Darja Boršič, Ph.D.
Jani Bekő, Ph.D.
University of Maribor
Faculty of Economics and Business

PURCHASING POWER PARITY IN THE CZECH REPUBLIC AND SLOVENIA: AN EMPIRICAL TEST

Pariteta kupne moči na Češkem in v Sloveniji: Empirično preverjanje

Abstract

UDC: 339.743(437.1/.2:497.4)

In this paper we test the theory of purchasing power parity for the Czech Republic and Slovenia in comparison to Austria, Germany, France and Italy by employing data from January 1992 to December 2001. Results of unit root tests indicate that the eight time series of the real exchange rates of the koruna and the tolar are integrated of order one. Though some cointegration was found among the nominal exchange rates and selected consumer price indices, the presented results do not support the theory of purchasing power parity in any of the two observed economies.

Key words: purchasing power parity, exchange rate, cointegration

Izvleček

UDK: 339.743(437.1/.2:497.4)

V tem prispevku preverjamo teorijo paritete kupne moči na Češkem in v Sloveniji v primerjavi z Avstrijo, Nemčijo, Francijo in Italijo na osnovi podatkov od januarja 1992 do decembra 2001. Rezultati testov enotnega korena so pokazali, da je vseh osem časovnih vrst realnega deviznega tečaja krone in tolarja integriranih prvega reda. Čeprav smo dokazali kointegracijo med nominalnimi deviznimi tečaji in indeksi cen življenjskih potrebščin, dobljeni rezultati ne potrjujejo teorije paritete kupne moči.

Ključne besede: pariteta kupne moči, devizni tečaj, kointegracija

JEL: E31, F31, F41

1 Introduction

In the last few decades, the validity of the theory of purchasing power parity (PPP) has been scrutinized in numerous empirical papers. Froot and Rogoff (1995), Sarno and Taylor (2002) and Taylor and Taylor (2004) present reviews of relevant literature. Explicit research on PPP theory has yielded varying results, partly as a result of the different estimation techniques, observation periods and data sets that have been employed; and partly because of factors that complicate the law of one price, such as obstacles to international trade, the inclusion of transaction costs, pricing-to-market strategy, discretionary exchange rate management and changes in the structure of price indices. Researchers, however, agree on two issues related to this exchange rate theory (Rogoff 1996): first, real exchange rates tend to converge on levels predicted by PPP in the long run; and second, short-run deviations from the PPP relationship are substantial and variable. While there is a great deal of empirical work on PPP theory for developed market economies, similar studies for transition countries are rather rare. Varamini and Lisachuk (1998) analyze the case of Ukraine for the period 1992–1996 and gain evidence in favor of PPP, despite some short-run deviations. Christev and Noorbakhsh (2000) deal with six Central European Countries (Bulgaria, the Czech Republic, Hungary, Poland, Romania, and Slovakia) in the period from 1991 to 1998. They find moderate proof of long-run equilibrium of prices and exchange rates, but conditions for the law of one price are violated. Pufnik (2002) and Payne et al. (2005) examine the Croatian economy, finding no support for PPP theory. Barlow (2004) also tests the theory for the Czech Republic, Poland and Romania using Johansen cointegration tests, but the conclusions for the time period 1994–2000 are mixed regarding different combinations of the exchange rates of selected countries.

The present paper aims to expand the investigation of PPP for two advanced transition countries: the Czech Republic and Slovenia. Considering different views on how the process of economic transformation since the beginning of the nineties and its effects on reforming countries' price mechanisms are compatible with rigorous assumptions of the theory of PPP (see Brada 1998), there is an obvious need for further empirical evaluation to supply clear-cut evidence on macroeconomic forces that govern the exchange rate behavior in the aforementioned economies. Because the majority of transition countries have undergone several phases of economic restructuring, these most likely also triggered shifts in their equilibrium real exchange rates. This suggests that, when comparing developed market economies with those still under economic reforms, the degree of a country's similarity, especially in terms of trade pattern, level of development and the structure of relative prices, could importantly affect the assessment of PPP. In order to provide detailed estimates, this study is based on separate testing of PPP in the Czech Republic and Slovenia with reference to their main trading partners from the EU–15, i.e. Austria, Germany, France and Italy. From 1992 to 2001, these four countries accounted for 56 percent of Slovenia's exports and imports. In the same period their share in Czech exports amounted to 48 percent and they also covered 51 percent of Czech imports on average.

The paper consists of three additional sections. In Section 2, after describing the general model of PPP and presenting the relevant data, the stationarity of real exchange rates is dissected. Section 3 proceeds with a search for cointegration among nominal exchange rates, domestic consumer prices and foreign consumer prices by relying on Johansen's methodology (1991). Concluding remarks are given in the final section.

2 The Model of PPP and Unit Root Tests of Real Exchange Rates

The general model of testing for PPP can be specified in the following form (Cheung and Lai 1993):

$$e_t = \alpha_0 + \alpha_1 p_t + \alpha_2 p_t^* + \xi_t \quad (1),$$

where e_t stands for nominal exchange rates, defined as the price of foreign currency in the units of domestic currency; p_t denotes domestic price index and p_t^* foreign price index; while ξ_t stands for the error term showing deviations from PPP. All the variables are given in logarithmic form. In the strictest version of PPP, there are the following assumptions: α_0=0, α_1=1, α_2=-1. The symmetry restriction applies such that absolute values of α_1 and α_2 are equal, whereas the limitation of being equal to one is called the proportionality restriction (Froot and Rogoff 1995).

Throughout this study we utilized monthly data series for Slovenia from January 1992 and for the Czech Republic from January 1993 to December 2001 (for both countries), when the euro was put into circulation. Primary data included monthly averages of nominal exchange rates and consumer price indices gathered from the central banks of individual countries. Each of the exchange rates has been defined as the koruna (CZK) or tolar (SIT) cost of a unit of foreign currency. Consumer price indices used in this study for Slovenia refer to January 1992, while for the Czech Republic they refer to January 1993.

The empirical analysis starts off with the most restrictive version of Equation 1, α_1=1, α_2=-1, that is, with testing the properties of real exchange rates. In the context of relative PPP, the movements in nominal exchange rates are expected to compensate for price level shifts. Thus, real exchange rates should be constant over the long run and their time series should be stationary (Parikh and Wakerly 2000). The real exchange rates are a function of nominal exchange rates and relative price indices in two observed economies. They are calculated from the nominal exchange rates using the consumer price indices:

$$RE_t = E_t\,(P_t^*/\,P_t) \quad (2),$$

where RE_t stands for the real exchange rate, E_t is the price of a foreign currency in units of domestic currency, and P_t^* and P_t represent the foreign price index and the domestic price index, respectively. Taking the logarithms of Equation 2, the real exchange rates are defined as:

$$re_t = e_t + p_t^* - p_t \quad (3).$$

The graph of a stationary time series is not supposed to reflect any kind of a time trend. Figure 1 presents the graphs of real exchange rates of the Czech koruna (CZK) and the Slovenian tolar (SIT) in comparison to the Austrian schilling (ATS), the German mark (DEM), the French franc (FRF) and the Italian lira (ITL). The graphs of exchange rates show that after an initial real depreciation, from 1996 onwards, the Slovenian tolar experienced a systematic real appreciation in comparison to the currencies of the selected market economies. As can be seen from Figure 1, the regular real appreciation of the Czech koruna against the currencies of the four developed market economies in the 1993–2001 period was partly interrupted only in 1997 reflecting exchange rate instability due to a domestic currency crisis. Such a pattern in real exchange rate movements is explained in the literature by a range of factors, including inherited macroeconomic imbalances in transition countries, mixed performance of chosen exchange rate arrangements, monetary difficulties arising from huge capital inflows, the inflationary impact of wage and price adjustments, and real exchange rate appreciation due to the catching-up process (Halpern and Wyplosz 1997; Brada 1998).

For checking the stationarity of real exchange rates, the augmented Dickey-Fuller (1979) test was used, taking into account the following equation:

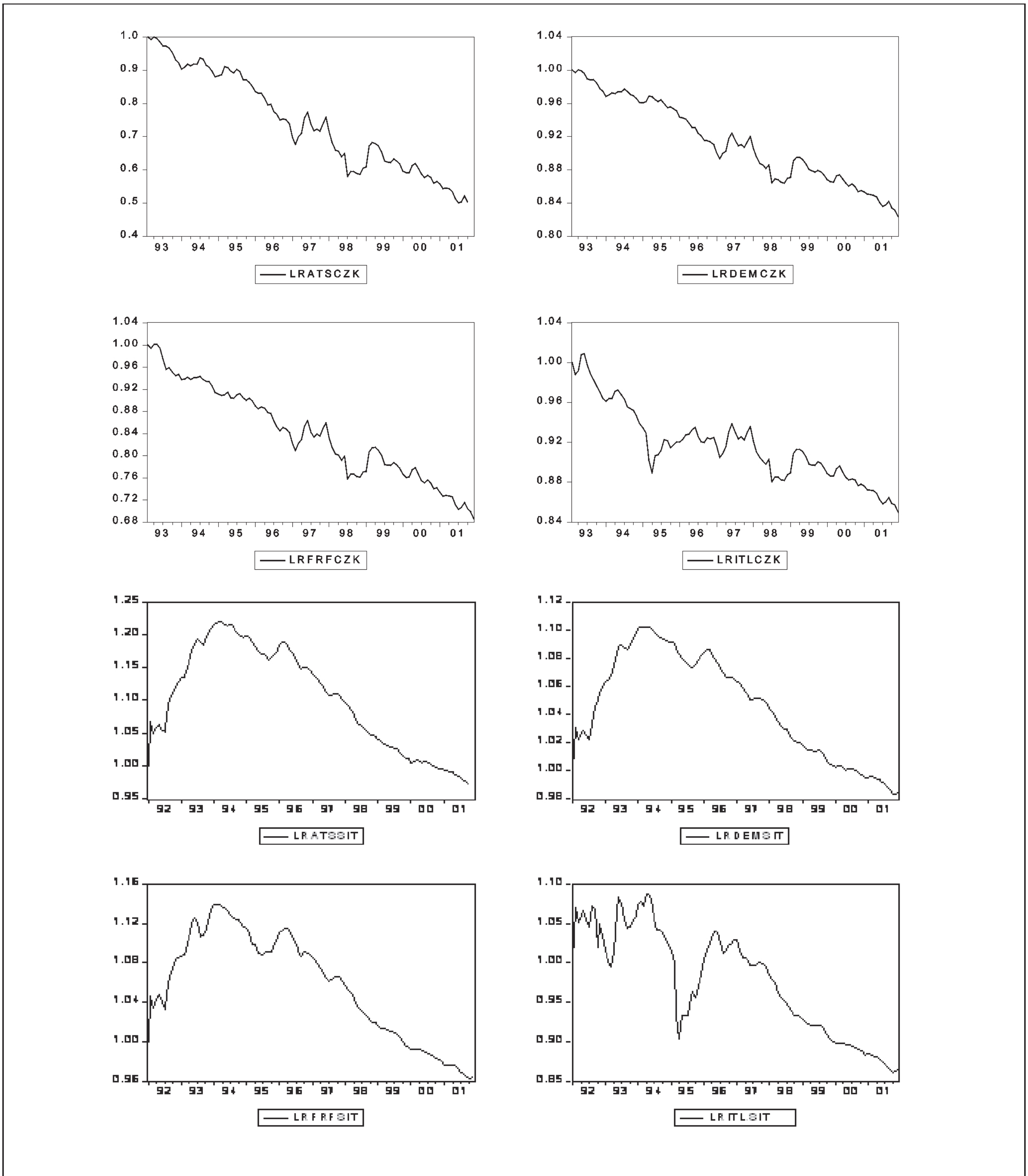
$$\Delta Y_t = \beta_1 + \beta_2 t + \delta Y_{t-1} + \sum_{i=1}^{m} \chi_i \Delta Y_{t-i} + \varepsilon_t \quad (4),$$

where β_1, β_2, δ and χ_i are parameters of the test, t is linear time trend, Y_t is the tested time series, $\Delta Y_{t-i}=Y_{t-i}-Y_{t-i-1}$ and m is selected so that the residuals (ε_t) are white noise. We test the null hypothesis H_0: δ=0, which implies that there is a unit root present and the time series is non-stationary.

Following Barlow (2004), Equation 4 was estimated assuming β_2=0. In order not to unnecessarily lose too many observations in a relatively short time series, the orders of augmentation were set to m=6 for all tests of unit root by using critical values according to MacKinnon (1991). Campbell and Perron (1991) prefer determining the time lags according to a t-test. They argue that a VAR with a maximum number of lags should be carried out. If the last included lag is statistically significant, it is appropriate to use it in ADF regressions. The number of lags should be reduced as long as the last included lag is statistically significant. Also Ng and Perron (1995) argue that information-based rules (AIC, SIC) tend to select too low truncation lags, while the t-test is supposed to provide results with more robust size properties in models. In the present analysis, the estimates are obtained on the basis of time lags which correspond to the minimum value of the Akaike Information Criterion (AIC) and are in line with the t-test approach.

Results of the augmented Dickey-Fuller test are shown in Table 1. Each calculation is stated twice, according to the time lag determined by the two approaches described above. Although AIC and the t-test select different time lags, the results of the ADF test using both selection criteria do not

Figure 1: *Real Exchange Rates of the Czech koruna and the Slovenian tolar*



Notes: L stands for logarithm, R for real; the next three letters (ATS, DEM, FRF, ITL) represent the currencies of Austria, Germany, France and Italy, respectively, while the last two letters (CZK, SIT) denote the currencies of the Czech Republic and Slovenia, respectively. 1992:01=100 for Slovenia, 1993:01=100 for the Czech Republic.

Source: The Czech National Bank and Bank of Slovenia.

Table 1: *Results of the ADF Test for Real Exchange Rates of the Czech koruna and the Slovenian tolar*

Variable	Level		First difference	
	AIC	t-test	AIC	t-test
LRATSCZK	-0.7843_1	-0.7679_6	-4.2397_6	-4.2397_6
LRDEMCZK	-0.4984_1	-0.5233_6	-4.1920_6	-4.1920_6
LRFRFCZK	-0.3030_6	-0.3030_6	-4.0726_6	-4.0726_6
LRITLCZK	-1.5060_6	-1.5060_6	-3.9556_6	-3.5755_5
LRATSSIT	-0.6400_6	0.0867_3	-2.9538_6	-2.9538_6
LRDEMSIT	-0.8162_6	0.0138_3	-3.2579_6	-3.2579_6
LRFRFSIT	-0.6003_6	-0.3864_2	-2.8850_6	-2.8850_6
LRITLSIT	-0.8123_4	-0.8123_4	-5.7593_3	-5.3557_4

Notes: L stands for logarithm, R for real; the next three letters (ATS, DEM, FRF, ITL) represent the currencies of Austria, Germany, France and Italy, respectively, while the last two letters (CZK, SIT) denote the currencies of the Czech Republic and Slovenia, respectively. Critical values: -3.4890 (1%), -2.8870 (5%) and -2.5802 (10%). The subscripts indicate the value of m in Equation 4.

contradict, but are rather similar. The figures show that the eight time series of the real exchange rates of the koruna and the tolar are integrated of order one, which means we cannot reject the hypothesis of the presence of the unit root. Thus, the ADF test confirms the graphical results of non-stationarity in the observed time series.

3 Cointegration Analysis and Comments on Results

When all restraints in Equation 1 are omitted ($\alpha_1 \neq 1$, $\alpha_2 \neq -1$), it becomes the least restrictive version of PPP. The only requirement that remains is the signs of the coefficients. This implies that we are looking for any linear relationship among the observed variables that has stationary properties. Taking into account the unstable characteristics of non-stationary time series, the existence of a stationary relationship among them is more important than deviations of coefficients from the strict theory of PPP (Liu 1992). If a cointegration among nominal exchange rates, domestic consumer prices and foreign consumer prices is found and it is presented by the cointegrating vector of $(1, \alpha_1, \alpha_2)$ (Equation 1), the validity of the theory of PPP is proven.

Since we are looking for a stationary linear combination of three variables, the Johansen cointegration test is appropriate to use. This method is based on a VAR and can be briefly described as follows (Johansen 1991):

$$Y_t = A_1 Y_{t-1} + \dots + A_m Y_{t-m} + BX_t + \eta_t, \quad (5),$$

where $A_1, \dots, A_m$ and B are matrices and the parameters of the model, t ranges from 1 to T, Y_t is a vector of k variables, which are integrated of the first order, X_t is vector of deterministic variables and η_t is a vector of innovations. VAR in Equation 5 can be also written as:

$$\Delta Y_t = \Pi Y_{t-1} + \sum_{i=1}^{m-1} \Gamma_i \Delta Y_{t-i} + BX_t + \eta_t \quad (6),$$

$$\text{where } \Pi = \sum_{i=1}^{m} A_i - \mathrm{I} \text{ and } \Gamma_i = -\sum_{j=i+1}^{m} A_j \quad (7).$$

Matrix Π contains information about long-run variation of the time series. According to the Granger representation theorem (Engle and Granger 1987; Johansen 1991), matrix Π can be divided into k x r matrices ρ and α with rank of r (r≤k-1), so that Π=ρα' if Π also has reduced rank r<k. Matrix ρ contains r linear cointegrating vectors, while matrix ρ presents adjustment coefficients of the error correction model.

The number of cointegrating vectors is assessed by two statistics. The trace statistic (LR_{tr}) tests H_0: the number of cointegrating vectors is less than or equal to r, against the H_1: the number of cointegrating vectors is k, where k is the number of endogenous variables for r=0, 1, ..., k-1. The trace statistic is specified as:

$$LR_{tr}(r \mid k) = -T \sum_{i=r+1}^{k} \log(1-\lambda_i) \quad (8),$$

where λ_i is the maximum eigenvalue of A_i in Equation 7. The maximum eigenvalue statistic (LR_{max}) checks H_0: the number of cointegrating vectors is equal to r, and H_1: the number of cointegrating vectors is equal to r + 1. LR_{max} can be calculated as follows:

$$\begin{aligned} LR_{\max}(r \mid r+1) &= -T\log(1-\lambda_{r+1}) = \\ &= LR_{tr}(r \mid k) - LR_{tr}(r+1 \mid k) \end{aligned} \quad (9),$$

where the abbreviations are the same as in Equation 8 and the text above.

Critical values for the Johansen cointegration test are stated in Johansen (1988) and Johansen and Juselius (1990). Osterwald-Lenum (1992) recalculated and extended them by handling the whole test sequence. Therefore, this study applies improved critical values of Osterwald-Lenum (1992). To undertake the Johansen cointegration test, an appropriate lag structure had to be found in order to remove serial correlation in the residuals. Estimation on the basis of VAR's Akaike Information Criterion (AIC) and Final Prediction Error (FPE) gave the same lag specification for all eight cases under consideration. Figures for time lags are quoted next to the individual countries' names in Table 4.

Prior to cointegration analysis, it is necessary to establish the compatible orders of integration of the employed variables. For this reason, ADF tests were conducted for individual nominal exchange rates, domestic consumer prices and foreign consumer prices following the procedure described in the previous section. Results of unit root tests for nominal exchange rates are presented in Table 2, while Table 3 summarizes the unit root tests for selected consumer price indices. Again, AIC and a t-test were used to determine the number of time lags in ADF regressions.

Table 2: *Results of the ADF Test for Nominal Exchange Rates of the Czech koruna and the Slovenian tolar*

Variable	Level		First difference	
	AIC	t-test	AIC	t-test
LNATSCZK	-2.2033_1	-2.2033_1	-6.1397_1	-6.1397_1
LNDEMCZK	-2.2096_1	-2.2096_1	-6.1650_1	-6.1650_1
LNFRFCZK	-2.1139_1	-2.1139_1	-6.2416_1	-6.2416_1
LNITLCZK	-2.2676_3	-2.2987_1	-5.7129_2	-7.3552_1
LNATSSIT	-1.4738_6	-1.4738_6	-3.9710_6	-3.9710_6
LNDEMSIT	-1.4646_6	-1.4646_6	-4.0363_6	-4.0363_6
LNFRFSIT	-2.1977_6	-4.0035_3	-3.6652_6	-3.6652_6
LNITLSIT	-0.9879_4	-0.8748_5	-6.2102_3	-5.2060_4

Notes: L stands for logarithm, N for nominal; the next three letters (ATS, DEM, FRF, ITL) represent the currencies of Austria, Germany, France and Italy, respectively, while the last two letters (CZK, SIT) denote the currencies of the Czech Republic and Slovenia, respectively. Critical values: -3.4890 (1%), -2.8870 (5%) and -2.5802 (10%). The subscripts indicate the value of m in Equation 4.

All the nominal exchange rates in level form are found to be non-stationary, except the exchange rate of the Slovenian tolar to the French franc, which is stationary according to lag specification by t-test. A glance at the figures in Table 3 reveals that stationarity of five consumer prices is achieved (at the 5% significance level at least) only after the series are transformed into first differences. The Czech consumer price index, however, is specified as I(2).

MacDonald (1993) claims that also in the case of different orders of integration, it is possible that the volatility of variables still implies a stationary linear combination among them. This is clearly impossible in the case of three variables being integrated of three different orders (Granger 1986). To retain consistency of the lag estimation criterion by ADF and cointegration tests, the integration orders of variables that enter cointegrating relations in our study were based upon the AIC. Table 4 and Table 5 list the results of the Johansen test by applying the basic model for PPP testing, i.e. Equation 1.

From Table 4 it can be seen that for the Czech Republic limited evidence on cointegration among the nominal exchange rates and consumer prices was found, but only in comparison to France and Italy. In both pairs of countries the estimated coefficients appear to be statistically significantly different from zero. According to Equation 1, the signs of the coefficients of domestic prices should be positive, while signs of the coefficients of foreign prices should be negative. Thus, the signs of all cointegrating coefficients invalidate the PPP theory on the Czech data.

Looking at Slovenia, values of LR_{tr} and LR_{max} show that there is cointegration among the nominal exchange rates and consumer price indices in comparison to Austria, Germany and Italy. In all three cases the coefficients of domestic prices are proven to be statistically significantly different from zero, while for coefficients of foreign prices the standard errors are too high to conclude the same. The signs of estimated cointegrating coefficients, reported in Table 4, are again wrong to confirm PPP. Only the coefficient of Austrian consumer prices tends to have the right sign. In reference to France, there is no proof of cointegration either. In addition, the estimated coefficients are statistically insignificant and only the coefficient of French consumer prices has a sign corresponding to PPP theory.

The presented results do not support the theory of PPP in any of the two observed economies. Such an outcome is in line with the rather weak empirical evidence on PPP reported for transition countries in the introductory part of this paper. The invalidity of PPP found in our study is also consistent with the real appreciation of the national currencies of the Czech Republic and Slovenia stated by, inter alia, Desai (1998) and Bole (1999). One part of real exchange rate appreciation can be attributed to the faster growth of domestic tradable prices compared to tradable prices of developed European economies, although this sort of real appreciation was substantially mitigated in Slovenia by monetary policy interventions on foreign exchange markets in order to preserve external competitiveness (Bole 1999). In the Czech Republic, on the other hand, the contribution of relative prices of tradables to the real exchange rate appreciation was preponderant (Kovács et al. 2002) and the domestic monetary authorities were, until 1997, obliged to sustain the exchange rate peg. The Czech example, therefore, corroborates the findings of Barlow (2004) that implementation of a more rigid exchange rate policy in conditions of still volatile inflation and price inertia is to blame for violating PPP.

Table 3: *Results of the ADF Test for Individual Consumer Price Indices in the Observed Countries*

Variable	Level		First difference	
	AIC	t-statistic	AIC	t-statistic
LCPIC	-2.4320_6	-2.4320_6	-1.8331_6 $(-12.9527_4)^1$	-1.6783_5 $(-12.9527_4)^1$
LCPIS	-1.7299_3	-2.4605_2	-75.0159_1	-75.0159_1
LCPIA	-1.1433_6	-1.8836_5	-4.5889_5	-5.3592_4
LCPIG	-2.2409_1	-2.2797_2	-6.5652_2	-6.5652_1
LCPIF	-1.5175_6	-1.6488_5	-4.3132_6	-4.3132_6
LCPII	-1.2901_1	-1.2901_1	-3.1141_3	-3.3836_4

Notes: L stands for logarithm, CPI for consumer price index; C, S, A, G, F and I denote the Czech Republic, Slovenia, Austria, Germany, France and Italy, respectively. Critical values: -3.4890 (1%), -2.8870 (5%) and -2.5802 (10%). The subscripts indicate the value of m in Equation 4. [1]Second difference.

Table 4: *Results of the Johansen Cointegration Test for the Czech Republic and Slovenia*

Number of cointegrating equations		Czech Republic		Slovenia	
		Statistic [1,2]		Statistic [1,2]	
Austria$_1$		α_1=-3.2421 (0.9084) α_2 =12.3977 (4.0192)		α_1=-0.7075(0.3732) α_2=-1.8258 (2.7453)	
H_0:	r=0 r≤1 r≤2	LR_{tr}	23.3980 9.0559 0.7835	LR_{tr}	**61.4312 13.6185 *4.7704
H_0:	r=0 r=1 r=2	LR_{max}	14.3421 8.2724 0.7835	LR_{max}	**47.8127 8.8481 *4.7704
Germany$_2$		α_1=-1.2001 (0.2886) α_2=4.8959 (1.3415)		α_1=-0.7090 (0.3012) α_2=2.2350 (2.4313)	α_1=-0.4601 (0.0461) α_2=0.0000
H_0:	r=0 r≤1 r≤2	LR_{tr}	23.9747 8.9509 0.4842	LR_{tr}	**42.2441 **20.7690 2.7287
H_0:	r=0 r=1 r=2	LR_{max}	15.0238 8.4667 0.4842	LR_{max}	*21.4751 *18.0402 2.7287
France$_1$		α_1=-2.8237 (0.5052) α_2=12.9935 (2.7350)		α_1=-0.3307 (0.3878) α_2=-4.4198 (3.5580)	
H_0:	r=0 r≤1 r≤2	LR_{tr}	29.5547 8.2806 0.7824	LR_{tr}	25.9711 6.7105 0.7913
H_0:	r=0 r=1 r=2	LR_{max}	*21.2741 7.4982 0.7824	LR_{max}	19.2606 5.9192 0.7913
Italy$_2$		α_1=-1.5897 (0.5011) α_2=4.1891 (1.2320)		α_1=-1.1967 (0.4644) α_2=0.9779 (1.8231)	
H_0:	r=0 r≤1 r≤2	LR_{tr}	*34.0048 *15.4926 2.2730	LR_{tr}	**37.7476 12.3176 *4.4377
H_0:	r=0 r=1 r=2	LR_{max}	18.5122 13.2196 2.2730	LR_{max}	*25.4300 7.8798 *4.4377

Notes: ** (*) denotes rejection of the null hypothesis at the 1% (5%) significance level, respectively; figures in parentheses are standard errors. [1]Critical values for LR_{tr} at the 5% level are 29.68 (r=0), 15.41 (rd”1), and 3.76 (rd”2); and at the 1% level are 35.65 (r=0), 20.04 (rd”1), and 6.65 (rd”2). [2]Critical values for LR_{max} at the 5% level are 20.97 (r=0), 14.07 (r=1), and 3.76 (r=2); and at the 1% level are 25.52 (r=0), 18.63 (r=1), and 6.65 (r=2).

In Slovenia, a far more important source of real exchange rate appreciation comes from faster growth of nontradable to tradable prices in comparison to relative prices of developed market economies. As documented in Kovács (2004), changes in relative labor productivity explain a considerable portion of nontradable/tradable relative price behavior in Slovenia in the 1992–2001 period. Besides this productivity-based real appreciation, relative mark-ups account for real appreciation in the case of Slovenia as well (see also Bole 2003). In addition, studies by Kutan and Dibooglu (1998) and Kovács (2004) imply that the variety of real shocks encountered by transition economies and expansive macroeconomic policies can significantly strengthen real exchange rate appreciation, the former via improving efficiency and boosting productivity, while the latter by increasing inflation differentials with respect to levels in developed market economies.

4 Concluding remarks

Testing for stationarity of real exchange rates of the Czech koruna and the Slovenian tolar showed no firm evidence in favor of PPP. After examining the stationarity of real exchange rates, the proportionality and symmetry restrictions were omitted. The Johansen cointegration technique was applied to find a long run linear relationship among chosen nominal exchange rates and individual time series of consumer prices. Although some cointegration was proven, the theory of PPP could not be confirmed. Regarding the low national price levels in both countries in question (see, for example, IEDP 2003; 2004) compared to levels in the EU–15, even after a decade of reforms, such a result is not unexpected. Another reason for failing to substantiate PPP could be the relatively short period of observation for such a long relationship to be detected among the observed variables.[1] Since the early nineties both countries had already pursued a strategy of more or less successful gradual disinflation. Managing low variations of nominal exchange rates during periods of excessive inflation could also imply deviations from PPP. However, the empirical work completed so far reveals that the underlying cause of real exchange rate appreciation in Slovenia stems from differences in relative productivity gains and from steady price increases due to inadequate competition in the nontradable sector.

[1] Rogoff (1996) stresses that it takes three to five years for one half of the exchange rate deviation from the PPP level to be completed.

References

1. Barlow, D. (2004). Purchasing Power Parity in Three Transition Economies. *Economics of Planning* 36 (3): 201–221.
2. Bole, V. (1999). Financial Flows to a Small Open Economy: The Case of Slovenia. In *The Mixed Blessing of Financial Inflows, Transition Countries in Comparative Perspective*, ed. J. G<cs, R. Holzmann and M.L. Wyzan, 195–238. Cheltenham: Edward Elgar.
3. Bole, V. (2003). Denarna politika v času odštevanja (Monetary Policy in the Time of Countdown). *Gospodarska gibanja* 346: 23–43.
4. Brada, J.C. (1998). Introduction: Exchange Rates, Capital Flows, and Commercial Policies in Transition Economies. *Journal of Comparative Economics* 26 (4): 613–620.
5. Campbell, J.Y., and P. Perron (1991). Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots. *NBER Working Paper Series: Technical Working Paper*, no. 100.
6. Cheung, Y., and K.S. Lai (1993). Long-Run Purchasing Power Parity During the Recent Float. *Journal of International Economics* 34 (1–2): 181–192.
7. Christev, A., and A. Noorbakhsh (2000). Long-Run Purchasing Power Parity, Prices and Exchange Rates in Transition. The Case of Six Central and East European Countries. *Global Finance Journal* 11 (1–2): 87–108.
8. Desai, P. (1998). Macroeconomic Fragility and Exchange Rate Vulnerability: A Cautionary Record of Transition Economies. *Journal of Comparative Economics* 26 (4): 621–641.
9. Dickey, D.A., and W.A. Fuller (1979). Distribution of Estimators for Autoregressive Time Series With a Unit Root. *Journal of the American Statistical Association* 74 (366): 427–431.
10. Engle, R.F., and C.W.J. Granger (1987). Co-integration and Error Correction – Representation, Estimation and Testing. *Econometrica* 55 (2): 251–276.
11. Froot K.A., and K. Rogoff (1995). Perspectives on PPP and Long run Real Exchange Rates. In *Handbook of International Economics Vol. III*, ed. G. Grossman and K. Rogoff, 1647–1688. Elsevier Science.
12. Granger, C.W.J. (1986). Developments in the Study of Cointegrated Economic Variables. *Oxford Bulletin of Economics and Statistics* 48 (3): 213–228.
13. Halpern, L., and C. Wyplosz (1997). Equilibrium Exchange Rates in Transition Economies. *IMF Staff Papers* 44 (4): 430–461.
14. Institute for Economic Diagnosis and Prognosis (IEDP) (2003). Slovenija in Avstrija – Ravni cen in plač (Slovenia and Austria – Levels of Prices and Wages. *Bilten EDP* 26 (2–3). Maribor.
15. Institute for Economic Diagnosis and Prognosis (IEDP) (2004). Slovenija in Češka – Ravni cen in plač (Slovenia and the Czech Republic – Levels of Prices and Wages). *Bilten EDP* 27 (1). Maribor.
16. Johansen, S. (1988). Statistical Analysis of Cointegration Vectors. *Journal of Economic Dynamics and Control* 12 (2–3): 231–254.
17. Johansen, S. (1991). Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. *Econometrica* 59 (6): 1551–1580.
18. Johansen, S, and K. Juselius (1990). Maximum Likelihood Estimation and Inference on Cointegration – with Applications to the Demand for Money. *Oxford Bulletin of Economics and Statistics* 52 (2): 169–210.
19. Kovács, M.A. (ed.); J. Benes; M. Klima; J. Borowski; M.K. Dudek; P. Sotomska-Krzysztofik; F. Hajnovic; and T. Žumer (2002). On the Estimated Size of the Balassa-Samuelson Effect in Five Central and Eastern European Countries. *NBH Working Paper* 5, July.
20. Kovács, M.A. (2004). Disentangling the Balassa-Samuelson Effect in CEC5 Countries in the Prospect of EMU Enlargement. In *Monetary Strategies for Joining the Euro*, ed. Gy. Szap<ry and J. von Hagen, 79–105. Cheltenham: Edward Elgar.
21. Kutan, A.M., and S. Dibooglu (1998). Sources of Real and Nominal Exchange Rate Fluctuations in Transition Economies. *The Federal Reserve Bank of St. Louis, Working Paper* 1998-022A.
22. Liu, P. C. (1992). Purchasing Power Parity in Latin America: A Co-Integration Analysis. *Weltwirtschaftliches Archiv* 128 (4): 662–679.
23. MacKinnon, J. (1991). Critical Values for Cointegration Tests. In *Long-Run Economic Relationships: Readings in Cointegration*, ed. R.F. Engle and C.W.J. Granger. Oxford: Oxford University Press.
24. MacDonald, R. (1993). Long-Run Purchasing Power Parity: Is It For Real? *The Review of Economics and Statistics* 75 (4): 690–695.
25. Ng, S., and P. Perron (1995). Unit Root Tests in ARIMA Models with Data-Dependent Methods for the Selection of the Truncation Lag." *Journal of American Statistical Association* 90 (429): 268–281.
26. Osterwald-Lenum, M. (1992). A Note with Quantiles of Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics. *Oxford Bulletin of Economics and Statististics* 54 (3): 461–472.
27. Parikh, A., and E. Wakerly (2000). Real Exchange Rates and Unit Root Tests. *Weltwirtschaftliches Archiv* 136 (3): 478–490.
28. Payne, J.; J. Lee; and R. Hofler (2005). Purchasing Power Parity: Evidence from a Transition Economy. *Journal of Policy Modeling* 27 (6): 665–672.
29. Pufnik, A. (2002). Purchasing Power Parity as a Long-Run Equilibrium: Co-Integration Test in the Case of Croatia (1991–1996). *Croatian Economic Survey 1996–1999*: 29–54.
30. Rogoff, K. (1996). The Purchasing Power Parity Puzzle. *Journal of Economic Literature* 34 (2): 647–668.
31. Sarno, L., and M.P. Taylor (2002). Purchasing Power Parity and the Real Exchange Rate. *IMF Staff Papers* 49 (1): 65–105.
32. Taylor, A.M., and M.P. Taylor (2004). The Purchasing Power Parity Debate. *NBER Working Paper* 10607.
33. Varamini, H., and H. G. Lisachuk (1998). The Application of Purchasing Power Parity to Ukraine by Using the Cointegration Approach. *Russian and East European Finance and Trade* 34 (3): 60–69.