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Purchasing Power Parity in Central and East European Countries

ABSTRACT: *This paper analyzes the validity of purchasing power parity (PPP) across twelve Central and Eastern European countries (CEECs). PPP is tested through linear and nonlinear unit root tests. The results suggest that when structural changes and nonlinearities are accounted for, the PPP hypothesis is supported for the CEECs.*

Long-run purchasing power parity (PPP), which suggests that differences in relative prices in two countries move together with nominal exchange rates in the long run, is one of the most important assumptions of many open economy macroeconomic models. The basis of the PPP proposition is the law of one price, which holds that the price of a given commodity (or bundle of commodities) shall be the same across all countries when expressed in terms of a common currency. Even though there may be small deviations in relative prices due to transportation costs and trade barriers, it is generally agreed that large deviations in relative prices will be traded away by arbitrageurs and prices equalize across countries in the long run (Rogoff 1996).

This paper examines the validity of the PPP proposition for twelve Central and East European (CEE) countries, namely, Bulgaria, Croatia, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Macedonia, Poland, Romania, Slovakia, and Slovenia. All these countries are EU members except Croatia and Macedonia, which are candidate countries. As Alba and Park (2005) and Bahmani-Oskooee et al. (2008) point out, it is important to examine the PPP hypothesis for these countries from

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several perspectives. First, measuring and comparing income across countries is usually based on PPP. If PPP does not hold, comparing income convergence among these countries and other EU countries may be misleading. Second, these countries are in the process of entering the euro zone, and thus, they need an estimate of equilibrium exchange rates before permanently linking to the euro. If PPP holds for these countries, then PPP rates as an equilibrium exchange rate measure may be used to estimate the appropriate exchange rates between the national currencies and the euro. Finally, the failure of PPP to hold may indicate exchange rate misalignments. Overvaluation of national currency relative to main trading partners will actually widen current account deficits and adversely affect the country's macroeconomic stability, a key precondition for entering the euro zone.

Although enormous empirical work examining PPP has accumulated to date, little work has been done for transition economies (for recent surveys see, e.g., Rogoff 1996; Taylor 2003). Apergis (2003), Barlow and Radulescu (2002), Bekő and Boršič (2007), Boršič and Bekő (2006), Doganlar (2006), Payne et al. (2005), Sideris (2006), and Thacker (1995) fail to find any evidence of PPP employing conventional univariate and multivariate cointegration tests. Bahmani-Oskooee et al. (2008), Barlow (2003), Cuestas (2009), and Koukouritakis (2009) conclude that PPP holds only for some of the countries under examination. On the other hand, Christev and Noorbakhsh (2000) and Varamini and Lisachuk (1998) find some weak evidence of PPP for transition countries. Using annual data for twenty-one transition countries and panel cointegration tests, Solakoglu (2006) concludes that PPP holds for these countries.

Our approach differs from previous studies on PPP in transition countries from several perspectives. First, except for Bahmani-Oskooee et al. (2008), all the above studies examine bilateral real exchange rates, whereas we use trade-weighted real effective exchange rates (REERs). We chose this particular measure because REER movements are crucial for studies of international trade flows. As Bahmani-Oskooee et al. (2008) argue, REER stationarity implies that PPP holds not only with respect to a country's bilateral trading partners, but also with respect to its many trading partners.[1] Second, all the above studies except Bahmani-Oskooee et al. (2008) and Cuestas (2009) test the validity of PPP within a linear framework. However, convergence of the real exchange rate to an equilibrium level might be nonlinear due to transaction costs. Therefore, the conclusions drawn from linear unit root and cointegration tests might be misleading. Third, though Bahmani-Oskooee et al. (2008) and Cuestas (2009) allow for nonlinear adjustment to the equilibrium level, they do not consider possible structural breaks in the data. Payne et al. (2005) and Koukouritakis (2009) allow for a structural break, but they fail to account for nonlinearities in the adjustment process. As Perron (1989) has shown, the power of unit root tests may decrease if there is a structural break, even if the data are indeed stationary. Considering specific features of real exchange rate series of transition countries, in addition to the conventional augmented Dickey–Fuller (ADF; Dickey and Fuller 1979) and Kwiatkowski et al. (1992; KPSS) tests, we apply the Kapetanios et al. (2003) nonlinear unit root test as well as the Sollis (2004) unit root test procedure allowing for asymmetric adjustment with smooth structural change in the data-generating process.

Our results suggest that allowing for nonlinearities and gradual structural changes in the data-generating process results in more rejection of the null hypothesis of unit root. Using linear unit root tests, we found that only five out of twelve series are stationary, consistent with the PPP proposition. Nonlinear unit root tests, on the other hand, imply that PPP holds for seven countries. However, after allowing for both gradual structural change and asymmetric adjustment, we conclude that the PPP proposition holds for all countries considered. These results point to the importance of taking account of both structural changes and nonlinear adjustment of real exchange rates in transition countries.

Methodology

Empirical investigations of PPP usually have been based on testing for the unit root in the real exchange rate series. The evidence is very controversial. Earlier studies on PPP that used conventional ADF tests reported results against PPP (e.g., Mark 1990; Meese and Rogoff 1988; Taylor 1988). The results were usually attributed to the low power of the ADF test in detecting stationarity in a small data sample. To overcome this problem, alternative frameworks to test for a unit root have been proposed. Frankel (1986) and Lothian and Taylor (1996) suggest using long-span data, in terms of years covered, to increase the power of unit root tests. But Hegwood and Papell (1998) and Papell and Prodan (2006) point out that very long spans of data cover both fixed and floating periods, and thus, may produce spurious results because of regime changes. Abuaf and Jorion (1990) and Frankel and Rose (1996) suggest using panel data to increase the number of observations and thus the power of tests. However, Taylor and Sarno (1998) and Taylor (2003) argue that testing PPP using panel unit root tests may be misleading because of heterogeneity problems. They argue that rejecting the null hypothesis of unit root in panel data implies that at least one of the series is mean reverting, but not that all the series under consideration are stationary. Another strand of the literature stresses the nonlinear nature of the dynamics of real exchange rates and claims that the PPP puzzle can be resolved once nonlinear adjustment toward equilibrium is properly modeled (Dumas 1992; Michael et al. 1997; Sarno and Taylor 2002; Sercu et al. 1995). Transaction costs, such as those of transportation and storage, render arbitrage unprofitable for small deviations from the equilibrium level. The arbitrageurs do not engage in trade if deviations from the equilibrium are small in size and therefore arbitrage is not profitable. If the deviations from equilibrium are large enough, however, arbitrageurs engage in profitable trading strategies, bringing the real exchange rate to equilibrium levels. Reversion of the real exchange rates to the equilibrium level, therefore, takes place only when the deviations from the equilibrium level are large. This implies that the dynamic behavior of real exchange rates differs according to the size of the deviation from the equilibrium, giving rise to asymmetric dynamics for real exchange rates. In addition to transaction costs, interaction of heterogeneous traders and diversity in agents' beliefs also may lead to persistent deviations from equilibrium.[2]

In the light of the above arguments, and considering that the CEE countries have undergone major structural changes during the analyzed period, in addition to the conventional ADF and KPSS tests, we also employ unit root tests that allow for nonlinearities and gradual structural change in the data-generating process.

Nonlinear Unit Root Test

The unit root test procedure proposed by Kapetanios et al. (2003) is based on the following exponential smooth transition (ESTAR) regression model:

$$\Delta q_t = \phi q_{t-1} + \gamma q_{t-1}\left[1-\exp\left(-\theta q_{t-1}^2\right)\right] + \varepsilon_t, \tag{1}$$

where q_t is the real exchange rate series. The transition function, $F(\theta, q_{t-1}) = 1 - \exp(-\theta q_{t-1}^2)$, is continuous, U-shaped around zero, and bounded from zero and one. The parameter θ measures the speed of transition between two regimes that correspond to extreme values of the transition function.

The ESTAR model has the nice property of allowing modeling different dynamics of series depending on the size of the deviations from the equilibrium level (e.g., Michael et al. 1997). As briefly discussed above, arbitrageurs do not engage in trade strategies if deviations from the equilibrium are small. If the deviations from equilibrium are large enough, however, arbitrageurs engage in profitable trading strategies, bringing the real exchange rate to its equilibrium level. In the context of the ESTAR model, this would imply that whereas $\phi \geq 0$ is possible, one must have $\gamma < 0$ and $\phi + \gamma < 0$ for the process to be globally stationary. Under these conditions, the process might display unit root for small values of q_{t-1}^2, but for larger values of q_{t-1}^2, it has stable dynamics, and as a result, is geometrically ergodic. As Kapetanios et al. (2003), show, the ADF test may not be very powerful when the true process is nonlinear yet globally stationary.

Imposing $\phi = 0$, which implies that q_t follows a unit root in the middle regime, the ESTAR model can be written as

$$\Delta q_t = \gamma q_{t-1}\left[1-\exp\left(-\theta q_{t-1}^2\right)\right] + \varepsilon_t. \tag{2}$$

The null hypothesis for the global stationarity of the process q_t can be written as H_0: $\theta = 0$ against the alternative H_1: $\theta > 0$. However, testing the null hypothesis directly is infeasible, as the parameter γ is not identified under the null. To circumvent this problem, Kapetanios et al. (2003) develop a t-type test statistic by replacing the transition function $F(\theta, q_{t-1}) = 1 - \exp(-\theta q_{t-1}^2)$ with its first-order Taylor approximation around $\theta = 0$, yielding the following regression model:

$$\Delta q_t = \delta q_{t-1}^3 + e_t, \tag{3}$$

where e_t contains not only original error term ε_t but also the error term resulting from Taylor approximation. The null hypothesis of unit root can be tested using t-statistics from that testing $\delta = 0$.

In the more general case, where errors in (3) are serially correlated, assuming that the serially correlated terms enter in a linear fashion, one may extend the auxiliary regression (3) to

$$\Delta q_t = \delta q_{t-1}^3 + \sum_{i=1}^{p} \beta_i \Delta q_{t-i} + e_t. \tag{4}$$

Although the Kapetanios et al. (2003) test procedure is convenient for testing the null hypothesis of unit root in the case of nonlinear adjustment, this test procedure does not account for possible structural breaks in the data-generating process. It is well known that the CEE countries have undergone major structural changes during the transition period (see, e.g., Fischer et al. 1996; Fischer and Sahay 2000; Foster and Stehrer 2007). The CEE countries had to implement a wide range of economic reforms, aiming at price and trade liberalization, privatization, demonopolization, and establishment of market institutions, to restructure their centrally planned economies to market economies. Such structural reforms caused an increase in the volume of international trade and reorientation of trade toward the European Union. Furthermore, integration of the CEE countries with the European Union during the accession period and aftermath intensified trade between these countries and older members of the European Union (e.g., Cheptea 2007; Fidrmuc 2005; Havrylyshyn and Al-Atrash 1998). Because the CEE countries have undergone several phases of economic restructuring during the transition and accession period, it is likely that equilibrium real exchange rates have shifted during the analyzed period (Bekő and Boršič 2007).

The failure to account for a possible structural break in the data-generating process may give rise to spurious results. Perron (1989) has shown that power of unit root tests may decrease if there is a structural change, even if the data are indeed stationary. To account for a structural break, Perron (1989) suggests adding dummy variables corresponding to a prespecified break date to the standard ADF regression. Subsequently, Perron (1997), Perron and Vogelsang (1992), and Zivot and Andrews (1992) have proposed new procedures that select the break date endogenously. Leybourne et al. (1998) argue that assuming instantaneous structural change may not be appropriate for economic time series and suggest a new test procedure that allows for gradual structural change. Sollis (2004) extends the test procedure of Leybourne et al. (1998) to allow for both gradual structural change and asymmetric adjustment. Accounting for the fact that the real exchange rate series of the CEE countries might exhibit structural breaks during the analyzed period, as well as the fact that adjustment of real exchange rates to equilibrium level might be nonlinear, we also apply the test procedure of Sollis (2004).

Unit Root Test Under Structural Break and Asymmetric Adjustment

Sollis (2004) combines the smooth transition methodology of Leybourne et al. (1998) with the threshold autoregressive methodology of Enders and Granger (1998) to develop unit root tests that allow for a smooth transition between deterministic linear trends, around which asymmetric adjustment may occur. Following Leybourne et al. (1998), Sollis (2004) models the structural change in the series by the following smooth transition regression model:

$$q_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 t S_t(\gamma, \tau) + \nu_t, \tag{5}$$

where ν_t is a zero-mean I(0) process and $S_t(\gamma, \tau)$ is the logistic smooth transition function, based on a sample size of T, $S_t(\gamma, \tau) = [1 + \exp\{-\gamma(t - \tau T)\}]^{-1}$. The transition function $S_t(\gamma, \tau)$ is a continuous function bounded between zero and one. The parameter τ determines the timing of the transition midpoint, as for $\gamma > 0$, we have $S_{-\infty}(\gamma, \tau) = 0$, $S_{+\infty}(\gamma, \tau) = 1$, and $S_{\tau T}(\gamma, \tau) = 0.5$. Thus, the transition function allows for two regimes, associated with the extreme values of the transition function, $S_t(\gamma, \tau) = 0$ and $S_t(\gamma, \tau) = 1$, whereas the transition from one regime to the other is gradual. The parameter γ determines the speed of transition between regimes. If γ is small, then $S_t(\gamma, \tau)$ takes a long period of time to traverse the interval (0, 1), and in the limiting case with $\gamma = 0$, $S_t(\gamma, \tau) = 0.5$ for all t. On the other hand, for large values of γ, $S_t(\gamma, \tau)$ traverses the interval (0, 1) very rapidly, and as $\gamma \rightarrow +\infty$, this function changes value from 0 to 1 instantaneously at time $t = \tau T$. Thus, if it is assumed that ν_t is a zero-mean I(0) process, then regression (5) implies that the series under investigation, q_t, is stationary around a nonlinear attractor, the mean of which changes from initial value α_1 to $\alpha_1 + \alpha_2$. The slope changes from initial value β_1 to $\beta_1 + \beta_2$ simultaneously and with the same speed of adjustment (Leybourne et al. 1998).

Whether the series under investigation q_t converges to a nonlinear attractor or not can be tested via a two-step procedure. In the first step, Equation (5) is estimated using a nonlinear least squares algorithm. In the second step, the following asymmetric ADF regression model is estimated for the residuals $\hat{\nu}_t$, obtained from the first step:

$$\Delta \hat{\nu}_t = I_t \alpha_1 \hat{\nu}_{t-1} + (1 - I_t)\alpha_2 \hat{\nu}_{t-1} + \sum_{i=1}^{p} \beta_i \Delta \hat{\nu}_{t-1} + \eta_t, \tag{6}$$

where $I_t = 1$ if $\hat{\nu}_{t-1} \geq 0$, $I_t = 0$ if $\hat{\nu}_{t-1} < 0$, and η_t is a zero-mean stationary process. If $\alpha_1 = \alpha_2 = 0$ in Equation (6), then $\hat{\nu}_t$ and therefore q_t contain a unit root, whereas if $\alpha_1 = \alpha_2 < 0$, q_t is a stationary smooth transition threshold autoregressive (ST-TAR) process with symmetric adjustment. If $\alpha_1 < 0$, $\alpha_2 < 0$, and $\alpha_1 \neq \alpha_2$, then q_t is a stationary ST-TAR process with asymmetric adjustment. Sollis (2004) suggests testing the null hypothesis of unit root in two different ways: using the F-statistic for testing $\alpha_1 = \alpha_2 = 0$ in (6) or the most significant of the t-statistics from those testing $\alpha_1 = 0$ and $\alpha_2 = 0$.

The next section provides the results of the above unit root tests for the real exchange rate series of CEE countries.

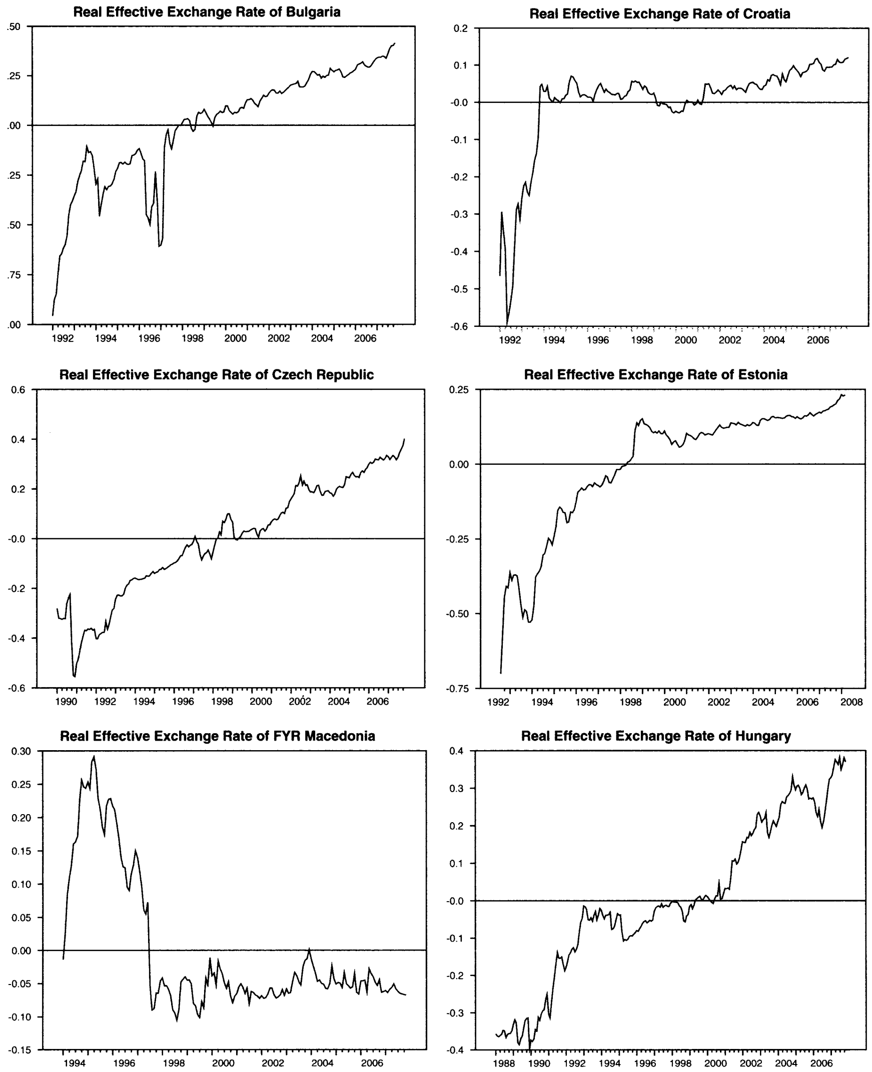
Data and Test Results

This paper uses monthly data on trade-weighted REERs. The data on REERs of Bulgaria, Croatia, the Czech Republic, Hungary, Macedonia, Poland, Romania, and Slovakia are taken from the International Monetary Fund's International Financial Statistics database. The data on Lithuanian and Estonian REERs were downloaded from the central banks of those countries, and the Latvian and Slovenian REERs were obtained from Eurostat. The sample period is 1992M1–2007M11 for Bulgaria and Croatia; 1990M1–2007M11 for the Czech Republic and Slovakia; 1992M8–2008M3 for Estonia; 1988M1–2007M11 for Hungary, Poland, and Romania; 1994M1–2006M12 for Latvia and Slovenia; 1993M7–2007M12 for Lithuania; and 1994M1–2007M11 for Macedonia. Figure 1 plots the logarithm of the REER series.

Perhaps one of the most prominent features of the real exchange rates of transition economies is that, as Figure 1 shows, they exhibit a trend appreciation following a short-run depreciation after the demise of the command economy (see, e.g., Brada 1998; Egert et al. 2006; Halpern and Wyplosz 1997). Trend appreciation of real exchange rates in transition countries was mainly attributed to, first, transition-specific adjustments in administered and regulated prices (Egert and Kutan 2005); second, correction to the initial sharp depreciation (De Broeck and Slok 2006; Halpern and Wyplosz 1997); third, rapid productivity gains due to economic restructuring and integration to the European Union (Brada 1998; Halpern and Wyplosz 1997; Kutan and Yigit 2007); and fourth, dramatic changes in quality and a progressive shift in both domestic and foreign consumers' preferences toward domestically produced goods (Egert and Kutan 2005; Egert et al. 2006).[3] Furthermore, Egert et al. (2006) argue that an increase in disposable income per capita may result in rising consumption, which falls increasingly on nontradable goods because of a high income elasticity of demand for nontradable goods. Therefore, demand-side pressures may also cause exchange rates to appreciate. In addition, an increase in income may cause exchange rates to appreciate by raising demand for domestic currency.

Accounting for the above-mentioned specific features of real exchange rates of the transition countries, we tested the stationarity of the REER series under investigation with and without a trend in the regression models. As Taylor (2003) and Bahmani-Oskooee et al. (2008) argue, if the real exchange rates revert to a trend, it can be considered as evidence of the Balassa–Samuelson effect.

Figure 1. **Real Effective Exchange Rates**

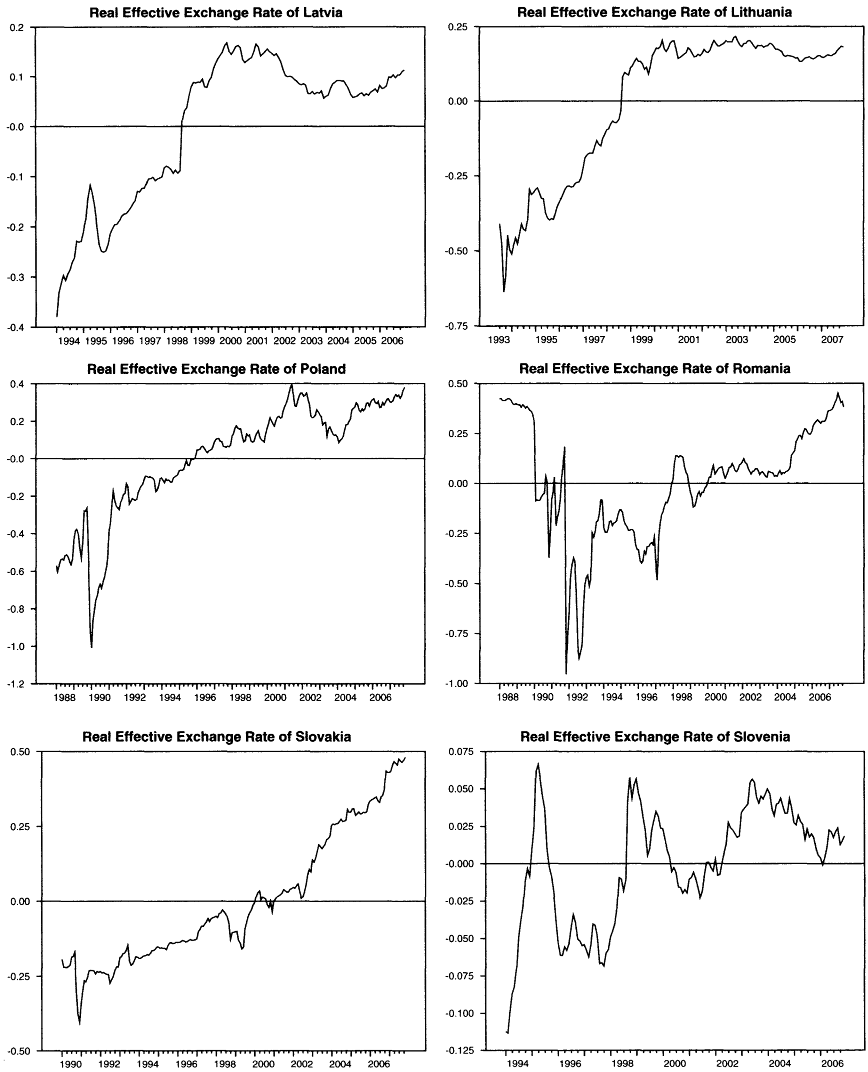


(continues)

Linear Unit Root Test Results

We first examine the stationarity of the REER data ignoring possible structural changes and nonlinearities in the data-generating process. We apply a conventional ADF test for this purpose. Frankel (1990) argues that, whereas a researcher may not

Figure 1. **(Continued)**



be able to reject the null hypothesis of unit root, it does not necessarily mean that the researcher must then accept that hypothesis. Therefore, in addition to the ADF test, we apply the KPSS test, which differs from the ADF test in that the former assumes that the series under investigation are stationary under the null hypothesis and the latter test assumes that series have a unit root under the null hypothesis. Table 1 contains the results of both tests.

Table 1

Linear Stationarity Test Results

	ADF test		KPSS test	
Country	Constant only	Constant and trend	Constant only	Constant and trend
Bulgaria	−2.660*	−4.347***	1.699	0.130*
Croatia	−2.184	−2.345	0.886	0.192**
Czech Republic	−0.177	−3.535**	1.809	0.156**
Estonia	−1.833	−2.482	1.486	0.388
Hungary	−0.548	−2.714	1.932	0.145*
Latvia	−2.094	−1.575	1.130	0.344
Lithuania	−2.017	−1.432	1.356	0.400
Macedonia	−1.550	−3.170*	0.877	0.270
Poland	−1.511	−2.987	1.788	0.327
Romania	−2.722	−3.192*	0.562**	0.286
Slovakia	1.022	−2.346	1.756	0.405
Slovenia	−3.408**	−3.569**	0.795	0.053***

*, **, and *** denote rejection of the null hypothesis of unit root at 10 percent, 5 percent, and 1 percent significance levels for the ADF test and no rejection of the null hypothesis of stationarity at 90 percent, 95 percent, and 99 percent significance levels for the KPSS test.

The results of the ADF and KPSS tests provide evidence of PPP only in a few cases. When no trend is included in the ADF regression, the null hypothesis of unit root is rejected for the Bulgarian and Croatian REER data, consistent with the PPP hypothesis. With a linear trend, however, the unit root null hypothesis is rejected in five out of twelve series, namely for the Bulgarian, Czech, Macedonian, Romanian, and Slovenian REER data. The KPSS test suggests that only Romanian REER data comprise a mean-reverting process, whereas Bulgarian, Croatian, Czech, Hungarian, and Slovenian REER data are trend-stationary processes. As neither conventional ADF nor KPSS tests account for nonlinear adjustment and structural breaks in data-generating processes, the conclusions drawn from these tests might be misleading. Therefore, we turn to the results of nonlinear unit root tests as well as unit root tests under structural break and asymmetric adjustment.

Nonlinear Unit Root Test Results

Table 2 presents the results of the nonlinear unit root test procedure proposed by Kapetanios et al. (2003). Demeaned and detrended data were obtained by first regressing the REER series of each country on a constant and on both a constant and a time trend, respectively, then saving the residuals.

Table 2

Nonlinear Unit Root Test Results

Country	Demeaned data	Detrended data
Bulgaria	–4.779***	–5.835***
Croatia	–4.284***	–4.894***
Czech Republic	–3.230**	–3.922**
Estonia	–3.435**	–4.145***
Hungary	–1.264	–2.879
Latvia	–2.110	–1.243
Lithuania	–2.991**	–1.798
Macedonia	–2.939**	–3.141*
Poland	–2.154	–1.706
Romania	–0.858	–1.269
Slovakia	0.718	–3.725**
Slovenia	–3.459**	–3.892**

***, ** and * denote rejection of the null hypothesis at 1 percent, 5 percent, and 10 percent significance levels. Critical values of the test statistic at 1 percent, 5 percent, and 10 percent significance levels are –3.48, –2.93, and –2.66 for the demeaned data and –3.93, –3.40, –3.13 for the detrended data, respectively (Kapetanios et al. 2003, p. 364).

As Table 2 reveals, allowing for nonlinear adjustment to the equilibrium level results in more frequent rejection of the unit root null hypothesis for the REER series. We can reject the null hypothesis of unit root in seven out of twelve demeaned REER data, namely for the Bulgarian, Croatian, Czech, Estonian, Lithuanian, Macedonian, and Slovenian data, consistent with the PPP hypothesis. Similarly, the unit root null hypothesis is rejected in seven cases, for the Bulgarian, Croatian, Czech, Estonian, Macedonian, Slovakian, and Slovenian detrended REER data. The results for the detrended REER data are consistent with Bahmani-Oskooee et al. (2008), except that they fail to reject the unit root null hypothesis for Estonian but rejected it for Romanian REER data. This discrepancy may be due to different time spans. All in all, our results from the nonlinear unit root tests imply that the REERs of most of the CEE countries adjust to the long-run equilibrium level nonlinearly, suggesting that real devaluations affect the international trade of those countries in a nonlinear fashion.

ST-TAR Unit Root Test Results

As Figure 1 reveals, there are breaks in the REER data of the CEE countries under investigation. Furthermore, the equilibrium real exchange rates of the CEE countries might have shifted as a result of economic restructuring during the transition and

Table 3

Results of the ST-TAR Unit Root Tests

Country	Date of structural break	*t*-max statistic	*F*-statistic
Bulgaria	1993:M12	–4.073**	8.312
Croatia	1994:M09	–4.223**	10.571
Czech Republic	1992:M03	–5.258***	21.200***
Estonia	1999:M09	–4.734***	15.844**
Hungary	1994:M12	–3.763*	12.035*
Latvia	1999:M03	–3.821*	11.823*
Lithuania	1998:M11	–4.047**	13.154**
Macedonia	1997:M06	–4.213**	13.229**
Poland	1991:M02	–3.833*	11.407*
Romania	1993:M03	–5.026***	12.806**
Slovakia	2003:M01	–5.689***	19.096***
Slovenia	2000:M05	–4.162**	10.739

Notes: The break date refers to the midpoint of the gradual structural change—that is, to $t = \tau T$ in the regression model (5). ***, **, and * denote rejection of the null hypothesis at 1 percent, 5 percent, and 10 percent significance levels. Critical values of the *t*-max and *F*-statistics at 1 percent, 5 percent, and 10 percent significance levels are –4.393, –3.937, –3.704, and 15.892, 12.787, and 11.349, respectively (Sollis 2004, p. 413).

accession periods. Thus, the failure to account for possible structural breaks in the real exchange rate series of these countries may lead to misleading results. For this purpose, we also apply the ST-TAR unit root test proposed by Sollis (2004), which allows for asymmetric adjustment to gradually changing trend function.[4] Table 3 presents the results of the ST-TAR test.

The results in Table 3 provide strong evidence of the stationarity of all the REER data under consideration. When judged by *F*-statistic, the null hypothesis of unit root is rejected in nine out of twelve cases, namely, the Czech, Estonian, Hungarian, Latvian, Macedonian, Polish, Romanian, and Slovakian REER data. However, the unit root hypothesis is rejected for all REER data by the *t-max* statistic. These results once again point to the importance of accounting for both structural changes and asymmetric adjustment in exchange rates of CEE transition countries, and may explain why other researchers failed to discover stationarity of real exchange rates in transition countries.

Another interesting issue is the timing of the structural breaks in the REER data of the CEE countries. The first column of Table 3 reports the estimated break dates. Structural breaks in the REER data for relatively small economies, such as

Estonia, Latvia, Lithuania, Slovakia, and Slovenia, occurred during the accession period, after the commencement of EU accession negotiations in December 1997, but before joining the European Union in May 2004. On the other hand, the break in the REER data occurred during the early 1990s for large transition countries, such as Bulgaria, the Czech Republic, Hungary, Poland, and Romania, as well as for relatively small economies, such as Croatia and Macedonia. These results imply that the break in the REER data of relatively large CEE countries occurred during the early transition period, possibly due to real shocks, such as the collapse of the socialist production and trade mechanisms and drastic changes in relative prices. The break in the data of relatively small countries occurred during the accession period, due to EU integration. This finding may reflect the fact that reforms and efforts aimed at preparation for EU membership might have shifted equilibrium real exchange rates in relatively small economies.

Conclusion

This paper analyzed the validity of the PPP proposition for twelve Central and East European transition countries. After the collapse of the command economy, these countries have implemented comprehensive reforms aimed at restructuring their economies and integration to the European Union. Considering that the CEE countries have undergone major structural changes during the transition period and that adjustment of exchange rates might be nonlinear, in addition to conventional unit root tests, we also applied unit root test procedures that allow for both structural breaks and asymmetric adjustment in the series under examination.

The results suggest that proper modeling of nonlinearities and structural changes in the data-generating process result in more rejection of the null hypothesis of unit root, consistent with the PPP proposition. When neither structural breaks nor nonlinearities are accounted for, we find some evidence in favor of PPP in only five countries. After accounting for nonlinearities in the data, we can reject the null hypothesis of unit root for seven real exchange rate series. However, after allowing for both structural breaks and asymmetric adjustment, we found that the PPP proposition holds for all countries.

Our results have clear implications. As the CEE countries are in the process of entering the euro zone, they need estimates of the equilibrium exchange rates before a permanent link to the euro. As we find that PPP holds for all countries under consideration, PPP rates may be used to estimate appropriate exchange rates between national currencies and the euro. Second, our findings suggest that REERs exhibit high nonlinearities, which, in turn, implies that real devaluations affect trade flows nonlinearly. Finally, our results indicate that structural breaks in the real exchange rate series of some relatively small economies occurred during the accession period. This finding implies that the accession efforts might have intensified trade flows between these countries and the European Union, and have shifted equilibrium real exchange rates in these countries. One must be cautious and account for both possible structural changes and nonlinearities when examining the exchange rate series of transition countries.

Notes

1. Cashin and McDermott (2003) and Bahmani-Oskooee et al. (2007) also tested PPP using trade-weighted REERs.

2. Taylor (2003) argues that there are at least three sources of nonlinear adjustment of real exchange rates: transportation costs and trade barriers, interaction of heterogeneous traders in the foreign exchange market, and official interventions in the foreign exchange market.

3. For a thorough discussion of possible reasons of trend appreciation of real exchange rates, see Halpern and Wyplosz (1997).

4. We also applied Zivot and Andrews's (1992) unit root test that allows for an instantaneous structural break in both the mean and trend of the data. The Zivot–Andrews (1992) test rejected the null of unit root only in seven out of twelve cases, namely, for Bulgarian, Czech, Latvian, Lithuanian, Macedonian, Romanian, and Slovakian REER data. The break dates chosen by the Zivot–Andrews (1992) test procedure were quite similar for all countries to those chosen by the Sollis (2004) test. To save space, we do not provide results of the Zivot–Andrews (1992) test, which are available upon request.

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