PURCHASING POWER PARITY IN EASTERN EUROPEAN COUNTRIES: FURTHER EVIDENCE FROM BLACK MARKET EXCHANGE RATES

Alper Aslan[1*] and Ferit Kula[2]
[1)] Nevsehir University, Nevsehir, Turkey
[2)] Erciyes University, Kayseri, Turkey

Abstract
The purchasing power parity (PPP) has been amongst the most tested theories in the international finance literature. The empirical findings from the extant literature for the PPP hypothesis are mixed. This article applies univariate and panel Lagrange Multiplier (LM) unit root tests with one and two structural breaks to real exchange rates for six Eastern European countries. Both univariate and panel LM tests with structural breaks strongly suggest that PPP is valid hypothesis for Bulgaria, Czech Republic, Hungary, Poland, Romania and Russia.

Keywords: PPP, Black Market Exchange Rates, Eastern European countries

JEL Classification: F31

Introduction

The analysis of Purchasing Power Parity (PPP hereafter) has probably been one of the most controversial topics of the last decades within internationa economics. Many studies examined the stationarity of real exchange rates based on data from the post Bretton-Woods Period. However, findings from tests regarding to the stationarity of real exchange rates are mixed for Eastern European countries. Although Mahdavi and Zhou (1994), Sideris (2006), Cuestas (2009) found strong evidence on the empirical fulfillment of PPP, Thacker (1995), Christev and Noorbakhsh (2000), Beirne (2007) and Acaavci and Ozturk (2010) found weak support for PPP.

Mahdavi and Zhou (1994) analyzed the empirical validity of PPP as a long-run equilibrium relationship in a sample of thirteen "high-inflation" countries by using co integration and error-correction model. They found empirical evidence in favor absolute or relative versions of PPP in former Yugoslavia and seven other non- European countries. Choudhry (1999) investigates PPP between United States and Poland, Romania, Russia and Slovenia by using fractional and Harris-Inder cointegration test methods. They provide evidence of (relative) PPP using only the Russian and the Slovenian data, and very little evidence of strict (absolute) PPP is also found.

[*] Corresponding author, **Alper Aslan -** alperleo@gmail.com

Sideris (2006) investigated the validity of long-run PPP for seventeen European economies in transition. The analysis is performed for each economy vis-à-vis the US, using the Johansen cointegration methodology and then for the whole set of countries using the Larsson *et al* panel cointegration technique. The analysis provides support for long-run equilibria, but the coefficients of the estimated cointegrating vectors violate the symmetry and proportionality hypotheses suggested by PPP. Cuestas (2009) analyzed the PPP hypothesis in Central and Eastern European Countries (Bulgaria, Croatia, the Czech Republic, Hungary, Macedonia, Poland, Romania and the Slovak Republic). They found stronger evidence on the empirical fulfillment of PPP in these countries, once account has been taken of these sources of nonlinearities.

On the other hand, Thacker (1995) used Philips-Perron cointegration and unit root tests to examine the validity of PPP for Poland and Hungary for the period 1981:1-1993:2. In this study, PPP is re-examined vis-à-vis three countries, the United States, United Kingdom and Germany. They found that the real exchange rates follow a random walk.

In a dynamic method, Christev and Noorbakhsh (2000) investigate the validity of PPP for Bulgaria, Czech Republic, Hungary, Poland, Romania and Slovak Republic for the period of 1990:1- 1998:11. They found weak support for PPP. In a different method, Barlow (2003) examined PPP for two advanced transition economies (Poland and the Czech Republic) and one lagging transition economy (Romania) for the period of 1994:4 to 2000:12. PPP is not rejected between the lagging reformer and developed market economies, but is rejected between the advanced reformers and the developed economies. In the case of the Croatian economy for PPP, Payne et al., (2005) investigated whether purchasing power parity holds or not by using the minimum LM unit root tests. They found that purchasing power parity is not supported in the case of the Croatian economy. By panel cointegration tests, Beirne (2007) examined PPP for Bulgaria, Czech Republic, Hungary, Poland, Romania, Slovak Republic and Slovenia and found that weak evidence in favour of PPP.

While a great deal of literature has emerged to testing PPP hypothesis in transition economies, the empirical results have been mixed. These unexpected results may be due to employ short period for testing PPP. Theoretically, the PPP theory is a long run model in determining the equilibrium exchange rate. Hence, the relationship might deviate in the short-run (Taylor, 2009).

This information encourages us to employ a long span of data. Therefore, black market exchange rates (BMR) which have more periods compared to official rates for Eastern European countries are used to analyze the validity of PPP.

In addition, since foreign exchange controls continue to be viewed as a viable policy instrument by policymakers in many developing countries, black market exchange rates in those countries continue to be very active. In transition countries where official market exchange rates have been distorted by controls, black-market rates constitute a suitable proxy for the free market Exchange rate. Thus, black-market exchange rates can be used to test exchange rate determination hypotheses such as the PPP (Baliamoune-Lutz, 2009).

The main objective of this study is to examine the validity of PPP from BMR to find out whether PPP holds or not in Eastern European countries such as Bulgaria, Czech Republic, Hungary, Poland, Romania and Russia.

The plan of the paper will be as follows: In the next section the analytical framework of PPP hypothesis will be given and in section 3, the data and empirical results will be described. The final section will contain the conclusion.

1. Analytical Framework

According to the strong form of PPP, the nominal exchange rate is proportional to the relative price so that the real exchange rate remains constant overtime.

$$ER_R = NE_r \frac{P^*}{P} \tag{1}$$

where ER_R is the real exchange rate, NE_r is the nominal exchange rate and P* and P are the foreign and domestic prices, respectively.

In logarithmic form, the real exchange rate can be represented by

$$\log(ER_R) = \log(NE_r) + \log(P^*) - \log(P) \tag{2}$$

Following equation shows the model of mean reverting real exchange rate

$$\log(ER_R)_t = \sigma + \phi \log(ER_R)_{t-1} + \varepsilon_t \tag{3}$$

where σ and ε are constant and error term, respectively. PPP suggest that real exchange rate series should be stationary. If real exchange rate is stationary, any percentage changes in the price level between two countries would be offset by an equal depreciation/appreciation of the nominal exchange rate. If there is a unit-root in the real exchange rate this implies that shocks to the real exchange rate are permanent and PPP does not exist between two countries.

2. Data and Empirical Results

The black market exchange rates data are taken from the study of Reinhart and Rogoff (2004). Price levels are defined as the logarithm of the price ratio generated by the each country's consumer price index (CPI) divided by the US CPI (IFS line 64) and taken from the International Monetary Fund's International Financial Statistics (IMF-IFS) database. Due to the lack of consistent data on the CPI index for Eastern European countries before 1969 M1 and unavailability of data beyond 1998 M12 for black market, the data spans from 1969M1-1998 M12. (See descriptive statistics in Table 1).

Table no. 1: Black Market Exchange Rates Descriptive Statistics

Countries	Mean	Std Dev.
Bulgaria	4.684	1.876
Czech Republic	3.449	0.222
Hungary	4.100	0.557
Poland	2.347	2.248
Romania	4.823	1.806
Russia	3.790	2.71

Since incorporating non-stationary or unit root variables in estimating the regression equations using OLS method give deceptive conclusions, the Augmented Dickey-Fuller (1979) test was widely used to test for stationarity. On the other hand, ADF type models are that they do not allow researchers to analyze the impact of structural changes in the economy. These structural changes, which could be by reason of shocks, have influence on macroeconomic variables. Perron (1989) also illustrated that failure to allow for an existing break leads to a bias that reduces the ability to reject a false unit root null hypothesis. This bias likely leads to spurious conclusions.

To cope with this problem, Perron (1989) proposed allowing for one known, or 'exogenous,' structural break in the augmented Dickey–Fuller (ADF, hereafter) unit root test. In particular, Perron (1989) identified three trend break models. These models are crash model a linear trend with an intercept shift; changing growth linear trend model with a change in slope of the linear trend and the two segments joined at the break date and finally a combined model with intercept and slope change. However, in these models, the break date was identified *ex ante* by economic information. It means that the Perron's method of assuming the break date as exogenously determined.

Following this studies in unit root literature, many authors including, Zivot and Andrews (1992) (hereafter ZA) and Perron (1997) suggested determining the break point 'endogenously' from the data. Lumsdaine and Papel (1997) modified the ZA model to accommodate two structural breaks.

Nevertheless, all these endogenous tests were criticized for their treatment of breaks under the null hypothesis. Given the breaks were absent under the null hypothesis of unit root there may be tendency for these tests to suggest evidence of stationarity with breaks (Lee and Strazicich, 2003). Lee and Strazicich (2003) propose a two break minimum Lagrange Multiplier (LM) unit root test in which the alternative hypothesis unambiguously implies the series is trend stationary (John et al, 2007). To avoid problems of bias and spurious rejections, the endogenous break LM unit root test derived in Lee and Strazicich (2003) is employed in PPP testing. In contrast to the ADF-type tests, size properties of the break LM test are unaffected by breaks under the null.

The break minimum LM unit root can be described as follows. According to the LM principle, a unit root test statistic can be obtained from the following regression:

$$\Delta r_{it} = \delta' \Delta Z_t + \phi \bar{S}_{t-1} + \mu_t \qquad (4)$$

Here, Δ is the first difference operator; $\bar{S}_t = r_t - \hat{\Psi}_x - Z_t \hat{\delta}_t$ t= 2,…..T; $\hat{\delta}$ are coefficients in the regression of Δr_t on ΔZ_t; $\hat{\Psi}_x$ is given by $r_t - Z_t \delta$.

If real exchange rate has a unit root for country *i* then $\phi t = 0$, which is the null hypothesis tested using the *t*-test against the alternative hypothesis that $\phi t < 0$. The panel LM test statistic is obtained by averaging the optimal univariate LM unit root t-test statistic estimated for each country. This is denoted as LM_i^{τ}

$$LM_{barNT} = \frac{1}{N}\sum_{i=1}^{N} LM_i^{\tau} \quad (5)$$

Im et al. (2005) constructed a standardized panel LM unit root test statistic by letting $E(L_T)$ and $V(L_T)$ denote the expected value and variance of LM_i^{τ} respectively under the null hypothesis. Im *et al.* (2005) then compute the following expression:

$$\Psi_{LM} = \frac{\sqrt{N}\,[LM_{barNT} - E(L_T)]}{\sqrt{V(L_T)}} \quad (6)$$

The numerical values for $E(L_T)$ and $V(L_T)$ are in Im *et al*. (2005). The asymptotic distribution is unaffected by the presence of structural breaks and is standard normal. We begin our empirical analysis by examining the univariate LM test without any structural breaks. These results are reported in table no. 2.

Two of the six reel black exchange rate series reject the unit root null at the 10% significance level. For Czech Republic and Poland, the null hypothesis of a unit root is rejected at the 5 percent level of significance.

However, for the remaining four countries, the null hypothesis could not be rejected, suggesting a lack of PPP hypothesis. In addition to individual LM statistics, we explore the panel version of the LM test to the group of six countries in our sample. The panel LM statistic obtained is -9.520, which is smaller than the critical value (-2.326) at the 1 percent level of significance.

Table no. 2: Univariate LM unit root test with no break

Countries	Univariate LM test statistic	k
Bulgaria	-0.040	8
Czech Republic	-3.465**	4
Hungary	-2.035	1
Poland	-3.125**	0
Romania	0.008	6
Russia	-2.519	5
Panel LM test statistic	-9.520***	

Note: The 1, 5 and 10% critical values for the LM test without a break are −3.63, −3.06, −2.77, respectively. The corresponding critical values for the panel LM test are −2.326, −1.645 and −1.282.

The failure to find stationarity may be due to the fact that univariate unit root tests have low power when structural breaks are ignored.

In order to avoid that bias exists against rejecting the unit root null when a particular time series is stationary around a structural break, the stationarity properties of the data by allowing for one structural break is examined in the data series.

These results are reported in table no. 3. Five of the six reel black exchange rate series reject the unit root null at the 10% significance level. For only Hungary, the null hypothesis of a unit root is not rejected at the 10 percent level of significance. The null hypothesis

could be rejected, suggesting a strong validity of PPP hypothesis for the remaining five countries. In addition to one break test, the LM univariate and panel tests are experimented further by conducting with two structural breaks to avoid the lack of ability to reject the unit root null hypothesis may be due to the failure to allow for more than one structural break.

Table no. 3: Univariate LM unit root test with one structural break

Countries	Univariate LM test statistic	k	Break date
Bulgaria	-4.095**	5	1990
Czech Republic	-3.702**	7	1990
Hungary	-2.766	4	1979
Poland	-3.367*	0	1986
Romania	-4.446***	0	1983
Russia	-7.524***	4	1990
Panel LM test statistic	-18.628***		

*Note: Critical values for the LM test statistic from Lee and Strazicich (2004) at the 10%, 5% and 1% significance levels are -3.211, -3.566, -4.239. *,**and *** denote statistical significance at the 10%, 5% and 1% levels respectively. And the 1%, 5% and 10% critical values for the panel LM unit root tests with structural breaks are -2.326, -1.645 and -1.282 respectively.*

In comparison with the results reported in Table 3, we find evidence of PPP for Hungary and more evidence for Poland in table no. 4.

Table no. 4: Univariate LM unit root test with two structural breaks

Countries	Univariate LM test statistic	k	Break date
Bulgaria	-4.338**	4	1992-1993
Czech Republic	-4.375**	7	1980-1990
Hungary	-3.808**	8	1979-1981
Poland	-4.371**	8	1989-1996
Romania	-4.606***	0	1988-1996
Russia	-7.967***	4	1979-1990
Panel LM test statistic	-4.378***		

Note: The 1, 5 and 10% critical values for the panel LM test with a break are −2.326, −1.645 and −1.282, respectively. The 1, 5 and 10% critical values for the minimum LM test with two breaks are −4.545, −3.842 and −3.504, respectively. (, **, ***) denote statistical significance at the 10, 5 and 1% levels, respectively.*

When the panel version of the LM test is explored for the group of six countries in sample with one and two breaks, the panel LM statistics found as -18.628 and -4.378, which are smaller than the critical value (-2.326) at the 1 percent level of significance. Taken

together, both univariate and panel tests with structural breaks strongly suggest that PPP is valid hypothesis for Bulgaria, Czech Republic, Hungary, Poland, Romania and Russia.

This study represents more evidence in support of PPP than studies such as Thacker (1995), Christev and Noorbakhsh (2000) and Beirne (2007) which found weak support for PPP hypothesis in transition countries.

The outcome of the break dates has some important tools for PPP hypothesis. In 1982, Hungary was admitted to the International Monetary Fund and received assistance from the World Bank. For Bulgaria, Communist domination ended in 1990. On the other hand, in 1991 the breakup of the Soviet Union transformed the Russian Federation into a single subject in international relations. In generally, the transition can be said to have begun in 1989 with Poland inaugurating its big bang stabilization and reform program on January 1, 1990.

Conclusion

While a great deal of literature has emerged to testing PPP hypothesis, the empirical results have been mixed. In this paper, the issue of PPP is revisited for Eastern European countries such as Bulgaria, Czech Republic, Hungary, Poland, Romania and Russia. Although there is a growing literature that tests evidence for PPP for Eastern European countries, there is an absence of (a) PPP test for black market in these group countries, and (b) an application of the recently developed panel LM unit root tests with structural breaks. The use of LM unit root tests ensured a comprehensive treatment of PPP in Eastern European countries with black market exchange rates which represents market forces in emerging economies. Both univariate and panel tests with structural breaks strongly suggest that PPP is valid hypothesis for Bulgaria, Czech Republic, Hungary, Poland, Romania and Russia. In addition to these findings, the transition can be said to have begun in 1989 with Poland inaugurating its big bang stabilization and reform program in 1990 which is the most often identified time period of breaks by examining break dates.

References

1. Acaravci, A. and Ozturk, I., 2010. Testing Purchasing Power Parity in Transition Countries: Evidence from Structural Breaks. *Amfiteatru Economic*, XII(27), pp. 190-198.
2. Baliamoune-Lutz, M., 2009. Black and official exchange rates in Morocco: an analysis of their long-run behaviour and short-run dynamics (1974-1992). *Applied Economics*, 42(27), pp.3481-3490.
3. Barlow, D., 2003. Purchasing power parity in three transition economies. *Economics of Planning*, 36(3), pp. 201- 221.
4. Beirne, J., 2007. Purchasing power parity in the European Union: A cointegrated approach and transition economies. In: European Economics and Finance Society, *6th Annual Meeting of the EEFS International Conference*. Sofia, Bulgaria, 31 May-3 June 2007.
5. Choudhry, T., 1999. Purchasing power parity in high-inflation Eastern European countries: evidence from fractional and Harris–Inder cointegration tests. *Journal of Macroeconomics,* 21(2), pp. 293–308.

6. Christev, A. and Noorbakhsh, A., 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), pp. 87- 108.
7. Cuestas, J.C., 2009. Purchasing power parity in Central and Eastern European countries: an analysis of unit roots and nonlinearities. *Applied Economics Letters*, 16(1), pp. 87-94.
8. Dickey, D.A. and Fuller, W.A., 1979. Distributions of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association,* 74(366), pp. 427–431.
9. Eliot, G., Rothenberg, T. Stock, J., 1996. Efficient Tests for an Autoregressive Unit Root. *Econometrica,* 64(4), pp. 813-836.
10. Im, K.S., Lee, J. and Tieslau, M., 2005. Panel LM unit root tests with level shifts. *Oxford Bulletin of Economics and Statistics,* 67(3), pp. 313–419.
11. John G., Nelson, P. and Reetu, V., 2007. Unit Root Tests and Structural Breaks: A Survey with Applications. *Journal of Quantitative Methods for Economics and Business Administration,* 3(1), pp. 63-79.
12. Lee, J. and Strazicich, M., 2003. Minimum LM unit root test with two structural breaks. *Review of Economics and Statistics,* 85(4), pp. 1082–1089.
13. Lee, J. and Strazicich, M.C., 2004. Minimum LM Unit Root Test with One Structural Break. *Working Paper*, Department of Economics, Appalachain State University.
14. Lumsdaine, R.L. and Papell, D.H., 1997. Multiple trend breaks and the unit root Hypothesis. *Review of Economics and Statistics*, 79(2), pp. 212- 218.
15. Mahdavi, S. and Zhou, S., 1994. Purchasing power parity in high inflation countries: further evidence. *Journal of Macroeconomics,* 16(3), pp. 403- 422.
16. Payne, J., Lee, J. and Hofler, R., 2005. Purchasing power parity: Evidence from a transition economy. *Journal of Policy Modeling,* 27(6), pp. 665-672.
17. Perron, P., 1989. The Great Crash, the Oil Price Shock and the Unit Root Hypothesis. *Econometrica*, 57(6), pp. 1361- 1401.
18. Perron, P., 1997. Further evidence on breaking trend functions in macroeconomic variables. *Journal of Econometrics*, 80(2), pp. 355- 385.
19. Reinhart, C.M. and Rogoff, K.S., 2004. The modern history of exchange rate arrangements: a reinterpretation. *Quarterly Journal of Economics,* 119(1), pp. 1–48
20. Sideris, D., 2006. Purchasing power parity in economies in transition: evidence from Central and East European countries. *Applied Financial Economics,* 16(1-2), pp. 135–143.
21. Taylor, M.P., 2009. Editorial: Long-run purchasing power parity and real exchange rates: introduction and overview. *Applied Economics Letters*, 16(1), pp. 1–4.
22. Thacker, N., 1995. Does PPP hold in the transition economies? The case of Poland and Hungary. *Applied Economics*, 27(6), pp. 477- 481.
23. Zivot, E. and Andrews, D., 1992. Further evidence on the great crash, the oil price shock, and the unit root hypothesis. *Journal of Business and Economic Statistics,* 10(3), pp. 251- 270.