Testing the Purchasing Power Parity Hypothesis for the New Member and Candidate Countries of the European Union: Evidence from Lagrange Multiplier Unit Root Tests with Structural Breaks

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ABSTRACT: This paper investigates the validity of purchasing power parity (PPP) for the eleven Central and East European transition countries and three market economy countries, Cyprus, Malta, and Turkey. Unlike previous studies on PPP, this study uses Lagrange multiplier (LM) unit root tests that incorporate structural breaks in the data series. The findings indicate that in cases of one and two structural breaks, for a U.S. dollar–based real exchange rate series, there is little evidence supporting the validity of PPP. For a deutsche mark–based real exchange rate series, for the cases of both one and two breaks, there is evidence of stationarity of real exchange rates for eight sample countries, which is consistent with PPP. The results also indicate that the estimated half-life of a shock to the real exchange rate ranges from 1.25 (15.05 months) to 2.72 (32.72 months) years across countries. The empirical findings may provide direction for policy makers to coordinate monetary policies for the process of European monetary integration.

KEY WORDS: Central and Eastern European countries, European monetary integration, purchasing power parity, structural breaks.

Purchasing power parity (PPP), which considers a particular relation between exchange rates and relative national prices, is one of the central theoretical concepts in determining long-run real exchange rates in international finance. The absolute version of PPP is based on the law of one price, which states that, in terms of a common currency, the price of an identical good or service should be the same in two countries. The relative PPP, however, states that a change in the exchange rates should equal the inflation differential and bilateral exchange rates should adjust to account for these differentials in the long run. PPP assumes that any differences in the goods market should be reflected in the currency market.1

The Central and East European (CEE) countries have undergone significant institutional and structural changes to create market-oriented economies in the 1990s. During the economic transformation, these countries focused on price and market liberalization, trade reform, various privatization programs, and competition policy. They adopted new
exchange rate regimes, developed financial institutions, and promoted the inflow of foreign direct investment. The overall purpose of their restructuring and liberalization processes has been to integrate progressively into the greater European political and economic structure. Most countries have successfully brought inflation under control, stabilized their currencies, and begun to experience productivity growth. Eight of these countries—the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia, and Slovenia—as well as Cyprus and Malta joined the European Union (EU) in May 2004. Bulgaria and Romania joined in January 2007, and Croatia is expected to follow by 2010. Turkey, which is not a transition economy or an EU member, started negotiations in October 2005. A further step for the twelve new EU member countries and two candidate countries is to satisfy the Maastricht convergence criteria to join the European Monetary Union (EMU). Apart from reforming their real economic structures, the countries must adopt similar monetary policies and exchange rate regimes to enlarge the euro area. Inflation convergence and nominal exchange rate stability among EU members are the two major convergence criteria required for monetary integration. These two criteria imply the stability of real exchange rates and, eventually, real exchange rate convergence. For PPP to hold in the long run, the real exchange rate has to revert to a constant level over time. One would eventually expect that achieving PPP occurs with the Maastricht convergence criteria. Therefore, testing the empirical validity of PPP for potential EMU accession countries has significant implications for policy makers in the recent enlargement, as real exchange rate convergence among these countries will strengthen their economic integration into the euro area.

This paper examines the empirical validity of PPP for the new CEE EU member countries, as well as Malta and Cyprus and the two candidate countries, Croatia and Turkey. Koedijk et al. (2004) specify three reasons to test PPP in the euro area: the opportunity to test PPP for a common currency, the importance of inflation convergence for policy makers, and the effect of real exchange rate risk on financial markets. It is also interesting to clarify the influence of real exchange rate behavior on the European economic integration process, particularly for the CEE countries. There are two stylized facts about these countries. First, the transition process leads to real productivity shocks, and thus long-run (permanent) deviations from PPP. Second, most of these countries have experienced massive monetary shocks that might have been caused by high inflationary pressures, and thus short-run (temporary) deviations from the PPP (Halpern and Wyplosz 1996). Moreover, foreign exchange market interventions to prevent the real appreciation of national currencies might cause inflation to be high and eventually affect the short-run adjustment to PPP, which, in turn, prohibits the fulfillment of the Maastricht criterion on price stability.3

The contribution of this paper is twofold. First, several studies, individually or partially, have investigated the empirical validity of PPP in the CEE countries, but none has examined the PPP hypothesis in the overall context of the potential EMU accession countries and emphasized the significance of real exchange rate convergence on European economic integration.4 Second, we use the minimum Lagrange multiplier (LM) unit root tests proposed by Lee and Strazicich (2003, 2004) that allow for a maximum of two endogenously determined structural breaks to test for the stationarity of fourteen real bilateral exchange rates against the numeraire currencies, the U.S. dollar (USD) and deutsche mark (DM). We consider the structural breaks, as these countries experienced significant monetary and real shocks during the transition process.
Methodology

We test the validity of PPP by analyzing the stationarity of the real exchange rate, which measures the deviation from PPP in the following equation:

\[ r_{ert} = e_t + p_{t}^{f} - p_{t}^{d}, \]  

where \( e_t, \) \( p_t^{f}, \) and \( p_t^{d} \) are the natural logarithm of the nominal exchange rate, the foreign price level, and the domestic price level, respectively. The stationarity of the real exchange rates supports the PPP proposition.

In this paper, we use unit root test with structural breaks to check whether real exchange rates are stationary. Perron (1989) shows that in the presence of a structural break, conventional testing procedures—augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) unit root tests—may erroneously fail to reject the null hypothesis that a series is integrated of higher order. He extended the ADF unit root test model by allowing for one known, or exogenous, structural break in the data series. The main idea behind introducing a structural break in the data series is to increase the power of the conventional ADF model to reject the unit root null hypothesis by providing the model with more information. Imposing a predetermined break, however, might induce incorrectly specified break points. Following Zivot and Andrews (1992; ZA) and Perron (1997), we determine the break point endogenously from the data. A potential problem of both tests is that they derive their critical values as they assume no break under the null hypothesis. Nunes et al. (1997) show that this assumption leads to size distortions in the presence of a unit root with break. Lee and Strazicich (2001) show that in the presence of a unit root with break, the ZA and Perron (1997) tests tend to select the break point at which bias and size distortions are the greatest. Lumsdaine and Papell (1997; LP) extend the ZA unit root test by incorporating two structural breaks in the data series. Lee and Strazicich (2003) show that, in the presence of two breaks, the same outcome occurs as in the one-break case. Hence, ADF-type endogenous break unit root tests might produce misleading results.

To avoid the potential problems of the ZA, Perron (1997), and LP tests, we use the endogenous one and two-break LM unit root tests developed by Lee and Strazicich (2001; 2003). In contrast to the previous tests, the size properties of the LM tests are unaffected by breaks under the null hypothesis. Therefore, the findings of the LM tests are more revealing, as rejecting the null hypothesis unambiguously implies stochastic convergence.

To test the PPP hypothesis for the sample countries, the following LM unit root test is used:

\[ rer = \delta'Z_t + \varepsilon_t, \quad \varepsilon_t = \alpha \varepsilon_{t-1} + u_t, \]  

where \( Z_t \) is a vector of exogenous variables. The test for the unit root is based on the parameter \( \alpha. \) In this paper, we consider a model equivalent to Perron’s (1989) Model C, which allows for a shift in intercept and change in trend slope under the alternative hypothesis. In this model, a structural break can be incorporated by specifying the vector of exogenous variables \( Z_t \) as \( Z_t = \{1, t, D_t, DT_t\}', \) where \( DT_t = t - T_B \) for \( t \geq T_B + 1 \) and zero otherwise. \( T_B \) is the time of break.

In the case of two structural breaks, the model can be described by \( Z_t = \{1, t, D_1, D_2, DT_1, DT_2\}', \) where \( D_1 \) and \( D_2 \) are dummy variables that capture the first structural break and the second structural break, respectively. Following Lee and Strazicich (2003), the unit root test statistic (LM test statistic) is specified as follows:
where $\Delta$ is the first-difference operator and $\tilde{S}_t$ is the detrended series. The testing procedure in Equation (3) involves $\Delta Z_t$ instead of $Z_t$. Hence, $\Delta Z_t$ is described by $[1, B_t, D_t]$, where $B_t = \Delta D_t$ and $D_t = \Delta DT_t$. Therefore, $B_t$ and $D_t$ correspond to a change in intercept and trend under alternative, and to a one-period jump and change in drift under the null hypothesis, respectively. In a similar manner, the case of two structural breaks, the series, $\Delta Z_t$, is described by $[1, B_{1t}, B_{2t}, D_{1t}, D_{2t}]$. If the real exchange rate has a unit root, then $\phi = 0$, which is the null hypothesis tested using the $t$-test against the alternative hypothesis that $\phi < 0$. The LM $t$-test statistic is given by

$$\bar{t} = t\text{-statistic testing the null hypothesis } \phi = 0.$$  

The location of the breaks ($T_\theta$) is determined by searching all possible break points for the minimum (i.e., the most negative) unit root test statistic. The LM unit test is specified as follows:

$$LM_{\xi} = \inf_{\lambda} \bar{t}(\lambda),$$  

where $\lambda = T_\theta/T$.

### Data and Empirical Results

#### Data

Monthly data on nominal exchange rates and the consumer price index (CPI) were obtained from the International Monetary Fund’s International Financial Statistics (IFS) database. We use monthly CPI-based real bilateral exchange rate data for fourteen countries in our sample. Two real exchange rate series are formed, based on two benchmark currencies, the USD and DM. Table 1 reports the sample period for each country.

#### Empirical Results

This section first examines the stationarity of the sample real exchange rates using the LM unit root test without structural breaks so as to establish a basis for comparison. The test is performed for both the USD-based and DM-based exchange rate series. Table 2 reports the results, which indicate that the null hypothesis of a unit root cannot be rejected. This is not consistent with the PPP hypothesis for all sample countries except Romania in the case of the USD-based real exchange rates. Similarly, in the case of the DM-based real exchange rates, the null of unit root cannot be rejected for all sample countries except Romania and Slovakia.

The failure to reject the null hypothesis of a unit root may be because traditional unit root tests have low power when structural breaks are ignored. Perron (1989) shows that in the presence of a structural break, traditional testing procedures may erroneously fail to reject the null hypothesis. Hence, the inability to reject nonstationarity may be due to the failure to account for a structural break in the data series. To allow for the possibility of a structural change, we use the minimum LM tests developed by Lee and Strazicich (2001; 2003). The results, based on one break of the Model C, are reported in Table 3, which shows, in the case of the USD-based real exchange rate series, that the null hypothesis of a unit root is rejected only for Romania and Turkey. Like the LM unit root test without structural breaks, we cannot reject the unit root null in most countries in the presence of
one structural break. Unlike the LM unit root test without structural breaks, the results for the case of the DM-based exchange rate series indicate that the null of a unit root is rejected for Croatia, Cyprus, Romania, Slovakia, Slovenia, and Turkey. Thus, real exchange rates for these countries are stationary, which is consistent with the PPP hypothesis.

Lee and Strazicich (2003) argue that the one-break test may lose power when there are two or more breaks in the data series. To address this problem, the stationarity of sample real exchange rates is investigated by using a unit root test that allows for two
Table 3. LM unit root test with one structural break

<table>
<thead>
<tr>
<th>Country</th>
<th>U.S. dollar–based real exchange rate</th>
<th></th>
<th>Break date</th>
<th>LM test statistic</th>
<th></th>
<th>Deutsche mark–based real exchange rate</th>
<th></th>
<th>Break date</th>
<th>LM test statistic</th>
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<tr>
<td></td>
<td></td>
<td>k</td>
<td></td>
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<tr>
<td>Latvia</td>
<td></td>
<td>12</td>
<td>1994:02</td>
<td>-1.958</td>
<td></td>
<td></td>
<td></td>
<td>12</td>
<td>1993:01</td>
</tr>
<tr>
<td>Lithuania</td>
<td></td>
<td>7</td>
<td>1996:02</td>
<td>-2.613</td>
<td></td>
<td></td>
<td></td>
<td>9</td>
<td>1996:03</td>
</tr>
<tr>
<td>Malta</td>
<td></td>
<td>1</td>
<td>2001:08</td>
<td>-3.783</td>
<td></td>
<td></td>
<td></td>
<td>6</td>
<td>1996:11</td>
</tr>
<tr>
<td>Poland</td>
<td></td>
<td>1</td>
<td>1991:12</td>
<td>-3.262</td>
<td></td>
<td></td>
<td></td>
<td>10</td>
<td>2002:10</td>
</tr>
<tr>
<td>Romania</td>
<td></td>
<td>12</td>
<td>1993:12</td>
<td>-4.962**</td>
<td></td>
<td></td>
<td></td>
<td>11</td>
<td>1997:12</td>
</tr>
<tr>
<td>Slovenia</td>
<td></td>
<td>8</td>
<td>1999:12</td>
<td>-2.887</td>
<td></td>
<td></td>
<td></td>
<td>10</td>
<td>1994:10</td>
</tr>
<tr>
<td>Turkey</td>
<td></td>
<td>1</td>
<td>2001:03</td>
<td>-4.858**</td>
<td></td>
<td></td>
<td></td>
<td>1</td>
<td>1994:11</td>
</tr>
</tbody>
</table>

Notes: The value $k$ denotes the optimum lag length; n indicates an insignificant break point. Critical values are based on Lee and Strazicich (2004). * and ** indicate statistical significance of a unit root with a structural break at 1 percent and 5 percent, respectively.
structural breaks. Table 4 reports the two-break minimum LM tests results for both the USD-based and DM-based exchange rate series. The results for the USD-based exchange series are similar to those of the one-break case. The null of unit root is rejected only for Romania and Turkey. The test results for the DM-based exchange rate series, however, are somewhat different than those of the one-break case. For the DM-based exchange rate series, contrary to the preceding tests results, the null of a unit root for Estonia and Bulgaria is rejected. Hence, the results of two-break tests reveal that the PPP hypothesis holds for Bulgaria, Croatia, Estonia, Romania, Slovakia, Slovenia, and Turkey. Because the second break date is not significant at conventional significance levels, the results of one-break tests are valid for Cyprus. The PPP hypothesis for Cyprus holds in the one-break case. In general, the results are consistent with the empirical literature, suggesting the validity of PPP for Romania (Barlow and Redulescu 2002), Slovenia (Choudhry 1999), and Turkey (Sarno 2000; Yazgan 2003).8

In the early years of transition, high-productivity growth in the traded goods sectors, in addition to general structural changes, caused significant appreciation of real exchange rates. Moreover, transition countries experienced several major changes in the last decade. The Bulgarian banking crisis caused the economy to enter a period of hyperinflation in late 1996, the Czech banking crisis of early 1997 led to a large depreciation, and the Romanian stabilization program started in early 1997 produced inflationary effect; consumer prices increased by 38 percent and 154 percent in 1996 and 1997, respectively. The transition countries also suffered the pressure of severe financial crises, such as the Asian and Russian crises. The Russian crisis of 1998 caused the continued vulnerability to external shocks for some countries, especially those with close trade linkages with Russia. The Turkish financial crises of 1994 and 2001 showed similar weaknesses. All these cases reflect adjustment in real exchange rates.

In general, significant break dates correspond to either a sharp increase in consumer prices (in Bulgaria, 1997; Romania, 1997; Slovakia, 2003; and Turkey, 1994) or a sharp decrease (in Croatia, 1994; the Czech Republic, 2002; Latvia, 1994; Lithuania, 1995 and 1996; and Poland, 1991). Similarly, appreciation (in Bulgaria, 1997; the Czech Republic, 1997; Lithuania, 1995; Poland, 1999; Romania, 1997; and Turkey, 2001) or depreciation (in Bulgaria, 1993; the Czech Republic, 2002; Latvia, 1994; Lithuania, 1996; and Slovakia, 2003) of national currencies obviously results in breaks in the real exchange rate series. Furthermore, among other countries, real exchange rates in Poland and Romania seem to be the most affected from exchange rate regime changes.

We also use an alternative approach to measure the persistence of real exchange rate deviations from its PPP long-run level. A commonly used measure of persistence in the literature is the half-life of PPP deviations.9 The half-life of a shock is defined as the expected duration of time for deviations from the long-run equilibrium arising from shocks to dissipate by 50 percent. Financial and monetary shocks affect nominal exchange rates, as well as real exchange rates, due to price stickiness. Persistent deviations from PPP would be difficult to explain by appealing to price stickiness, which accounts for rigid prices for one to two years. Rogoff (1996) and Taylor and Taylor (2004) report that the half-life of a real exchange rate shock is about three to five years for most countries, implying a slow rate of adjustment to the parity condition, despite the observed high short-term real exchange rate volatility. However, Murray and Papell (2002) have criticized these findings, arguing that the techniques used to measure persistence in PPP deviations were not appropriate. Hence, following Baharumshah et al. (2008), we use a test that is more powerful than the ADF, namely the Dickey–Fuller generalized least squares (DF-GLS)
<table>
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<th>Country</th>
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<td>$k$ Break date statistic</td>
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</tr>
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Notes: $k$ denotes the optimum lag length; n indicates an insignificant break point. Critical values are based on Lee and Strazicich (2004). *, **, *** indicate statistical significance of a unit root with a structural break at 1 percent, 5 percent, and 10 percent, respectively.
to compute the half-lives.\textsuperscript{10} We use the estimated value of the coefficient $\alpha$ from the estimation of the DF-GLS regression to compute the half-life according to the formula $h = \ln(0.5)/\ln\alpha$. Following Rossi (2005), we construct a 95 percent confidence interval associated with the above half-time statistics as follows: $h \pm 1.96\hat{\sigma}_{\hat{\alpha}}(\ln(0.5)/\hat{\alpha}[\ln(\hat{\alpha})]–2)$, where $\hat{\sigma}_{\hat{\alpha}}$ is the standard error of $\hat{\alpha}$.

Table 5 reports the estimated half-lives and the 95 percent confidence intervals (measured in months) for both the USD-based and DM-based real exchange rate series. The half-lives are calculated only for the real exchange rate series that are stationary according to the two-break LM unit root test.\textsuperscript{11} The results indicate a wide range of half-life point estimates across countries. The half-life estimates range from 1.25 years (15.05 months) to 2.72 years (32.72 months). On average, it takes 1.9 years to correct half of any PPP deviation. Hence, all the half-life estimates are shorter than the consensus of three to five years reported in Rogoff (1996) and Taylor and Taylor (2004). We also obtain relatively narrow confidence intervals for Bulgaria, Croatia, Cyprus, and Estonia compared to those of Romania, Slovakia, Slovenia, and Turkey. The lower bounds of the confidence intervals reported in Murray and Papell (2002) and Rossi (2005) are mostly less than two years, whereas the upper bounds are generally infinite. The lower bounds are consistent with the theory of PPP; the upper bounds are consistent with no convergence to long-run PPP. The results reported in Table 5 indicate that the lower bounds of the confidence intervals are all less than two years, whereas upper bounds are mostly less than three years.

### Conclusion

This study investigates the validity of PPP among all the potential EMU accession countries, eleven transition economies of Central and Eastern European (CEE) countries, and three market economy countries, Cyprus, Malta, and Turkey. Of the fourteen countries, twelve have joined the European Union in recent years. Apart from integration into the
European economic structure, monetary stability and nominal exchange rate stability are required for monetary integration. These two Maastricht convergence criteria imply real exchange rate stability and, eventually, the achievement of PPP. Therefore, empirical evidence on PPP may provide directions for economic integration in the euro area. The recent history of most of the sample countries, however, is especially interesting to test the validity of PPP, as there have been several economic structural reforms since the beginning of the 1990s. The restructuring and liberalization process was associated with several monetary and real shocks, and thus, short-run and long-run deviations from PPP, respectively.

To test whether the PPP holds for potential EMU accession countries, the stationarity of the USD-based and DM-based real exchange rate series were tested. The test results reveal that the real exchange rate series are stationary for only Romania and Turkey in case of the USD-based series. In the case of the DM-based series, however, the series are stationary for Bulgaria, Croatia, Cyprus, Estonia, Romania, Slovakia, Slovenia, and Turkey, which is consistent with the PPP hypothesis. Empirical evidence suggests that, in these countries, deviations from the parity can be caused by sudden changes in exchange rates, high inflationary pressures, or monetary shocks; do not persist over time; and allow PPP to hold in the long run.

We also use an alternative approach to measure the persistence of real exchange rate deviations from its PPP long-run level. A commonly used measure of persistence in the literature is the half-life of PPP deviations. The results indicate a wide range of half-life point estimates across countries. The half-life estimates range from 1.25 years (15.05 months) to 2.72 years (32.72 months). Hence, it takes on average 1.9 years to correct half of any PPP deviation.

Overall, we conclude that the evidence for PPP is strong in the case of the DM-based real exchange rate series for eight of the fourteen sample countries. There have been significant implications on these countries' prospects for achieving the two major Maastricht convergence criteria. In turn, the real exchange rate convergence may indicate their possible integration with the EMU. Broadly speaking, this result may provide some ground for policy makers to coordinate monetary policies and, consequently, strengthen their economic integration into the euro area.

The other six potential EMU accession countries, in which we failed to find evidence of the PPP hypothesis, need to catch up with others in satisfying the two major convergence criteria. These countries need to assure macroeconomic stability, which is inevitably accompanied by large shocks to fundamentals. In this respect, they must adopt an extensive range of measures to enact monetary and fiscal policies that would suit the needs of stable currencies and low levels of inflation. Apart from these policies, macroeconomic stability should be supported with a higher degree of reforms. It is also crucial to know whether or not the deviations from the PPP caused by the real disturbances are self-reversing over time for policy makers. Furthermore, the failure of the PPP to hold might indicate exchange rate misalignment. Thus, policies aiming at preventing excess over (under) valuation of national currencies may help the fulfillment of the Maastricht criteria on exchange rate stability.

Notes

1. There are several economic implications of the validity of PPP. First, deviations from PPP, caused by monetary or real shocks, can be used to explain real exchange rate movements. Second, PPP is used to determine the degree of misalignment. Third, the validity of PPP is a necessary as-
sumption in open economy macroeconomics. Fourth, PPP helps to smooth the price differentials for national income comparisons. Finally, PPP is used to set off the exchange rate parities (Rogoff 1996; Sarno and Taylor 2002).

2. Dibooglu and Kutan (2001) find that real exchange rate movements in Poland and Hungary can be explained by the nominal and real shocks. Borghijs and Kuijs (2004) show that the variability of nominal exchange rates in the Czech Republic, Hungary, Poland, Slovakia, and Slovenia can be explained mainly by nominal shocks.


4. An extensive literature has emerged in the search for evidence of PPP among industrial countries, and in recent years, testing the empirical validity of PPP among potential EMU accession countries has received considerable attention. These studies in the literature take the empirical approach of unit root testing, cointegration, and nonlinear modeling techniques, and provide mixed results. Some of these studies are by Alba and Park (2005), Barlow (2003), Barlow and Radulescu (2002), Choudhry (1999), Christev and Noorbakhsh (2000), Ertl (2003; 2004), Lopez and Papell (2007), Mahdavi and Zhou (1994), Payne et al. (2005), Sarno (2000), Sideris (2006), Telatar and Kazdagli (1998), and Yazgan (2003).

5. \[ S_t = rer_t - \bar{\mu} - Z_t \delta \], where \( \delta \) are the coefficients in the regression of \( \Delta rer_t \) on \( \Delta Z_t \), and \( \bar{\mu} \) is the restricted MLE of \( \mu \) given by \( rer_1 - Z_1 \delta \).

6. The ADF and Kwiatkowski et al. (1992; KPSS) unit root tests were also used to test whether the sample time series has a unit root. For both the U.S. dollar–based real exchange rates and the deutsche mark–based real exchange rates, the null hypothesis of the ADF test is not rejected at the conventional levels for all sample series. As for the KPSS tests, which have a null hypothesis of stationarity, the null is rejected in favor of the unit root alternative for almost all exchange rate series. There are a few exceptions. In the case of the U.S. dollar–based real exchange rate, we find that the real exchange rate series for Romania is stationary, consistent with the PPP hypothesis. However, in the case of the DM-based real exchange rates, we find that the exchange rate series for Romania and the Slovakia are stationary. To conserve space, the results of both unit root tests are not reported but are available from the authors upon request.

7. Since Model A and Model C with both one and two structural breaks produce different results, we focus on Model C, following the suggestions of Sen (2003a), who argues that Model C is preferable to Model A when the break date is treated as unknown. Using Monte Carlo simulations, Sen (2003b) shows that Model C yields more reliable estimates of the breakpoint than does Model A. Hence, our empirical analyses are based on Model C.

8. In case of the univariate unit root test, the null of unit root is rejected for about one-half of the sample countries. A possible reason for the failure to reject the unit root might be the time period of the data. As stated in Shiller and Perron (1985), univariate unit root tests generally have low power when the sample size is small. To address this problem, we used the panel LM unit root test with zero, one, and two structural breaks. For this test, we used data for the period January 1993–September 2006, which is the longest time period for which data are available for all fourteen sample countries. Only the two-break panel LM unit root test results indicate evidence for panel stationarity for the full panel in the case of the U.S. dollar–based exchange rates series. The test results for the DM-based exchange rates series, however, are somewhat different from those of the U.S. dollar–based exchange rate series. The results suggest strong evidence that the real exchange rate series is mean reverting for the full panel based on the LM unit root test with one and two structural breaks. However, the panel unit root test has some limitations. Rejecting the null does not imply that every real exchange rate series in the panel is stationary. The results of the panel unit root tests are available from the authors upon request.

9. We would like to thank an anonymous referee who pointed out the importance of using speed of half-life as a measure of persistence of real exchange rate deviations from its long-run level.

10. The DF-GLS test is based on the following regression:

\[ q_t = \alpha q_{t-1} + \sum_{i=1}^{k} \lambda_i q_{t-1} + \epsilon_t, \]

where \( q_t \) is the GLS-demeaned real exchange rate. The modified Akaike information criterion, which allows for the best combination of size and power in the presence of the GLS transformation, is used to choose the optimum lag for the DF-GLS.
11. We also calculate a half-life for Cyprus, as the real exchange rate series for Cyprus is stationary in the one-break case. The second break date is not significant.

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