Evidence on the Purchasing Power Parity in Panel of Cities

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Abstract

We provide evidence on the PPP hypothesis using a sample of fifty Spanish cities for a long time period through the application of panel data unit root tests. Although results suggest non-rejection of the PPP, short-run deviations –as measured by half-lives– indicate that real factors might be causing a slow rate of convergence to a common price index, even in highly integrated economies.

Keywords: Convergence in prices, cities, PPP, half-life, panel data unit root tests

JEL codes: F31, F37, F41

1 Introduction

The Purchasing Power Parity (PPP) hypothesis has raised a lot of interest in both theoretical and applied economic analysis. Some theoretical models on international economics have been developed assuming some kind of PPP hypothesis –strong or weak version– although it is in the empirical arena that testing the PPP hypothesis has became very popular. Froot and Rogoff (1995) review the literature and distinguish up to three different stages on the empirical PPP hypothesis testing: (i) old tests in which the PPP holds under the null hypothesis, (ii) unit root tests, in which under the null hypothesis it is assumed a permanent deviation from the PPP and (iii) cointegration tests, in which deviations from the PPP are analysed using a linear combination of prices. Note that the second and third stages involve the analysis of the stochastic properties of time series of prices. Actually, most of the recent evidence to the PPP hypothesis has been obtained applying the fruitful frameworks of the unit root, stationarity and cointegration techniques. Some of the outstanding contributions to this field are Frankel (1986), Meese and Rogoff (1988), Corbae and Ouliaris (1991), Lothian and Taylor (1996), and Montañés and Clemente (1999). The main limitation of these analyses come from the lack of power shown by the unit root and cointegration tests. Some practitioners have advocated for the use of longer time series as a way to solve this shortcoming, but this gives rise to another problem, *i.e.* the fact that long time series encompass periods in which nominal exchange rates regimes shifted from floating to fixed and back again.

The univariate analysis of the PPP has found a natural extension into the panel data framework. The idea behind the use of these tools is that the combination of the cross-section and time series information increases the power of the statistical inference which, in turn, allows practitioners to deal with time series corresponding to an homogeneous exchange rate regime. Some of the relevant papers that address the PPP testing in a panel data framework are Abuaf and Jorion (1990), Oh (1996), Papell (1997, 2002) , Culver and Papell (1999), O'Connell (1997), Papell and Theodoridis (1998) and Fleissig and Strauss (2000), to mention few. All of them point to the analysis of the PPP at an international level –samples of G7, industrialised OECD countries and European countries.

The study of the PPP hypothesis has not only been restricted to international economics. Recently, the PPP literature has experienced the rising of a broad number of empirical works aimed to testing for the PPP at an intranational level. In fact, there is wide consensus in that the PPP hypothesis should be most easily satisfied at an intranational level than when it is analysed at an international level. Among the reasons for that are the higher markets integration, the absence of trade barriers –such as tariffs and quotas– and the absence of exchange rate volatility. Although there are transportation costs that prevents arbitrage, they are presumably smaller within than between countries. Finally, price indices within a country are expected to be more homogeneous than price indices between countries since, first, they are collected by the same statistical institution and, second, the basket of goods is more homogeneous.

Most of the empirical evidence on intranational PPP has focused on the U.S. and Canadian cities – see Parsley and Wei (1996), Jenkins (1997), Culver and Papell (1999), Levin, Lin, and Chu (2002), Cecchetti, Mark, and Sonora (2002), and Chen and Devereux (2003). In contrast to this evidence for the North-American cities, there are less empirical studies that address the PPP hypothesis testing on other geographical areas. Exception are, for instance, Nenna (2001) and Esaka (2003) for twenty one Italian and for the major Japanese cities respectively, concluding in favour of the PPP hypothesis, though noting the high persistence of the deviations. It should be stressed that price convergence across economies has recently become a crucial issue for the implementation of the EU common currency area. In this regard, the aim of this paper is to extend the empirical evidence on the PPP hypothesis to a sample of fifty Spanish cities. The long time period covered by the monthly time series –from 1939:7 to 1992:12– allows us to apply some of the panel data unit root tests that have been recently proposed in the literature. Our results supports the PPP hypothesis among the Spanish cities, though in some cases deviations are highly persistent, even lasting more than ten years.

The outline of the paper is as follows. Section 2 briefly presents the PPP hypothesis and describes the panel data unit root statistics used to test it. Section 3 reports results for the sample of Spanish cities while Section 4 discusses the persistence of shocks affecting city prices. Finally, Section 5 concludes.

2 The PPP hypothesis testing

Let $p_{i,t}$ denote the log of price level of the economy i, $p_{j,t}$ the log of price level of economy j, and $tc_{ij,t}$ the log of nominal exchange rate that relates the currencies of both economies, $i, j = 1, \ldots, N$, $i \neq j$. The PPP hypothesis establishes the equalization of price levels for both economies once they have been expressed in the same currency. Deviations between prices might exist, though they should be temporary if the PPP is satisfied. This implies that the log of the real exchange rate:

$$q_{i,t} = p_{i,t} - p_{j,t} - tc_{ij,t}, \tag{1}$$

is a stationary stochastic process. Therefore, the PPP can be tested assessing the order of integration of $q_{i,t}$, which can be performed using unit root tests. Thus, non-rejection of the unit root is interpreted as evidence against the PPP.

Among the most relevant arguments to explain the evidence against the PPP are the existence of trade barriers –such as tariffs or quotas– and bureaucratic barriers, the cost of collecting information, the different nature of goods –tradeable and non-tradeable–, differences in productivity, firms exercising local monopoly through differential pricing policies, transportation costs, sticky nominal price level adjustment arising from imperfect markets where price changes are costly, and the fact that nominal exchange rates do not adjust to relative-price shocks –see Cecchetti, Mark, and Sonora (2002). All these factors might prevent arbitrage that would fully remove price differentials.

Notice that when working with price levels of economies that share a common currency, the effect of the nominal exchange rate in (1) disappears. Thus, the real exchange rate is given by:

$$q_{i,t} = p_{i,t} - p_{j,t}. \tag{2}$$

Moreover, when they belong to the same country there are not a direct border effect since there are no tariffs or quotas that restrict the arbitrage. Although the literature has suggested some indirect border effects, as for instance those caused by different systems of taxation, the evidence for U.S. cities points to a minimal influence when explaining the deviations from PPP –Parsley and Wei (1996) and Jenkins (1997). We expect this effect to be even more negligible for the Spanish case provided the prevalence of a centralized taxation system. In all, conducting the analysis within a country, as in this paper, increases the probability of finding evidence in favour of the PPP –we get rid of the nominal exchange rate fluctuations and formal border effects–, though it is still possible to find permanent price deviations if factors that hinder the arbitrage exist within a country.

In order to test the PPP hypothesis amongst the sample of Spanish cities, we apply the panel data unit root based tests in Maddala and Wu (1999) –hereafter MW–, Levin, Lin, and Chu (2002) –LLC– and Im, Pesaran, and Shin (2003) –IPS–, provided that they accommodate panel data sets with moderate number of individuals (N) and large number of time periods (T). All these tests specify the null hypothesis of unit root, but Levin, Lin, and Chu (2002) restrict the model under the alternative hypothesis to be a stationary autoregressive process with a common autoregressive parameter for all the individuals, whereas Maddala and Wu (1999) and Im, Pesaran, and Shin (2003) allow for heterogeneity in the autoregressive parameters under the alternative hypothesis. Levin, Lin, and Chu (2002) propose to test the null hypothesis of $H_0 : \delta = 0$ against the alternative hypothesis of $H_1 : \delta < 0$ using:

$$\Delta q_{i,t} = \alpha_{mi} d_{mt} + \delta q_{i,t-1} + \sum_{k=1}^{p} \gamma_k \Delta q_{i,t-k} + \varepsilon_{i,t}, \tag{3}$$

where d_{mt} denotes the deterministic components and $\varepsilon_{i,t}$ is assumed to be independently distributed across i and t, $i = 1, \ldots, N$, $t = 1, \ldots, T$. The normalised bias and the pseudo t-ratio that corresponds with the pooled OLS estimation of δ in (3), once they have been properly normalised, converge to a standard Normal limit distribution as $N \to \infty$, $T \to \infty$ in a way that $\sqrt{N}/T \to 0$.

The test in Im, Pesaran, and Shin (2003) is based on the estimation of (3) but replacing δ with δ_i. The null hypothesis is given by $H_0 : \delta_i = 0$ $\forall i$, whereas the alternative hypothesis is $H_1 : \delta_i < 0$ $i = 1, \ldots, N_1$; $\delta_i = 0$ $i = N_1 + 1, \ldots, N$. Therefore, the null is rejected if there is a subset (N_1) of stationary individuals. The first test that they propose is the standardised group-mean Lagrange Multiplier (LM) bar test statistic:

$$\Psi_{\overline{LM}} = \frac{\sqrt{N}\left[\overline{LM} - N^{-1}\sum_{i=1}^{N} E\left(LM_i\right)\right]}{\sqrt{N^{-1}\sum_{i=1}^{N} Var\left(LM_i\right)}}, \tag{4}$$

with $\overline{LM} = N^{-1}\sum_{i=1}^{N} LM_i$, where LM_i denotes the individual LM test for testing $\delta_i = 0$ in (3), and $E\left(LM_i\right)$ and $Var\left(LM_i\right)$ are obtained by Monte Carlo simulation. The second test is a standardised group-mean t bar test statistic, $\Psi_{\overline{t}}$, with an expression similar to (4) but replacing $\overline{LM}$ and LM_i by $\overline{t}$ and t_i, respectively. We define $\overline{t} = N^{-1}\sum_{i=1}^{N} t_i$, where t_i denotes the individual pseudo t-ratio for testing $\delta_i = 0$ in (3), and $E\left(t_i\right)$ and $Var\left(t_i\right)$ are obtained using Monte Carlo simulation. The authors show that as $N \to \infty$, $T \to \infty$ and $N/T \to k$, the limiting distribution of both test statistics is standard Normal. Finally, Maddala and Wu (1999) propose to combine the individual p-values (π_i) associated to the pseudo t-ratio for testing $\delta_i = 0$ in (3). The test is given by

MW$= -2\sum_{i=1}^{N} \ln(\pi_i)$, which under the null hypothesis is distributed according to MW$\sim \chi^2_{2N}$.[1]

It should be stressed that the limiting distributions of all these tests rely on the assumption of cross-section independence. But this is a strong assumption when testing the PPP hypothesis –see for instance O'Connell (1997). Thus, following Maddala and Wu (1999) here we apply Bootstrap techniques to compute the empirical distribution of the tests –we have performed 1,000 replications for the parametric Bootstrap. This allow us to deal with the presence of cross-section correlation when performing the unit root testing in the panel data framework.

3 Results

The data used in this paper is provided by the Spanish National Institute of Statistics (INE) and refers to the monthly aggregate CPI for fifty Spanish cities from 1939:7 to 1992:12. From 1993:1 onwards there was a change in the methodology used to compute the CPIs that avoids to include more recent observations in the analysis. As price indexes are used, only the weak version of the PPP is tested. The cities considered in the analysis correspond with the capitals of the Spanish provinces (NUTS3 EUROSTAT regional breakdown of the EU), that shared the same currency all over the period. This defines a panel data set with $N = 50$ individuals and $T = 642$ time periods, that is used to test the presence of a unit root in (2).

Here we will supply results for the case in which the national CPI is used as the numerarie. It should be mentioned that the selection of the numeraire can be a controversial question when working at an international level, as different results can be drawn depending on it –see Papell and Theodoridis (2001). Although dependence on the choice of the numeraire is supposed to be mitigated when the analysis is carried out at an intranational level, because of the absence of the nominal exchange rate effect, we tried with alternative numeraires (*i.e.* the cross-section mean and the price in each city in the sample) and the results were substantially the same.[2] Additionally, using the national CPI as the benchmark is in accordance with the common practice of governmental agencies when reporting price differentials within a country.

[1] In order to facilitate computation of π_i we have carried out 100,000 replications to obtain the empirical percentiles for the ADF test for a DGP given by a random walk without drift. Then a response surface has been estimated to approximate the corresponding p-values using the logistic functional form given by

$$\pi_i = \frac{\exp\{x_i\beta\}}{1+\exp\{x_i\beta\}},$$

where $x_i\beta = \beta_0 + \beta_1 x_i + \beta_2 x_i^2 + \beta_3 x_i^3 + \beta_4 x_i^4$, with x_i being the value of the ADF test and π_i the corresponding percentile.

[2] These results are available from the authors upon request. Further discussion on the choice of the numeraire can be found in *e.g.* O'Connell (1997), Culver and Papell (1999) and Cecchetti, Mark, and Sonora (2002).

Table 1: Panel data unit root tests for the Spanish cities

Panel A. Cross-section independence							
$\Psi_{\overline{LM}}$	p-val	$\Psi_{\bar{t}}$	p-val	MW	p-val	LLC	p-val
7.869	0.00	-6.077	0.00	201.705	0.00	0.688	0.75

Panel B: Empirical distributions (Bootstrap)	5%	10%	25%	50%	75%	90%	95%
$\Psi_{\overline{LM}}$	-2.12	-1.82	-1.09	-0.28	0.51	1.42	1.88
$\Psi_{\bar{t}}$	-1.17	-0.73	0.02	0.82	1.63	2.42	2.85
MW	68.02	71.96	81.08	91.26	102.50	113.61	121.75
LLC	-0.39	0.15	0.92	1.81	2.70	3.56	3.99

Table 1 presents results for the LLC, MW and IPS tests including individual fixed effects in the deterministic component of (3), since this is the specification consistent with the PPP hypothesis. Following Ghysels, Lee, and Noh (1994), who analyse the performance of unit root tests applied to seasonal time series, we have set the order of the autoregressive correction equal to $p = 12$. Panel A reproduces the test statistics and the corresponding p-values from the standard distributions, that is, imposing cross-section independence. The MW and IPS tests lead to the rejection of the null hypothesis of unit root, thus supporting the PPP for the Spanish cities. However, the LLC test does not reject the null of non-stationarity.

As mentioned above, independence across individuals is likely to be an unrealistic assumption so that we have computed the empirical distributions of the tests using Bootstrap techniques. Percentiles for each test are reproduced in Panel B of Table 1. Comparison of the test statistics in Panel A with these critical values lead to the same conclusions, that is, we reject the null hypothesis of a non-stationary panel data set when using the MW and IPS tests, whereas the LLC test does not provide evidence in favour of the PPP –the hypothesis testing is performed on the left tail of the distribution for the $\Psi_{\bar{t}}$ and LLC tests and on the right tail for the $\Psi_{\overline{LM}}$ and MW tests.

To sum up, three of the four panel data unit root tests that we have computed provide evidence in favour of the PPP in the sample of Spanish cities. The reason that might be behind the contradictory results of the IPS and MW tests, on the one hand, and the LLC test, on the other, is likely to be the amount of parameter heterogeneity that they allow in their underlying data generating process. Note that under the alternative hypothesis the MW and IPS tests specify a different autoregressive parameter for each individual, while the LLC assumes a common parameter. Moreover, Maddala and Wu (1999) show that the MW and IPS tests encompass the LLC test in terms of empirical size and power. In all, we can conclude that the evidence that has been showed here points to the PPP hypothesis, a conclusion that seems to be robust, first, to different choices of the numerarie and, second, to cross-section dependence in the sample.

4 Persistence of the PPP deviations

Results above allow to reject permanent deviations from the PPP in the panel of Spanish cities over a long time period. It is now interesting to measure the speed of convergence towards the PPP when city prices suffered transitory shocks. The most popular measure of persistence of a shock is its half-life, which is defined as the number of time periods required for a unit impulse to dissipate by one half. There is some consensus in the literature on half-lives of deviations from the PPP of around 3-5 years, which is based on analyses using long-horizon data sets and appliying univariate methods –see Rogoff (1996). Although application of panel data methods slightlty decreases persistence to around 2.5 years –see Papell (1997)–, estimates of half-lives are far from negligible. Accordingly, it has been suggested that departures from the PPP should be driven by factors other than nominal rigidities.

Estimates of half-lives are not free of criticisms. First, the picture based on point estimates is incomplete as they do not come with confidence intervals. Second and even more striking, these half-lives are not appropiately computed, as it has been common practice to estimate them as $\ln(0.5) / \ln(\alpha)$ where α is the autoregressive coefficient in the Dickey-Fuller (DF) equation, $q_{i,t} = \alpha\ q_{i,t-1} + \varepsilon_{i,t}$. The problem arises when $q_{i,t}$ follows an autoregressive process of order higher than one. In this case, the ADF-type equation should be used to test the null hypothesis of unit root, and the half-life should no longer be computed as $\ln(0.5) / \ln(\alpha)$. Instead, it ought to be computed from the impulse response function –see Cheung and Lai (2000) and Murray and Papell (2002). Moreover, these half-life point estimates can be supplemented with Bootstrap confidence intervals, which offer a measure of precision for the point estimates. Murray and Papell (2002) still outline a third pitfall when computing half-lives. It is related to the bias of estimates of the autoregressive parameters in finite samples. Since the impulse response function is build upon these biased estimates, the half-life might turn out to be a biased measure. In this regard, Murray and Papell (2002) follow Andrews (1993) and Andrews and Chen (1994) and compute median-unbiased estimations of the autoregressive parameters and the half-lives. Here we have decided not to follow this suggestion since our time series are long enough to assume that the bias effect should not significantly alter our half-lives estimations.

Table 2 presents point estimates and Bootstrap 95% confidence intervals for half-lives in each city in our sample. It should be stressed the large heterogeneity in the rate of adjustment of the Spanish cities to the national CPI. While shocks in some of them vanish quite rapidly –half-lives of around 2 years–, for others they last more than a decade. The mean point estimate of half-lives is 4.5 years. But the mean confidence interval is rather wide –lower bound of 1.2 years and upper bound of 9.2 years. To prevent the effect of outliers, we have also computed the median of the estimated half-lives. In this case, the point estimate is 3.6 years, with a shorter range for the confidence interval –1.1 and 5.9 years, for the lower and upper bounds respectively.

Our estimates are thus in accordance with the 3-5 years consensus found with international data. This casts doubts on the idea that we should expect lower deviations from the PPP in highly integrated economies. In addition, confidence intervals show that the lower limit is below 1.5 years, a threshold sometimes used to conclude that deviations from PPP are due to nominal rigidities. However, upper bounds are large enough to rule out real rigidities –differentials in productivity or different sectorial economic structures– as a potential determinants of the persistence of these deviations –see Engel and Rogers (2001) and O'Connell and Wei (2002).

5 Concluding remarks

Using a set of long time series for a sample of Spanish cities this paper has shown evidence in favour of the PPP. The unit root hypothesis is strongly rejected through the application of panel data tests that allow for heterogeneity in the parameters for each city. In performing the tests we have taken into account cross-section correlation that can arise when dealing with real exchange rates. In addition, results are robust to the choice of the numerarie.

Although the unit root hypothesis is rejected, this does not prevent to find large persistence of short-run deviations from the PPP, particularly in some of the cities. Estimation of half-lives, based on the impulse response function, are of a magnitude comparable to estimates obtained using international data. In addition, confidence intervals show that these deviations might be long enough to be caused just by nominal rigidities. Thus, our results suggest that further research should be done in order to identify real factors that are behind the slow rates of price convergence in economies such as the EU in which formal barriers between member states have almost vanished and share a common currency.

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Table 2: Point and confidence interval estimates of half-lives

	Half-Lives (years)				Half-Lives (years)		
		95% C. I.				95% C. I.	
	HL	L	U		HL	L	U
A Coruña	6.8	1.4	11.7	León	3.3	1	5.3
Alacant	1.7	0.8	2.8	Logroño	6.4	1.4	10.5
Albacete	2.3	0.8	3.7	Lugo	3.8	1.3	6.1
Almería	1.6	0.8	2.8	Lleida	2.6	0.9	3.8
Ávila	3.3	1.1	5.6	Madrid	2.9	1	4.2
Badajoz	1.8	0.6	3	Málaga	2.8	1	4.6
Barcelona	3.6	1.1	5.8	Murcia	1.7	0.5	2.7
Bilbo	3.8	2.1	81.5	Ourense	3.6	1.1	5.8
Burgos	3.8	1.2	6	Oviedo	1.6	0.8	2.8
Cáceres	13.3	2	35.8	Palencia	2.4	0.8	3.5
Cádiz	13.3	1.8	21.2	P. Mallorca	6.5	1.6	12.4
Castelló	9.4	1.7	14.2	Pontevedra	3.6	1.2	5.9
C. Real	4.9	1.5	7.9	Salamanca	2.9	1	4.7
Córdoba	3	1	4.9	Santander	2.6	0.8	4.2
Cuenca	3.1	1	5.2	Segovia	3.7	1.2	6.1
Donosti	15	1.8	27.2	Sevilla	3.7	1.1	6.3
Gasteiz	11.2	1.7	15.8	Soria	6.3	1.4	11.6
Girona	5.6	1.4	9.1	Tarragona	4.1	1.2	6.7
Granada	2.2	0.8	3.7	Tenerife	2.8	0.9	4.5
Guadalajara	1	0.4	1.6	Teruel	8.9	1.6	14.3
Huelva	0.8	0.4	1.3	Toledo	5.1	1.4	8.4
Huesca	2	0.8	3.2	València	1.2	0.6	1.8
Iruña	4.6	1.3	7.9	Valladolid	7.2	1.4	11
Jaén	5.7	1.4	9.5	Zamora	7.6	1.6	12
Las Palmas	3.9	1.2	6.7	Zaragoza	2.9	1	4.6

		95% C. I.	
	HL	L	U
Mean	4.5	1.2	9.2
Median	3.6	1.1	5.9

The column labelled as HL reproduces the point estimate of the half-life, while the columns labelled as L and U denote the lower and the upper bounds respectively of the 95% confidence interval.

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