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Purchasing Power Parity in CEE and Post-War Former Yugoslav States*

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Abstract
In this paper we investigate purchasing power parity in the CEE and post-war former Yugoslav countries during the EU integration process in 1994–2006. This work stems from longer-term tests of real exchange rate convergence in the former Yugoslavia. This period is of interest on two fronts: first, it investigates real exchange dynamics in the aftermath of a war financed in part through seigniorage; and second, we investigate the level of economic integration with the European Union following the breakup of the former Yugoslavia. Given the short-run nature of the available data we use both univariate and panel unit root tests with and without structural breaks. The results suggest that there is statistical evidence that real exchange rates between the eight transition countries and Germany are stationary when breaks are accounted for. Given the size of nominal shocks in the region, particularly in the early 1990s, estimates indicate that convergence to the long-run equilibrium is relatively quick.

1. Introduction

The goal of this paper is to investigate the convergence to purchasing power parity (PPP) during the transition and EU accession process in Croatia, Macedonia, Serbia, and Slovenia (former Yugoslav countries) and the Czech Republic, Hungary, Poland, and Slovakia (Central European countries) employing the Im, Lee, and Tieslau (2005) two-break LM panel unit root test. The research originates from the vast literature on the long-run validity of the PPP hypothesis and from the growing literature that has emerged investigating the unprecedented appreciation of real exchange rates in transition countries during the last 18 years (Egert, MacDonald, and Halpern 2006).

In addition to testing for PPP in transition countries, we hope to shed light on the way in which the war and the disintegration of the former Yugoslavia affected convergence to PPP in the post-war period. Until recently, a lack of data precluded the full inclusion of Serbia and Macedonia in PPP research. A newly constructed data set allows us to analyze how economic forces function in an environment of post-war normalization and trade redirection over the period 1994–2006.

Transition economy real exchange rate behavior after the introduction of market reforms (post-communist) is unique when compared to either developed or other

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developing countries.[1] One of the stylized facts of transition economies is initial undervaluation of real exchange rates caused by low price levels when compared to similarly developed market economies because of economic planning and the relative isolation of these countries from global goods and capital markets. As demonstrated by Egert et al. (2006) this effect was particularly pronounced in the Czech Republic, Hungary, Poland, and Slovakia.

On the other hand, real exchange rates in the former Yugoslav countries (FYC) were less undervalued. The "self-management" version of communism in the former Yugoslavia was more open to global economic forces, and price levels at the beginning of the transition were more closely aligned with similarly developed market economies. Clearly, the abandonment of central planning in favor of self-management (as early as 1948) resulted in much earlier (pre-transition) price level convergence (Pertot, 1971; Egert et al., 2006).

In the late 1990s, after the initial period of transition, strong appreciation trends in central European (CEE) countries were also attributed to Harrod-Balassa--Samuelson (HBS) related phenomena. Several studies, such as Halpern and Wyplosz (2001), De Broeck and Slok (2001), and Lojschova (2003), have even claimed that it is improbable for transition countries to simultaneously converge in terms of GDP per capita and in terms of European Monetary Union (EMU) inflation targets.

In the former Yugoslav countries there is less evidence in favor of an HBS effect. Relative productivity data for Serbia and Macedonia are not available for much of the period, but Mihaljek and Klau (2004) did manage to find evidence of an HBS effect in Slovenia, though there is little evidence of it in Croatia. Throughout the entire period of transition Croatia, Macedonia, Serbia, and Slovenia have had much smaller (if any) appreciation trends of relative real exchange rates, and during the war and the disintegration of Yugoslavia (1991–1995) real exchange rates experienced unprecedented volatility – when compared to the other four transition countries.[2]

With this in mind, it is clear that the economic environment of transition economies presents a challenge for real exchange rate stationarity. Therefore, a panel Lagrangian multiplier (LM) unit root test, formulated by Im, Lee and Tieslau (2005), with up to two endogenously determined breaks, is employed to accommodate structural shifts and appreciation trends in real exchange rates during transition. Given the long-run nature of real exchange rate behavior, the panel approach decreases the power problem of univariate tests, and the panel LM method with structural breaks should reinforce the power of the test in a volatile economic environment which includes various institutional shocks.

In order to highlight the effect of economic disintegration in the former Yugoslavia, two panels are employed, one including former Yugoslav countries and the other including other Central European countries over the sample period 1994.01–2006.12.

Using panel unit root tests with and without structural breaks we find evidence of PPP convergence in the four former Yugoslav countries and in the other Central European countries in the sample. In addition, when accounting for structural breaks, we also find evidence for convergence in some univariate real exchange rates.

[1] See Egert, Macdonald, and Halpern (2006) for a comprehensive list of transition-specific theories of real exchange rate determination.

[2] Due to several technical reasons pre-1994 data is omitted from our econometric analysis, as discussed below.

The remainder of the paper is divided into four sections. Section 2 provides a theoretical justification for real exchange rate convergence and the econometric methodology. Section 3 discusses the data and provides an overview of the two unit root break tests. Section 4 presents and discusses the results and Section 5 concludes.

2. PPP in Transition Countries

The quality and length of many economic time series complicate analysis of long-run economic hypotheses such as PPP in transition economies. However, several studies have found evidence of univariate convergence to PPP in Eastern European countries. Amacher and Hodgson (1974) were able to find evidence for PPP between the Yugoslav dinar, the German mark, and the Italian lira in the 1950s and 1960s. Using data from 1952–2003 Tica (2006) rejected the null hypothesis of a random walk for the Croatian real exchange rate with respect to Germany, the United States, and Italy. Sideris (2006) performed long-run PPP tests for 17 transition economies using a panel cointegration test. The analysis provided support for long-run equilibrium, but the cointegrating vectors violated the symmetry and proportionality hypotheses suggested by PPP.

In a relatively early paper, Thacker (1995) was unable to reject the null hypothesis of a unit root in the real exchange rates of Hungary and Poland. Similarly, Barlow (2005) employed cointegration methodology to test for PPP in Poland, the Czech Republic, and Romania, and found no evidence of PPP vis-à-vis developed economies. Payne, Lee, and Hofler (2005) employed a battery of unit root tests with structural breaks in order to test short-span PPP in Croatia. As expected, their findings do not demonstrate a mean-reverting process in Croatian real exchange rates. Giannellis and Papadopoulos (2006) managed to reject the null hypothesis in six out of eight real exchange rates in four transition economies.

There are several studies which use cointegration methods to test for PPP in developing and transition economies. Using Jöhansen VECM cointegration tests Mahdavi and Zhou (1994) find evidence for PPP in high-inflation countries, including the former Yugoslavia. Results with quarterly data indicate the existence of either relative or absolute PPP in the former Yugoslavia and seven other non-European countries.

Choudhry (1999) investigated PPP between the USA and Poland, Romania, Russia, and Slovenia and provided evidence for relative PPP only in Slovenia and Russia. Christev and Noorbakhsh (2000) considered PPP in Bulgaria, the Czech Republic, Hungary, Poland, Romania, and Slovakia and though they found some evidence supporting a long-run equilibrium, the estimated cointegrating vector contravenes the values suggested by PPP.

Recently, several papers with nonlinear econometric tests increased the power of the stationarity tests and resulted in stronger evidence for the PPP hypothesis. Cuestas (2009) employed two tests to control for the sources of nonlinearities in eight transition countries. The results indicate that PPP holds in most of these countries once nonlinear deterministic trends and exponential transition have been taken into account. Bahmani-Oskooee, Kutan, and Zhou (2008) tested the null of non-stationarity versus an alternative hypothesis of non-linear stationarity in 88 developing countries including transition economies. The nonlinear model supported the PPP theory in twice as many developing countries compared to the ADF test.

Despite the large number of PPP studies, the relationship between real exchange rate appreciation and the HBS effect, together with several other transition hypotheses, has dominated the discussion about real exchange rate movements during transition in the Czech Republic, Croatia, Hungary, Poland, Slovakia, and Slovenia. However, the studies of relative price behavior have excluded Bosnia and Herzegovina, Serbia, or Serbia-Montenegro, and Macedonia due to a lack of data.

The stylized facts in most Central European and Baltic countries follow a more "traditional" transition pattern than those in the ex-Yugoslavian countries. Halpern and Wyplosz (1997) showed that Slovenia experienced a mild appreciation trend and was not initially undervalued. Croatia has been depicted as an exception in terms of initial undervaluation and in terms of the HBS effect. Egert et al. (2006) and Mihaljek and Klau (2004) find little evidence supporting HBS and no proof of initial undervaluation. Data for Serbia and Macedonia were not used in previous studies, although they did provide evidence that the former Yugoslav countries demonstrate "peculiar" real exchange rate movements during transition.

The most probable explanation of the peculiar transition in the former Yugoslavia lies in its self-managed and non-aligned communist system. The self-management system in the former Yugoslavia was more open to international trade, international capital movements, and even international labor movements than other Eastern European countries.[3]

In addition, the former Yugoslav countries are idiosyncratic because of the war[4] and the disintegration of Yugoslav tariff, monetary, and economic integration, followed by the European integration processes.[5] It might be suggested that such an environment results in specific macroeconomic behavior.

3. Data and Statistical Analysis

3.1 Data

Our monthly inflation and nominal exchange rate sample begins in January 1994, about four years into the transition for the Eastern European economies, because it is from this date that we can collect the most consistent data, and ends in 2006.12. We use the consumer price index (CPI) for most of the countries and for the majority of the period. However, some transition countries did not use the CPI and/or switched relatively late in the transition process. In such cases, we use the retail price index (RPI) rather than the CPI.

The price data for Serbia are the RPI for the entire period 1994.01–2006.12. In the case of Croatia, the CPI was introduced in January 1998 and we use the RPI prior to 1998. Macedonia introduced the CPI in January 1997, thus the RPI is used

[3] As early as 1948 the former Yugoslav type of communism diverged from Soviet-style communism. Planning was abandoned and economic decisions were decentralized at the level of companies, which were run by employees (self-management) and controlled by the party. Price liberalization happened in 1965, and price levels began to converge prior to transition (Pertot, 1971).

[4] The disintegration of Yugoslavia was initiated by various Balkan wars: Slovenia (summer 1991), Croatia (1991–1995), Bosnia and Herzegovina (1992–1996).

[5] EU integration occurred in parallel with the war and the disintegration of the former Yugoslavia. Slovenia is a member of the European Union (EU), and Croatia is expected to join the EU in 2012. Macedonia is a candidate. Bosnia and Herzegovina, Kosovo, Montenegro, and Serbia all lag behind in terms of integration processes.

from 1994.01 to 1996.12. In the Czech Republic, Germany, Hungary, Poland, Slovenia, and Slovakia the CPI is available for the entire period.

Compiling the end-of-month nominal exchange rates vis-à-vis Germany is also complicated due to the introduction of the euro. The German mark nominal exchange rate is used prior to the introduction of the euro in January 1999 and the euro is used thereafter. We use the December 1998 mark-euro conversion rate – 1.95538DM to the euro – after the introduction of euro.

Data for Croatia, the Czech Republic, Hungary, Poland, Slovenia, and Slovakia were acquired from The Vienna Institute for International Economics Studies (2007). Collecting data for Serbia and Macedonia is the most problematic. Data for both countries are not readily available and most of the data were received through direct communication with their national banks. The collapse of the Yugoslav monetary system between October 1991 and April 1992 and the resulting hyperinflation make it difficult to construct meaningful real exchange rates prior to 1993. Furthermore, Serbia uses anachronistic data collection methodology and Macedonia modernized data collection in the late 1990s, prohibiting the extension of our sample prior to 1994.

We conduct our analysis on the real exchange rate, q_t, defined as

$$q_t = \ln\left(E_t \frac{P_t^*}{P_t^i} \right) \tag{1}$$

where P_t^i is the price level in country $i = 1,\ldots,8$; P_t^* is the price level in the numeraire country, Germany; and E_t is the euro-x_i exchange rate, where x_i represents country i 's currency. If PPP holds in the long run, then the series q_t is stationary.

Figure 1 shows the consumer price indices and *Figure 2* the real exchange rate in each of the sample countries. Stabilization programs had finished in all the transition countries by January 1994. Therefore, price level growth was much slower compared to the early 1990s. During the sample period inflation was by far the highest in Serbia, moderate in Hungary, Poland, Slovenia, and Slovakia, and lowest in Macedonia, Croatia, and the Czech Republic.

In terms of real exchange rate movements the former Yugoslav countries exhibited smaller appreciation trends during the post-1994 (war) period than the CEE countries. All countries, with the exception of Serbia, experienced constant real exchange rate appreciation throughout the period. Serbian real exchange rates experienced behavior similar to the so-called "exchange rate based stabilization syndrome" (see Kiguel and Liviatan, 1992). The other four Central European countries slowly tamed inflation and their real exchange rates exhibited stronger appreciation trends due to initial undervaluation of absolute price levels (in PPP terms).

Table 1 presents the descriptive statistics of monthly inflation and the real exchange rate for each of the transition economies from 1994–2006.[6] Given the war--influenced hyperinflation in Serbia in the early 1990s it is not surprising that its economy endured the highest level of inflation and price instability. Similarly, Croa-

[6] Descriptive statistics for nonstationary data are difficult to interpret. However, we present them to give the reader a flavor of the idiosyncrasies of each series.

Figure 1

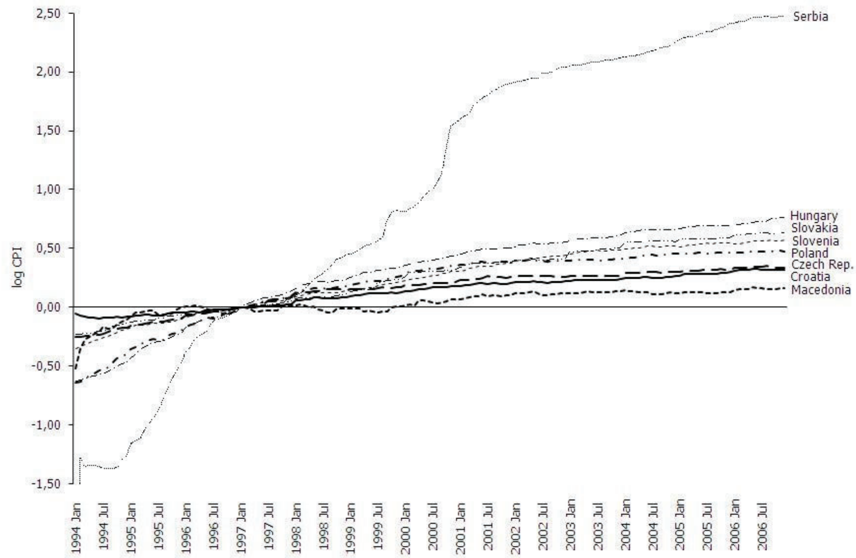
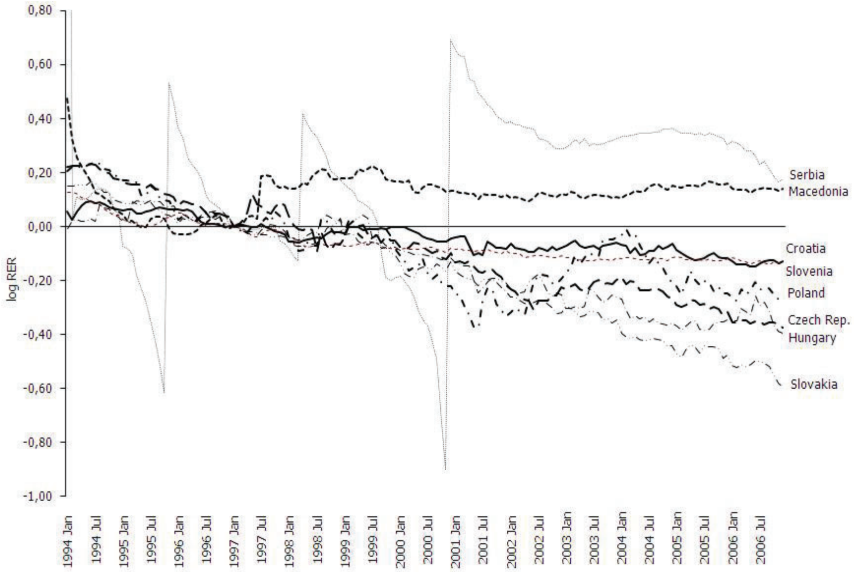


Figure 2



tia and Macedonia, whose economies were also influenced by the war, experienced a substantial degree of instability. The remaining four Eastern European countries had far less price and real exchange rate instability.

Tables 2 and *3* show the inflation and real exchange rate correlations, respectively, over the same period. As can be seen, and perhaps unexpectedly, inflation

Table 1 Descriptive Statistics

	Inflation				Real exchange rate			
	Mean	SD	Min	Max	Mean	SD	Min	Max
Serbia	0.044	0.252	-0.077	3.131	-0.019	0.302	-3.124	1.591
Croatia	0.002	0.005	-0.014	0.028	-0.001	0.011	-0.039	0.049
Slovenia	0.006	0.006	-0.008	0.024	-0.002	0.006	-0.016	0.015
Macedonia	0.004	0.019	-0.029	0.172	-0.002	0.022	-0.150	0.156
Poland	0.007	0.009	-0.009	0.044	-0.003	0.026	-0.053	0.099
Hungary	0.009	0.009	-0.004	0.043	-0.002	0.019	-0.058	0.070
Czech	0.004	0.006	-0.008	0.039	-0.004	0.019	-0.077	0.073
Slovakia	0.006	0.009	-0.004	0.056	-0.005	0.017	-0.066	0.037

Note: Data is monthly, 1994–2006.

Table 2 Inflation Correlation

	Serbia	Croatia	Slovenia	Mace-donia	Poland	Hungary	Czech	Slovakia
Serbia	1.000							
Croatia	-0.227	1.000						
Slovenia	0.116	0.103	1.000					
Macedonia	0.702	-0.155	0.328	1.000				
Poland	0.059	0.211	0.613	0.287	1.000			
Hungary	0.073	0.282	0.539	0.187	0.786	1.000		
Czech	-0.015	0.259	0.275	0.038	0.462	0.428	1.000	
Slovakia	0.010	0.309	0.269	0.069	0.230	0.322	0.373	1.000

Note: Data is monthly, 1994–2006.

Table 3 Real Exchange Rate Correlation

	Serbia	Croatia	Slovenia	Mace-donia	Poland	Hungary	Czech	Slovakia
Serbia	1.000							
Croatia	0.211	1.000						
Slovenia	0.046	0.093	1.000					
Macedonia	0.431	0.134	0.211	1.000				
Poland	-0.028	0.158	0.067	0.079	1.000			
Hungary	-0.104	0.172	0.061	0.067	0.429	1.000		
Czech	0.028	0.104	-0.054	-0.019	0.344	0.173	1.000	
Slovakia	0.067	0.052	0.149	0.047	0.327	0.311	0.274	1.000

Note: Data is monthly, 1994–2006.

within the former Yugoslavian countries exhibits low correlation, with the exception of prices between Serbia and Macedonia, which might be expected given the close relations between these two countries over the sample period. Also of interest is the negative correlation between Croatian and Serbian inflation rates. On the other hand, Slovenia appears to be more closely aligned with the four non-Yugoslav countries than it is with its former co-states. The four non-Yugoslav economies display much higher correlation than the former Yugoslav countries. Obviously, disintegration and war created divergent price behaviors.

Similar results pop out of the correlations with respect to real exchange rates. In non-Yugoslav countries real exchange rates exhibit higher correlation compared to the rest of the sample. The real exchange rate correlation rates between CEE and FYC as well as within the four FYC are low and in some cases even negative.

In terms of real exchange rate movements the former Yugoslav countries are heterogeneous, and slightly higher correlations can be found between Croatia and Serbia and between Macedonia and Slovenia (*Table 3*).

3.2 Statistical Methodology

Perron (1989) was the first to demonstrate that structural breaks in data might be misinterpreted as a permanent stochastic process. He considered three models which explain changes in a deterministic process. In Model "A" the time series undergoes a single level shift; Model "B" exhibits a change in slope; and Model "C" nests both processes. While his test was successful at rejecting unit roots in the standard Nelson and Plosser (1982) data, the test itself requires rather savvy use of the eyeball metric by the econometrician to exogenously choose the break point.

We employ the Im, Lee, and Tieslau (ILT, 2005) panel LM unit root method to test for real exchange rate stationarity in the eight transition economies in our sample. The test is the panel analog of the Schmidt and Phillips (1992) univariate LM unit root test, allowing for up to two endogenously determined structural breaks, and augments the Amsler and Lee (1995) and Lee and Strazicich (2003) one- and two-structural-break univariate LM tests, respectively. Consider the following DGP:

$$y_t = \delta' Z_t + e_t, e_t = \beta e_{t-1} + \varepsilon_t \tag{2}$$

where Z is a vector of exogenous variables. In Model "A" we allow for two level shifts, $Z_t = (1, t, D_{1t}, D_{2t})'$ for $D_{jt} = 0$ for $t < TB_j$ and 1 otherwise. In Model "C" $Z_t = (1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t})'$, where D_{jt} is defined as above and $DT_{jt} = t - TB_j$ for $t \geq TB_j$ and 0 otherwise, a change in the slope coefficients. With this specification, the DGP breaks under the null, $\beta = 1$, and the alternative, $\beta < 1$, hypotheses.

The two-break test is estimated using the LM specification as

$$\Delta y_t = \delta' \Delta Z_t + \varphi \tilde{S}_{t-1} + \sum_{j=1}^{p} \rho_j \Delta \tilde{S}_{t-j} + u_t \tag{3}$$

where $\tilde{S}_t = y_t - \hat{\psi} - Z_t \hat{\delta}$; $\hat{\delta}$ is the estimated coefficient from the regression of Δy_t on ΔZ_t and $\hat{\psi}$ is given by $y_1 - Z_1 \hat{\delta}$, where y_1 and Z_1 are the first-period observations of y and Z. Under the null, $\phi = 0$, which is tested using the Studentized *t*-statistic τ. The number of lagged $\tilde{S}$ is chosen using the standard method of starting with p_{max} (6 months) and working backwards. As is standard in the literature, the two breakpoints are chosen from the interval $[0.1T, 0.9T]$ to avoid endpoints. The no-break and single-break LM unit root tests of Schmidt and Phillips (1992) and Amsler and Lee (1995), respectively, are special cases of the two-break test.

The panel LM unit root test is simply a panel analog of the tests described above and takes as its starting point a regression of the form

$$\Delta y_{it} = \gamma + \delta_i Z_{it} + \beta_i \tilde{S}_{t-1} + \sum_{j=1}^{p_i} \rho_{i,j} \Delta \tilde{S}_{i,t-j} + \nu_{it} \tag{4}$$

where $\tilde{S}_t = y_{it} - \tilde{\gamma} t - \tilde{\delta} D_{it}$ and $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]'$. The variable Z for the two-break test defined nests an additional two models, a no-break test $Z_t = [1, t]'$, and one with a single break $Z_t = [1, t, D_{1t}, DT_{1t}]'$. The LM Studentized t-statistic is calculated under the null hypothesis of $\beta_i = 0$, that is, a unit root denoted $\tau_{LM,i}$ for each i series and its average denoted $\bar{\tau}_{LM}$. Defining τ_{LM} as the t-statistics with *no* structural shifts, the panel LM statistic with breaks is given by

$$\Gamma_{LM} = \frac{\sqrt{N}\left[\bar{\tau}_{LM} - \frac{1}{N}\sum_{i=1}^{N} E(\tau_{LM})\right]}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} Var(\tau_{LM})}} \Rightarrow N(0,1) \tag{5}$$

unless N/T diverges as $N, T \to \infty$.

In cases which contain breaks the minimum LM test uses a "grid search" to endogenously find each of the breaks, defined by $\lambda_j = T_{Bj}/T,\ j = 1,2$, given by $LM_\tau = \inf_{\lambda} \tilde{\tau}(\lambda)$.

4. Results

In total, five versions of the univariate LM test (for each country) and two versions of the panel LM test (for each group of countries) are estimated. The univariate LM tests are estimated with no break; Models A and C are conducted with both one and two breaks. The Panel LM test is estimated without breaks, with one intercept break, and with two intercept breaks.[7] We examine two different panels of countries: panel FYC contains the four former Yugoslavian countries Serbia (SER), Croatia (CRO), Slovenia (SLO), and Macedonia (MAC); and CEE refers to Poland (POL), Hungary (HUN), the Czech Republic (CZE), and Slovakia (SLK).

Table 4 presents the univariate and panel LM unit root test statistics for each of the tests. Univariate estimates of the $AR(1)$ parameter and the optimal lag length are in the columns denoted $\hat{\beta}(p^*)$, and $LM - t$ -statistics are in parenthesis below; the tests with breaks also include the break date. For Model "A", which only includes a level shift, the associated t-statistics are in parenthesis. For Model "C", each break date includes a superscript: *a* denotes a level break statistically significant at the 10% level or better; *b* denotes a trend shift; and *a,b* means both are significant.[8]

[7] The ILT (2005, p. 398) test supports the panel unit root test for Model "A" only.

[8] Detailed results on Model "C" are available upon request.

Table 4 LM Unit Root Tests – 1994–2006

	No break	One Break				Two Breaks					
	Univariate LM Tests										
		Model A		*Model C*		*Model A*			*Model C*		
	$\hat{\beta}(p^*)$	$\hat{\beta}(p^*)$	Break	$\hat{\beta}(p^*)$	Break	$\hat{\beta}(p^*)$	Break 1	Break 2	$\hat{\beta}(p^*)$	Break 1	Break 2
SER	-0.038(0)	-0.040(0)	2000.11	-0.374(0)***	1994.04^{b}	-0.041(0)	1998.03	2000.11	-0.374(0)***	1995.04^{b}	2000.11
	(-1.726)	(-1.772)	(0.981)	(-5.892)		(-1.787)	(0.362)	(0.976)	(-5.856)		
CRO	-0.113(6)*	-0.154(5)*	1996.04*	-0.152(1)***	2001.07^{a,b}	-0.163(5)	1996.04	2000.08	-0.204(1)	1998.02^{a}	2001.07^{b}
	(-2.827)	(-3.353)	(-1.884)	(-3.926)		(-3.436)	(-1.886)	(-0.796)	(-4.671)		
SLO	-0.029(0)	-0.034(2)	2002.05	-0.150(4)***	1998.08^{b}	-0.039(2)	1996.02	2001.11	-0.331(4)***	1995.08^{b}	1998.06^{b}
	(-1.890)	(-2.043)	(-1.058)	(-4.168)		(-2.206)	(-1.957)	(-2.329)	(-6.188)		
MAC	-0.007(5)	-0.010(5)	1997.09**	-0.046(1)	1995.04^{b}	-0.012(5)	1996.10	1997.09	-0.145(5)	1995.04^{b}	1998.06^{a}
	(-0.759)	(-1.011)	(-2.347)	(-1.763)		(-1.223)	(-2.317)	(-2.544)	(-3.215)		
POL	-0.049(1)	-0.062(1)	2003.08***	-0.120(1)**	2002.05^{b}	-0.079(1)	2003.02	2003.08	-0.203(1)	2002.03^{b}	2004.06^{b}
	(-2.198)	(-2.460)	(-2.752)	(-3.343)		(-2.769)	(-1.898)	(-2.809)	(-4.439)		
HUN	-0.044(0)	-0.060(0)	2001.04***	-0.147(0)**	2001.05^{a,b}	-0.072(0)	1998.12	2001.04	-0.200(1)	1996.02	2001.12
	(-1.874)	(-2.161)	(-2.955)	(-3.464)		(-2.382)	(-2.855)	(-3.012)	(-4.054)		
CZE	-0.122(0)*	-0.144(6)**	1997.05	-0.268(6)***	2003.05^{b}	-0.170(6)	1998.09	2002.06	-0.287(6)	2003.04^{b}	2004.07^{b}
	(-3.062)	(-3.390)	(-1.557)	(-4.416)		(-3.685)	(-2.177)	(0.759)	(-4.612)		
SLK	-0.124(2)**	-0.163(2)***	1998.08**	-0.212(6)***	1998.06^{b}	-0.170(2)	1999.02	1999.04	-0.395(6)***	1998.06^{b}	1999.11^{b}
	(-3.523)	(-3.914)	(-2.493)	(-4.288)		(-4.276)	(-3.282)	(-2.519)	(-6.078)		
	Panel LM Tests										
FYC	0.544	-6.780***				-10.404***					
CEE	2.420**	-6.626***				-9.790***					

Notes: ***, **, and * represent rejection of the null hypothesis at the 1%, 5%, and 10% level respectively; $\hat{\beta}(p^*)$ is the estimated AR(1) coefficient and the selected lag length from equation (4). Univariate *LM–t*-unit root statistics is in parenthesis. Critical values are from Schmidt and Phillips (1992) for no break test, Amsler and Lee (1995) for one break test and Lee and Strazicich (2003) for two break test. For the panel statistics, *FYC*: Serbia, Croatia, Slovenia, and Macedonia; and *CEE*: Poland, Hungary, the Czech Rep., and Slovakia. For Model 'A', statistics in parenthesis are *t*-statistics on level breaks; for Model 'C' the break year superscript nomenclature is: *a* if the break is a level break and *b* if it is a trend break, or both, *a,b*, at the 10% level or better. Panel *LM* statistics are: *N*(0,1).

Considering first the results of the no-break test, we can see that only a couple of series are univariate stationary (second column, *Table 4*). And the panel unit root tests display stationarity in the non-former Yugoslavian countries *FYC*. Stationarity tests for the *CEE* countries cannot reject a unit root process.[9]

When a single break is included, we can reject a unit root process in both groups of countries at the 1% level of significance (columns 3–6, *Table 4*). Indeed, we can see that including a break increases the number of rejections of the null in the univariate tests as well. Most success is achieved using Model "C", which accounts for a slope break. Several countries have significant breaks in slope, while Croatia and Hungary display both statistically significant level and trend shifts.

The results from the two-break univariate tests display few rejections when using Model "C" and no rejections of the null with Model "A". On the other hand, the panel unit root tests comfortably reject $I(1)$ processes in both groups of countries (columns 7–12, *Table 4*).

The endogenously estimated breaks can be partly attributed to the numerous institutional shifts, exchange rate regime changes, and political shocks (especially in Serbia). In Serbia the breaks are dominated by nominal shocks. Nominal exchange rate devaluations might explain the breaks in 1995.04, 1998.03, and 2000.11. Also, the end of the war in Croatia might explain the estimated break at the beginning of 1995 in Serbia and that in early 1996 in Croatia. In the case of Croatia most of the shocks from 1998 on might be attributed to the introduction of VAT and the following depreciation. Later shocks are more of an institutional nature. Institutional reforms sped up after the change in government in 2000 and the WTO and EU accession that followed. In Slovenia most of the breaks are estimated in the mid-1990s during the gradual stabilization program. However, the June 2004 membership in the ERM-II is not captured as a shock. Most of the breaks in Macedonia were estimated in 1997–8, as the earliest market-oriented reforms can be traced to that period. The short war of 2001 did not have any major effects on real exchange rate movements.

In the CEE countries the reforms and institutional changes were more homogenous. In Poland four out of the six estimated breaks are within a year from the EU expansion date. The switch to a floating exchange rate regime in 2000.04 helps explain the two breaks in 2002 as a lagged reaction to the exchange regime switch. In Hungary, a change in the currency basket composition can explain four of the estimated breaks – using models "A" and "C" – in 2001. In the Czech Republic the break in 1997.05 can be explained by the switch to a floating exchange rate regime and that in 1998.09 is linked to the introduction of a monetary policy of inflation targeting by the Czech National Bank. The break in the summer of 2004 is linked to EU expansion. The breaks in the Slovakian real exchange rate at the end of 1998 and the beginning of 1999 are probably connected with the start of major market-oriented reforms and the switch to floating in October 1998.[10]

[9] A well known objection to standard unit root tests, under the null of nonstationarity, is that they are subject to Type 1 errors, especially if a break is present in the data. We conducted the Hadri (2000) panel unit root test under the null of $I(0)$, which was able to reject stationarity without breaks. The results are available upon request.

[10] Information on the timing of crucial policy changes in CEE countries is taken from Frömel and Schobert (2003).

In addition to institutional breaks it is interesting to consider our results through the lens of the HBS theory – the major drivers of structural breaks in a transition economy's real exchange rates are due to productivity shocks in the tradable goods sector. Growth of relative productivity in the tradable sector may induce an increase in relative prices of nontradables and create appreciation of the real exchange rate.

On the other hand, it is also possible that investment in research and development and human capital boosted the non-price competitiveness of Eastern European firms in monopolistically competitive markets, or that pricing to market inflated mark-ups in the nontradable sector and augmented prices in the tradable sector as well.

5. Conclusion

This paper uses the real exchange rates of eight transition countries in order to test the PPP hypothesis during 13 years of transition. Three univariate LM unit root tests and the Im, Lee, and Tieslau (2005) panel LM unit root test with structural breaks are employed in order to circumvent problems associated with the power problem, the initial undervaluation of absolute price levels, the strong appreciation trends in CEE countries, and the volatility of the former Yugoslav countries prior to the disintegration of the common country. We find statistical evidence for PPP between the transition countries and Germany when breaks are accounted for. Furthermore, stationarity of the real exchange rates of several countries is implied when using the single-break univariate test. The evidence of stationarity with only 13 years of data using univariate tests, and the low AR(1) coefficient estimates, as well as the rejection of the null in all panels with breaks, are evidence of relatively fast post-war convergence rates in FYC real exchange rates. The economic meaning of these statistical results is rapid movement of macroeconomic relationships vis-à-vis the EU in Central and Southeast Europe toward Western European economies.

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