RE-EXAMINATION OF THE PURCHASING POWER PARITY IN CENTRAL AND EASTERN EUROPEAN ECONOMIES

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Abstract

The empirical record on PPP tests for transition economies is far from being so opulent as it is for the developed market economies. This paper attempts to fill this gap by scrutinizing the theory of PPP on a sample of Central and Eastern European economies. The paper has two main advances with respect to previous PPP studies. First, it employs a monthly data base on real exchange rates for a panel of 12 Central and Eastern European economies by testing the theory separately with respect to US dollar and with respect to Euro. Second, we utilise a panel unit root test that involves the estimation of the ADF regression in a SUR framework. Since our study found support for the validity of PPP in some reforming European economies, special attention should be devoted to individual country-specific factors that cause PPP deviations.

Key words: *Purchasing power parity, Real exchange rates, Stationarity test*

1. INTRODUCTION

The debate about the validity of purchasing power parity (PPP) has a long history in economic theory (Taylor, 2006) and is accompanied by an extensive empirical research. Although the majority of empirical tests has produced rather mixed outcomes, researchers generally agree that real exchange rates tend to converge toward levels predicted by PPP in the long-run, yet short-run deviations from the PPP relationship could be substantial (Rogoff, 1996). The relative instability in real exchange rate movements of transforming European economies since the beginning of the nineties, which might be in conflict with propositions of PPP theory, is explained in the literature by a range of factors, including inherited macroeconomic imbalances in these countries, mixed performance of chosen exchange rate arrangements, monetary difficulties arising from huge capital inflows, the inflationary impact of wage and price adjustments, and real exchange rate appreciation due to the catching-up process (Égert et al., 2006). Despite of growing interest for PPP in transition economies, the empirical evidence for this group of countries is far from being so comprehensive as it is for developed market economies. Examples of studies on PPP for European transition countries include inter alia Christev and Noorbakhsh (2000), Payne et al. (2005), Barlow (2004), Sideris (2006), Solakoglu (2006) and Koukouritakis (2009), while an in-depth survey of relevant empirical results for these economies can be found in Bahmani-Oskooee and Hegerty (2009).

This paper aims to expand the investigation of PPP for a group of 12 Central and Eastern European economies with respect to US dollar and Euro by using a battery of panel unit root tests. The paper proceeds as follows. In Section 2, after describing the general model of PPP and presenting the relevant data, the methodology of testing for stationarity of real exchange rates is elaborated. Section 3 reports the stationarity properties of the examined real exchange rates. Concluding remarks are given in the final section of the paper.

2. THE METHODOLOGY OF TESTING THE PPP

The basic model of testing for relative PPP can be derived in the following form (Froot and Rogoff, 1995):

$$e_t = \alpha_0 + \alpha_1 p_t + \alpha_2 p_t^* + \xi_t \qquad (1),$$

where e_t stands for nominal exchange rates, defined as the price of foreign currency in the units of domestic currency; p_t denotes domestic price index and p_t^* foreign price index; while ξ_t stands for the error term showing deviations from PPP. All the variables are given in logarithmic form. The strict version of PPP contains two types of restrictions imposed on the parameters. Under α_0=0, the symmetry restriction applies such that α_1 and α_2 are equal in absolute terms, whereas the limitation of α_1 and α_2 being equal to 1 and -1, respectively, is called the proportionality restriction.

In the present study we relied on relevant monthly data frequency covering the period of January 1994–December 2008 for the following countries: Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Macedonia, Poland, Romania, Slovakia and Slovenia. Primary data included monthly averages of nominal exchange rates and consumer price indices gathered from the central banks of individual countries, from the European Central Bank, Eurostat, and from national statistical offices of individual countries. Each of the exchange rates has been defined as the number of units of domestic currency for the US dollar and for the Euro. Consumer price indices used in this study for all countries refer to January 1994.

The empirical analysis consists of testing the characteristics of real exchange rates (strict version of Equation 1). Following relative PPP, the movements in nominal exchange rates are expected to compensate for price level shifts. Thus, real exchange rates should be constant over the long-run and their time series should be stationary (Parikh and Wakerly, 2000). The real exchange rates are a function of nominal exchange rates and relative price indices in two observed economies. They are calculated from the nominal exchange rates using the consumer price indices:

$RE_t = E_t\,(P_t^*/\,P_t)$

(2),

where RE_t stands for the real exchange rate, E_t is the price of a foreign currency in units of the domestic currency, and P_t^* and P_t represent the foreign price index and the domestic price index, respectively. Taking the logarithms of Equation 2, the real exchange rates are defined as:

$$re_t = e_t + p_t^* - p_t \quad (3).$$

The general model of testing for the presence of unit root takes into account the following AR(1) process for panel data:

$$y_{i,t} = \rho_i y_{i,t-1} + X_{i,t}\delta_i + \varepsilon_{i,t} \quad (4),$$

where i represents N cross-section units observed over periods $t=1, 2, ..., T_i$, $X_{i,t}$ are exogenous variables in the model (any fixed effects or individual trends), ρ_i are autoregressive coefficients, while errors ($\varepsilon_{i,t}$) are assumed as mutually independent idiosyncratic disturbances. If absolute value of autoregressive coefficients is less then 1, y_i is said to be weakly stationary. If the absolute value of autoregressive coefficients is 1, y_i contains a unit root. There are two assumptions about the autoregressive coefficients in panel unit root tests: first, persistence parameters are common across cross-sections ($\rho_i=\rho$) for all i, and second, ρ_i vary across cross-sections. Among tests with common unit root processes we utilized the test by Levin, Lin and Chu (2002), while Im, Pesaran and Shin (2003), Fisher ADF and Fisher PP (Madala and Wu, 1999; Choi, 2001) tests assume individual unit root processes.

Levin, Lin and Chu test (Levin et al. 2002) is based on ADF specification:

$$\Delta y_{i,t} = \alpha y_{i,t-1} + \sum_{j=1}^{p_i} \beta_{i,j}\Delta y_{i,t-j} + X'_{i,t}\,\delta + \varepsilon_{i,t} \quad (5),$$

where a common $\alpha=\rho-1$ is assumed, while the lag order for difference terms (p_i) varies across cross-sections. Under the null hypothesis (H_0: $\alpha=0$), there is a unit root. Under the alternative hypothesis (H_1: $\alpha<0$), there is no unit root. Levin et al. (2002) estimate α from proxies for $\Delta y_{i,t}$ and $y_{i,t}$ that are standardized and free of autocorrelations and deterministic components:

$$\tilde{y}_{i,t} = \alpha\tilde{y}_{i,t-1} + \eta_{i,t} \quad (6).$$

Levin et al. (2002) derive modified t statistics (t*) for resulting $\hat{\alpha}$ and show that it is asymptotically normally distributed:

$$\boldsymbol{t_a * = \frac{t_a - (N\widetilde{T})S_N \hat{\sigma}^{-2} se(\hat{\alpha})\mu_{m\widetilde{T}} *}{\sigma_{m\widetilde{T}} *} \rightarrow N(0,1)} \tag{7}$$

where t_α is standard t-statistic for $\hat{\alpha} = 0$, $\hat{\sigma}^2$ is estimated variance of the error term η, se(α) is standard error of $\hat{\alpha}$ and

$$\widetilde{T} = T - (\sum_i p_i / N) - 1 \tag{8}$$

Abbreviations $\mu_{m\widetilde{T}} *$ and $\sigma_{m\widetilde{T}} *$ in Equation 7 refer to adjustment terms for mean and standard deviation, while S_N is standard deviation ratio estimated by kernel-based techniques. In our testing procedure, number of lags used in each cross-section ADF regression (p_i) was defined by Schwarz information criterion using maximum 12 lags. Computation of S_N was conducted by Newey-West bandwidth selection using Bartlett kernel.

Im, Pesaran and Shin (2003) base their test on the assumption of individual unit root processes and they estimate individual ADF regression for each cross-section:

$$\boldsymbol{\Delta y_{i,t} = \alpha_i y_{i,t-1} + \sum_{j=1}^{p_i} \beta_{i,j} \Delta y_{i,t-j} + X'_{i,t} \delta + \varepsilon_{i,t}} \tag{9}$$

where the null hypothesis is

$$H_0 : \alpha_i = 0, \quad \text{for all i} \tag{10}$$

while the alternative hypothesis is defined as:

$$\boldsymbol{H_1 :} \begin{cases} \boldsymbol{\alpha_i = 0} & \text{for} \quad \mathrm{i} = 1,2,\ldots,\mathrm{N}_1 \\ \boldsymbol{\alpha_i < 0} & \text{for} \quad \mathrm{i} = \mathrm{N}_1 + 1, \mathrm{N}_1 + 2, \ldots, \mathrm{N} \end{cases} \tag{11}$$

The average t-statistics for α_i from separate ADF regressions is adjusted (standardized) to derive the desired test statistics (W). Number of lags was in this testing procedure again defined by Schwarz information criterion using maximum 12 lags.

Maddala and Wu (1999) and Choi (2001) propose to use Fisher's (1932) results to model tests that combine p-values from separate unit root tests. If π_i is defined as p-value from individual unit root test (ADF and PP) for cross-section i, then there is asymptotic result distributed as:

$$-2\sum_{i=1}^{N} log(\pi_i) \rightarrow \chi^2_{2N} \tag{12}$$

Additionally, if Φ^{-1} is the inverse of the standard normal cumulative distribution function, Choi Z-statistic is distributed normally:

$$Z = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1}(\pi_i) \rightarrow N(0,1) \tag{13}$$

The null and alternative hypotheses correspond to Im, Pesaran and Shin test (Equations 10 and 11). Number of lags used in each cross section ADF regression was specified by Schwarz information criterion using maximum 12 lags, while for the PP form of the test Newey-West bandwidth selection using Bartlett kernel was applied.

The common characteristic of the panel unit root tests presented above is that they deliver conclusions only about stationarity properties of the panel as a whole and do not allow to detect how many and which one of the series in the panel satisfies the stationarity hypothesis. Following the shortcoming of previous panel unit root tests we employed the seemingly unrelated regressions augmented Dickey-Fuller test (SURADF) proposed by Breuer et al. (2001, 2002). The test is based on the system of ADF equations which can be represented as:

$$\Delta y_{1,t} = \alpha_1 + \beta_1 y_{1,t-1} + \sum_{j=1} \varphi_j \Delta y_{1,t-j} + u_{1,t}$$

$$\Delta y_{2,t} = \alpha_2 + \beta_2 y_{2,t-1} + \sum_{j=1} \varphi_j \Delta y_{2,t-j} + u_{2,t}$$

$$\Delta y_{N,t} = \alpha_N + \beta_N y_{N,t-1} + \sum_{j=1} \varphi_j \Delta y_{N,t-j} + u_{N,t} \tag{14}$$

where $\beta_j = (\rho_j - 1)$ and ρ_j is the autoregressive coefficient for series *j*. This system is estimated by SUR procedure and the null and the alternative hypotheses are tested individually as

$$H_0^1 : \beta_1 = 0; \qquad H_A^1 : \beta_1 < 0$$

$$H_0^2 : \beta_2 = 0; \qquad H_A^2 : \beta_2 < 0$$

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$$H_0^N : \beta_N = 0; \qquad H_A^N : \beta_N < 0 \tag{15}$$

with the test statistics computed from SUR estimates of system (14), while the critical values are generated by Monte Carlo simulations. The procedure posed several advantages of, first, by exploiting the information from the error covariances and allows for autoregressive process, it produce efficient estimators over the single equation methods. Second, the estimation also allows for heterogeneity lag structure across the panel members. Third, the SURADF panel integration test allows us to identify which members of the panel contain a unit root.

As this test has non-standard distributions, the critical values of the SURADF test must be obtained through Monte Carlo simulations. In the simulations, the intercepts, the coefficients on the lagged values for each series were set equal to zero. In what follows, the lagged differences and the covariances matrix were obtained from the SUR estimation on the actual data. The SURADF test statistic for each of the 12 series was computed as the t-statistic calculated individually for the coefficient on the lagged level. To obtain the critical values, the experiments were replicated 10,000 times and the critical values of 1%, 5% and 10% are tailored to each of the 12 panel members.

3. EMPIRICAL RESULTS

The results of the panel unit root tests are summarized in Table 1 and Table 2. All the estimations were performed with constant as well with constant and trend variable.

Table 1: Results of panel unit root tests for US dollar rates

Test	Constant	Constant and trend
Levin, Lin and Chu (t*)	-1.155 (0.124)	-1.631 (0.051)
Im, Pesaran and Shin (W-stat)	1.340 (0.910)	0.818 (0.793)
Fisher ADF (χ^2)	15.288 (0.912)	17.450 (0.829)
Fisher ADF (Choi Z-stat.)	1.408 (0.920)	0.909 (0.819)
Fisher PP (χ^2)	20.733 (0.654)	20.754 (0.653)
Fisher PP (Choi Z-stat.)	1.103 (0.865)	0.749 (0.773)

Note: Figures in parentheses are p-values.

Unequivocally, the null hypothesis cannot be rejected for the case when US dollar is the base currency, while in the case of Euro rates one can reject the null and confirm the PPP theory in the panel of observed countries. Furthermore, with the single exception of Levin, Lin and Chu test for the US dollar as the numeraire currency, the empirical results in Table 1 and Table 2 appear to be insensitive to considering the time trend in the models.

Table 2: Results of panel unit root tests for Euro rates

Test	Constant	Constant and trend
Levin, Lin and Chu (t*)	-4.996 (0.000)	-12.863 (0.000)
Im, Pesaran and Shin (W-stat)	-1.856 (0.032)	-7.374 (0.000)
Fisher ADF (χ^2)	54.815 (0.000)	131.841 (0.000)
Fisher ADF (Choi Z-stat.)	-1.858 (0.032)	-6.596 (0.000)
Fisher PP (χ^2)	104.015 (0.000)	104.185 (0.000)
Fisher PP (Choi Z-stat.)	-2.856 (0.002)	-5.282 (0.000)

Note: Figures in parentheses are p-values.

The empirical findings from SURADF test for the panel with the US dollar as the numeraire currency reveal that 7 out of 12 countries' real exchange rates are stationary which is consistent with assumption of PPP theory (Table 3). In addition, the figures in Table 3 testify that the hypothesis about the unit root process can be rejected for the same set of countries irrespective of whether a trend variable is excluded or included into the estimation procedure.

Table 3: Results from the SURADF and the critical values (US dollar rates)

Country	Test statistics	Critical values			Test statistics	Critical values		
	Constant	0.01	0.05	0.10	**Constant and trend**	0.01	0.05	0.10
Bulgaria	-5.108 (4)***	-3.888	-3.252	-2.933	-5.130 (4)***	-3.902	-3.319	-2.995
Croatia	-2.595 (7)	-4.602	-3.997	-3.632	-2.589 (7)	-4.258	-3.638	-3.292
Czech Republic	-0.442 (1)	-4.181	-3.536	-3.210	-1.333 (1)	-4.595	-3.981	-3.645
Estonia	-5.737 (3)***	-5.261	-4.573	-4.187	-5.747 (3)***	-4.616	-4.013	-3.653
Hungary	-3.715 (5)	-5.003	-4.288	-3.929	-3.700 (5)	-4.651	-4.053	-3.705
Latvia	-5.030 (3)***	-4.785	-4.101	-3.738	-5.018 (3)***	-4.369	-3.696	-3.325
Lithuania	-4.067 (4)**	-4.275	-3.634	-3.275	-4.201 (4)***	-3.881	-3.287	-2.977
Macedonia	-4.304 (7)**	-4.399	-3.759	-3.413	-4.151 (7)**	-4.228	-3.584	-3.242
Poland	-4.004 (5)*	-4.703	-4.015	-3.643	-4.002 (5)**	-4.591	-3.929	-3.583
Romania	-4.555 (4)***	-4.020	-3.436	-3.083	-4.956 (4)***	-4.000	-3.404	-3.062
Slovakia	-2.460 (11)	-4.842	-4.167	-3.798	-2.830 (11)	-3.988	-3.339	-3.019
Slovenia	-3.586 (7)	-5.451	-4.788	-4.414	-3.507 (7)	-4.618	-3.950	-3.610

Note: The estimated critical values are tailored by the simulation experiments based on 187 observations for each series and 10,000 replications, following the work by Breuer et al. (2002). The error series were generated in such a manner to be normally distributed with the variance-covariance matrix given from the SUR estimation of the 12 countries panel structures. Each of the simulated real exchange rates was then generated from the error series using the SUR estimated coefficients on the lagged differences. (***), (**) and (*) denotes statistically significance at the 0.01, 0.05 and 0.10 level, respectively. Figures in parentheses indicate the lag length. The estimations and the calculation of the SURADF were carried out in RATS 5.02 using the algorithm kindly provided by Myles Wallace.

The empirical findings from SURADF test for the panel with the US dollar as the numeraire currency reveal that 7 out of 12 countries' real exchange rates are stationary which is consistent with assumption of PPP theory (Table 3). In addition, the figures in Table 3 testify that the hypothesis about the unit root process can be rejected for the same set of countries irrespective of whether a trend variable is excluded or included into the estimation procedure.

When the Euro is used as the numeraire currency, the SURADF tests indicate that we are able to reject the null hypothesis of unit root for 5 out of 12 cases (Table 4). In other words, the validity of PPP is confirmed for 5 real exchange rates with respect to the Euro. The stationarity of real exchange rates against the Euro in case of Bulgaria, Croatia, Hungary, Poland and Slovakia holds also when the presence of the time trend is considered in the model (Table 4); similarly, this is true for seven currencies with respect to the US dollar (Table 3).

Table 4: Results from the SURADF and the critical values (Euro rates)

Country	Test statistics	Critical values			Test statistics	Critical values		
	Constant	0.01	0.05	0.10	**Constant and trend**	0.01	0.05	0.10
Bulgaria	-4.578 (4)***	-4.114	-3.484	-3.149	-4.756 (4)***	-4.134	-3.476	-3.161
Croatia	-7.507 (2)***	-5.384	-4.849	-4.544	-7.742 (2)***	-5.661	-5.125	-4.799
Czech Republic	-3.528 (4)	-5.294	-4.756	-4.433	-2.718 (4)	-5.511	-4.897	-4.593
Estonia	-0.129 (6)	-5.475	-4.918	-4.616	-0.019 (5)	-5.713	-5.154	-4.848
Hungary	-9.514 (1)***	-5.356	-4.811	-4.515	-9.761 (1)***	-5.570	-4.997	-4.692
Latvia	-0.397 (12)	-4.069	-3.477	-3.162	-1.035 (12)	-4.204	-3.559	-3.216
Lithuania	-0.812 (7)	-5.320	-4.774	-4.449	-2.304 (4)	-5.496	-4.911	-4.585
Macedonia	-0.113 (7)	-5.465	-4.894	-4.597	-0.379 (3)	-5.692	-5.141	-4.836
Poland	-7.923 (2)***	-5.344	-4.765	-4.461	-8.324 (2)***	-5.479	-4.881	-4.588
Romania	-1.192 (7)	-5.042	-4.510	-4.212	-1.531 (7)	-5.226	-4.634	-4.298
Slovakia	-7.541 (2)***	-5.382	-4.820	-4.510	-7.833 (2)***	-5.621	-5.058	-4.750
Slovenia	-0.023(8)	-5.410	-4.878	-4.567	-0.906 (4)	-5.674	-5.113	-4.800

Note: The estimated critical values are tailored by the simulation experiments based on 180 observations for each series and 10,000 replications, following the work by Breuer et al. (2002). The error series were generated in such a manner to be normally distributed with the variance-covariance matrix given from the SUR estimation of the 12 countries panel structures. Each of the simulated real exchange rates was then generated from the error series using the SUR estimated coefficients on the lagged differences. (***), (**) and (*) denotes statistically significance at the 0.01, 0.05 and 0.10 level, respectively. Figures in parentheses indicate the lag length. The estimations and the calculation of the SURADF were carried out in RATS 5.02 using the algorithm kindly provided by Myles Wallace.

4. CONCLUSION

The results of empirical investigations on PPP for European transition economies has been pretty mixed, comprising studies that clearly reject this exchange rate theory as well studies that provide unreserved support for PPP proposition. In this paper we applied a range of panel unit root tests to re-examine the validity of PPP in a sample of 12 Central and Eastern European economies and consequently to extend the list of PPP studies for these group of countries.

According to the results from panel unit root tests that rely on a single statistic about the presence of stationarity, the theory of PPP is verified for the panel of real exchange rates with respect to the Euro, whereas the stationarity of real exchange rates in panel against the US dollar could not be confirmed. Our results, derived from the SURADF estimates, however, show that the PPP proposition holds approximately for half of the countries in the analyzed panel with respect to the US dollar as well to the Euro. Two basic conclusions can be derived from our research. First, the concept of PPP is corroborated for some, but not for all Central and Eastern European economies; whatever generalization about the validity of PPP theory for the group of reforming European economies is therefore unjustified. Second, judgment on the validity of parity conditions for individual country remains also conditioned by the choice of the numeraire currency. Country-cases, where a clear rejection of PPP assumption was found, might reflect exchange rate misalignment. On the other hand, cases of Euro-based series, where the PPP rule holds, provide an argument for an increasing coordination of national monetary and exchange rate policies and for a faster integration of these economies with the euro area.

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