A possible solution for the purchasing power parity puzzle: panel cointegration

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PPP (purchasing power parity) explaining the long-run behaviour of nominal exchange rates is one of the most fundamental theories in international economics. However, its empirical validity is controversial and thus, referred to as the PPP puzzle in the literature (*Rogoff* [1996]). While many possible improvements of the baseline model have been suggested, the importance of the appropriate empirical methodology in testing PPP has been given less attention. Since PPP describes a long-run equilibrium relationship, the proper testing method involves testing for cointegration.

As the efficiency of panel estimations is greater because of the higher number of observations, this study investigates the empirical validity of PPP with three cointegrated panel estimation methods, with the DFE (dynamic fixed-effects), the MG (mean-group), and the PMG (pooled mean-group) estimations. PPP is confirmed in all three samples using the DFE estimator. The results most accurately matching the theoretical expectations are found in the 1985–2011 sample.

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PPP (purchasing power parity) explaining the long-run behaviour of nominal exchange rates is one of the most fundamental theories in international economics. However, its empirical validity is controversial and thus, referred to as the PPP puzzle in the literature (*Rogoff* [1996]). Although there is a wide range of literature on PPP, the results reported in these papers vary as sometimes they support the theory and sometimes they reject the empirical validity of the model. Among these *Krugman* [1978], *Frenkel* [1981], and *Harris–Leybourne–McCabe* [2005] failed to find evidence of PPP. In different intervals, different methods can be observed in the testing procedures: linear regressions for the nominal exchange rate and the domestic and foreign price levels (*Frenkel* [1978]); applying unit root tests for real exchange rates (*Hakkio* [1984]); the use of cointegration tests to investigate the existence of the long-run equilibrium relationship between the nominal exchange rate and price levels (*Mark* [1990]); and also the testing of panel datasets (*Narayan* [2008]). Several Hungarian studies deal with the topic as well including *Erdey–Földvári* [2009], *Győrffy* [2009], and *Sebestyén* [1998]. Although PPP is a classical model developed in the 1920s, testing its empirical validity today still remains a hot topic (*Lothian* [2016], *Bahmani-Oskooee–Chang–Lee* [2016], *Huang–Yang* [2015], *Robertson–Kumar–Dutowsky* [2014], *He–Ranjbar–Chang* [2013], *Wu–Lee–Wang* [2011], *Chang et al.* [2011]). *Bahmani-Oskooee–Chang–Lee* [2016] investigated 11 developing countries between January 1994 and March 2003 and presented strong evidence for PPP. *Lothian* [2016] analysed PPP in three historical periods using panel datasets and reached favourable results. *Wu–Lee–Wang* [2011] and *Chang et al.* [2011] also succeeded in finding evidence for PPP but *Huang–Yang* [2015] did not find strong support for its empirical validity in the case of 11 Eurozone countries. *He–Ranjbar–Chang* [2013] achieved positive results only in half of the examined countries: Hungary, Czech Republic, and Russia. Clearly, the more recent results are also heterogeneous.

On the one hand, various ideas have emerged for improving PPP as well as for explaining the empirical failure of the model. These include: applying inappropriate price indices; incorporating the Balassa–Samuelson effect and other variables into the model (for example, government spending on public services [*Froot–Rogoff* [1991], *Ricci–Milesi-Ferretti–Lee* [2013]] and the current account position [*Bayoumi et al.* [1994]]); examining the fulfilment of the assumptions; and investigating the integrity of the international goods markets. On the other hand, as PPP is a long-run equilibrium model, applying the adequate methodology is not a negligible question.

Until the end of the 1980s, the investigation of the long-run equilibrium relationship between non-stationary variables caused problems. Later, *Engle–Granger* [1987] presented a breakthrough when they introduced the definition of cointegration.[1] Since then, the long-run relationship between non-stationary variables can be examined if cointegration exists between them. As applying inadequate techniques can lead to spurious regression, concentrating on the choice of adequate methodological processes is important to solving the PPP puzzle. Since the power of the time series unit root tests and the cointegration tests is low, in the 1990s panel data testing began to spread in investigating of PPP because the accuracy of the tests and the estimations could be enhanced by increasing the number of observations. Thus, in this study the empirical validity of PPP is investigated by cointegrated panel estimation methods, assuming that applying the adequate methodological process (cointegration methods) is one approach that can solve the PPP puzzle.

1. PPP and the PPP puzzle

PPP was the first model to explain the behaviour of nominal exchange rates. The theory was developed at the end of 1920s by *Gustav Cassel* [1921], [1922], [1928]. Although the classics such as *John Stuart Mill*, *Viscount Goschen*, *Alfred Marshall*, and *Ludwig von Mises* also dealt with the topic Cassel was the first to revise the PPP model in an empirically applicable form (*Rogoff* [1996]).

The fulfilment of the LOP (law of one price) is one important assumption of PPP. The LOP states that the same product will have the same price in different countries if its price is expressed in terms of a common currency. If a difference is realised between the prices of the same product in different countries, an arbitrage opportunity occurs in the market. As this will be exploited immediately by market participants, the prices will, thereby, equalise. Consequently, the goods arbitrage in the market forces product prices to equalise internationally. This concept is the absolute LOP model:

$$P_i = EP_i^* , \qquad /1/$$

where P_i is the price of the i-th product in the domestic country in terms of the domestic currency, P_i^* is the price of the i-th product in the foreign country in terms of the foreign currency, and E is the nominal exchange rate (the price of the foreign currency in terms of the domestic currency). The relative LOP model means a weaker assumption:

$$\frac{P_{it+1}}{P_{it}} = \frac{E_{it+1}}{E_{it}} \cdot \frac{P_{it+1}^*}{P_{it}^*}$$

[1] Information about cointegration in Hungarian can be found in *Kovács* [1989], *Kőrösi–Mátyás–Székely* [1990], and *Darvas* [2004].

or in another form:

$$\frac{E_{it+1}}{E_{it}} = \frac{P_{it+1}}{P_{it}} \cdot \frac{P_{it}^{*}}{P_{it+1}^{*}}, \qquad /2/$$

where t denotes the time.

According to equation /2/, the nominal exchange rate changes proportionately with the change of the relative price of the i-th product. It is apparent that goods arbitrage plays a central role in the PPP model. However, the fulfilment of the LOP requires more assumptions such as: transaction costs do not exist in the market (for example, transportation costs, taxes, customs, fees, and other non-tariff barriers); and the traded goods and services are homogeneous, that is, quality differences cannot be discovered between them. If these assumptions fulfil a broad basket of goods and services, then the price levels in the given countries in terms of a common currency are equal or otherwise the nominal exchange rate can be determined as the ratio of the price levels of the two countries (*Rogoff* [1996]). The price levels are reached through the aggregation of a broad basket of goods and services:

$$\sum_{i=1}^{N} \alpha_i P_i = E \sum_{i=1}^{N} \alpha_i P_i^{*}, \qquad /3/$$

where $P = \sum_{i=1}^{N} \alpha_i P_i$ is the domestic price level, $P^{*} = \sum_{i=1}^{N} \alpha_i P_i^{*}$ is the foreign price level, and the sum of the weights is: $\sum_{i=1}^{N} \alpha_i = 1$ (*Marsh–Passari–Sarno* [2012]). Originally, *Cassel* [1921] took into consideration the observation that the exchange rate is the relative price of two currencies; in equilibrium, the relative values of currencies should reflect their relative purchasing power. The purchasing power of the domestic currency is $\frac{1}{P}$ and the purchasing power of the foreign currency is $\frac{1}{P^{*}}$, thus, the exchange rate should be $E = \frac{P}{P^{*}}$ (*Mark* [2001]). This relationship is referred to as the absolute PPP in the literature. Considering the logarithms of the variables, the absolute PPP can be written as follows:

$$e = p - p^{*}, \qquad /4/$$

where the lowercase letters denote the logarithms of the variables. Since the conditions of absolute PPP are strict, which are rarely met in the real economy, and various problems emerged during the testing process (such as international variations of the price indices or lack of data), the relative PPP has gained attention through empirical testing. The relative PPP explains the change in the nominal exchange rate with the difference of the change in the corresponding price levels (that is, with the difference of corresponding inflations). The relative PPP requires only that the growth rate of the exchange rate offset the difference between the growth rates of the domestic and foreign price indices (*Rogoff* [1996]). The relative PPP can be written as follows:

$$\Delta e = \Delta p - \Delta p^{*}, \qquad /5/$$

where the Δ-s denote the changes (for example, $\Delta e_t = e_t - e_{t-1}$).

Initially, PPP was treated as a short-run analytical tool but later it was used for long-run real and nominal exchange rate examination. Some studies have investigated the co-movement of the nominal exchange rate and price levels; others have examined the reversion of the real exchange rate to a long-run equilibrium level. In fact, in such cases, the departure from the PPP is tested as:

$$q = e - p + p^{*}, \qquad /6/$$

where q denotes the logarithm of the real exchange rate. Thus, when there is no departure from PPP, $q = 0$. Therefore, the stationarity of the real exchange rate should be examined (*Chinn* [2012]). In this study, PPP will be tested using the nominal exchange rate.

Although the empirical testing of PPP is extensive in the literature, in several cases, the analyses do not support its empirical validity. The empirical failure of the testing, namely, the extreme volatility of the exchange rate in the short run and the slow reversion of the exchange rate to the PPP level, is referred to as the PPP puzzle in the literature thanks to *Rogoff* [1996].

2. Methodology: testing of PPP with cointegrated panel estimation

Here, we test whether the conjectures of the PPP can be justified, that is, whether the PPP puzzle can be solved when the adequate methodology is chosen. Since PPP describes a long-run equilibrium relationship between the exchange rate and corresponding price levels, cointegration methods are required to capture the long-run effects. A panel dataset was used here because the time series unit root tests and the time series cointegration tests have less power than similar panel tests (*Shiller−Perron* [1985], *Otero−Smith* [2000]); thus, this approach enhances the accuracy of the estimations. After presenting the empirical model for the testing strategy, the applied methods will be introduced.

2.1. The model

According to PPP, the domestic and foreign price levels determine the nominal exchange rate in the long run; so the tested empirical model is as follows:

$$e_{it} = \beta_1 p_{it} + \beta_2 p_t^* + u_{it}, \qquad /7/$$

where p_{it} is the logarithm of the price level of the i-th country at time t, p_t^* is the logarithm of the US (United States) price level at time t, and u_{it} is a white noise process. The strong version of PPP requires that $\beta_1 = 1$ and $\beta_2 = -1$, that is, it presumes unit price elasticity in the case of the domestic and foreign price indices, which means a kind of symmetry. During our testing process to justify PPP, we settle for whether the signs of the estimated coefficients correspond to the theory and their sizes converge to the expected extent. The estimated panel error correction model for PPP is:

$$\Delta e_{it} = \gamma(e_{it-1} - \beta_1 p_{it-1} - \beta_2 p_{t-1}^*) + \alpha_1 \Delta p_{it} + \alpha_2 \Delta p_t^* + u_{it}, \qquad /8/$$

where γ is the adjustment parameter (the error correction coefficient), the β-s denote the long-run effects of the price levels on the nominal exchange rate, and the α-s denote the short-run effects.

2.2. Testing strategy

The long-run equilibrium relationship between the nominal exchange rate and the monetary macro-fundamentals can be captured by cointegration methods. The definition of cointegration[2] was introduced by *Engle–Granger* [1987]. Two non-stationary processes are cointegrated – that is, long run equilibrium relationship exists between

[2] The definition of integration and cointegration had already been identified by *Granger* [1981].

the variables – if there is a linear combination of them that is stationary[3]. However, cointegration can only be exist among non-stationary processes; therefore, we must investigate the order of the integration of our variables.

Since the unit root tests are, in general, very sensitive, we applied more tests to check the robustness of our results such as the Fisher-ADF (augmented Dickey−Fuller), the Fisher-PP (Phillips–Perron), and the Hadri tests (*Maddala–Wu* [1999], *Hadri* [2000]). The Hadri test is the only one that has the null hypothesis of stationarity (the alternative hypothesis is that some units in the panel contain a unit root); the other two tests are panel unit root tests. The Fisher type tests combine the *p*-value of the unit root tests of the *i*-th cross-section unit. Their null hypothesis assumes unit roots in the panel data (*Maddala–Wu* [1999], *Baltagi* [2008]). In the case of the Fisher-ADF test, the number of the lags in the auxiliary regression is determined automatically by the Schwarz information criteria; in the case of the Fisher-PP and Hadri tests, the Bartlett kernel is applied to correct possible autocorrelation. In addition, further modelling feasibilities were investigated to see the robustness of the results: *1*. the time series includes an individual intercept; and *2*. the time series includes an individual intercept and trend. The selection of the tests was influenced by their assumptions. There are tests that presume there is an identical autoregressive structure at each cross-section unit, but this assumption is far from reality. The selected tests have no such assumptions; they permit a different autoregressive structure of the pooled time series, except for the Hadri test (but this is the only panel stationarity test supported by software packages).[4]

Since we apply cointegrated panel estimation methods, we presume the existence of a long-run equilibrium relationship between the examined variables. This can be investigated by panel cointegration tests; however, if we do not receive significant negative adjustment parameters during the estimations, this indicates the absence of cointegration. Thus, we do not necessarily need to test for cointegration: on the one hand, we may obtain contradictory results in the estimations; on the other hand, the study is not aiming to test PPP in a weak conception.

Three cointegration methods were engaged for all the panels to test PPP: DFE (dynamic fixed-effects), MG (*Pesaran–Smith* [1995]), and PMG (*Pesaran et al.* [1999]) estimations. Thus, the robustness of the results can be checked. During the justification of the models we do not consider as an assumption the fulfilment of the

[3] Since the economic time series based on empirical studies are usually integrated of order one ($I(1)$) (*Hendry–Juselius* [2000]), and in the case of cointegration their linear combination should be a lower order of integration, thus, the linear combination of two $I(1)$ processes should be stationary. A more general formulation is as follows: let us have two $I(d)$ ($d > 0$) processes. The two processes are cointegrated if there is such a linear combination, where $I(d-b)$ ($d \geq b > 0$), that is, the order of the integration of the constructed process is lower (*Engle–Granger* [1987], *Darvas* [2004]).

[4] A short introduction to the panel unit root tests is also included in *Szabó* [2014].

proportionality and the symmetry hypotheses, however, we investigate them. If the signs in the estimated cointegrated vector are the same as the signs expected by the theory and the sizes of the coefficients of the variables converge[5] to the expected sizes, then we accept the empirical validity of PPP. The DFE, MG, and PMG methods estimate a panel error correction model and they estimate the error correction parameter. Therefore, we accept the empirical validity of PPP if there is cointegration – the adjustment to the long-run equilibrium is revealed –, the signs of the coefficients of the variables in the cointegrated vector are in accordance with the theory, and the size of the coefficients is near to the expected size. If cointegration exists but the signs are wrong, that is, the exchange rates accommodate a cointegrated vector, which does not reflect the conjectures of PPP, then we establish the empirical failure of PPP. Hence, if we succeed in estimating a proper cointegrated vector but the exchange rates do not adjust to this, we cannot confirm the empirical validity of PPP.

The DFE, MG, and PMG methods are applied primarily for estimating such non-stationary heterogeneous panels where both the time dimension and the number of cross-section units are large. The DFE is a traditional fixed-effects estimator, while the MG and PMG are relatively new methods, developed by *Pesaran–Smith* [1995] and *Pesaran–Shin–Smith* [1999]. The three methods differ in terms of the freedom of the estimated parameters across cross-section units. The DFE allows only the heterogeneity of the intercepts across the cross-section units (for example, $\Delta e_{it} = \alpha_{i0} + \gamma\,(e_{it-1} - \beta_1 p_{it-1} - \beta_2 p^*_{t-1}) + \alpha_1 \Delta p_{it} + \alpha_2 \Delta p^*_t + u_{it}$). The MG estimator fits the average of the parameters of N cross-section units (for example, $\Delta e_{it} = \alpha_0 + \gamma\,(e_{it-1} - \beta_1 p_{it-1} - \beta_2 p^*_{t-1}) + \alpha_1 \Delta p_{it} + \alpha_2 \Delta p^*_t + u_{it}$); thus, in this case, the intercepts, the slope coefficients, and the variances of the residuals can differ as well across groups. The PMG estimator assumes an identical cointegrated vector for all cross-section units – that is, only one common cointegrated vector is estimated – but the adjustment parameter and the short-run effects (the intercepts, the short-run coefficients, and variances of the residuals) can differ across groups (for example, $\Delta e_{it} = \alpha_{i0} + \gamma_i (e_{it-1} - \beta_1 p_{it-1} - \beta_2 p^*_{t-1}) + \alpha_{i1} \Delta p_{it} + \alpha_{i2} \Delta p^*_t + u_{it}$). The PMG estimates the parameters with the maximum likelihood method (*Pesaran–Smith* [1995], *Pesaran–Shin–Smith* [1999]). Unfortunately, the DFE and the PMG estimators do not provide consistent results in all cases. The DFE estimator forces the identity of the slope coefficients on the model but when, in fact, the slope coefficients are heter-

[5] In the case of time series analyses, we could face two-digit coefficients during the estimations, therefore we consider as a positive result that we estimated only one-digit coefficients in all cases during the panel estimations. The smallest coefficient in absolute value – for the examined model to still be considered justified – was 0.798; the largest was 2.097. However, we also obtained coefficients quite close to one in absolute value, thus, even if we evaluated our estimations through stricter criteria, we could find empirical support for the PPP, just in fewer cases. In turn, if both the proportionality and the symmetry hypotheses were required by the evaluation of the results, then we could accept the empirical validity of the PPP in one case only.

ogeneous, the DFE estimator may lead to inconsistent and misleading results. In addition, DFE models are usually subject to simultaneous equation bias from the endogeneity between the residual and the lagged dependent variable. The PMG estimator forces homogeneity of the long-run effects on the data. If this restriction is true, the PMG estimator provides efficient and consistent estimations. However, if the long-run effects are heterogeneous, the PMG estimations will be inconsistent as well. The MG estimations will be consistent even if the coefficients of the true model are heterogeneous or homogeneous. The Hausman test can be applied to determine which model fits the data best (*Blackburne III–Frank* [2007]).

3. Results

In the following sections, we present the data, the unit root tests results for each variable, and the results of the cointegrated panel estimations.

3.1. Data

To collate our panel data, we used the OECD (Organisation for Economic Co-operation and Development) database of statistics of MEI (Main Economic Indicators). Three panels were arranged, which differ in the number of cross-section units and length of time: the long-time span (1973, Q1–2011, Q4) contains less cross-section units; whereas, the shorter time spans (for example, 1996, Q1–2011, Q4) contain more cross-section units. Table 1 presents the panels where we show the names of the countries instead of their currencies.

In the panel containing the most cross-section units, 15 OECD countries' (14 countries and the Eurozone) US dollar exchange rates were analysed using quarterly data. During the sample period, the exchange rate policy of the examined countries was characterised primarily by floating exchange rates. The frequency of the original data was monthly; however, in this study we investigated quarterly data. This was for two reasons: one, we could test one more cross-section unit as the CPI (consumer price index) of Australia was only available as quarterly data; second, several studies with positive test results have used quarterly data. We averaged the data to achieve quarterly frequency. As the PPP was tested, our variables are the nominal exchange rate and the corresponding CPIs. The exchange rates were average period values, so the monthly values were averaged to calculate the quarterly data. The CPI was available for the longest period and for most countries, so we used this variable as is common in most studies. The CPI is seasonally unadjusted because, according to the OECD, seasonal effects are not too significant[6], so we set this aside. The base year of the CPI was 2005 and from the product groups we chose the 'all items' category. We used Stata 13 for the analysis.

Table 1

The panels of OECD countries' US dollar exchange rates

Number of cross section units	Panel		
	1973, Q1 –	1985, Q1 –	1996, Q1 –
	– 2011, Q4		
1.	Canada	Australia	Australia
2.	Norway	Canada	Canada
3.	Sweden	Denmark	Czech Republic
4.	Switzerland	Japan	Denmark
5.		Mexico	Eurozone
6.		Norway	Hungary
7.		Sweden	Japan
8.		Switzerland	Korea
9.		Turkey	Mexico
10.		United Kingdom	Norway
11.			Poland
12.			Sweden
13.			Switzerland
14.			Turkey
15.			United Kingdom
Number of observations	624	1,080	960

Note. The panels show the names of the countries instead of their currencies. Here and in the following tables, OECD: Organisation for Economic Co-operation and Development.

3.2. Results of panel unit root tests

Since cointegration exists only among non-stationary variables[7], it is necessary to test the order of the integration of the variables. The empirical model contained the

[6] *ILO et al.* [2004].

[7] In general, two non-stationary variables are cointegrated if their linear combination is integrated of a lower order (*Darvas* [2004]). Empirical studies indicate that most economic time series are unit root processes (*Hendry–Juselius* [2000]), thus, they will be cointegrated if their linear combination is stationary. Thus, we expect that the examined variables will be $I(1)$ processes when we check the unit root test results.

logarithm of the examined variables and in every case the logarithm of the variables was tested for the integration order. Three tests were applied: Fisher-ADF, Fisher-PP, and the Hadri test (*Maddala–Wu* [1999], *Hadri* [2000]). For the robustness of the results, all model possibilities were tested. As the test results were ambiguous in several cases, the graphs of the time series were taken into consideration as well to determine the order of the integration of the variables.[8] Since the examined price levels for almost all cases showed a trend identified in the graphs, stationarity was excluded. Furthermore, in these cases, we did not run unit root tests on the levels of the logarithms of the price levels. The logarithm of the price level of the US was investigated using time series unit root tests for the different periods. The applied ADF and Ng–Perron unit root tests results and the KPSS (Kwiatkowski−Phillips−Shmidt−Shin) stationarity test results are presented in Table 2 (*Dickey−Fuller* [1979], *Ng−Perron* [2001], *Kwiatkowski et al.* [1992]).

Table 2

The results of the ADF and the Ng−Perron unit root tests and the KPSS stationarity test on the price level of the United States

Variable	ADF test		KPSS test		Ng–Perron test	
	A	B	A	B	A	B
	1973, Q1–2011, Q4					
Δp_t^*	–2.795*	–3.684**	1.054***	0.159**	–4.165	–19.199**
$\Delta^2 p_t^*$	–	–	0.008	0.008	–295.708***	–287.030***
	1985, Q1–2011, Q4					
Δp_t^*	–5.232***	–5.680***	0.458*	0.064	–6.314*	–9.501
$\Delta^2 p_t^*$	–	–	0.059	0.044	–0.340	–198.124***
	1996, Q1–2011, Q4					
Δp_t^*	–8.856***	–8.789***	0.171	0.150**	–91.075***	–89.650***
$\Delta^2 p_t^*$	–	–	0.188	0.146**	–	–

Note. Here and hereinafter, ADF: augmented Dickey–Fuller; KPSS: Kwiatkowski−Phillips−Shmidt−Shin. In the case of the Ng−Perron test, the MZ_α test statistic was taken into account. Here and in Table 3, A: the time series contains a constant; B: the time series contains a constant and a trend. * at 10%, ** at 5%, and *** at 1% significance levels, the null hypothesis can be rejected.

The logarithm of the price level of the US shows an ambiguous picture. However, as a trend stands out in the graph, stationarity is excluded. Although, the ADF test

[8] The graphs of the times series can be provided on request.

showed an unambiguous $I(1)$ process for all periods, the KPSS and the Ng–Perron tests were more uncertain. The KPSS test showed second order integration ($I(2)$) for the longest period, while for the other two periods it showed $I(1)$ or $I(2)$. The Ng–Perron test indicated the same for the periods of 1973, Q1–2011, Q4 and 1985, Q1–2011, Q4 but in the case of 1996, Q1–2011, Q4, it showed an $I(1)$ process. In the first difference of the process, an outlier value was observed, which could cause the uncertainty of the tests. Thus, the process was evaluated as $I(1)$.

The 1973, Q1–2011, Q4 panel included four (the Canadian dollar/, the Norwegian krone/, the Swedish krona/, and the Swiss franc/) US dollar exchange rates. The logarithms of the exchange rates of the examined OECD countries appeared to be $I(1)$ processes according to all three tests. The Fisher-ADF and the Fisher-PP tests showed $I(1)$ for the logarithms of the price levels of the examined OECD countries. Stationarity could be excluded because trends were detected in the time series. However, the Hadri test indicated $I(2)$, which may be caused by the outlier values in the time series and the possible breakpoints (Canada, Norway, and Sweden). Thus, the price levels of the examined OECD countries were evaluated as $I(1)$ as well.

In the 1985, Q1–2011, Q4 panel, six additional exchange rates could be added: the Australian dollar/, the Danish krone/, the British pound sterling/, the Japanese yen/, the Mexican peso/, and the Turkish lira/US dollar exchange rates. Thereby, the number of observations could be enhanced compared to the previous panel. We obtained unfavourable results for the logarithms of the exchange rates of the examined OECD countries. The Fisher tests indicated stationarity and the Hadri test showed uncertainty. However, it seems that most of the exchange rates had trends; thus, they presumably could not be stationary. The noticeable outlier values in the differences of the logarithms of the exchange rates could be the cause of the uncertainty of the Hadri test. Thus, in this sample, we also considered the exchange rates as $I(1)$ processes. The results for the price levels of the examined OECD countries were similar to the results of the previous panel: the Fisher tests indicated $I(1)$ – since stationarity could be excluded on the basis of the apparent trends in the graphs of the logarithms of the time series –, while the Hadri test stood by the $I(2)$ result, similar to the previous sample. Based on the above, the price levels of the examined OECD countries were considered as $I(1)$.

In the sample, beginning from 1996, Q1, the logarithms of the exchange rates of the examined OECD countries were $I(1)$ or $I(0)$ according to the Fisher tests; the Hadri test showed $I(3)$. In turn, it was obvious that in the processes, several outlier values are present; the Hadri test can be sensitive to this. It is unlikely that the processes were stationary because the majority of the exchange rates had a trend, so we supposed that the examined nominal exchange rates were $I(1)$. The logarithms of the price levels of the examined OECD countries were $I(1)$ according to the Fisher tests; however, the Hadri test indicated $I(2)$ or $I(3)$. The stationarity could be excluded on the basis of the observable trends in the time series; therefore, we did not run any test for the levels. In the first differences in these series, we noted not only outlier values but also the possible existence of some breakpoints (Mexico, Poland, and Turkey). This could cause the Hadri test to present somewhat unrealistic results. In this case, we supposed again that the processes were $I(1)$. Table 3 details the results.

Table 3

The results of the Fisher-ADF and the Fisher-PP unit root tests and the Hadri stationarity test for the US dollar panels of the OECD countries

Variable	Fisher-ADF test		Fisher-PP test		Hadri test	
	A	B	A	B	A	B
	1973, Q1–2011, Q4					
e_{it}	11.029	7.886	7.701	5.885	7.268***	5.530***
Δe_{it}	235.982***	216.368***	229.508***	209.749***	0.076	–0.335
Δp_{it}	14.001*	20.663***	185.191***	259.390***	12.809***	4.141***
$\Delta^2 p_{it}$	–	–	–	–	–1.699	–1.426
	1985, Q1–2011, Q4					
e_{it}	83.403***	46.307***	74.012***	41.911***	8.620***	8.653***
Δe_{it}	372.337***	338.429***	406.379***	368.626***	2.193**	4.586***
$\Delta^2 e_{it}$	–	–	–	–	1.398*	10.549***
Δp_{it}	99.196***	130.301***	487.763***	489.352***	9.012***	7.215***
$\Delta^2 p_{it}$	–	–	–	–	–1.938	–1.164
	1996, Q1–2011, Q4					
e_{it}	33.298	47.306**	43.237	30.525***	10.108***	8.716***
Δe_{it}	342.912***	292.416***	310.789***	252.381***	2.264**	3.761***
$\Delta^2 e_{it}$	–	–	–	–	3.480***	17.343***
Δp_{it}	276.651***	246.490***	555.328***	538.776***	6.754***	11.269***
$\Delta^2 p_{it}$	–	–	–	–	1.259	7.529***

Note. PP: Phillips–Perron. In the case of the Fisher-ADF and Fisher-PP tests, the χ^2 statistic was applied; in the case of the Hadri test, the heteroscedastic consistent *Z*-statistic was applied. * at 10%, ** at 5%, and *** at 1% significance levels, the null hypothesis can be rejected.

The order of the integration of the variables was determined by considering the tests results and the graphs of the time series. To investigate the price level of the US, time series unit root tests were applied. However, although the results were somewhat heterogeneous, eventually the US price level was evaluated as an $I(1)$ process. The variables of the examined OECD countries did not show an unambiguous picture in all cases, but the graphs of the time series helped us in the analysis. Based on the graphs, we excluded the stationarity for the times series that had an apparent trend. In the case of the differences of the logarithms of the time series, several outlier values and possible breakpoints also could be observed, which could cause the uncertainty of the tests. Eventually, considering the available information, the examined processes were evaluated as $I(1)$. Thus, cointegration can exist among the variables, so we could investigate whether the assertions of the PPP were realised.

3.3. DFE, MG, and PMG estimation results

Since the power of the pure time series cointegration tests is lower than the power of the panel cointegration tests, increasingly, more studies are applying the panel technique for testing PPP. The panel estimation has been prevalent since the end of the 1990s. *Pedroni*'s 1996 study is an earlier version of his 2001 work; in it, not only were simulations run, but the PPP was estimated by the FM-OLS (fully modified ordinary least squares) method. Monthly and annual IFS (International Financial Statistics) data were also estimated between 1974 and 1993. For the different analyses, the panels included the variables of 20-25 countries. Even though Pedroni was not satisfied with the results because the applied test in most cases rejected that the coefficient of the price difference is one, the estimated coefficients approximated one, which was not an entirely negative result. *Pedroni* [2001] estimated PPP not only using the FM-OLS method but also using the DOLS (dynamic ordinary least squares) method. Similar data were examined as in his earlier [1996] work: monthly IFS data between June 1973 and November 1993 for 20 countries. US dollar exchange rates were investigated, and the price level was captured in the CPI. The results were negatively appreciated again because the null hypothesis – that the estimated slope coefficients were one – could have been rejected, although the coefficients were near to one. *Robertson–Kumar–Dutkowsky* [2014] examined the PPP not solely in a weak concept but also in a cointegrated panel estimated for the Mexican peso/US dollar exchange rate between January 1982 and February 2010. The panel was estimated by the FM-OLS and DOLS methods. However, although the estimated coefficients approximated one, the *t*-test rejected that the coefficients were accurately one, which was why the authors evaluated their results as negative. At the same time, PPP had empirical support in a strong sense for actively traded goods.

During our panel estimations, we reached relatively favourable results. Here, we discuss the results of the DFE, MG, and PMG estimations. These methods estimated a panel-error correction model. If the adjustment parameter in the error correction model was insignificant, then, we could not speak about cointegration, that is, PPP could not be justified. Therefore, beyond the long-run effects, we also report the error correction coefficients. However, the PMG estimation allowed the adjustment speeds to differ across the cross-section units, we report only their averages.

The adjustment coefficients were significantly negative in all cases. The significant adjustment coefficients confirm the existence of cointegration, as they imply that the system eliminated the deviation from the long-run equilibrium level. However, according to our criteria, the presence of correct cointegrating vectors will be necessary as well for the justification of PPP. Considering the whole analysis, we obtained a conflicting sign with the theoretical expectations by a significant variable in only one case: the 1996, Q1–2011, Q4 panel during the MG estimation for the domestic price level. Usually, such favourable results are not typical in the case of time series estimations. As all other signs of the significant variables were in line with expectations, we, thus, do not detail the signs in the presentation of the results.

We have succeeded in empirically justifying PPP in the case of all three samples with at least one cointegrated panel estimation method. Considering the overall analysis, four of the nine cases could be evaluated positively. In these cases, the exchange rate adjusted to an acceptable cointegrating vector, namely, the signs in the cointegrating vector were in line with the theoretical expectations, and the size of the coefficients also converged to the expected size. The most favourable results were in the 1985, Q1–2011, Q4 panel, which was probably due to the high number of observations.

The proportionality and symmetry hypotheses were also examined by each panel. In 44% of the estimated significant coefficients, the proportionality hypothesis could not be rejected; and, in 33% of the estimated significant coefficients, the symmetry of the coefficient of the domestic and foreign variable could not be rejected. Table 4 details the results.

The best (matches to the most criteria) estimation for PPP was obtained in the 1973, Q1–2011, Q4 panel. The cointegrating vector, in the case of the DFE estimation, converged to the theoretical expectations so the Wald test did not reject the proportionality and the symmetry hypotheses, and the exchange rates adjusted to the estimated cointegrating vector as well. The Hausman test showed the DFE estimation was efficient. However, in case of the comparison of the MG and PMG estimations, the assumptions of the Hausman test were not met; thus, it was not run. The variables in the cointegrating vectors of the MG and PMG estimations were not significant; however, the exchange rates adjusted to these vectors but to no purpose. In this sample, we found strong evidence for the empirical validity of PPP in the DFE estimation.

Table 4

MG, PMG, and DFE estimations for PPP for the OECD countries' US dollar panels

Variable/Test	MG	PMG	DFE
	estimation		
		1973, Q1–2011, Q4	
p_{it}	–0.194 (0.120)	0.293 (0.411)	1.052*** (0.304)
p_t^*	–0.163 (0.228)	–0.506 (0.389)	–1.090*** (0.298)
Error correction coefficient	–0.178*** (0.016)	–0.084*** (0.012)	–0.083*** (0.009)
Wald test ($H_0 : \beta_p = 1$)	–	–	0.03
Wald test ($H_0 : \beta_{p*} = -1$)	–	–	0.09
Wald test ($H_0 : \beta_p = -\beta_{p*}$)	–	–	0.11
Hausman test (MG-PMG)		–	
Hausman test (MG-DFE)		0.12	
		1985, Q1–2011, Q4	
p_{it}	0.955 (0.643)	0.872*** (0.038)	0.938*** (0.019)
p_t^*	–0.934 (0.596)	–0.798*** (0.081)	–0.857*** (0.092)
Error correction coefficient	–0.127*** (0.022)	–0.104*** (0.022)	–0.094*** (0.011)
Wald test ($H_0 : \beta_p = 1$)	–	11.41***	10.39***
Wald test ($H_0 : \beta_{p*} = -1$)	–	6.24**	2.40
Wald test ($H_0 : \beta_p = -\beta_{p*}$)	–	1.27	0.89
Hausman test (MG-PMG)		0.07	
Hausman test (MG-DFE)		0.00	
		1996, Q1– 2011, Q4	
p_{it}	–3.896** (1.599)	0.454 (0.281)	0.961*** (0.098)
p_t^*	–1.499 (0.999)	–1.826*** (0.282)	–2.097*** (0.196)
Error correction coefficient	–0.145*** (0.017)	–0.088*** (0.010)	–0.083*** (0.009)
Wald test ($H_0 : \beta_p = 1$)	9.37***	–	0.16
Wald test ($H_0 : \beta_{p*} = -1$)	–	8.58***	31.51***
Wald test ($H_0 : \beta_p = -\beta_{p*}$)	–	–	46.66***
Hausman test (MG-PMG)		6.53**	
Hausman test (MG-DFE)		0.00	

Note. MG: mean-group; PMG: pooled mean-group; DFE: dynamic fixed-effects; PPP: purchasing power parity. Only the χ^2 statistic was reported in the case of the Wald tests. * at 10%, ** at 5%, and *** at 1% significance levels, the null hypothesis can be rejected.

In the 1985, Q1–2011, Q4 panel, PPP was estimated in a shorter period with the addition of six currencies' (the Australian dollar/, the Danish krone/, the British pound sterling/, the Japanese yen/, the Mexican peso/, and the Turkish lira/) US dollar exchange rates; thus, the number of observations increased. In this sample, we received good results. The PMG and DFE estimations also justified the conjectures of PPP. The symmetry between the domestic and foreign price levels could not be rejected by either of the estimations. In the DFE estimation, the proportionality hypothesis seems to be fulfilled in the case of the foreign price level. In the case of the MG estimation, the exchange rates adjusted to the estimated cointegrating vector but none of the price levels was significant; therefore, this estimation did not support the conjectures of PPP. According to the Hausman test, the PMG and DFE estimations were efficient.

In the 1996, Q1–2011, Q4 panel, PPP was successfully justified empirically with the DFE estimation. The foreign price level was not significant in the cointegrating vector in the MG estimation and the sign of the coefficient of the significant domestic price level was contrary to the expectations; however, the exchange rates adjusted to this vector. In the PMG estimation, the exchange rates also showed adjustment, but the domestic price level was not significant in the cointegrating vector. In contrast, in the case of the DFE estimation, both variables were significant, the size of the coefficients did not differ significantly from one; and the exchange rates also adjusted to the estimated cointegrating vector. The proportionality hypothesis could not be rejected by the domestic price level. In addition, the Hausman test showed that the DFE estimation was efficient; further, the MG estimation was preferred over the PMG estimation.

Based on the results, we can conclude that there is empirical support for PPP primarily by the DFE estimation. Furthermore, using this estimation process, we can justify the empirical validity of PPP in all three samples.

4. Conclusions

PPP explaining the long-run behaviour of the nominal exchange rates is one of the most fundamental theories in international economics. However, although its theoretical role is substantial, its empirical validity is controversial, referred to as the PPP puzzle (*Rogoff* [1996]). While many possible improvements in the baseline model have been suggested, the importance of using appropriate empirical methodology in testing PPP has been given less attention. Before the 1980s, the possibilities were limited to investigating the long-run relationships between non-stationary processes based on existing methodology. The breakthrough occurred in the *Engle–Granger* [1987] study where the definition of cointegration was introduced.

Since then, the long-run equilibrium relationships between non-stationary processes could have been examined if they are cointegrated.

Since the PPP describes a long-run equilibrium relationship between the nominal exchange rate and the adequate price levels, the proper testing method involves testing for cointegration. In the lack of proper testing method, we could receive spurious regression results. Thus, it is important to stress that applying a proper methodological process is one approach that could help solve the PPP puzzle. However, in several cases, applying cointegration techniques using time series estimations to find empirical support for PPP has also failed, the precision of the estimations and the power of the tests can be enhanced by increasing the number of the observations. Thus, since the 1990s, the application of panel cointegration techniques has become increasingly dominant in testing PPP.

In the study, we examined the empirical validity of PPP in the case of three panels, due to a lack of data[9], by three cointegrated panel estimation methods (DFE, MG, and PMG). After we verified on the basis of unit root tests that the processes were likely to be unit root processes, we ran the estimates. To justify PPP, we did not consider the fulfilment of the proportionality and symmetry hypotheses as essential assumption, although we investigated them. In the DFE estimation, we found support for PPP in every sample; we obtained the best results (which met most of the theoretical assumptions) in the 1985, Q1–2011, Q4 panel. Presumably, the reason for this was that this sample contained the most observations. In this panel, we confirmed the empirical validity of PPP also with the PMG estimation. In some cases, the Wald test did not reject the proportionality and symmetry hypotheses but only in the case of the 1985, Q1–2011, Q4 panel using the DFE estimation was there the chance to confirm both hypotheses.

Although we did not obtain unambiguous results, for all three samples the empirical validity of the PPP was successfully justified using at least one estimating process, which cannot be considered a negative result. We can conclude that a suitable methodological (cointegration) process is necessary to capture long-run equilibrium relationships. Panel data estimations provide a more accurate picture of the empirical validity of PPP. Thus, applying the cointegrated panel estimation methods may contribute to solving the PPP puzzle.

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[9] If we had had time series with identical length, we would have estimated one large balanced panel.

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