Purchasing Power Parity among developing countries and their trade-partners. Evidence from selected CEECs and Implications for their membership of EU.

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Abstract

The purpose of the paper is twofold. Firstly, we test the validity of the PPP hypothesis for selected CEEC (Czech Republic; Hungary; Poland and Slovak Republic). Secondly, we attempt to define those countries' trade linkages between Euro Area; US and the rest of the world. By applying univariate unit root tests as well as a multivariate cointegration test, we find stronger evidence of PPP from the latter test. Moreover, any failure to accept PPP cannot be attributed to structural breaks, apart from one case (between Czech Republic and EU). In overall, there is evidence of strong-form PPP in 6 out of the 8 cases, while for the rest two, weak-form PPP is accepted. Thus, we confirm PPP as a long run equilibrium baseline for these exchange rates per EURO. Furthermore, the fact that PPP holds between these countries and Euro Area indicates absence of trade frictions and other barriers. The implied well-developed trade relations are consistent with those countries' entry into EMU.

Key words: PPP, Unit Root, Structural Breaks, Cointegration, Developing Countries
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I. Introduction

Purchasing Power Parity hypothesis does not hold in the short-run because prices adjust very slowly. Hence, the empirical literature is focused on whether this hypothesis is valid in the long-run, i.e. when prices become flexible. When PPP among developed countries is examined, the main task is the speed of adjustment toward equilibrium. However, many researchers have found evidence of convergence to PPP in the long run with high measures of "half-life" - 3 to 5 years – (see for example, Rogoff, 1996 and Obstfeld & Rogoff, 2000).

Recently, there is an increasing interest in PPP hypothesis for developing countries. Some studies apply univariate unit root tests on real exchange rates, while others apply more powerful panel unit root tests. For example, Bahmani-Oskooee (2000) employ a KPSS test on real effective exchange rates of 20 developing countries and find supportive evidence against PPP. Alba & Park (2003) by employing a panel unit root test (Levin et al, 2002) on 15 developed and 65 developing countries find that PPP is more easily accepted for more open; high inflation and low-growth economies. In line with this finding, Holmes (2000) shows that PPP is strongly accepted for high inflation but rejected for low inflation countries. Using a similar methodology for 88 developing countries, Oh (1996) finds supporting evidence of PPP when the whole period (1950-1990) is examined. In contrast, when this period is split into fixed and flexible exchange rate regimes, PPP cannot be accepted.

Some researchers apply univariate (Engle-Granger, 1987); multivariate (Johansen, 1988) and panel cointegration techniques on the relationship between nominal exchange rates and relative prices. Mahdavi & Zhou (1994) find evidence of weak-form PPP in 8 out of the 13 developing countries and state that stronger evidence exists in relatively high inflation countries. Similarly, Salehizadeh & Taylor (1999) confirm weak-from PPP for 27 developing countries (exchange rates per US dollar). Moreover, Nagayasu (1998) and Boyd & Smith (1998) support PPP in developing countries by employing panel cointegration tests. On the other hand, Drine & Rault (2003) and Basher & Mohsin (2004) fail to confirm PPP in developing countries by panel cointegration tests. The former study shows that while the exchange rate regime does not matter, PPP is more possible to hold in high inflation countries.

Others examine the validity of PPP hypothesis under the presence of structural breaks. Aggarwal et al (2000) and Sabate et al (2003) apply univariate unit root tests with structural breaks and find supporting evidence of PPP for 7 Asian currencies against Japanese Yen and the peseta-sterling exchange rate, respectively. In contrast, Payne et al (2005) cannot establish PPP hypothesis in the case of Croatia by modeling two breaks in unit root tests. Besides to unit root tests, Zurbruegg & Allsopp (2004) apply multivariate cointegration techniques, by allowing the presence of structural breaks, to test PPP in Asian countries in a period including the financial crisis (1997). They conclude that under the presence of significant structural breaks, PPP is found to be a valid long run relationship in 5 out of the 8 countries. In line with the presence of structural breaks, some studies show that convergence to PPP equilibrium may be a non-linear instead of a linear mean reverting process. Indeed, this is confirmed by Sarno (2000) for 11 Middle Eastern countries and by Liew (2003) for Asian developing countries.

The present study concentrates on four Central & Eastern European Countries (Czech Republic; Hungary; Poland and Slovak Republic), which recently became the new country-members of EU. The purpose of this paper is twofold. Firstly, we seek whether PPP is a valid long run relationship in the case of these developing countries. Secondly, we attempt to define those countries' trade linkages between Euro Area; US and the rest of the world. For this reason we examine 3 types of exchange rates. For each country, we estimate 2 bilateral exchange rates (against EURO and US dollar) and the effective exchange rate. In other words, this paper contributes on understanding whether PPP holds as a groundwork of equilibrium exchange rate. Namely, in line with their entry into EU, we expect strong trade linkages with former EU country-members. By establishing PPP hypothesis we can argue that these trade relations exist, indicating no trade frictions and other barriers. Therefore, a normal entry into EMU requires PPP to be valid between these countries and former EU members.

The following section describes the data used in this study, while sections 3 & 4 illustrate evidence of PPP from univariate unit root and multivariate cointegration-based tests, respectively. A final section concludes by evaluating the estimation output.

II. Data

The dataset consists of four bilateral (nominal and real) exchange rates against EURO and four bilateral (nominal and real) exchange rates against US dollar. Real Exchange Rates are computed based on Consumer Price Indices of Czech Republic; Hungary; Poland; Slovak Republic; Euro Area and US. The above rates are taken from OECD statistical database. Exchange rates per EURO stand for cross exchange rates, while the EURO/US dollar exchange rate is estimated by the OECD methodology, in which prior to 1999 rates stand for ECU rates. The data sample includes monthly observations for all variables from 1991:1 to 2003:8 for Czech Republic and Hungary; 1995:1-2003:8 for Poland; and 1993:1 to 2003:8 for Slovak Republic.

Finally, the dataset includes four real (CPI-based) effective exchange rates provided by IFS statistical database (1990:1 to 2004:6). The effective exchange rate is an indicator of the domestic economy's international competitiveness in terms of its foreign exchange rate. It is a measure of the value of the domestic currency against a basket of other currencies. It is calculated as a weighted average of exchange rates and it is expressed as an index (base year 2000 = 100). As a consequence, the effective exchange rate is applied to capture the domestic country's trade linkages with the rest of the world. All variables are presented in natural logarithms.

III. Unit Root Tests

Here we apply two alternative univariate unit root tests (ADF & PP) on bilateral real exchange rates as well as real effective exchange rates. PPP can be accepted only by rejecting the unit root hypothesis. This is because even if the Law of One Price (LOP) does not hold, PPP will be valid if the real exchange rate follows a mean reverting process. In other words, deviations from PPP equilibrium must be only transitory. This is confirmed by establishing the stationary nature of the real exchange rate. Equation (1) expresses the ADF test (Dickey-Fuller, 1981), when both a constant and a trend are included.

$$\Delta Z_t = \gamma + \delta \cdot t + (p-1)Z_{t-1} + \sum_{j=1}^{l} \psi_j \Delta Z_{t-j} + u_t \qquad (1)$$

The problem here is the selection of the appropriate lag length. If "l" is too small, the test will not be asymptotically valid and if "l" is too large, the test will suffer from low power. The Akaike Information Criterion (AIC) provides a useful test to manage this problem. We can choose this lag length, which is associated with the lowest value of the AIC statistic.

To confirm robustness we apply one more unit root test, which has its origins in Phillips (1987) and Phillips-Perron (1988). As in ADF, the P-P test is expressed by equation (2), when both exogenous terms are included.

$$Z_t = \gamma + \delta(t - T/2) + pZ_{t-1} + u_t \qquad (2)$$

Phillips-Perron test computes test statistics suitable for testing the null hypothesis (p=1). For the most restricted case (no exogenous term) these statistics have the following form:

$$K(p) = T(p-1) - \frac{1}{2}(S_{Tl}^2 - S_u^2)(T^{-2}\sum_{t=2}^{T} Z_{t-1}^2)^{-1} \qquad (3)$$

$$K[t(p)] = (S_u / S_{Tl})t(p) - \frac{1}{2}(S_{Tl}^2 - S_u^2)[S_{Tl}(T^{-2}\sum_{t=2}^{T} Z_{t-1}^2)^{1/2}]^{-1} \qquad (4)$$

where $S_u^2 = T^{-1}\sum_{t=1}^{T} \hat{u}_t^2$ and $S_{Tl}^2 = T^{-1}\sum_{t=1}^{T} \hat{u}_t^2 + 2T^{-1}\sum_{j=1}^{l}\sum_{t=j+1}^{T} \hat{u}_t \hat{u}_{t-j}$.

In both ADF and PP tests, the null hypothesis states that the real exchange rate contains a unit root (i.e. p=1). Rejection of the null states that the real exchange rate is mean reverting, indicating that PPP holds in the long-run. Table 1 shows the statistics and the probabilities of accepting the unit root hypothesis. The two alternative tests provide quite similar results. This confirms robustness of our tests. The results show that in the case of Czech Republic both bilateral real exchange rates are non-stationary. Only the real effective exchange rate seems to be stationary (at 5% and 10% significance level according to ADF and PP tests, respectively). Even worse is the evidence for Hungary. There is strong evidence against stationarity in all types of real exchange rates.

Table 1: ADF and P-P Unit Root Tests

Real Exchange Rate	Augmented Dickey-Fuller		Phillips-Perron	
	Exogenous Term (lags)	Statistic (probability)	Exogenous Term (bandwidth)	Statistic (probability)
Czech/EURO	none (1)	1.75 (0.98)	none (2)	2.13 (0.99)
Czech/US	none (1)	1.09 (0.92)	none (4)	1.35 (0.95)
Czech Effective	c & t (2)	-3.70 (0.02)	c & t (6)	-3.30 (0.06)
Hungary/EURO	none (8)	1.06 (0.92)	none (8)	4.57 (1.000)
Hungary/US	none (5)	2.08 (0.99)	none (8)	4.01 (1.000)
Hungary Effective	c & t (1)	2.62 (0.99)	none (10)	3.02 (0.99)
Poland/EURO	constant (2)	-2.94 (0.04)	constant (23)	-3.79 (0.04)
Poland/US	constant (5)	-3.33 (0.01)	constant (6)	-3.33 (0.01)
Poland Effective	constant (4)	-4.003 (0.001)	c & t (0)	-6.12 (0.000)
Slovak/EURO	none (1)	3.25 (0.99)	none (2)	3.83 (1.000)
Slovak/US	none (1)	2.09 (0.99)	none (5)	2.32 (0.99)
Slovak Effective	c & t (1)	-4.90 (0.000)	c & t (10)	-3.76 (0.002)

*MacKinnon (1996) one-sided p-values.

On the other hand, there is strong evidence that real Polish zloty/US dollar and real zloty effective exchange rate are stationary. Weaker evidence, but sufficient, exist for the real Polish zloty/EURO exchange rate (stationary at 5% and 10%). Similarly, non-stationarity is strongly rejected for the Slovak crown real effective exchange rate. In contrast, both bilateral Slovak real exchange rates are found to be non-stationary.

To sum up our findings, when it comes to bilateral exchange rates we found supporting evidence of PPP only in the case of Poland. In line with this, we found that the Polish zloty real effective exchange rate is stationary as well. Thus, the implied consistency with PPP - found in bilateral exchange rates – is incorporated in the real effective exchange rate, which illustrates the external relations of Polish economy with the rest of the world. However, this does not hold in the rest of our estimated exchange rates. While by examining the Slovak and the Czech real effective exchange rates we are able to confirm PPP as valid long run relationship, the bilateral real exchange rates are not mean reverting. Namely, it seems that those countries have more developed trade relations with other countries rather than US and EU. Finally, when Hungary is the case, PPP cannot be accepted in any exchange rate form.

Unit Root Test with Structural Breaks

However, conventional unit root tests may be inappropriate when structural breaks are present in real exchange rates. Kocenda (2001) examines the presence of breaks in the currencies of 11 developing countries against US dollar and Deutsche mark (1991-1997) by the Vogelsang's (1997) approach.[1] The evidence is strong in Balkan and Baltic countries, while this phenomenon is less usual in Central European Countries. In general, structural breaks in exchange rates are present in less stable economies. Under the presence of breaks conventional unit root tests are biased against rejecting non-stationarity. For this reason we apply Perron's (1997) unit root test, which allows the presence of structural breaks in real exchange rates. When it comes to the PPP hypothesis, the presence of structural breaks in exchange rates is alone a negative sign for the validity of this hypothesis. On the other hand, rejection of a unit root in real exchange rates, when breaks exist, implies a mean reverting process. These two findings are indeed contradictory. The above contradiction yields to a new version of PPP, which is called by Hegwood & Papell (1998) as "quasi-long run PPP" – henceforth quasi PPP.[2]

Hence, we test for quasi PPP in those exchange rates which were found non-stationary. The methodology is based upon Perron (1997).[3] Perron (1989) presents three alternative break specification models. The first model, named "Innovational Outlier Model 1", allows only a change in the intercept under both the null and the alternative hypotheses. It has the following form:

$$y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{5}$$

where μ is a constant, DU is a dummy variable which captures the effect on the real exchange rate when the break occurs, t is a time trend and $D(T_b)$ is a dummy variable

[1] This method allows for detecting a break at an unknown date, without imposing any restrictions on the nature of the data.
[2] Quasi PPP is referred to a situation in which the breaks create only transitory shocks.
[3] This test has its origins in Perron (1989). The present test differs from the Perron (1989) in the way the break point is determined. In Perron (1989), the break point was set exogenously. On the contrary, Perron (1997) test determines the break point endogenously.

which captures the effect on the α-coefficient when the break occurs. The term $\sum_{i=1}^{k} c_i \Delta y_{t-i}$ is included in order to "soak up" autocorrelation. The second model, "Innovational Outlier Model 2", allows for both a change in the intercept and the slope at time T_b and has the following form:

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{6}$$

where the dummy DT captures the change in the slope. The third model, "Additive Outlier Model", allows a change in the slope but both segments of the trend function are joined at the time of break. Firstly, the series are de-trended by the regression (7), and finally the test is performed in regression (8)

$$y_t = \mu + \beta t + \gamma DT^* + \tilde{y}_t \tag{7}$$

$$\tilde{y}_t = \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} c_i \Delta \tilde{y}_{t-i} + e_t \tag{8}$$

The main advantage of the Perron (1997) unit root test is that both the time of the break and the k-lag length are treated as unknown. These are identified endogenously to the system. The k-lag length is selected by the "general to specific" procedure instead of any information criteria, such as Akaike and Schwarz. When it comes to the selection of the break date, there are two alternative methods. First, T_b is selected as the value which minimizes the t-statistic for testing α=1. Secondly, T_b is this value which minimizes either the t-statistic on the parameter associated with the change in the intercept (IO1 model), or the t-statistic on the change in the slope (IO2 & AO models). In the present paper we perform this test by the Colletaz & Serranito (1998) procedure for RATS. While the k-lag length is selected by the general to specific method, the break date is selected by minimizing the t_α-statistic. The following table resumes this test's output.

Table 2: Unit Root Test with Breaks

Real Exchange Rate	Model	Sample	Break Time	k	μ	θ	β	γ	δ	α	t_a
Czech/EURO	AO	1993:1-2003:8	1999:07	8	3.21	------	0.006	-0.009	---	0.70	-4.92**
					(564.92)		(70.32)	(-37.01)		(11.82)	
Czech/US	IO2	1993:1-2003:8	2000:05	1	0.36	0.30	9.40	-0.002	-0.09	0.87	-4.05
					(4.10)	(3.70)	(3.94)	(-3.87)	(3.67)	(29.16)	
Hungary/EURO	IO2	1993:1-2003:8	1994:11	3	0.16	0.82	0.001	-0.001	-0.03	0.45	-4.005
					(4.28)	(5.65)	(3.92)	(-4.37)	(-2.18)	(91.46)	
Hungary/US	IO2	1993:1-2003:8	2000:05	1	0.24	0.28	0.001	-0.002	-0.07	0.93	-2.84
					(3.09)	(2.99)	(2.77)	(-3.15)	(-3.35)	(39.55)	
Hungary Effective	AO	1990:1-2004:6	2002:04	11	4.36	------	0.002	0.004	------	0.92	-2.72
					(483.39)		(23.21)	(4.48)		(35.17)	
Slovak/EURO	IO1	1993:1-2003:8	1998:06	1	0.43	0.02	3.76	------	-0.03	0.87	-4.43
					(4.57)	(4.52)	(2.40)		(-2.17)	(31.83)	
Slovak/US	AO	1993:1-2003:8	2003:07	1	3.17	-------	0.008	-0.25	------	-0.92	-2.67
					(163.44)		(32.67)	(-2.33)		(33.23)	

** means rejection of the null at 5% significance level

The results show a significant change only in the constant for the Slovak/EURO (1998:06), while a significant change in both the slope and the constant is found in Czech/US (2000:05); Hungary/EURO (1994:11) and Hungary/US (2000:02) exchange rates. Finally, a significant change only in the slope is found for the Czech/EURO (1999:07); Slovak/US (2003:07) and the Hungarian forint effective exchange rate (2003:04). The break points in Czech exchange rates are not linked with the exchange rate regime switch (1997:05). Furthermore, the break dates in the Hungarian exchange rates do not match with 1991:09, when the exchange rate was fixed to a central parity against EURO. Finally, the exchange rate regime switch, for the case of Slovakia, does not affect the observed breaks because it happens after the end of the estimated period (2004).

When it comes to the unit root hypothesis test, non-stationarity can be rejected in a unique only case. Thus, by allowing the presence of structural breaks we failed to confirm that the failure of rejecting the unit root hypothesis can be attributed to structural breaks. Quasi-PPP is accepted only between Czech Republic and EU. This implies that although the Czech/EURO real exchange rate was stationary, a break (happened in 1999:07) caused a significant change in the slope, which was responsible for deriving misleading results.

In overall, we have found strong evidence that PPP holds for the case of Poland. Besides, PPP hypothesis is accepted among Czech Republic; EU and the rest of its trade partners apart from US. While real Slovak effective exchange rate is stationary, which implies that PPP holds, both bilateral real exchange rates are non-stationary. Finally, there is no sign that PPP is established between Hungary and any of its trade partners. But, can we make valid implications based only on unit root tests? Many researchers argue that univariate unit root tests suffer from low power. They can increase power either by using longer span of data or by employing panel unit root tests. Therefore, it is necessary to re-test the PPP hypothesis before we state that this does not hold when non-stationary real exchange rates are found. Below, we apply a more powerful multivariate cointegration test, based on Johansen's (1988) technique.

IV. Multivariate Cointegration Analysis

At a first stage we need to establish a valid long run relationship among the nominal exchange rate, the domestic and the foreign price levels. This is confirmed by finding at least one cointegrating vector. This is the necessary condition for PPP to hold in the long run. If this is confirmed, the sufficient condition states that the domestic and the foreign CPI's should be proportional. Namely, the proportionality condition requires that if p=1, then p* = -1. If only the necessary condition holds, PPP is accepted in its weak-form. Furthermore, if both the necessary and the sufficient conditions hold, strong-from PPP is accepted.

We start with estimating 8 VAR models in levels, in which the endogenous vector includes 3 variables (nominal exchange rate; domestic CPI; foreign CPI). The appropriate lag length, which "soaks up" autocorrelation, is selected by the Akaike Information Criterion.[4]. Furthermore, we test the specification of each of the VAR models in order to confirm robustness. Specifically, we apply the Lagrange Multiplier test for autocorrelation; the White's heteroskedasticity test and the Jargue-Bera test for normality.

Table 3: Diagnostics

Model	Lags	LM test statistic (probability)	White test statistic (probability)	Jargue-Bera test statistic (probability)
Czech/EURO	2	15.73 (0.07)	104.86 (0.06)	7862.3 (0.0000)
Czech/US	3	4.39 (0.88)	131.46 (0.22)	3900.2 (0.0000)
Hungary/EURO	2	8.11 (0.52)	92.56 (0.24)	1187.1 (0.0000)
Hungary/US	3	6.73 (0.66)	153.8 (0.02)	504.36 (0.0000)
Poland/EURO	7	8.88 (0.44)	257.2 (0.60)	14.19 (0.027)
Poland/US	9	3.14 (0.95)	325.3 (0.65)	50.31 (0.001)
Slovak/EURO	1	7.44 (0.59)	68.7 (0.02)	952.25 (0.0000)
Slovak/US	2	3.26 (0.95)	107.4 (0.04)	746.91 (0.0000)

[4] This statistic is given by $AIC = T\log|\Sigma| + 2N$, where T= number of observations, N = total number of parameters, and $|\Sigma|$ stands for the determinant of the variance/covariance matrix of the residuals. We select this number of lag which fits with the lowest value of the AIC statistic.

The residuals are not serially correlated as the no autocorrelation hypothesis is strongly accepted. When it comes to homoskedasticity, there is strong evidence in 5 out of the 8 models. In 3 models the homoskedasticity hypothesis is rejected at 10% and 5% significance levels, but it is accepted at 1%. Table 3 provides strong evidence against normality. In all cases, except Poland/EURO model, there is strong evidence that errors are not normally distributed. However, this is not really a problem. Since our sample size is quite large, estimators are approximately Normal (Central Limit Theorem). Thus, the presence of Non-normality does not affect the validity of our estimation output.

As we have verified that our VAR models are not misspecified, we can estimate those models in first differences (VECM) to test for cointegration. This is performed by the well-known Johansen Likelihood Ratio test. This test determines the rank of matrix Π ($\Pi=\alpha\beta'$) by computing two test statistics: the Trace and the max-eigenvalue test statistics. Based on the trace statistic we find evidence of cointegration in all the cases. These are shown below[5]:

Table 4: Cointegration Test

Model	**Cointegration Sub-model**	**Cointegrating Vectors**	**Likelihood Ratio Statistic**	**Probability**
Czech/EURO	2	1	0.72	0.69
Czech/US	2	1	0.72	0.69
Hungary/EURO	1	1	32.67	0.00
Hungary/US	2	2	7.13	0.03
Poland/EURO	2	2	7.88	0.02
Poland/US	1	2	2.46	0.29
Slovak/EURO	1	1	16.10	0.00
Slovak/US	1	1	5.51	0.06

MacKinnon-Haug-Michelis (1999) p-values

[5] The column cointegration sub-model corresponds to the cointegration specification. Sub-model 1 does not include any deterministic component in the data. In sub-model 2 there are no linear trends in the data but, a constant term is included to the model.

The last two columns in table 3 represent the test of the proportionality condition. When the cointegrating vector is normalized, we assume that the domestic price is equal to one and the foreign price is equal to minus one. This can be tested by restricting the coefficients in the following way: (s, p, p*) = (1, 1, -1). This hypothesis cannot be accepted in two cases.

In general, we have found evidence of cointegration in all models. However, the proportionality restriction holds in 6 out of the 8 models. This implies that for these 6 models strong-form PPP is confirmed, while for the rest two models, PPP holds only in its weak version. This happens in the Hungary/EURO and Slovak/EURO models. In contrast, the corresponding evidence (when US is the reference country) shows that strong-form PPP is accepted. This fact illustrates that for these two countries and during the estimated period, there are stronger trade linkages between those and US rather than EU.

This points out the significant influence of the US economy on these countries, which are new EU members and potential members of EMU. Does this imply that, at this moment, these countries are more oriented toward US rather than EU? In addition, can we imply that these countries have currently better trade relations with US even though their entry into EU? In our point of view, the answer in both questions is negative. We cannot safely state that these countries have now more developed trade linkages with US. The above contradictory finding is because our data sample describes a past situation instead of the current one. Thus, we need more data (observations) in order to be able to capture the increase of trade linkages with EU and their consequences.

V. Conclusion

This paper tests the validity of the Purchasing Power Parity hypothesis for four Central & Eastern European Countries – members of the European Union (Czech Republic; Hungary; Poland and Slovak Republic). Through the examination of this hypothesis we seek to define how well-developed are the trade relations between those countries and their trade partners. In doing so, we employ three types of exchange rates: two bilateral national rates per US dollar and EURO and a national effective exchange rate. While the bilateral rates capture the trade linkages between the domestic county and the US and EU respectively, the effective exchange rate captures trade relations with the rest of their trade partners.

By applying two univariate unit root tests (ADF, PP) we found evidence of PPP for the cases of Czech Republic; Poland and Slovak Republic (between those and the rest of the world). When it comes to bipartite relations, we found evidence of PPP between Poland & US and Poland & Euro Area. Next, we performed Perron's (1997) unit root test, which allows the presence of a structural break (endogenously determined) in real exchange rates. While we failed to find evidence of PPP between Czech & EU by conventional unit root tests, this test manages to accept quasi-PPP. However, in the rest of the real exchange rates, non-stationarity cannot be rejected although we found significant break points. This implies that any failure to accept PPP cannot be attributed to structural breaks, apart form only one case. Furthermore, we failed to find evidence of PPP between Slovakia & EU and Slovakia & US, even though we found that PPP holds between Slovakia and the rest of the world. This may mean that Slovakia has more developed trade relations with other trade partners rather than US and EU.

However, this contradictory finding may be due to the low power of univariate unit root tests. To confirm our estimation, we employ a more powerful cointegration test. We found evidence of strong-form PPP in 6 out of the 8 cases. Weak-form PPP is accepted between Hungary and EU and Slovakia and EU. The lack of strong-form PPP in these two cases could mean that Hungary and Slovakia are more oriented toward US rather than EU. But, carefully analyzing this output we state that we need more observations in order to see if this is really true.

To sum up, by comparing the results from unit root and multivariate cointegration tests, the latter provides stronger evidence of PPP. Moreover, any rejection of the PPP hypothesis cannot be charged to structural breaks. This happens in only one case. So, focused more on cointegration analysis we confirm PPP as a long run equilibrium baseline for these exchange rates per EURO. This entails a positive implication for the introduction of those countries into EMU. Furthermore, the fact that PPP holds between these countries and Euro Area implies that well-developed trade linkages exist between CEEC and EU. As a consequence, this paper provides supportive evidence that the entry of those countries into EMU is going to be normal.

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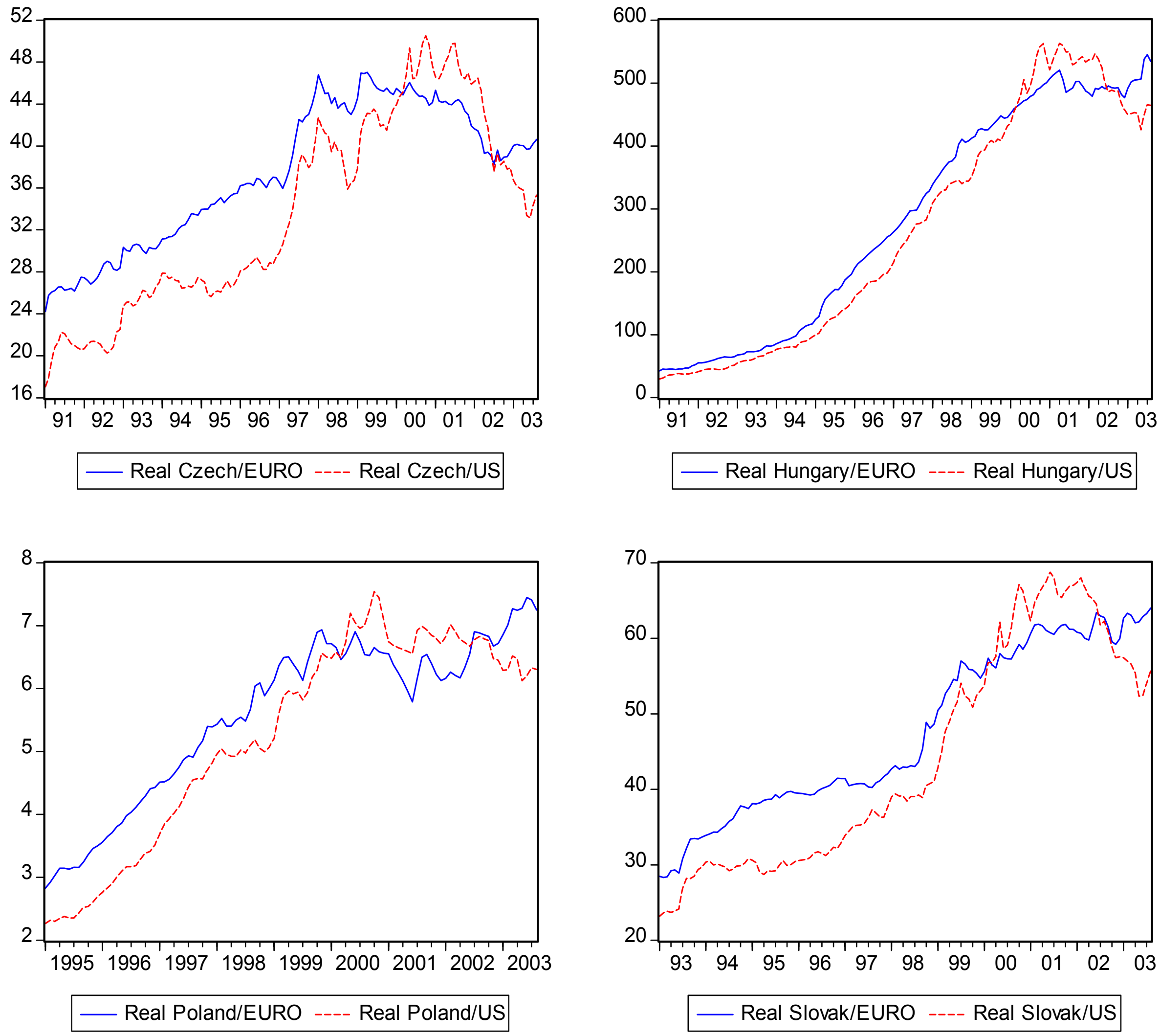
DATA APPENDIX

Variable	Sample	Source	Code
Czech crown/US dollar	1991:1-2003:8	OECD	777003D
Hungarian forint/US dollar	1991:1–2003:8	OECD	807003D
Polish zloty/US dollar	1995:1–2003:8	OECD	817003D
Slovak crown/US dollar	1993:1–2003:8	OECD	797003D
EURO(ECU)/US dollar	1991:1–2003:8	OECD	OL7003D
Czech Real Effective Exchange Rate	1990:1–2004:6	IFS	935..RECZF
Hungarian Real Effective Exchange Rate	1990:1–2004:6	IFS	944..RECZF
Polish Real Effective Exchange Rate	1990:1–2004:6	IFS	964..RECZF
Slovak Real Effective Exchange Rate	1990:1–2004:6	IFS	936..RECZF
Czech Consumer Price Index	1991:1–2003:8	OECD	775241K
Hungarian Consumer Price Index	1991:1–2003:8	OECD	805241K
Polish Consumer Price Index	1995:1–2003:8	OECD	815241K
Slovak Consumer Price Index	1993:1–2003:8	OECD	795241K
Euro Area Consumer Price Index	1991:1–2003:8	OECD	OM5241K
US Consumer Price Index	1991:1–2003:8	OECD	425241K

<u>FIGURES</u>

Bilateral Real Exchange Rates



Real Effective Exchange Rates