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THE PERSISTENCE OF REAL EXCHANGE RATES IN THE CENTRAL AND EASTERN EUROPEAN COUNTRIES

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Abstract. This paper investigates the mean reversion in real exchange rates for Central and Eastern European countries. In contrast to previous studies, we use the local-persistent model to measure the half-life. We find that the adjustment to purchasing power parity is more rapid after accounting for structural breaks, taking less than 18 months to be cut in half. The empirical evidence shows that there is no clear-cut difference in the speed of adjustment to shocks between the transition economies and the larger member countries of the European Union. The narrow confidence intervals for the half-lives that accord with the standard sticky-price models provide strong support for purchasing power parity. The purchasing power parity puzzle does not seem to hold in these transition countries. The practical implication of our findings is that the transition countries have successfully adopted trade policies that mimic those of the European Union, with a view to alignment in readiness for European Union membership.

Keywords: half-lives; local persistence; structural breaks; real exchange rate; PPP puzzle; transition economies.

JEL Classification: C0, F21, F36.

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Introduction

Perhaps, the most frequently asked question in the empirical research on international macroeconomics is the validity of purchasing power parity (PPP) as a long-run equilibrium relationship. This is because PPP is the theoretical foundation and major building block of many open macroeconomic models. PPP postulates that the real exchange rates (RERs) revert to a mean and most studies in the past have attempted to provide supporting evidence by applying unit-root/stationary tests on RERs. Despite all these research efforts over the years, empirical evidence on the international parity condition over the long-term remains mixed.[1] Lengthy half-lives for RERs in the presence of a high degree of exchange rate volatility have been considered as one of the most puzzling empirical regularities in international economics and have been dubbed the PPP puzzle (Rogoff 1996). One usual explanation for the puzzle has been based on the low power of the conventional unit-root tests, which is aggravated by the high persistence of RERs. The difficulty of rejecting the unit-root null is often associated with the small samples typically used in verifying the parity condition. The issue is more apparent in the transition countries, which are the main focus of this study, where a long span of data is unavailable for empirical analysis. In addition, macroeconomic instability in the last three decades may have significant influence on the behavior of RERs in the transition economies that make testing for PPP a challenging task.

The bulk of the literature has focused mainly on the industrial countries (see Taylor, Taylor 2004). Recently, the research effort has shifted somewhat to the emerging market economies by applying a wide range of methodologies. Studies of the transition economies, particularly in Central and Eastern Europe (CEE), are quite limited. The

[1] The different econometric methods have yielded mixed results about the validity of PPP. Recent studies have concentrated on panel unit-root tests, unit-root tests with multiple breaks, and nonlinear unit-root tests.

behavior of RERs in these countries is relevant to the study for at least two reasons. First, most of the countries are in the process of entering the eurozone, and in need of an estimate of the equilibrium exchange rate before they can connect their currency to the euro. If a hypothesis fails to hold in these countries, then the PPP equilibrium rates may not provide an appropriate exchange rate between the new European Union (EU) members and the euro. Hence, knowledge of whether PPP holds for these countries may help in assessing the readiness of these countries to join the European Monetary Union (EMU). Second, the confidence that international financial markets placed in the transition economies with regard to trade and investments depend on the behavior of RERs. If PPP holds, this would imply almost no risk associate with price level convergence (Zhou *et al.* 2008).

The main purpose of this study is to ascertain whether long-run PPP is upheld in the CEE countries. To this end, we examine the validity of PPP between each of the transition economies and their major trading partners, namely the United States and the Eurozone-member countries. The former is to account for the role of the United States in the foreign exchange market and in the reconstruction of the countries under review.[2] The choice of the eurozone is justified on the basis of the increasing integration between the CEE countries and the eurozone. Eight out of the CEE countries (the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia, and Slovenia) joined the EU in May 2004, and these countries have completed the transition process. Two other states that acceded in 2007 are Bulgaria and Romania.

Previous studies, to a large extent, restrict their investigation to an $I(0)/I(1)$ RER process. Analysis based on this line does not shed much light on the speed of adjustment of PPP to its equilibrium value. We depart from those studies by considering an intermediate process—local persistent model—to calculate the point estimates of the half-life and their confidence intervals (CIs). A major attraction of this model (also referred as an in-between process) is that it tends to produce narrower CIs for half-life deviations than those calculated from Augmented Dickey-Fuller (ADF) and local-to-unit root models (Kim, Lima 2010). We also compare the speed of convergence to PPP between the transition economies and the core EU countries, noting that a near unit-root (highly persistent) process best characterizes RERs in both the transition and developed economies. Our analysis reveals an interesting result: the Rogoff's puzzle does not apply for our data that include the CEE countries and some core EU countries.

The plan of the paper is as follows. The next section provides a brief overview of the PPP literature. This is followed by the econometric methodology regarding the computation of half-life derived from the local-persistent model. Next, we describe the data and present the empirical results. The final section concludes the paper.

1. Literature review

The PPP literature on the major industrialized countries has highlighted a puzzling result—the outcomes of the unit-root tests are broadly inconsistent with the high persistence characteristic of RERs (in other words long half-life). Existing point estimates of half-life deviation from PPP based on the ADF and local-to-unity models largely confirm Rogoff's (1996) PPP puzzle: that while PPP holds for the majority of the countries, the speed of

[2] The level of trade integration between the CEE countries and the EU market is high, and in the majority of cases it is even higher that of the EU-15.

reversion of RERs to parity is, in many cases, rather slow (around 3-5 years). In addition, critics have pointed out that the wide CIs reported in the past provide little information regarding the speed of convergence. Several papers, such as, Maican and Sweeney (2013), Rossi (2005), and Sekioua (2008) have attempted to address the PPP puzzle. Sekioua (2008) notes that 'it is possible that unit root tests reject the nonstationarity hypothesis but the process is still persistent in the sense that deviations are slow to die out.' (p. 78). Meanwhile, Maican and Sweeney (2013) justifiably argue that the models used to test PPP in past studies are not specifically designed to capture parameter shifts in RERs associated with a volatile economic and financial history. Based on data from 10 CEE countries, the authors report strong support of PPP using models that specifically allow for breaks in the data. A striking feature of their findings is that break models provide substantial faster mean reversion than the other models tested, including nonlinear models.

The widespread failure to reject unit-root due to structural breaks is also taken up by Holmes *et al.* (2012) with a broader set of Organisation for Economic Co-operation and Development (OECD) countries. Apparently, they show that the panel-based stationary test that accounts for cross-sectional dependence (CD) and structural breaks performs better than the standard model in terms of power and size properties. The authors employ the Pesaran and Shin (1998) generalized impulse response functions to arrive at half-life estimates of four quarters. Holmes *et al.* (2012) rely solely on point estimates of the impulse response (neglecting CIs) for statistical inference, and hence are not very informative in terms of the speed of convergence. Basher and Carrion-i-Silvestre (2013) in analyzing the persistence of shocks that affect OECD countries point out the importance of proper modeling of breaks in the data generating process (DGP). Their results based on various estimation methods show that ignoring breaks lead to an upward bias in the autoregressive (AR) parameters, and an overestimation of the half-lives.[3] The literatures mentioned above highlight the importance of exploring models that accommodates structural breaks as a means to obtain tighter CIs and stronger evidence for long-run PPP. This point is taken up in the present paper.

The empirical literature on PPP in transition economies has grown significantly recently. Important contributions include Bahmani-Oskooee *et al.* (2008), Cuestas (2009), Kasman *et al.* (2010), Boršič *et al.* (2011), Chang *et al.* (2012), Liu *et al.* (2012), and Payne *et al.* (2005). Payne *et al.* (2005), who looked at data from Croatia, provide evidence against the hypothesis, despite the use of a battery of unit-root tests with break(s). The authors conclude that the lack of evidence on the validity of PPP could be closely connected to the appreciation of the exchange rate due to growth in productivity and real wages. Bahmani-Oskooee *et al.* (2008), who apply both the linear and nonlinear ADF unit-root tests to 88 developing counties, find mixed results. Their findings based on the real effective exchange rate (REER) reveal that the nonlinear approach provides more support for PPP than does the widely used linear ADF unit-root model. In the sample from the transition economies, they find evidence of mean reversion to a trend in only four countries based on the standard approach. For three other countries, REER reverts to a trend in a nonlinear fashion. Meanwhile, Liu *et al.* (2012) provide new insights by showing that the adjustment to PPP for a sample of eight transition countries may not be characterized by a nonlinear

[3] To evaluate the half-lives, Basher and Carrion-i-Silvestre (2013) provide the median-unbiased method of Andrews and Chen (1994), the bootstrap method in Kilian (1998) and grid bootstrap procedure in Hansen (1999) to show the robustness of their findings.

stationary process. In addition, Maican and Sweeney (2013), who apply nonlinear models and compare them with piecewise linear break models, show that their linear break models appear superior in detecting mean reversion in the PPP. To summarize, two arguments stand out in the literature. First, during the catching-up phase of the transition, equilibrium exchange rates should exhibit an upward trend as the countries experience growth in productivity and real wages. Second, large capital flows due to capital account liberalization might appreciate RERs, thus invalidating PPP.

Most of the above-mentioned papers focus on whether the RER series is an $I(0)/I(1)$ process (exceptions being Kasman *et al.* 2010; Chang *et al.* 2012; Maican, Sweeney 2013) and do not consider half-life deviations from PPP estimates as complementary to their unit-root tests. In particular, Chang *et al.* (2012) who apply the panel seemingly unrelated regression augmented Dickey-Fuller (SURADF) method (with Fourier function) find the RER in the transition economies (the sole exception is Russia) is characterized by an $I(0)$ process. Based on this preferred specification, they report half-life estimates between 7-64 months, shorter that those without breaks (10-133 months). Uncertainty surrounds the half-life estimates provided in Chang *et al.* (2012) as the authors do not provide the CIs for their half-life estimates. The speed of convergence is faster in using the panel SURADF with a Fourier function, to specifically address the issue on incompatibility of lengthy half-lives with the observed high degrees of exchange rate volatility. He and Chang (2013) also places emphasis on nonlinearities and structural breaks, but ignore uncertainty surrounding their estimates.

Although empirical evidence obtained in these limited studies is mixed, the findings assembled by recent studies based on advanced time-series and panel-data approaches, including nonlinear methods (for example Cuestas 2009) seem to support the PPP hypothesis as a long-run relationship in the CEE countries. However, this evidence is inconsistent with the long half-lives reported in some papers. The bulk of the literature suggests that standard measures of half-life might be problematic and spuriously unfavorable empirical results may be obtained on the parity condition when highly persistent processes are unaccounted for in the model.[4] Least square estimates of half-lives of parity reversion based on the unit-root model in small samples tend to be lengthy, and the cumulative work based on a different set of countries typically replicates Rogoff's (1996) consensus.[5] Moreover, CIs of the half-life based on conventional methods assuming the RER is an $I(1)/I(0)$ process tend to be very wide (including the possibility of an infinite half-life), thereby providing virtually no information on the size of the half-lives. As illustrated in Rossi (2005), the standard sampling methods used in constructing CIs might be unreliable when the variables are highly persistent and the sample size is too small.

Give the poor performance of modeling the RER as an $I(0)/I(1)$ process, Kim and Lima (2010) suggest the estimation of a local-persistent model (Phillips *et al.* 2001), a process between a stationary and a unit-root process. This intermediate process allows the RER to show persistence over a range of time, but eventually the effects of shocks will dissipate in

[4] Murray and Papell (2002) have shown that if the DGP was a linear $AR(1)$, least square estimates of the half-life is biased downward in a small sample, with bias growing as the AR parameters approach unity. To address the problem, they applied Andrews and Chen's (1994) median-unbiased estimation method. They find although the point estimates are in the range of between 3-5 years, the CIs provide virtually no information regarding the estimates.

[5] Lengthy half-lives that give rise to the PPP puzzle are at odds with the implications of sticky-price models, which imply that the half-life of a shock to RER should be between one and two years.

support of PPP. An important finding that emerges from this study is that the PPP puzzle in some major industrialized countries is less pronounced than commonly thought.

The present study combines two strands of the literature. First, it applies the local-persistent model to the transition economies of Europe. Second, it accounts for breaks in the data given the findings of recent studies that incorrect specification (including the actual number of breaks) may bias the analysis toward a finding of slow convergence. The allowance for breaks in the data for the transition economies is justified as the occurrence of major economic reforms and changing macroeconomic regimes may give a false impression of the speed of adjustment toward long-run equilibrium. The combination of both strands in the literature undertaken in this paper allows us to obtain much shorter half-lives and narrower CIs (as well as, low upper bounds that are less than 18 months) than those reported earlier.

2. Half-life measure

Rossi (2005) proposed a local-to-unit root model that is robust to the presence of high persistence in small samples to construct CIs for the point estimates of the half-life. After applying the model to the industrialized countries, she concluded that while the lower bound (4-8 quarters) of the CI for the half life is comparable to nominal price and wage stickiness, the upper bound is infinity for all countries. That is, the wide CIs from the local-to-unit root model fails to provide sufficient evidence to refute Rogoff's PPP puzzle. In this paper, we consider instead the local-persistent model proposed by Phillips *et al.* (2001) and extended by Lima and Xiao (2007) to model the in-between process, which lies between the conventional stationary and unit-root process, to capture the behavior of RERs in the transition countries included in our study.

Following Kim and Lima (2010), the half-life property of local persistence is computed as $\ln(0.5b(\hat{1}))/(-1/n^{\hat{d}})$, where the estimate local persistence parameter $\hat{d} = -\ln(1-\hat{\phi})/\ln(n)$, and $b(\hat{1}) = 1-\sum_{j=1}^{k}\hat{\phi}_{j-1}^{*}$ is the correction factor. In the above expressions, n is the number of observations, and $\hat{\phi}$ is a consistent estimator of the largest root measuring persistence and determined from the break models. The delta method is used to compute the two-sided 95 percent CIs. Accordingly, the CIs can be obtained by $\hat{h}_{0.50} \pm 1.96se(\hat{\phi})([-\ln 0.5/\hat{\phi}][\ln(\hat{\phi})]^{-2})$, where $se(\hat{\phi}) = \sqrt{2}/(n^{\frac{1}{2}+\frac{\hat{d}}{2}})$.[6] If d lies between 0 and 1, the series is considered the standardized local persistence process. Two interesting special cases may be considered: when $d = 1$, the series is identical to the local-to-unit root process proposed by Rossi (2005) and when $d = 0$, the time-series process has short-memory dynamics. Sekioua (2008) argues that the unit-root null can seldom be rejected for highly persistent variables; hence, a misleading conclusion with regard to PPP will be drawn if the persistence of deviation from the parity is ignored.

2.1 Testing for local persistence parameter—*d*

[6] Readers may refer to the original papers by Lima and Xiao (2007) and Kim and Lima (2010) for details.

To determine whether a series follow to what Phillips *et al.* (2001) referred as a "locally persistent" behavior, we consider the tests suggested in Kim and Lima (2010). Under the null hypothesis of covariance stationary ($H_0 : d = 0$) against $H_1 : 0 < d < 1$ (that is, local persistence), the test statistic is given by

$$Q_n = \underset{1 \le \upsilon \le n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\hat{\omega}_y} \left| \sum_{i=1}^{\upsilon} y_i - \frac{\upsilon}{n} \sum_{i=1}^{n} y_i \right|. \tag{1}$$

The test statistic in Eq. (1) is computed in the presence of an intercept in the model. However, if the real time series is not directly observable in the presence of the intercept, Eq. (1) can be evaluated at the demeaned value $\hat{y}_i$, that is

$$\hat{Q}_n = \underset{1 \le \upsilon \le n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\hat{\omega}_y} \left| \sum_{i=1}^{\upsilon} \hat{y}_i - \frac{\upsilon}{n} \sum_{i=1}^{n} \hat{y}_i \right|, \tag{2}$$

where $\hat{y}_i$ is the demeaned value of the time series (say RERs) and it being computed as the difference between the observed (y_i) and average of the series ($\bar{y}$).

Nonparametric kernel smoothing is used to estimate the long-run variance parameter, given by $\hat{\omega}_y^2 = \sum_{v=-q}^{q} k(v/q)\hat{\gamma}(v)$, where q is the bandwidth parameter that satisfying $q \to \infty$ as the sample size, $n \to \infty$, $k(.)$ is the lag window defined on [-1,1], and $\hat{\gamma}(v)$ is the sample variance computed as $\sum_{i \ge 1}^{i+v \le n} \hat{y}_i \hat{y}_{i+v} / n$. The $\hat{\omega}_y^2$ is a consistent estimator of ω_y^2 under the H_0 as detailed in Kim and Lima (2010). The $k(x) = 1 - |x|$ is used to get the Barlett estimator [$\hat{\omega}_y^2 = \sum_{v=-q}^{q} (1 - (|v|/q))\hat{\gamma}(v)$].

Under the null hypothesis of covariance stationary and based on theorems discuss in Lima and Xiao (2007), the asymptotic behavior of the test statistics for Q_n and $\hat{Q}_n$ are defined as:

- Theorem 1:

  $$Q_n = \underset{1 \le \upsilon \le n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\hat{\omega}_y} \left| \sum_{i=1}^{\upsilon} y_i - \frac{\upsilon}{n} \sum_{i=1}^{n} y_i \right| \Rightarrow \sup_{0 \le r \le 1} |W(r) - rW(1)|,$$

  where y_i is a process without a time trend and $W(.)$ is a standard Brownian motion.

- Theorem 2:

  $$\hat{Q}_n = \underset{1 \le \upsilon \le n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\hat{\omega}_y} \left| \sum_{i=1}^{\upsilon} \hat{y}_i - \frac{\upsilon}{n} \sum_{i=1}^{n} \hat{y}_i \right| \Rightarrow \sup_{0 \le r \le 1} \left| W(r) - rW(1) + 6(1-r) \left\{ \frac{1}{2} W(1) - \int_0^1 W(s)ds \right\} \right|,$$

  where y_i is a process with a time trend and $\hat{y}_i$ is the detrended value of y_i.

The test statistics Q_n for y_i and $\hat{Q}_n$ for $\hat{y}_i$ have been proved to converge to a functional Brownian motions (Lima, Xiao 2007). Accordingly, we generate the critical values (CVs)

of the Q_n and $\hat{Q}_n$ for sample size 70, 100, 200, 500 and 1,000. Table A.1 reports the simulated CVs based on 5,000 numbers of replications.

3 Data and empirical results

3.1 Data

This paper investigates the PPP relationship over the period 1996:M1 to 2011:M10. It extends previous studies to include the effect of the global financial crisis (GFC) and the subsequent sovereign debt crises of the euro countries. The transition countries include Bulgaria (BG), Croatia (CR), Czech Republic (CZ), Estonia (EE), Hungary (HU), Latvia (LV), Lithuania (LT), Poland (PL), Romania (RO), Slovakia (SK), and Slovenia (SL). Only three of these countries (EE, SK, and SL) have adopted the euro as their common currency. Latvia joined the euro in January 2014. The sample period was dictated by the availability of the harmonized consumer price index (HCPI).[7] The main source of the data is the Wiener Institut für Internationale Wirtschaftsvergleiche (WIIW) database. Because the WIIW monthly database does not cover all the series for the countries under review, the national currency/euro exchange rate series for LV, LT, and EE were obtained from the Eurostat database, and the U.S.-based series come from the International Monetary Fund's *International Financial Statistics* (IFS). The consumer price index data (CPI, 2005 = 100) and HCPI (2005 = 100) were collected from *IFS,* the WIIW, and the Eurostat database. We also incorporate data on five EU countries, including the three largest—Belgium (BEL), France (FRA), Germany (DEU), Italy (ITA), and the Netherlands (NLD) for comparison. As in the past studies, these series are not seasonally adjusted, and all the variables are expressed in logs.

3.2 Unit-root tests

In our preliminary analysis, we applied an array of unit-root tests. Unreported results from standard unit-root tests reject the PPP hypothesis in all but a few cases. Next, we consider two other unit-root tests allowing for breaks advocated by Carrion-i-Silvestre *et al.* (CDL, 2005) and Narayan and Popp (NP, 2010). The former approach considers panel stationarity tests that account for both CD and multiple structural breaks, while the latter approach is a two-break unit-root test. The single time-series two-break test (unreported) advocated by NP yields some support for PPP.[8] By contrast, the results from CDL (see Table 1) that have the power advantages of panel unit-root tests fail to reject the stationary null, that is PPP holds as a long-run relationship for the same set of data. These results are quite robust as they apply for the two panels (the CEE panel and the CEE panel augmented with 5 EU countries) and two different choices of breaks (two and five). The advantage of using the CDL panel test is that if the null of the CDL test is not rejected, then we can conclude that all of the series in the panel are stationary. We attribute this finding to the use of a model

[7] The sample also covers a period where capital controls had been abolished in the EU area.
[8] The test rejects the unit-root null in two (CZ and EE) out of 11 countries for the euro-based. The unit-root null is rejected in only one country for the U.S.-based (BG).

which allows for both structural breaks and CD. The support for the PPP relationship appears to be insensitive to the choice of numeraire currency.

[Insert Table 1]

The timing of the first break identified by the Bai and Perron's (2003) dynamic algorithm is concentrated around 1999-2001, which coincides with the 1998 Russian financial turmoil and the launching of the euro. The location of the second break during the period 2004 to 2006 may be associated with joining the EU (CR, CZ, HU, SK and EE). For the dollar rates, the break occurs during the 2007-08 period, which coincides with the depreciation of the euro during the GFC.[9] These break dates are taken into consideration in the empirical analysis that follows.

3.3 Half-life

Having shown that each RER series is stationary, we now turn to the speed of adjustment to PPP, an issue that has not received as much attention. The fact that stationarity is not rejected cannot be interpreted as confirmation of mean reversion. As noted in Rossi (2005) and others, the outcome from the stationarity test may suggest mean reversion, but the speed of reversion to parity is rather slow to support the hypothesis. The literature based mainly in the industrialized countries has recorded lengthy half-life estimates for RERs in the presence of a high degree of nominal and real exchange rate volatility. To capture these findings, Rogoff (1996) coined the term PPP puzzle which refers to the unreasonably lengthy interval of three to five years for prices to adjust to their equilibrium. Does the PPP puzzle—which asserts that while PPP holds for the majority of countries, the speed of reversion of RERs to parity is rather slow—hold in the CEECs? To provide an answer to this question, we apply the local-persistent model instead of using the ADF or the local-to-unit root model which was commonly used in the past. The reason is that RERs are typically affected by monetary and financial shocks. Table 2 (Panel A) reports the degree of local persistence of the stochastic process (*d*), the corresponding half-life estimates, and the associated 95 percent CIs for 11 CEE countries. Similar information on the monthly RERs of five EU countries vis-à-vis the euro and the U.S. dollar is reported in Panel B.

The results reported in Table 2 (columns 2–9) are quite interesting. First, the estimated local persistence parameter (*d*) varies from 0.189 (CZ) to 0.512 (LV) for euro-based exchange rates, and from 0.193 (BG) to 0.506 (LV) for the U.S. dollar-based exchange rates. Second, for the euro-based exchange rates, we obtain point estimates of half-lives that range from 1.87 months (CZ) to 10.19 months (LV). Similarly, for the US dollar exchange rates our half-life estimates range from 1.91 months (BG) to 9.83 months (LV). Third, the estimated CIs show that the half-life of PPP deviation varies widely across our sample countries, but none of the upper bounds of these intervals exceed 18 months, irrespective of the reference currency. Fourth, similar results regarding the estimated half-lives and their CIs apply for the five EU countries. In the majority of countries, the upper bound of the CIs is less than one year. The only exceptions are Belgium and the Netherlands (for the dollar exchange rate), where the upper bound is approximately fifteen months. These results contrast sharply with previous literature using a different methodology. Specifically, we find that the RER reverts to its equilibrium at a rate much faster than in previous studies, as the shocks have been removed with the inclusion of

[9] The full results are not reported here but are available upon request from the authors.

dummy variables in the model, according to the conclusions of the break tests. Our findings on the relatively small half-lives in the CEEC echo the view expressed in Maican and Sweeney (2013). The CEEC typically had 10 years or less to make the huge adjustments that are required for accession to the EU. It would be interesting to compare the estimated half-lives and the respective CIs for the CEE and EU countries with those obtained from the traditional ADF model. With very few exceptions, the half-life estimates are much higher and the CIs much wider than those obtained from the NP model (columns 10–17).

[Insert Table 2]

Finally, to address the issue of the impact of the recent GFC, we carried out the above computations for a reduced sample ending in 2006:M12. Experiments with a shorter sample sub-period (1996-2006) for the U.S. dollar- and euro-based rates point to substantially lower persistence, a finding rather expected given the GFC started in 2007 (not reported). The estimated median half-lives are 4.20 and 4.01 months for the euro- and U.S.-based RERs, respectively. We note that the median half-life is shorter than for the full sample that cut across the break dates.[10] For most countries in our sample, the half-life point estimates are less than six months with upper bounds not exceeding 12 months with a few exceptions. The median upper bounds are 6.34 and 6.70 months for the U.S. dollar and euro-based RERs, respectively. The differences in the speed of mean reversion between the long and short sample periods suggest changes in the speed of mean reversion over time. We formally test ($\hat{Q}_n$) the null $H_0 : d = 0$ against $H_1 : 0 < d < 1$ and the results (Table 3) show that all of the $\hat{d}$ are significantly different from zero. As in Kim and Lima (2010), we find support of the locally persistent process in the RER being examined.

[Insert Tables 3]

Table 4 summarizes the proportion of upper-bound estimates of half-lives for three time intervals (less than 6 months, from 6 to 12 months, and more than 12 months) for 11 CEEC, EU-5, and EU-16 (11 CEE plus EU-5) countries. We note that the majority of the upper bounds of the CIs fall in the first and second categories and only a small proportion fits into the third category of more than 12 months. Specifically, around 64 percent (55 percent) of our half-life estimates for CEE countries are less than 12 months for euro- (U.S. dollar) based rates. As shown in Table 4, the EU countries also have short-lived estimates—80 percent and 60 percent are less than 12 months for the euro-based and U.S. dollar-based, respectively. The last two lines of Table 4 report the mean and median of the upper bounds of the half-lives in each country group. The median upper bound for the CEE countries is only two months larger than the corresponding figure for the EU countries. These findings indicate that the proportions of the upper bounds of the half-lives, as well as, the median of the upper bounds in the two groups are not very different after all. Many of the transition economies adopted trade policies that mimic those in member countries of the EU. Kutan and Yigit (2004) find increasing similarities in real and monetary developments between the eurozone and its CEE partners. As their growth and inflation rates converge to those of the developed economies, we can expect the persistence of the RERs to match that of the advanced EU countries.

[Insert Table 4]

[10] The median half-life for the full sample is 6.76 for the dollar-based rates while it is 6.09 for the euro-based RERs.

Conclusions

This paper explores the mean reversion in RERs for the CEE countries with the objective to provide an additional perspective on the PPP puzzle debate. Past literature on PPP based on unit-root tests provides little (if any) evidence of its validity, thereby casting serious doubt on the ability of the model to determine the behavior of RERs. We applied the local-persistent model and obtained a number of interesting results. First, we find that the dynamics of RERs are described by a transient process that reverts to its fixed mean after major economic shocks, a characteristic that accords with the PPP hypothesis, with persistence estimates lower than in the standard stationarity process. Second, we find that, accounting for structural breaks, the estimated half-life is less than 12 months with an estimated CI upper bound less than 18 months. It supports the notion that the PPP puzzle disappears when both the persistence process and structural breaks are allowed in the data. We take these findings as evidence to suggest that the PPP puzzle reported in previous studies may have been overstated.

A third and rather striking finding is that after breaks are introduced, the array of point half-life estimates, as well as, the upper bounds of the CIs are well within the vicinity of the larger EU countries (below 18 months). This means that the degree of economic integration in the CEE countries is no different from that of their neighboring core members. Our findings endorse the view that the CEE countries have successfully adopted trade policies that mimic those of the EU, with a view to alignment in readiness for EU membership. The reform process that started in the 1980s and intensified over time led to a reduction in persistent shocks to international parity. Finally, the evidence found in this study is robust to the choice of the reference currency.

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Table 1. CDL (2005) panel stationarity test with breaks

	Euro-based				U.S.-based			
	m=2		*m=5*		*m=2*		*m=5*	
				Panel A: 11 CEEC				
Homogeneity	−2.171	(0.985)	−2.515	(0.994)	−3.589	(0.999)	−2.764	(0.997)
Heterogeneity	1.243	(0.107)	−0.723	(0.765)	−1.592	(0.944)	−2.059	(0.980)
				Panel B: 11 CEEC + EU-5				
Homogeneity	−2.627	(0.996)	−3.207	(0.999)	−4.243	(0.999)	−3.281	(0.999)
Heterogeneity	1.275	(0.105)	1.033	(0.151)	−1.398	(0.919)	−1.241	(0.893)

Notes: Values in ( ) refers to *p*-value and *m* indicates the number of breaks allowed. The trimming is taken on the interval [0.1T, 0.9T]. Newey-West bandwidth selection is using a Bartlett kernel.

Table 2. Half-life and degree-of-local-persistence estimates

	Narayan and Popp Model								ADF Model							
	Euro-based				U.S.-based				Euro-based				U.S.-based			
	$\hat{d}$	HL(M)	$se(\hat{\phi})$	95% CI	$\hat{d}$	HL(M)	$se(\hat{\phi})$	95% CI	$\hat{d}$	HL(M)	$se(\hat{\phi})$	95% CI	$\hat{d}$	HL(M)	$se(\hat{\phi})$	95% CI
Panel A: 11 CEEC																
BG	0.423	6.37	0.034	[2.48,10.25]	0.193	1.91	0.062	[1.26,2.56]	0.344	3.86	0.042	[1.76, 5.95]	0.416	5.79	0.034	[2.11, 9.47]
CR	0.305	3.44	0.046	[1.89,4.99]	0.470	8.16	0.030	[2.51,13.80]	0.434	6.39	0.033	[2.15, 10.64]	0.624	17.94	0.020	[0.00, 36.88]
CZ	0.189	1.87	0.063	[1.24,2.50]	0.358	4.54	0.040	[2.20,6.88]	0.385	4.86	0.037	[1.98, 7.75]	0.551	12.13	0.024	[1.48, 22.77]
HU	0.323	3.78	0.044	[1.99,5.56]	0.471	8.18	0.030	[2.51,13.86]	0.534	11.02	0.025	[1.73, 20.31]	0.562	12.84	0.024	[1.24, 24.44]
PL	0.370	4.82	0.039	[2.26,7.38]	0.304	3.42	0.046	[1.88,4.95]	0.504	9.40	0.027	[2.02, 16.78]	0.506	9.46	0.027	[2.01, 16.91]
RO	0.464	7.91	0.030	[2.53,13.29]	0.435	6.76	0.033	[2.50,11.03]	0.466	7.63	0.030	[2.17, 13.08]	0.498	9.09	0.028	[2.06, 16.11]
SK	0.249	2.56	0.053	[1.56,3.55]	0.484	8.79	0.029	[2.49,15.08]	0.522	10.37	0.026	[1.89, 18.85]	0.638	19.41	0.019	[0.00, 40.61]
SL	0.414	6.09	0.035	[2.46,9.72]	0.459	7.69	0.031	[2.52,12.85]	0.489	8.66	0.029	[2.11, 15.22]	0.607	16.33	0.021	[0.00, 32.82]
EE	0.461	7.58	0.032	[2.40,12.77]	0.412	5.89	0.036	[2.34,9.44]	0.615	16.53	0.021	[0.00, 33.73]	0.617	16.71	0.021	[0.00, 34.20]
LV	0.512	10.19	0.027	[2.33,18.06]	0.506	9.83	0.027	[2.36,17.30]	0.695	26.10	0.017	[0.00, 59.04]	0.626	18.07	0.020	[0.00, 37.23]
LT	0.511	10.15	0.027	[2.34,17.96]	0.416	6.13	0.035	[2.45,9.82]	0.597	15.54	0.021	[0.20, 30.87]	0.555	12.38	0.024	[1.37, 23.38]
Median		6.09		[2.33,9.72]		6.76		[2.45,11.03]		9.40		[1.89, 16.78]		12.84		[1.24, 24.44]
Mean		5.89		[2.13,9.64]		6.48		[2.28,10.69]		10.94		[1.45, 21.11]		13.65		[0.93, 26.80]
Panel B: EU-5																
BEL	0.476	8.41	0.030	[2.52,14.30]	0.485	8.85	0.029	[2.49,15.21]	0.400	5.31	0.036	[2.05, 8.57]	0.643	19.89	0.019	[0.00, 41.88]
FRA	0.198	1.96	0.061	[1.28,2.63]	0.377	5.02	0.038	[2.30,7.74]	0.318	3.31	0.045	[1.61, 5.02]	0.650	20.68	0.019	[0.00, 43.98]
DEU	0.425	6.45	0.034	[2.49,10.41]	0.377	5.02	0.038	[2.30,7.73]	0.467	7.69	0.030	[2.18, 13.19]	0.641	19.65	0.019	[0.00, 41.24]
ITA	0.167	1.66	0.066	[1.13,2.19]	0.375	4.97	0.038	[2.29,7.65]	0.505	9.47	0.027	[2.03, 16.90]	0.659	21.64	0.018	[0.00, 46.54]
NLD	0.372	4.89	0.039	[2.27,7.50]	0.480	8.61	0.029	[2.50,14.72]	0.647	20.30	0.019	[0.00, 42.95]	0.626	18.11	0.020	[0.00, 37.28]
Median		4.89		[2.72,7.50]		5.02		[2.30,7.74]		7.69		[2.03, 13.19]		19.89		[0.00, 41.88]
Mean		4.67		[1.94,7.40]		6.49		[2.38,10.61]		9.21		[1.57, 17.33]		19.99		[0.00, 42.19]

Notes: HL(M) is the half-life for the local-persistent model measured in months by $\ln(0.5b(\hat{1}))/(-1/n^{\hat{d}})$, and half-life for the ADF model measured in months by $\ln(0.5b(\hat{1}))/\ln(1-\hat{\phi})$. The persistence parameter $\hat{d} = 1-\hat{\phi} = n^{-d}$, where $0 < d < 1$. The two-sided 95 percent confidence intervals (CIs) measured monthly are constructed according to $\hat{h}_{0.50} \pm 1.96 se(\hat{\phi})([-\ln 0.5/\hat{\phi}][\ln(\hat{\phi})]^{-2})$ where $\hat{\phi}$ is drawn from model 2 in Narayan and Popp (2010) and ADF model, respectively.

Table 3. Test statistics for Q_n and $\hat{Q}_n$

	Q_n		$\hat{Q}_n$	
	Euro-based	U.S.-based	Euro-based	U.S.-based
Panel A: 11 CEEC				
BG	1.627	3.044[a]	4.977[a]	5.992[a]
CR	0.147	0.518	5.860[a]	5.463[a]
CZ	0.282	0.521	5.701[a]	6.055[a]
HU	0.158	0.282	5.933[a]	6.149[a]
PL	0.345	0.868	4.399[a]	5.990[a]
RO	0.774	1.386	5.231[a]	6.189[a]
SK	3.804[a]	5.846[a]	6.149[a]	6.281[a]
SL	5.001[a]	3.512[a]	5.010[a]	5.292[a]
EE	0.196	0.455	5.149[a]	5.788[a]
LV	1.288	1.853	4.647[a]	5.915[a]
LT	0.651	1.225	5.140[a]	6.110[a]
Panel B: EU-5				
BEL	0.007	0.191	2.198[b]	4.645[a]
FRA	0.031	0.343	4.982[a]	4.250[a]
DEU	0.211	1.070	5.863[a]	3.974[a]
ITA	0.011	0.099	5.066[a]	5.080[a]
NLD	0.115	1.064	4.672[a]	4.894[a]

Notes: The test statistics Q_n and $\hat{Q}_n$ for each country are computed under the null hypothesis of covariance stationary. ([a]), ([b]) and ([c]) denote statistical significance at 0.01, 0.05 and 0.10 based on the critical values of Q_n and $\hat{Q}_n$. The critical values of Q_n and $\hat{Q}_n$ are tabulated in Table A.1.

Table 4. Proportion of the upper bound of the half-life

	CEEC	EU-5	CEEC + EU-5
	$n = 11$	$n = 5$	$n = 16$
HL < 6 months	36.36%	40.00%	37.50%
	(18.19%)	(0.00%)	(12.50%)
6 months ≤ HL ≤ 12 months	27.28%	40.00%	31.25%
	(36.36%)	(60.00%)	(43.75%)
12 months < HL	36.36%	20.00%	31.25%
	(45.45%)	(40.00%)	(43.75%)
Mean (months)	9.64 (10.69)	7.40 (10.61)	8.94 (10.66)
Median (months)	9.72 (11.03)	7.50 (7.74)	8.61 (10.42)

Notes: The values in parentheses refer to the proportion of the half-lives for the U.S. dollar-based rate. Both the mean and the median refer to the upper bound of the CI of the half-life computed from Table 2 (Narayan and Popp Model).

Appendix

Table A.1. Upper tail critical values for Q_n and $\hat{Q}_n$

Sample size n	Level of significance	Critical value	
		Q_n	$\hat{Q}_n$
70	0.10	1.90	1.92
	0.05	2.18	2.19
	0.01	2.76	2.76
100	0.10	1.90	1.91
	0.05	2.16	2.16
	0.01	2.68	2.68
200	0.10	1.92	1.92
	0.05	2.19	2.20
	0.01	2.81	2.81
500	0.10	1.91	1.91
	0.05	2.21	2.21
	0.01	2.73	2.74
1,000	0.10	1.96	1.96
	0.05	2.24	2.24
	0.01	2.77	2.77

Notes: The critical values of test statistics Q_n and $\hat{Q}_n$ are computed under the null hypothesis of covariance stationary. The results are taken at upper tail of the level of significance: 0.10, 0.05 and 0.01 for the sample size (n) = 70, 100, 200, 500 and 1,000, and it based on 5,000 numbers of replications.