivia[4].

[4] This limited evidence of structural shifts is similar to that reported by Luintel for the eight Asian countries. To assess the implications for the panel unit root tests, we have also calculated the relevant panel statistics by excluding the two countries for which we found significant trend-break stationarity (i.e. S. Lanka and Bolivia). The Im et al (1997) and bootstrap panel unit root statistics are 2. 18 and 34.538 (CV5%=49.76), respectively. Hence both these statistics provide strong evidence of a unit root in the black market real exchange rate for the group of eighteen countries, even after allowing for a structural break.

7. Results from Panel Cointegration Tests

Before using cointegration analysis to test for a long-run relationship between nominal exchange rate and relative prices, we perform unit root tests on each variable entering in the PPP equation (1).

The Im *et al* (1997) t-bar test, shown in Table 7, suggests that both the nominal exchange rate and domestic price are nonstationary. We also apply the bootstrap Π_λ-test and the results are reported in Table 8. The panel test confirms the results already obtained by the Im *et al* (1997) test, that is, both the nominal black market exchange rate and the domestic price for the full panel of emerging markets are I(1) stochastic processes, so they can enter into a cointegrating relationship. The ADF test for the USA consumer price (p^*) over the same period was –3.211. This implies a stationary process at the 5% level, but nonstationarity at the 1% significance level. This result seems rather strange, given the international evidence on the nonstationarity of consumer prices. In view of this evidence and the borderline value of the ADF statistic, we do not believe that the US-CPI is an I(0) process, so we treat it as nonstationary.

We specify the DOLS regression as follows:

$$s_{it} = \alpha_i + \beta_0 p_{t-1} - \beta_1 p^*_{t-1} + \sum_{j=-k_i}^{k_i} \phi_{ij} \Delta p_{i,t-j} + \sum_{j=-k_i}^{k_i} \Delta p^*_{i,t-j} + u_{it} \qquad (16)$$

$i = 1,\ldots N$, k_i = leads and lags of Δp_i and Δp^*_i

The number of leads and lags in equation (16) was chosen with the Akaike criterion. We use DGLS when significant evidence of residual autocorrelation is present in our data. The McCoskey and Kao (1998) panel cointegration statistic is shown in

Table 9. This is smaller than its critical value, which implies that the null hypothesis of cointegration for the whole panel cannot be rejected.

The estimates of the Pedroni (1999) panel cointegration statistics are also reported in Table 9. All seven statistics are well blow their respective critical values, so they reject the null hypothesis of no cointegration in the panel.

Hence, both panel cointegration tests strongly support the existence of a long-run equilibrium relationship between the nominal black market exchange rate and domestic and foreign prices for the full panel of emerging market economies. These findings, favouring PPP, are in sharp contrast with those obtained by panel unit root tests for the real exchange rate. We believe that one possible explanation could be that the joint symmetry and proportionality restriction imposed on unit root tests of the real exchange rate is too restrictive. We shall investigate this issue further in the next section.

The estimates of the long-run PPP relationship (1) obtained with the DOLS/DGLS estimators are exhibited in Table 10. The adjusted coefficient of determination suggests high explanatory power for all countries, while the *Pr*[*Fa]* values indicate absence of autocorrelation for all countries. The intercept is significant in almost all countries except five (Sri-Lanka, Venezuela, Ethiopia, Algeria, Egypt). It is interesting to notice that the foreign price has the wrong sign in most of the countries in the panel and is insignificant in six countries. On the other hand, the coefficient on the domestic price displays the expected sign in thirteen out of twenty countries and is significant in five countries.

Finally, we test for cointegration by using the new panel contegration test suggested by Larsson *et al* (2001). We include an intercept in the VAR to account for potential measurement errors, as in equation (1). The number of lags for each country was chosen on the basis of the Akaike criterion. The results are reported in Table 11. As the Johansen trace test shows, for most of the countries in the panel the maximum rank is two. The null of no cointegration is rejected only in Algeria. The panel cointegration rank trace statistic, shown at the bottom of the table, suggests a common rank of two in the panel. Hence the Larsson *et al* (2001) panel test favours the presence of two cointegrating vectors among the variables in equation (1) for the full panel of emerging market economies.

7.1 Interpreting Cointegrating Vectors

As we have obtained two cointegrating vectors from our PPP framework, we would like to give them an economic meaning by imposing a structure on them as suggested by Johansen (1995. We impose a structure on the two cointegrating vectors by implementing the joint symmetry and proportionality restriction that is implicitly incorporated in equation (3). In this way we have a cointegrating vector such as (1, -1, 1). We also impose a structure on the second cointegrating vector by assuming the US-CPI (p^*) to be an I(0) process, that is (0, 0, 1). In this case the USA price index would be itself a cointegrating vector. Finally we use a likelihood ratio test as in Johansen (1995) to test for the validity of these restrictions. Furthermore, following Larsson *et al* (2001) we extend that test to a panel context[5].The results are displayed in Table 12. On the basis of the individual statistics, we reject the null of valid restrictions in only five countries out

[5] However results from this panel test should be interpreted with caution since this panel test requires cross-sectional independence, an assumption that is rather too strong for the group of countries under consideration.

of twenty. The panel test (PLR) shown at the bottom of Table 12 suggests that the null hypothesis of valid restrictions is strongly rejected. The rejection of the null here confirms our view that the US consumer price index is likely to be a unit root process. However, we also reject the joint symmetry and proportionality restriction, indicating that this assumption is too restrictive.

Taking into account these results and the fact that the Larsson *et al* (2001) test tends to over-estimate the true number of cointegrating vectors, we decided to restrict the rank to be the same and equal to one across different countries. We impose a structure on the cointegrating vector by testing the joint-symmetry and proportionality restriction and using a likelihood ratio test for over-identifying restrictions. Furthermore, we extend this test to a panel context. The results are reported in Table 13. The individual country statistics imply acceptance of the null in seven out of twenty countries. But the panel statistic again strongly rejects the null hypothesis of valid restriction for the full panel of emerging market economies.

8. Conclusions

This paper examines the PPP hypothesis using a unique panel of black market exchange rates for twenty emerging market economies. This is the first empirical study of the PPP using black market exchange rates of such large dimension. We use a battery of new heterogeneous panel unit root and cointegration tests that have been shown in the literature to have greater power than the time series tests normally employed in empirical studies of the PPP.

The empirical evidence from panel unit root tests does not favour mean reversion in the black market real exchange rate. This result is not affected by structural breaks.

After extensive investigation using sequential tests, we found that the problem of structural breaks in black market real exchange rates is not widespread in our sample. But even after allowing for mean and trend shifts in the country ADF tests, and excluding from the group the two countries with significant trend-break stationarity in the real exchange rate, the evidence on the non-rejection of the unit root null does not alter. Furthermore, these findings on the PPP from unit root tests are in line with those obtained in Cerrato and Sarantis (2002) and other empirical studies for OECD countries. In contrast, all panel cointegration tests strongly favour cointegration between the nominal exchange rate and relative prices, thus providing strong support for the PPP hypothesis in the full panel of emerging market economies.

We also tested the joint symmetry/proportionality restriction using likelihood ratio tests and found this restriction not to be supported by our set of data. This result could have noticeable relevance for applied research on the PPP. Since this restriction is implicitly imposed on the unit root tests of the real exchange rate, failure of this restriction could be one of the reasons why unit root tests fail to reject the unit root hypothesis in the real exchange rate. Therefore, unit root tests on the exchange rate may be biased towards finding no mean reversion and rejecting the PPP. This could be an explanation of why we failed to reject the unit root null hypothesis in the real black market exchange rate.

The overall empirical findings from the black market exchange rates seem to provide support for the weak form but not the strong form of the PPP hypothesis in the emerging market economies.

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Table 1

Descriptive Statistics for Black Market Exchange Rates and Relative Price

Country	Δq_t		Δs_t		$\Delta(p_t/p^*_t)$	
	Mean	Stdv.	Mean	Srdv.	Mean	Stdv.
Nepal	0.015	0.018	0.051	0.066	0.013	0.011
Pakistan	0.009	0.015	0.029	0.053	0.008	0.009
Phil.	0.009	0.001	0.029	0.047	0.008	0.012
S.Lanka	0.013	0.075	0.421	0.181	0.009	0.011
Thail.	0.008	0.008	0.025	0.028	0.005	0.005
Turkey	0.004	0.004	0.005	0.049	0.035	0.076
Venez.	0.006	0.013	0.035	0.088	0.013	0.018
Indon.	0.005	0.019	0.062	0.569	0.008	0.098
Kenya	0.010	0.011	0.041	0.044	0.011	0.015
Korea	0.006	0.020	0.041	0.119	0.006	0.007
Malaysia	0.023	0.026	0.016	0.018	0.005	0.038
Etiopia	0.037	0.057	0.064	0.122	0.019	0.017
Ghana	0.014	0.024	0.095	0.237	0.039	0.047
Hungary	0.012	0.008	0.051	0.041	0.011	0.013
India	0.011	0.094	0.032	0.029	0.008	0.007
Algeria	0.014	0.013	0.054	0.061	0.021	0.021
Bolivia	0.011	0.057	0.088	0.295	0.065	0.151
Colomb.	0.003	0.036	0.023	0.028	0.014	0.009
D.Rep.	0.012	0.021	0.033	0.081	0.014	0.016
Egypt	0.151	0.411	0.071	0.161	0.017	0.017

Note: Δq_t is the monthly rate of change in the real exchange rate (in log), Δs_t is the monthly absolute rate of change of the nominal exchange rate and $\Delta(p_t/p^*_t)$ is the monthly absolute rate of change of the relative price ratios.

Table 2

Real Exchange Rate Volatility in OECD Countries

	Mean	Stdv
Aust.	0.011	0.009
Dan	0.014	0.011
Belg.	0.008	0.006
Fra.	0.015	0.013
Germ.	0.054	0.049
Ita.	0.003	0.003
NL	0.045	0.038
Norw.	0.012	0.01
Port.	0.005	0.004
Spa.	0.006	0.006
Swed.	0.013	0.012
Switz.	0.071	0.074
Can.	0.029	0.026
UK	0.034	1.58
NZ	0.043	0.05
Jap.	0.005	0.005
Gre.	0.005	0.004
Finl.	0.016	0.015
IR	0.006	0.005
Mex.	0.027	0.066

Table 3

Im *et al* (1997) Unit Root Test for the Real Exchange Rate

Country	Lag	t-stat
Algeria	4	-0.87
Colomb.	6	-0.67
D.Repub	5	-2.07
Egypt	5	0.89
Ethiopia	1	-1.85
Ghana	5	-2.29
Hungary	6	-1.65
India	0	-0.78
Indon.	1	1.14
Kenya	0	-0.98
Korea	5	-3.86
Malaysia	1	-0.55
Nepal	5	-1.5
Pakistan	1	-1.69
Philip.	1	-2.29
S.Lanka	4	-3.38
Thayl.	1	-1.71
Turk.	4	-1.42
Venez.	6	-1.35
Boliv.	0	-0.74
t-bar		**2.04**

Table 4

Bootstrap Panel Unit Root Test (π-values) for the Real Exchange Rate

Country	**π**	**logπ**
Alger.	0.671	-0.39899
Boliv	0.701	-0.35525
Colomb	0.796	-0.22753
D.Rep	0.258	-1.35286
Egyp.	0.958	-0.04239
Ethio.	0.300	-1.20231
Ghana	0.177	-1.73161
Hung.	0.451	-0.79629
India	0.754	-0.28236
Indon.	0.981	-0.01918
Kenya	0.788	-0.23826
Korea	0.010	-4.55638
Malaysia	0.207	-1.57262
Nepal	0.506	-0.68023
Pakistan	0.724	-0.32227
Philip.	0.178	-1.72317
S.Lank.	0.018	-3.57555
Thay.	0.423	-0.86038
Turk.	0.504	-0.68518
Venez.	0.561	-0.57714
		-21.1999
Π_λ-test		**42.3998**
CV5%		55.759
CV1%		63.691

Table 5
Sequential Tests of Structural Breaks

Country	mean-shift t-min(ADF)	trend-shift t-min(ADF)	mean and trend shift t-min(ADF)
Algeria	-2.12 [82:04]	-2.78 [80:08]	-3.11 [76:01]
Colombia	-2.39 [85:01]	-2.02 [75:12]	-2.31 [75:12]
D.Rep.	-3.42 [85:01]	-3.67 [83:07]	-3.81 [83:11]
Egypt	-3.09 [85:01]	-3.87 [80:12]	-3.71 [80:12]
Ethiopia	-3.99 [78:02]	-4 [80:12]	-4.23 [75:12]
Ghana	-7.21 [75:12]	-4.81 [78:02]	-7.52 [75:12]
Hungary	-3.91 [89:12]	-3.29 [86:01]	-4.03 [82:05]
India	-4.12 [86:11]	-3.98 [89:12]	-4.58 [86:01]
Indonesia	-2.26 [75:12]	-2.71 [78:02]	-5.64 [78:02]
Kenya	-3.16 [89:12]	-3.64 [89:12]	-3.71 [89:12]
Korea	-4.6 [82:05]	-4.67 [87:07]	-5.02 [85:01]
Malaysia	-4.27 [86:01]	-4.34 [89:12]	-4.52 [86:01]
Nepal	-4.62 [85:01]	-4.76 [75:12]	-4.93 [86:01]
Pakistan	-4.44 [75:12]	-4.2 [78:02]	-4.97 [85:01]
Philip.	-4.81 [82:05]	-3.81 [89:12]	-4.63 [82:05]
S.Lanka	-4.75 [78:02]	-5.38 [78:02]	-5.39 [78:02]
Thailand	-4.63 [83:11]	-4.24 [89:12]	-4.2 [86:11]
Turkey	-2.67 [80:12]	-3.38 [83:11]	-3.51 [82:05]
Venezuela	-3.84 [82:05]	-2.63 [87:07]	-3.73 [82:05]
Bolivia	-14.4 [86:11]	-2.46 [82:05]	-13.84 [86:11]
C-V 5%	**-4.8**	**-4.39**	**-5.08**

Table 6

Estimates of the Mean and Trend Shift Model (15)

Country	T_B	α	θ	β	η	δ
Ghana	1975:12	2.01	0.48	-0.006	0.006	0.68
		[7.21]	[1.02]	[-1.09]	[1.97]	[-7.52]
Korea	1985:01	2.79	0.07	0.0001	-0.001	0.59
		[1.03]	[1.35]	[0.6]	[-3.03]	[-5.02]
Nepal	1986:01	0.6	-0.03	0.0008	0.0006	0.81
		[4.9]	[-1.65]	[3.8]	[1.87]	[-4.93]
Indonesia	1978:02	0.52	0.05	-2E-04	0.0004	0.84
		[4.52]	[3.19]	[-0.55]	[0.14]	[-5.64]
S.Lanka	1978:02	2.07	0.08	0.007	-0.007	0.37
		[5.14]	[0.78]	[2.64]	[-2.56]	[-5.39]
Bolivia	1986:11	7.14	-6.89	0.003	-0.002	0.52
		[13.53]	[-14.34]	[2.44]	[-0.57]	[-13.84]
CV 5%	**-5.08**					

Table 7

Im *et al* (1997) Unit Root Test

Country	s_t	p_t
Alger.	0.36	1.38
Boliv	-1.83	0.25
Col.	1.26	0.031
D.Rep	-0.08	2.63
Egypt	-2.58	0.76
Ethy.	-1.83	-1.38
Ghana	-1.17	-2.52
Hung.	-0.51	5.37
India	0.32	-0.71
Indon.	1.77	-2.06
Kenya	2.01	2.27
Korea	-3.98	-1.38
Malay.	-2.84	-4.79
Nepal	-0.39	0.28
Pak.	-1.21	-3.28
Philip.	-0.55	-2.48
S.Lank.	-5.92	0.64
Thay.	-2.54	-5.17
Turk.	1.87	4.06
Venez.	0.83	7.50
t-bar	**3.55**	**8.35**

Table 8

Bootstrap Panel Unit Root Test (π-values)

	s_t	log-s_t	p_t	log-p_t
Alger.	0.429	-0.846	0.223	-1.501
Boliv	0.429	-0.846	0.598	-0.514
Col.	0.901	-0.104	0.516	-0.662
D.Rep	0.842	-0.172	0.858	-0.153
Egypt	0.332	-1.103	0.845	-0.168
Ethiopia	0.429	-0.846	0.223	-1.501
Ghana	0.324	-1.127	0.142	-1.952
Hung.	0.842	-0.172	0.994	-0.006
India	0.919	-0.084	0.416	-0.877
Indon.	0.391	-0.942	0.152	-1.884
Kenya	0.999	-0.001	0.953	-0.048
Korea	0.039	-3.244	0.223	-1.501
Malaysia	0.146	-1.924	0.011	-4.605
Nepal	0.858	-0.153	0.671	-0.401
Pakistan	0.588	-0.531	0.042	-3.171
Philip.	0.736	-0.306	0.147	-1.917
S.Lank.	0.005	-5.298	0.718	-0.331
Thay.	0.311	-1.172	0.009	-4.710
Turk.	0.408	-0.897	0.968	-0.032
Venez.	0.957	-0.044	0.993	-0.0070
		-19.82		-25.941
Π_λ-test		**39.625**		**51.881**
CV5%		55.759		
CV1%		63.691		

Table 9

Panel Cointegration Tests

Pedroni (1997)-Statistics	
Panel v-stat.	4.99
Panel rho-stat.	-2.43
Panel pp-stat.	-1.79
Panel ADF-stat.	-2.61
Group rho-stat.	-5.93
Group pp-stat.	-3.57
Group ADF-stat.	-4.86
McCoskey and Kao (1998)	
LM^{*}	-3.41

Note: (a) The LM^{*} test is one-sided with a critical value of 1.64 (i.e. LM*>1.64 implies rejection of the null hypothesis of cointegration). The mean and variance used for calculating the McCoskey and Kao (1998) statistic are respectively 0.0850 and 0.0055 (MacCoskey and Kao, 1998, Table 2).

(b) The mean and variance used for calculating the Pedroni statistics were obtained from Pedroni (1999, Table 2). The number of lag truncations was set to 1. The Pedroni tests is a one-sided test. All statistics, with the exception of the v-statistic, have a critical value of –1.64 (i.e.κ<-1.64 implies rejection of the null of no cointegration). The v-statistic has has a critical value of 1.64 (i.e. κ>1.64 suggests rejection of the null of no cointegration).

Table 10

Long-Run Equilibrium PPP (eq.1): DOLS/DGLS Estimates

Country	α	β_o	β_1	**Adj*R*^{2}**	**Akaike**	***Pr*[*Fa*]**	**Lead/Lag**
Nepal	1.61 [7.57]	1.45 [23.50]	-0.86 [-0.82]	0.985	-1.23	0.96	2
Pak.	1.16 [8.27]	1.2 [17.7]	-0.66 [-7.36]	0.977	-0.74	0.54	1
Philipp.	3.38 [12.89]	1.17 [24.29]	-1.13 [-11.18]	0.988	-0.56	0.38	2
S.Lanka	-0.69 [-1.09]	0.44 [3.18]	0.6 [2.36]	0.798	0.87	0.23	3
Thay	2.1 [36.39]	-0.78 [-6.56]	1.06 [8.48]	0.899	1.12	0.57	1
Turkey	4.66 [12.28]	0.84 [62.81]	0.46 [5.13]	0.965	0.99	0.78	1
Venez	-3.01 [-0.49]	0.58 [2.05]	1.16 [0.76]	0.995	-2.31	0.22	2
Indon.	3.13 [0.27]	-0.94 [-0.97]	1.625 [0.532]	0.947	1.01	0.908	1
Kenya	5.81 [14.97]	1.52 [22.53]	-1.77 [-12.96]	0.947	0.96	0.97	2
S.Korea	7.1 [16.24]	0.98 [7.19]	-1.02 [-4.47]	0.937	1.27	0.87	3
Malay.	1.49 [4.80]	-0.49 [-0.49]	0.38 [3.36]	0.862	-1.13	0.21	1
Ethiopia	-2.53 [-1.03]	0.106 [0.26]	0.91 [1.15]	0.924	-1.35	0.89	2
Ghana	-25.2 [-3.01]	-0.13 [-0.43]	7.12 [3.36]	0.993	-0.59	0.34	2
Hung.	2.1 [12.01]	0.39 [12.74]	0.23 [4.39]	0.942	1.56	0.981	1
India	1.33 [11.02]	1.41 [29.05]	-0.92 [-14.27]	0.911	0.871	0.898	3

Alger.	1.2 [1.34]	0.59 [3.21]	-1.1 [-1.56]	0.891	-0.91	0.68	3
Bolivia	10.15 [2.27]	-0.043 [-0.633]	-1.49 [-1.41]	0.641	-0.881	0.871	1
Col.	11.56 [22.64]	1.53 [38.45]	-2.38 [-16.68]	0.932	-0.78	0.58	1
D.Rep.	0.99 [45.09]	-0.51 [-7.61]	0.72 [3.03]	0.899	1.34	0.75	1
Egypt	-0.45 [-1.34]	-0.913 [-20.64]	0.67 [6.13]	0.897	-0.845	0.75	1

Note: Numbers within [..] below regression coefficients are t-values. Akaike is the information criterion used for determining the number of leads and lags in the model. *Pr*[*Fa*] is the probability value of an F version of the Breusch-Godfrey test for first-order autocorrelation. The equations for Venezouela [AR(1}], Indonesia [AR(1)], Ethiopia [AR(2)] and Ghana [AR(1)] were estimated with the DGLS method. All other estimates are DOLS.

Table 11

Larsson *et al* (2001) Panel CointegrationTest

Country	Lags	r=0	r=1	r=2	r_i
Algeria	7	21.39	9.44	0.13	0
Col.	7	73.18	13.91	5.63	3
D.Rep.	3	66.42	26.37	0.29	2
Egypt	4	85.37	35.47	5.45	3
Ethiopia	2	53.46	10.53	0.01	1
Ghana	6	40.08	9.251	0.09	1
Hung.	5	51.62	16.21	1.04	2
India	3	40.87	14.31	0.87	2
Indon.	2	80.07	8.361	1.91	1
Kenya	2	61.12	28.98	1.91	2
Korea	7	29.11	13.47	0.83	2
Malay.	4	32.61	8.758	0.11	1
Nepal	6	37.05	14.16	1.98	2
Pakistn	3	49.67	14.58	0.03	2
Phili.	3	47.39	14.44	0.57	2
S.Lanka	3	88.21	38.61	1.56	2
Thail.	2	63.98	20.56	2.65	2
Turk.	2	55.49	22.67	0.04	2
Venez.	4	57.36	26.83	2.23	2
Boliv.	2	89.15	26.99	0.11	2
		56.179	18.694	1.3726	
Y_{LR} -test		**37.07**	**10.07**	**0.71**	

Note: The critical values for $E(Z_k)$ and $VAR(Z_k)$ were obtained from Larsson *et al* (2001, Table 1). These are respectively 14.955 and 24.733 for r = 0; 6.086 and 10.535 for r = 1; 1.137 and 2.212 for r = 2.

Table 12

Johansen (1995) Likelihood Ratio Test

Country	**LR-test**
Algeria	10.0101
Col.	0.87196
D.Rep.	0.6911
Egypt	22.2445
Ethiopia	6.5263
Ghana	7.9276
Hung.	1.8144
India	10.0597
Indon.	1.1491
Kenya	3.2395
Korea	14.04
Malay.	7.6194
Nepal	7.3967
Pak.	8.8372
Phili.	7.1733
S.Lanka	6.3645
Thail.	12.268
Turk.	6.3615
Venez.	12.6794
Boliv.	9.2603
PLR-test	**156.53**

Note: The individual country statistic follows a χ^2 distribution with 1 d.f. The panel PLR test follows a χ^2 distribution with d.f. 1N, where N is the cross section dimension.

Table 13

LR-Test for Over-Identifying Restrictions

Country	**LR-test**
Nepal	3.02*
Pakistan	16.16
Philip.	15.03
S.Lanka	25.49
Thay	12.27
Turkey	3.99*
Venez	1.41*
Indon.	40.25
Kenya	4.11*
S.Korea	3.88*
Malaysia	17.59
Etiopia	16.74
Ghana	17.59
Hungary	7.93
India	8.42
Alger.	4.57*
Bolivia	18.35
Col.	0.88*
D.Rep.	17.51
Egypt	40.89
PLR	276.08

Note: The panel PLR statistic follows a χ^2 distribution with d.f. 2N (40). Asterisks indicate significant statistics.