TESTING FOR LONG-RUN PURCHASING POWER PARITY IN THE POST BRETTON WOODS ERA: EVIDENCE FROM OLD AND NEW TESTS

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ABSTRACT

This article uses the most recent tests available to carry out a detailed empirical analysis of the validity of Purchasing Power Parity (PPP) for the exchange rates of 21 industrialized countries in the post-Bretton Woods period. It looks at the stationarity properties of the real exchange rates which are required for PPP to hold, and it also analyses the presence of cointegration between the nominal exchange rate and domestic and foreign prices. Overall, the results provide evidence in favor of PPP.

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I. INTRODUCTION

Purchasing Power Parity (PPP) is one of the most thoroughly studied propositions, but even though it has received such attention by the literature, there is not yet clear agreement between the scholars about its empirical validity. In particular, the debate on the validity of long-run PPP for the floating period that followed the collapse of the Bretton Woods system in 1973 has not been concluded. In this article we will study long-run PPP in the post-Bretton Woods era by applying the most recent tests available.

The validity of long-run PPP has been assessed by analysing whether the real exchange rate is stationary. The real exchange rate (q_t) is defined as the nominal exchange rate minus the difference between the domestic price index (p_t) and the foreign price index (p_t^*) as follows:

$$q_t \equiv s_t - p_t + p_t^* \quad (1)$$

where s_t is defined as units of domestic currency per foreign currency and all variables are in logs. If PPP holds in the long-run, then the log of the real exchange rate, q_t, should be zero, that is, the log of the nominal exchange rate, s_t, should be equal to the difference in the price levels (in logs). Therefore, a necessary condition for PPP to hold in the long run is that the real exchange rate is mean reverting or, in the terminology of time series analysis, that it does not contain a unit root.

The literature that has analysed long-run PPP has adopted different approaches. A first approach to analysing long-run PPP has been through the examination of the temporal behaviour of the real exchange rate as defined in (1). At the end of the 80s, following on from the development of techniques specifically designed to test for unit roots, a substantial number of studies tested –and failed to reject– whether q_t contained a unit root[1]. However, as shown by Lothian and Taylor (1997), Sarno and Taylor (2002) and Shiller and Perron (1985), the power of these tests when using a reduced number of years was low, and so authors started to look at other avenues of analysis of the long-run behaviour of the real exchange rate.

One such avenue has complemented the univariate analysis of unit roots with panel data tests given that Banerjee (1999) and Baltagi and Kao (2000), among others, show that unit root tests based on panel data are more powerful than those based on individual data. Initially, many of the studies that applied panel

[1] These studies applied the augmented Dickey-Fuller (1979) test, the non parametric test of Cochrane (1988) and the techniques of fractional integration from Diebold *et al.* (1991) and Cheung and Lai (1993a). For a survey, see Taylor (2003).

unit root tests to a number of real exchange rates series over the recent float rejected the unit root hypothesis. Nonetheless, as pointed by Taylor and Sarno (1998), the null hypothesis of these tests is generally that all the series are generated by unit-root processes, and therefore the probability of rejection of the null hypothesis may be quite high, as one only needs that one of the series is stationary to reject the null. Recently, taking into account this criticism, Sarno and Taylor (1998) and Coakley and Fuertes (2000), for instance, using panel unit root tests find support for the long-run PPP.

Another avenue of study of PPP has been to examine whether the real exchange rate does in fact present mean reversion, but in a non-linear way or with high persistence. For instance, Taylor *et al.* (2001) and Kapetanios *et al.* (2003) apply the smooth transition autoregressive (STAR) model to real bilateral dollar exchange rates and show that they are well characterized by nonlinearly mean reverting processes. On the other hand, Cheung and Lai (1993a) and Gil-Alana and Toro (2002) show that real exchange rates are mean reverting but they exhibit significant persistence in the short run.

Finally, long-run PPP has also been analysed by looking at the presence of a cointegrating relationship between the nominal exchange rate and the prices. By using the panel cointegration methods that allow to test the null of no cointegration without imposing homogeneity of the cointegrating vector, Canzoneri *et al.* (1999) and Pedroni (2004) find support for the weak version of PPP. On the other hand, Pedroni (2001) tests directly that the cointegrating vector is homogeneous and equal to one (strong version of PPP) and it is clearly rejected.

The objective of this article is to study long-run PPP for 21 OECD countries from the end of the Bretton Woods era by applying a wide range of the econometric techniques available. To this end, we will benefit from the latest developments in the analysis of unit roots by applying a variety of panel tests, such as those developed by Levin-Lin-Chu (2002), Breitung (2000), Hadri (2000), Im-Pesaran-Shin (2003), Maddala-Wu (1999), and Pesaran (2005), and the PANIC/PASIC decomposition proposed by Bai and Ng (2004 a, b).

We will also complement the non-stationary analysis by considering, first, if there is non-linearity and second, if there is persistence in the real exchange rates' behaviour. To that end, we will first use non-linear techniques like the smooth transition regressions (STR), where we will consider a logistic transition function. Taylor *et al.* (2001) and Kapetanios *et al.* (2003) have applied an exponential transition function justified in the fact that the adjustment of exchange rates towards the equilibrium should be symmetric. In order to provide a complementary analysis, and also to take advantage of the fact that the test will, according to the data, choose whether there is a symmetric or an asymmetric behaviour, we will choose a logistic LSTR model. Second, we will also analyse the persistence of the series by applying ARFIMA models.

Finally, we will use the tests recently developed by Pedroni (1999, 2004), McCoskey-Kao (1998), Westerlund (2005a,b,c) and Larsson-Lyhagen-Löthgren (2001) to look for the presence of cointegration among the nominal exchange rate and the prices. When there is cointegration, then there is evidence in favour of the weak version of PPP. In the case that there is cointegration, we will also test for the fulfilment of the strong version of PPP.

The article is structured as follows. In section 2 we report the unit root analysis through univariate and multivariate tests. Section 3 presents the non-linearity and persistence analysis of real exchange rates. In section 4 the panel cointegration tests are discussed and, finally, section 5 concludes.

2. UNIT ROOT ANALYSIS

In this section we will test the stationarity of the real exchange rates considered in our sample. To calculate the real exchange rate as in (1), we will use quarterly nominal exchange rates (end of period and expressed as domestic currency per US dollar) and consumer price indexes for 21 industrialised OECD countries and the United States. The period of analysis will be the 31 post-Bretton Woods years since 1973Q1 to 2004Q4[2]. All data is available from the *International Financial Statistics Online Service* of the International Monetary Fund.

In table 1a we present the results of different individual unit root tests. columns 1 and 2 of table 1a report the standard augmented Dickey-Fuller (1979) ADF test, whose null hypothesis is the existence of a unit root, and the Kwiatkowski-Phillips-Schmidt-Shin (1992) KPSS test, whose null hypothesis is the stationarity of the variable analysed. The reason for using both tests lies, as pointed by Maddala and Kim (1998), in the well-known low power of the ADF test in small samples and in the tendency of the KPSS test for over-rejection in this type of samples. Columns 3 to 5 of Table 1a show the results from three types of tests that, according to the literature (Elliott *et al.*, 1996; Ng and Perron, 2001), are more powerful than the ADF in the analysis of non-stationarity: the augmented Dickey-Fuller (DF^{GLS}), the modified Sargan-Bhargava (MSB^{GLS}) and the modified Elliot-Rothenberg-Stock (MP_T^{GLS}) tests under GLS detrending[3]. It can be seen that, even though the ADF test only rejects the presence of a unit root in 4 cases (and only at the 90% significance level in 3 of

[2] Given that Belgium and Luxembourg shared a common currency during this period (before and after the introduction of the euro), Luxembourg has been omitted from the analysis.

[3] The optimal lag length has been selected with the modified Akaike information criterion (MAIC), starting with a maximum length of K = 12 quarters. For technical details and properties of this criterion of lag selection see Ng and Perron (2001).

them), all the other tests point towards the stationarity of the majority of the real exchange rates. Thus, the KPSS indicates the presence of only 5 non-stationary variables, the DF^{GLS} test rejects the presence of unit roots in 12 variables, and for the MSB^{GLS} and MP_T^{GLS} tests the rejection of the unit root hypothesis is fairly generalised (in 18 of the 21 variables).

Table 1a

INDIVIDUAL UNIT ROOT TESTS FOR THE LOG-LEVEL OF THE REAL EXCHANGE RATES, 1973:1-2004:4

	ADF	KPSS	DF^{GLS}	MSB^{GLS}	MP_T^{GLS}	STR	ARFIMA	ARMA ROOTS	
United Kingdom	-2.02 (2)	0.36*	-0.88 (9)	0.14***	1.60***	LSTR2, *d*=2	-0.08 (0.15)	0.89***	-0.25***
Austria	-2.28 (0)	0.12	-1.33 (4)	0.17***	1.91**	LINEAR	0.07 (0.10)	0.91***	-0.14
Belgium	-2.36 (3)	0.19	-2.11 (3)**	0.19**	1.88**	LINEAR	0.11 (0.09)	0.93***	-0.19**
Denmark	-2.58 (3)*	0.09	-1.50 (3)	0.19**	2.18**	LSTR2, *d*=4	0.10 (0.10)	0.90***	-0.22**
France	-2.01 (0)	0.17	-1.84 (0)*	0.26*	3.41*	LSTR2, *d*=4	0.11 (0.10)	0.91***	-0.18**
Germany	-2.05 (0)	0.15	-2.09 (3)**	0.19**	1.84**	LINEAR	0.07 (0.10)	0.92***	-0.15
Italy	-1.96 (0)	0.11	-1.98 (0)**	0.24**	3.37*	LINEAR	0.14 (0.11)	0.92***	-0.15
Netherlands	-2.14 (0)	0.14	-1.82 (3)*	0.18**	1.98**	LINEAR	0.08 (0.10)	0.90***	-0.18*
Norway	-2.36 (0)	0.18	-1.52 (0)	0.23**	3.10**	LSTR2, *d*=2	0.04 (0.10)	0.89***	-0.15
Sweden	-2.28 (3)	0.49**	-2.24 (3)**	0.22**	2.53**	LSTR2, *d*=1	0.17 (0.10)*	0.92***	—
Switzerland	-2.86 (0)**	0.22	-0.83 (0)	0.30	6.00	LINEAR	0.04 (0.10)	0.88***	-0.10
Canada	-2.19 (6)	0.86***	-1.78 (6)*	0.29	4.17*	LINEAR	0.14 (0.08)*	0.96***	—
Japan	-2.38 (0)	0.73**	-0.56 (5)	0.38	7.91	LINEAR	0.13 (0.12)	0.92***	-0.19**
Finland	-2.67 (3)*	0.29	-1.58 (5)	0.20**	2.07**	LSTR1, *d*=4	0.17 (0.10)*	0.90***	—
Greece	-1.75 (0)	0.15	-1.95 (5)**	0.16***	2.08**	LINEAR	0.02 (0.09)	0.95***	0.05
Iceland	-2.63 (0)*	0.15	-1.68 (0)*	0.24*	3.92*	LINEAR	0.07 (0.12)	0.89***	-0.04
Ireland	-1.98 (2)	0.20	-2.10 (3)**	0.18**	2.55**	LSTR2, *d*=4	0.06 (0.11)	0.89***	-0.15
Portugal	-1.54 (0)	0.26	-1.43 (4)	0.18**	2.47**	LINEAR	0.05 (0.09)	0.95***	-0.05
Spain	-1.80 (0)	0.13	-1.27 (3)	0.20**	2.80**	LINEAR	0.17 (0.09)*	0.91***	—
Australia	-1.71 (0)	0.83***	-1.64 (0)*	0.38**	7.21	LINEAR	0.06 (0.09)	0.95***	-0.01
New Zealand	-1.88 (0)	0.07	-1.79 (0)*	0.26*	3.80*	LSTR1, *d*=4	0.15 (0.09)	0.93***	-0.08

Notes: 1) The logarithm of the real exchange rate was computed as $q = s + p^{USA} - p$, where p^{USA} is the US aggregate log-price level; 2) *ADF* is the augmented Dickey-Fuller unit root *t* test (only with intercept) and the number between parenthesis is the lag order of the corresponding regression (based on *MAIC* criterion using a step-down procedure starting from *K*=12); 3) *KPSS* is the Kwiatkowski-Phillips-Schmidt-Shin unit root test; the computed bandwidth was 9 (using the Newey-West approach and a Barlett kernel as spectral estimation method); 4) DF^{GLS}, MSB^{GLS} and MP_T^{GLS} are the augmented Dickey-Fuller, modified Sargan-Bhargava and modified Elliot-Rothenberg-Stock tests under GLS detrending (based on *MAIC* criterion using a step-down procedure starting from $K=12$); 5) *STR* is the Smooth Transition Regression model (with a Logistic transition function) estimated for each variable to analyze nonlinear real exchange rate behaviour; 6) *ARFIMA* is the estimate of the *d* parameter (and of the standard error between parentheses) of a Fractionally Integrated ARMA model estimated for each variable to analyze long-run dependence in real exchange rate behaviour; 7) *ARMA ROOTS* are the estimated AR and MA parameters of an ARMA/ARFIMA adjusted for each series; 8) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values.

In order to complement the univariate study, we have also carried out a panel analysis of the real exchange rate series. The extension to panel analysis is justified by the results from recent studies (see Banerjee (1999) or Baltagi and Kao (2000), among others), which suggest that unit root tests based on panel data are more powerful than those based on individual data. Further, Karlsson and Löthgren (2000) analyse, through a Monte Carlo simulation, the power of some of the unit root tests for panel data used here and conclude that for panels with a considerable temporal dimension ($T > 100$) there is a risk of over-rejecting non-stationarity, whereas the opposite is true for panels with a small temporal dimension. To minimise this risk, they propose a simultaneous analysis, as the one carried out in this article, of the individual and panel tests. Before we present the results in table 1b, we would like to point out a few methodological notes about these tests.

All panel tests to be used are based on the null hypothesis of the presence of a unit root in the series, with the exception of Hadri's (2000) test, whose hypothesis is that the series are stationary. The tests differ from each other on the restrictions imposed on the autoregressive process of each of the panel series. Thus, the tests of Levin-Lin-Chu (2002), Breitung (2000) and Hadri impose a common persistence parameter to all the series –therefore, if the null is rejected, the alternative would be that all the series are simultaneously stationary for the first two tests and non-stationary for the latter. On the other hand, the tests of Im-Pesaran-Shin (2003), the Fisher-type tests suggested by Maddala and Wu (1999), and the Pesaran's (2005) CADF test allow for the autoregressive parameter to change freely among the different cross-sectional variables under consideration. Therefore, the alternative hypothesis in these cases is the presence of a non-null proportion of stationary series of the total[4]. The latter set of tests seem more adequate from an empirical point of view as they impose less restrictions on the data generating process.

[4] For an extensive theoretical discussion of the tests see Banerjee (1999) or Baltagi and Kao (2000). For a more detailed and recent review of the literature on unit roots and cointegration in panels see Breitung and Pesaran (2005).

All above mentioned panel tests, with the exception of Pesaran's, assume that there are no short-run or long-run cross-correlations among the autoregressive processes that govern the behaviour of each time series. In particular, all these tests are based in the absence of cross-correlation or cointegration among the variables of the panel. However, O'Connell (1998) and Banerjee *et al.* (2003, 2005) have demonstrated that all tests are affected when this property is missing. This will lead to less reliability as the null hypothesis will be rejected more often than it should be according to the confidence level prefixed. Nevertheless, it is most likely that in practice there are significant cross dependencies among the real exchange rates of the different countries given the presence of common components. For example, when using the US dollar as the base currency, not only independent changes in the dollar value and in the US price index will be included in the real exchange rate, but also any other type of variable or global shock that is common to all or some of the countries from the sample. For the 21 real exchange rates of our sample, the cross-correlations oscillate from a minimum value of -0.317 to a maximum of 0.996, which reveals the relevance of the cross-dependency problem. This fact brings about a potentially important bias in the standard tests, which we have tried to lessen in two ways.

In the first place, we have extracted a specific time effect which would collect all the contemporaneous factors that are common to all exchange rates. In practice this implies working with the time-demeaned real exchange rates, which, as shown by Luintel (2001), does not eliminate all the present correlation, but it does reduce it considerably. Second, we have dealt with the cross-dependency problem through the implementation of two tests that take into account the presence of cross-correlation and/or cointegration: the cross-dependence modified ADF test (CADF) suggested by Pesaran (2005) and the decomposition procedure put forward by Bai and Ng (2004 a,b)[5].

Table 1b reports the panel tests results. With the exception of the Pesaran test, under each standard panel test we report the corresponding version time demeaned, which will be less affected by the cross-dependency problem. The tests reported in Table 1b present, on balance, evidence in favour of PPP. The null hypothesis of non-stationarity is rejected in all tests with the exception of those of Levin-Lin-Chu and Hadri. Nonetheless, Levin-Lin-Chu and Hadri´s tests both have some limitations. Levin-Lin-Chu imposes strong parametric restrictions which imply that under the alternative hypothesis all series must be

[5] The most recent literature labels the panel tests that take into account the cross-dependency problem as "second generation tests" (see the surveys of, among others, Hurlin and Mignon (2004) and Breitung and Pesaran (2005)). Among the second generation tests, we have decided to use those suggested by Pesaran and Bai-Ng based on the results of Gengenbach *et al.* (2004), Baltagi *et al.* (2005), Gutierrez (2005) and Moon and Perron (2005) who, through Monte Carlo simulations, have shown that both tests keep good size and power properties under different specifications of the underlying model.

stationary and have the same autoregressive parameter. Hadri's test, on the other hand, has a tendency to over-reject the null hypothesis as it is based on KPSS tests and, as shown by Caner and Kilian (2001), the KPSS statistics tend to reject the stationarity hypothesis more often than they should at the specified significance level. Therefore, we can conclude that there is evidence that at least a significant proportion of the 21 real exchange rates are stationary.

Table 1b

PANEL UNIT ROOT TESTS FOR THE LOG-LEVEL OF THE REAL EXCHANGE RATES, 1973:1-2004:4

	Statistic	**Prob.**
Null: Unit root *(assumes common unit root process)*		
Levin-Lin-Chu		
Standard:	0.37	0.64
Time demeaned:	-0.10	0.46
Breitung		
Standard:	-4.92	0.00***
Time demeaned:	-3.19	0.00***
Null: Unit root *(assumes individual unit root process)*		
Im-Pesaran-Shin		
Standard:	-3.39	0.00***
Time demeaned:	-3.84	0.00***
Maddala-Wu ADF-Fisher		
Standard:	66.01	0.01***
Time demeaned:	84.03	0.00***
Maddala-Wu PP-Fisher		
Standard:	82.38	0.00***
Time demeaned:	98.90	0.00***
Null: No unit root *(assumes common unit root process)*		
Hadri		
Standard:	4.19	0.00***
Time demeaned:	18.24	0.00***
Pesaran *Null: Unit root* *(assumes individual unit root process)* Allows for cross-sectional dependence]		
CADF:	-2.51	(1% Critical value: -2.36)***

Notes: 1) The probabilities for the Fisher tests have been computed using an asymptotic Chi-square distribution (all the other tests assume asymptotic normality); 2) Time-demeaned statistics have been demeaned with respect to common time effects to accommodate for some forms of cross-sectional dependency; 3) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level.

Our conclusion is given further support when we use the methodology suggested by Bai and Ng (2004 a, b). Bai and Ng put forward an approach, which they call PANIC/PASIC[6], which consists of decomposing a time series panel in two components, a part that is common to all the series and a part with idiosyncratic components. Then, unit root and stationarity tests are carried out separately on each component[7]. The advantage of using this methodology lies in the fact that the decomposition allows for the construction of panel tests that verify the cross-independence hypothesis, so that the tests applied to the estimated components will display better statistical properties than those based on the original observed series.

To be more precise, the PANIC/PASIC methodology assumes that the real exchange rate observed series, q_{it}, can be decomposed in the form $q_{it}=\alpha_i+\lambda_i' F_t+e_{it}$, where α_i is the deterministic component of the series, F_t is a *k*-vector of common factors and e_{it} is a specific term. F_t and e_{it} are unobserved elements that must be estimated using the information from the complete panel. This factor model makes clear that for a real exchange rate to be stationary, the common and idiosyncratic components must also be stationary. Non-stationarity can arise from the presence of a unit root in the common factors or in the specific component of each series.

We have applied the principal components method put forward by Bai and Ng to our empirical analysis, estimating the factor model $q_{it}=\alpha_i+\lambda_i' F_t+e_{it}$. We selected one single common factor using the $IC_1(K)$ information criterion suggested by Bai and Ng (2002), which accounted for over 71% of the total variation of the data[8]. Next we estimated the idiosyncratic component of each series and, finally, we applied the unit root and/or stationarity tests to the two estimated components. The results from this analysis are shown in table 2a. In the first place, it can be seen that four out of the five tests applied to the common factor $\hat{F}_t$ point toward its stationarity. On the other hand, four panel tests (with the exception, again, of the Levin-Lin-Chu and Hadri tests) accept

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6 PANIC and PASIC refer to Panel Analysis of Non-Stationarity in Idiosyncratic and Common Components and Panel Analysis of Stationarity in Idiosyncratic and Common Components, respectively.

7 The decomposition method of the CADF test suggested by Pesaran (2005) is similar, but it is applied to the residuals of the autoregressive model used for each series and allows for the presence of only one common factor.

8 The second factor accounted for only an additional 5.7% of the variation, a similar percentage to that of the third factor (5.3%).

the alternative hypothesis for the idiosyncratic errors, and therefore there is evidence of the stationarity of a non-null proportion of these components. Taken together, both results suggest –given the additive nature of the factor model used– that at least a significant part of the original real exchange rate series presents mean reversion and so PPP is verified for this significant part.

Table 2a

THE PANIC/PASIC APPROACH TO PANEL TESTING OF REAL EXCHANGE RATES: GLOBAL RESULTS

	Statistic	Prob.
COMMON FACTOR $(\hat{F})$		
ADF *(Null: Unit root)*	-1.78	—
KPSS *(Null: No unit root)*	0.09	—
DF^{GLS} *(Null: Unit root)*	-1.81^{*}	—
MSB^{GLS} *(Null: Unit root)*	0.26^{*}	—
MP_T^{GLS} *(Null: Unit root)*	3.89^{*}	—
IDIOSYNCRATIC COMPONENTS $(\hat{e}_i)$		
Panel unit root tests		
Null: Unit root (assumes common unit root process)		
Levin-Lin-Chu	-2.82	0.00^{***}
Breitung	-1.30	0.09^{*}
Null: Unit root (assumes individual unit root process)		
Im-Pesaran-Shin	-2.86	0.00^{***}
Maddala-Wu ADF-Fisher	57.15	0.06^{*}
Maddala-Wu PP-Fisher	75.00	0.00^{***}
Null: No unit root (assumes common unit root process)		
Hadri	18.92	0.00^{***}

Notes: 1) The probabilities for the Fisher tests have been computed using an asymptotic Chi-square distribution (all the other panel tests assume asymptotic normality); 2) To test the (non)stationarity of the idiosyncratic components, the Levin-Lin-Chu, Breitung and the Maddala tests have been computed in a model with no deterministic term; 3)An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level.

In table 2b we study the relative weight of each of the estimated components[9]. It can be seen that the specific component dominates only for the real exchange rates of Canada, Japan, Iceland, Australia and New Zealand,

[9] Note that the selection applied depends on the cutting point chosen –0.6 in our case– and thus the following statements should be taken with caution.

whereas in the remaining cases the common factor explains the majority of the variations. This result indicates an additional value added from the PANIC/PASIC approach, as it shows that for the majority of the real exchange rates analysed the time dynamics are dominated by a common factor with a clear European profile.

Table 2b

THE PANIC/PASIC APPROACH TO PANEL TESTING OF REAL EXCHANGE RATES: INDIVIDUAL RESULTS FOR THE IDIOSYNCRATIC COMPONENTS

	$\mathrm{Var}(\Delta\hat{e}_i)/\mathrm{Var}(\Delta q_i)$	$\sigma(\hat{\lambda}_i\hat{F})/\sigma(\Delta\hat{e}_i)$	**Dominant component**
United Kingdom	0.44	1.09	*Common*
Austria	0.06	4.14	*Common*
Belgium	0.07	2.76	*Common*
Denmark	0.07	5.95	*Common*
France	0.09	3.96	*Common*
Germany	0.07	3.52	*Common*
Italy	0.28	1.94	*Common*
Netherlands	0.06	3.45	*Common*
Norway	0.19	2.96	*Common*
Sweden	0.32	1.13	*Common*
Switzerland	0.20	2.26	*Common*
Canada	0.98	0.10	*Idiosyncratic*
Japan	0.61	0.62	*Idiosyncratic*
Finland	0.26	1.28	*Common*
Greece	0.33	1.87	*Common*
Iceland	0.65	1.56	*Idiosyncratic*
Ireland	0.15	1.86	*Common*
Portugal	0.23	1.45	*Common*
Spain	0.28	1.85	*Common*
Australia	0.88	0.34	*Idiosyncratic*
New Zealand	0.70	0.95	*Idiosyncratic*

Notes: 1) $\mathrm{Var}(\Delta\hat{e}_i)/\mathrm{Var}(\Delta q_i)$ is the ratio of the standard deviation of the idiosyncratic component to the standard deviation of the differenced data; 2) $\sigma(\hat{\lambda}_i\hat{F})/\sigma(\Delta\hat{e}_i)$ is the ratio of the standard deviation of the common factor to the idiosyncratic component; 3) We have selected a *Idiosyncratic* dominant component if the $\mathrm{Var}(\Delta\hat{e}_i)/\mathrm{Var}(\Delta q_i)$ statistic exceeds 0.6 and a *Common* dominant component otherwise.

The set of individual and panel tests carried out in this section provide, taken as a whole, evidence in favour of the stationarity of (at least) some of the 21 real exchange rates analysed, and thus they are indicating that the long-run PPP proposition is verified.

3. NON-LINEARITY AND PERSISTENCE ANALYSIS

To complement the analysis of individual and panel stationarity, we have also studied the non-linearity, time-dependence and persistence properties of the real exchange rates, which have recently appeared as alternative or complementary ways of analysing long-run PPP in the literature.

In the first place, we have used non-linear techniques to assess the validity of the conclusions from recent theoretical models which predict a non-linear adjustment of the real exchange rates towards their long-run equilibrium values (see, for instance, the discussions of Taylor *et al.* (2001), Sarno and Taylor (2002) and Taylor (2003)). In other words, these studies point towards the presence of non-linear stationary autoregressive processes for real exchange rates. In contrast with non-linear stationary models, the maintained hypothesis of linear stationary models is that the adjustment towards equilibrium in exchange rates happens continuously and at constant speed, not taking into account the deviations from equilibrium at each point in time.

Non-linear behaviour can be characterised through Smooth Transition Regression (STR) models[10]. In particular, we have used a general class of models of the type $q_t = \phi' w_t + \theta' w_t G(\gamma, c, s_t) + u_t$, where $w_t' = (1, q_{t-1}, \ldots, q_{t-p})'$ is the autoregressive component of the model and $G(\gamma, c, s_t)$ is the transition function, which in our case is given by a general logistic function of the type $G(\gamma, c, s_t) = \left(1 + \exp\left(-\gamma \prod_{k=1}^{K} (s_t - c_k)\right)\right)^{-1}$. In this function, $\gamma > 0$ is the parameter that controls the slope of the function, $c = (c_1, c_2, \ldots, c_K)'$ is the vector of location parameters (with $c_1 \leq c_2 \leq \ldots \leq c_K$) and s_t is the transition variable, which in our case is given by y_{t-d}, where *d* is the lag parameter of the transition function. The most common choices for K are K=1 and K=2, which generate the so-called

[10] For a general revision of these models see, among others, Granger and Teräsvirta (1993), van Dijk *et al.* (2002) or Teräsvirta (1998, 2004). There are also alternative models suggested by the literature. One of them, the Self-Exciting Threshold Autoregressive (SETAR) model, and specially, the 3-regime SETAR put forward by Bec *et al.* (2004) are particularly interesting because of its consistency with PPP theories with transaction costs.

LSTR1 and LSTR2 models. In the LSTR1 model, the parameters $\phi+G(\gamma,c,s_t)$ change monotonously (and asymmetrically) from the initial values ϕ to the final values $\phi+\theta$, whereas in the LSTR2 model, those parameters change symmetrically around the average point $(c_1+c_2)/2$, where the function G has its minimum value. When $\gamma=0$, in both models the transition function is constant and thus both models become a linear autoregressive specification.

The results from the non-linearity tests of the 21 real exchange rates analysed are reported in column 6 of table 1a. As suggested by Teräsvirta (2004), in the STR modelling process we have followed a successive application of the specification, estimation and evaluation stages and concluded afterwards whether to apply a linear model or one of the two non-linear models LSTR1 or LSTR2[11]. There is evidence in favour of the non-linear adjustment towards equilibrium hypothesis for more than a third of the sample, in particular for 8 of the 21 real exchange rates analysed, and in two of them (Finland and New Zealand) the adjustment is asymmetric (LSTR1). This result raises two interesting issues. First, it demonstrates the difficulty for conventional unit root tests to detect mean reversion in exchange rates, as these tests are based on linear processes for the variables which, as shown, are not always linear. Second, it shows the importance of recent theories of the PPP which predict different behaviours for the exchange rates depending on the size of the deviation from their long-run equilibrium positions.

Another approach proposed by the literature to analyse whether real exchange rates have a unit root or are mean reverting has been the extension of the standard linear autoregressive models (ARIMA) to include more general specifications. For instance, the ARFIMA models are used for fractionally integrated variables, whose highly persistent behaviour make standard ARIMA models –and the stationarity tests based on them– inadequate. For our analysis, these models allow for a higher degree of persistence in the temporal dynamics of exchange rates and thus, for less restrictive processes of mean reversion.

ARFIMA models take the form $\Phi(L)(1-L)^d q_t=\Theta(L)\varepsilon_t$, where L is the lag operator, $\Phi(L)$ and $\Theta(L)$ are p and q order polynomials in L, and *d* is the fractional integration parameter. For $d=0$ the ARFIMA model becomes a stationary ARMA model, and for $d=1$ the process is non-stationary and it will not be mean reverting. The values of d determine the stochastic properties of the series q_t: for the series to be stationary it is required that $d<0.5$, whereas the behaviour of the series will be non-stationary when $d\geq0.5$ even though, as long

[11] The complete estimations of each model can be obtained from the corresponding author upon request.

as $d<1$, there will be long-run mean reversion. Therefore, in testing for PPP, we will have to identify and estimate the corresponding ARFIMA model for each real exchange rate series and test, in the first place, whether the fractional estimated parameter $\hat{d}$ is statistically different from zero and, in the second place, whether it is smaller than one. If $\hat{d}$ was smaller than one, this would indicate that there is a long-run mean reverting process toward the parity value, and the speed of adjustment would be higher or lower depending on whether the estimated value is smaller or bigger than 0.5.

We have estimated the 21 ARFIMA (*p*, *d*, *q*) models for the real exchange rate series by using the maximum likelihood method (Sowell, 1992), and we report the estimates of the parameter *d* in column 7 of table 1a[12]. It can be seen that the estimated parameter $\hat{d}$ is statistically different from zero at a 90% significance level in only four cases (Sweden, Canada, Finland and Spain), but in all four cases the value is significantly lower than 0.5[13]. This would imply that all the series analysed are stationary: 17 of them would follow stationary ARMA processes and 4 of them would (marginally) follow stationary ARFIMA processes. To complete the analysis, the last 2 columns of table 1a present the estimations of the autoregressive and moving average parameters of the ARMA(1,1) models of the 17 series for which the hypothesis $d=0$ could not be rejected, and the estimations of the ARFIMA (1, *d*, 0) models for the remaining 4 series[14]. In all cases the estimated parameter of the autoregressive component of the model is statistically significant and has a high value. This would indicate a high degree of temporal dependence in the evolution of exchange rates, and further, would prove the difficulty of standard unit root tests to detect the non-stationarity of the series, given the low power of these tests when the autoregressive parameter is near unity.

To summarise, the results obtained from the analysis of stationarity of the real exchange rates through a variety of techniques applied in section II and III indicate that there is strong evidence that the 21 real exchange rates considered in this article are stationary and therefore this provides support for the existence of PPP in its strong version.

[12] We started with a maximum value of 2 for *p* and *q* and we chose the final model according to the Akaike (AIC) and Schwarz (SBC) information criteria. In all cases we obtained models of the type ARFIMA(1,*d*,1) or inferior, which justifies that in columns 8 and 9 we only present the results of the estimation of the first order model for the exchange rates. All estimations are available from the authors upon request.

[13] For the four cases, the null hypothesis of a value above 0.5 was rejected through a Wald statistic $(\hat{d}-d)/\hat{\sigma}_{\hat{d}}$.

[14] A substantial number of the ARMA(1,1) models for the 17 series are over parameterised. However, we have preferred to present the general model in order to show the magnitudes of the corresponding autoregressive and moving average parameters.

4. PANEL COINTEGRATION TESTS

In this section we will study the weak version of long-run PPP, which relaxes the hypotheses of symmetry and proportionality that underlie the analysis of real exchange rates. In particular, we will look at expressions of the type $s_{i,t} = \beta_{0,i} + \beta_{1,i} p_t^{USA} + \beta_{2,i} p_{i,t} + \varepsilon_{i,t}$, that do not impose the restrictions $\beta_{1,i} = -1$ and $\beta_{2,i} = 1$ implicit in the strong version of PPP. This type of equations must be interpreted as long-run equilibrium relationships and, for this, it is required that there is cointegration among the variables. If cointegration is present, we will test for the strong version of PPP (i.e., whether $\beta_{1,i} = -1$ and $\beta_{2,i} = 1$).

The analysis to be developed next will be based on testing for the cointegration hypothesis and we will apply the econometric techniques developed in the recent literature. These techniques exploit the panel dimension of the data, considerably improving the statistical properties of standard cointegration tests used in the analysis of individual time series and allow for a higher degree of heterogeneity in the parameters and in the time dynamics of the series.

Before we proceed with the cointegration analysis, we will look at whether the nominal exchange rates and the domestic and foreign prices are unit root processes. Table 3 shows the results from two standard unit-root tests: the augmented Dickey-Fuller (1979) ADF test and the Kwiatkowski-Phillips-Schmidt-Shin (1992) KPSS test. The ADF test (first column of table 3) does not reject the unit root hypothesis for any of the nominal exchange rates, whereas the KPSS test (second column) accepts the unit root hypothesis for only 12 of the 21 cases analysed. On the other hand, the ADF test accepts the stationarity of 5 of the price indexes considered (Denmark, the Netherlands, Japan, Finland and Ireland), whereas the KPSS test rejects in all cases the stationarity hypothesis. From these results we can conclude that there is general evidence in favour of the presence of a unit root both in the nominal exchange rates and in the consumer price indexes considered[15].

[15] We also investigated the presence of a second unit root in the series and in all cases it was rejected. Further, we also applied alternative tests, such as those of Table 1a for real exchange rates, and in all cases the conclusions were similar to those of the ADF and KPSS. These complementary results are available from the authors upon request.

Table 3
UNIT ROOT TESTS FOR THE LOG-LEVEL OF THE ORIGINAL VARIABLES, 1973:1-2004:4

	s		*p*	
	ADF	*KPSS*	*ADF*	*KPSS*
United Kingdom	-1.83 (0)	0.15**	-2.81 (5)	0.32***
Austria	-2.42 (0)	0.11	-2.89 (4)	0.32***
Belgium	-2.43 (0)	0.09	-3.00 (2)	0.33***
Denmark	-1.60 (0)	0.12*	-3.33 (0)*	0.34***
France	-1.45 (0)	0.14*	-2.33 (3)	0.34***
Germany	-2.41 (0)	0.11	-2.07 (4)	0.27***
Italy	-1.21 (0)	0.17**	-1.81 (1)	0.34***
Netherlands	-2.23 (0)	0.10	-3.54 (4)**	0.25***
Norway	-2.14 (0)	0.09	-1.11 (4)	0.35***
Sweden	-2.45 (3)	0.10	0.05 (1)	0.35***
Switzerland	-2.99 (0)	0.13	-2.50 (4)	0.29***
Canada	-0.67 (0)	0.10	-2.36 (1)	0.34***
Japan	-2.03 (0)	0.17**	-6.39 (5)***	0.30***
Finland	-2.76 (3)	0.07	-3.80 (6)**	0.34***
Greece	0.56 (0)	0.25***	1.06 (9)	0.33***
Iceland	-0.09 (2)	0.33***	-1.02 (4)	0.35***
Ireland	-1.53 (0)	0.15**	-3.19 (4)*	0.33***
Portugal	0.01 (0)	0.30***	-0.31 (1)	0.35***
Spain	-0.90 (0)	0.16**	-3.11 (4)	0.34***
Australia	-1.33 (0)	0.17**	-2.68 (2)	0.35***
New Zealand	-0.72 (0)	0.23***	-1.25 (2)	0.35***
United States (*numeraire* country)	—	—	-3.01 (3)	0.32***

Notes: 1) *s* is the logarithm of the nominal exchange rate (relative to the US dollar) and *p* is the logarithm of the aggregate price level (*CPI*); 2) *ADF* is the augmented Dickey-Fuller unit root *t* test (with intercept and time trend) and the number between parentheses is the lag order of the corresponding regression (based on *MAIC* criterion using a step-down procedure starting from *K*=12); 3) *KPSS* is the Kwiatkowski-Phillips-Schmidt-Shin unit root test; the computed bandwidth was 9 (using the Newey-West approach and a Barlett kernel as spectral estimation method); 4) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values.

We will apply four groups of tests that have been proposed, respectively, by Pedroni (1999, 2004), McCoskey-Kao (1998), Westerlund (2005a,b,c) and Larsson-Lyhagen-Löthgren (2001)[16]. The tests of Pedroni, McCoskey-Kao and Westerlund are based on the residual estimates of the individual cointegration relationships, whereas the Larsson-Lyhagen-Löthgren test is based on the analysis of multiple cointegrating vectors. All these tests allow for a high degree of individual heterogeneity, so that the coefficients of each cointegrating relationship can vary freely for each exchange rate. It is interesting to note that the tests of Pedroni, Larsson-Lyhagen-Löthgren and two of the tests proposed by Westerlund have a null hypothesis of absence of cointegration among the variables of each equation whereas in the LM tests of McCoskey-Kao and the CUSUM tests of Westerlund the null hypothesis is stationarity of the residuals – and so the presence of cointegration among the variables. Finally, of all these tests, only the Durbin-Hausman-type tests proposed by Westerlund allow explicitly for the presence of dependency among the panel data.

Pedroni has developed seven different cointegration statistics, all of them based on the least squares residuals of long-run equations of the type $s_{i,t}=\beta_{0,i}+\beta_{1,i}p_t^{USA}+\beta_{2,i}p_{i,t}+\varepsilon_{i,t}$ and on the null hypothesis of no cointegration. Four of these tests have the panel test feature (within dimension), as they are constructed adding separately the numerator and the denominator over the cross-sectional dimension of the panel. At the same time, each of these 4 tests can be constructed weighted –using an estimation of the long-run variances as weights– or non-weighted, so that actually there are 8 different tests. The remaining 3 tests are group average tests (between dimension), constructed dividing first each numerator and denominator and afterwards adding to the number of panel units. In any case, the standardised distributions of the panel and group statistics are given by $\left(\chi_k-\mu_k\sqrt{N}\right)/\sigma_k \Rightarrow N(0,1)$, where χ_k is the corresponding statistic, and μ_k is their expected mean and σ_k^2 is their expected variance, which are tabulated in Pedroni (1999).

Table 4 reports the estimations of the cointegration tests proposed by Pedroni. We present the results for the original series and also for the time-demeaned series, with the aim to address the cross-dependency problem mentioned above. For the original series, the panel statistics $\nu-\text{stat}$ and $\text{ADF}-\text{stat}$ clearly reject the null hypothesis of non-cointegration, both in the weighted and non-weighted versions; of the group tests, only the $\text{ADF}-\text{stat}$ rejects the null hypothesis. Further, the cross-dependency corrected series all strongly reject the absence of cointegration, for any type and version of the tests applied. These results clearly point towards the presence of a cointegration relationship among

[16] See Gutierrez (2003) for a Monte Carlo analysis of the statistical properties of some of the cointegration tests proposed in the literature.

the nominal exchange rate and price variables where there could be a long-run equilibrium relationship of the type $s_{i,t}=\beta_{0,i}+\beta_{1,i}p_t^{USA}+\beta_{2,i}p_{i,t}+\varepsilon_{i,t}$ for a non-null proportion of the exchange rates analysed.

Table 4

PEDRONI'S PANEL AND GROUP COINTEGRATION TESTS FOR THE PROCESS $\left(s,p^{USA},p\right)$, 1973:1-2004:4

	ν−stat	ρ−stat	PP−stat	ADF−stat
Weighted Panel stats				
Standard:	5.27***	-1.27	-0.69	-4.12***
Time demeaned:	5.86***	-4.37***	-3.88***	-2.28**
Unweighted Panel stats				
Standard:	5.27***	-1.37*	-0.87	-4.11***
Time demeaned:	7.07***	-4.91***	-4.18***	-2.58***
Group-mean stats				
Standard:	—	0.59	0.57	-4.15***
Time demeaned:	—	-4.60***	-4.52***	-2.86***

Notes: 1) All of the panel and group statistics have been standardized by the means and variances given in Pedroni (1999) so that all reported values are distributed as $N(0,1)$ under the *null hypothesis of no cointegration*; 2) The panel-stats weighted statistics are weighted by long run variances (Pedroni, 1999, 2004). 3) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values (1.28, 1.64 and 2.33, respectively). 3) For the semiparametric *PP* tests we have used the Newey-West (1994) rule for truncating the lag length for the kernel bandwidth, and for the parametric *ADF* tests we have used a step-down procedure starting from *K*=12; 4) Panel and group mean time-demeaned statistics have been demeaned with respect to common time effects to accommodate for some forms of cross-sectional dependency.; 5) The residuals have been estimated using the least squares estimator.

McCoskey-Kao´s test is a panel version of the Lagrange multipliers test proposed by Harris and Inder (1994) and Shin (1994) for the individual analysis of cointegration. The expression for this test is $\overline{LM}=\frac{1}{N}\sum_{i=1}^{N}\left[\left(\frac{1}{T^2}\sum_{t=1}^{T}S_{it}^{+2}\right)\Big/s^{+2}\right]$, where S_{it}^{+2} represents the partial sums of the residuals $\left(S_{it}^{+2}=\sum_{j=1}^{t}\hat{\varepsilon}_{ij}\right)$, s^{+2} is a consistent estimation of the residual variance given by $s^{+2}=\frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\varepsilon}_{it}^2$, and $\hat{\varepsilon}_{it}$ represents the estimated residuals. The $\overline{LM}$ test is the weighted mean of the individual Lagrange tests of each equation, and McCoskey and Kao (1998) have shown that the standardised version of such average statistic is given by

$\overline{LM}^{*}=\left(\sqrt{N}\left(\overline{LM}-\mu_{v}\right)\right)/\sigma_{v}\Rightarrow N(0,1)$, where μ_v and σ_v^2 are, respectively, the expected mean and variance of the $\overline{LM}$ statistic.

The estimated residuals can be obtained by the dynamic OLS (DOLS) method –as proposed by Harris and Inder (1994)– or by the fully modified OLS (FMOLS) method –as proposed by Shin (1994). Kao and Chiang (2000) have demonstrated that the DOLS estimator has better properties than the FMOLS both in homogeneous panels and in heterogeneous panels as the one used in our analysis. Therefore, we have applied the generalised DOLS method in the estimation of each one of the relations $s_{i,t}=\beta_{0,i}+\beta_{1,i}p_t^{USA}+\beta_{2,i}p_{i,t}+\varepsilon_{i,t}$.

Table 5 reports the individual statistics and the (standardised and non-standardised) panel statistics proposed by McCoskey and Kao, obtained from the dynamic generalised least squares (DGLS) method suggested by Stock and Watson (1993) applied to the nominal exchange rates regressions. This DGLS method generalises the DOLS approach suggested by Saikkonen (1991) as it includes not only the lead and lagged explanatory variables to correct for their endogeneity, but it also includes an autoregressive process for the errors of the model in order to obtain autocorrelation-free residuals. In our case, we started with a general model for each equation with 4 lags and 4 leads of the variables p^{USA} and p_i and an AR(2) process for the errors ε_i, and have simplified according to the statistical significance of the parameters. It can be seen from table 5 that, at the individual level, the null hypothesis is only marginally rejected in two cases (Belgium and the Netherlands), whereas the panel statistic $\overline{LM}^{*}$ lies clearly in the non-rejection area of the null hypothesis of cointegration. Therefore, these results give further support to the conclusion obtained with Pedroni's tests, in the sense that there are stable long-run equilibrium relationships between the nominal exchange rate and the domestic and foreign (US) prices for each country.

Table 5

MCKOSKEY-KAO'S PANEL COINTEGRATION TEST FOR THE PROCESS $\left(s,p^{USA},p\right)$, 1973:1-2004:4

	***LM**$_i$* **statistics**
United Kingdom	0.046
Austria	0.098
Belgium	0.173*
Denmark	0.063
France	0.080
Germany	0.123

(Sigue)

(Continuación)

	***LM_i* statistics**
Italy	0.087
Netherlands	0.208*
Norway	0.114
Sweden	0.088
Switzerland	0.114
Canada	0.029
Japan	0.052
Finland	0.081
Greece	0.108
Iceland	0.079
Ireland	0.064
Portugal	0.163
Spain	0.083
Australia	0.054
New Zealand	0.100
$\overline{LM}$	0.096
μ	0.1219
σ^2	0.0099
$\overline{LM}^*$ ***panel test***	**-1.21**

Notes: 1) The panel test $\overline{LM}^*$ is $\sqrt{N}$ times the standardized version of the $\overline{LM}$ statistic (using the mean and variance given in McCoskey and Kao, 1998), so the reported value is distributed as $N(0,1)$ under the *null hypothesis of cointegration*; 2) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values (for the $\overline{LM}$ statistic these values are 1.28, 1.64 and 2.33, respectively; for the LM_i statistics they are 0.16, 0.22 and 0.38, respectively). 3) The residuals have been estimated using the generalized dynamic least squares estimator proposed by Stock and Watson (1993).

The cointegration tests put forward by Westerlund are all applied to the estimated residuals of the potential equilibrium relationship, even though the null and the alternative hypotheses differ according to the version of the test. Thus, the CUSUM test for panel data tests the null of cointegration, whereas the variance ratio tests –VR– and the Durbin-Hausman tests have the null of absence of cointegration amongst the variables.

The CUSUM test is an extension for panel data of the test proposed by Xiao (1999) and Xiao and Phillips (2002) for time series. This test tries to measure the extent of the fluctuations in the estimated residuals $\hat{\varepsilon}_{it}$ through the statistic

$$PC=\frac{1}{N}\sum_{i=1}^{N}\left[\max_{s=1,\dots,T}\left(\sqrt{\frac{1}{\hat{\varpi}_{i1,2}^{2}T}}\left|\sum_{t=1}^{s}\hat{\varepsilon}_{it}\right|\right)\right]$$, where $\hat{\varpi}_{i1,2}^{2}$ is a consistent estimation of the long-run variance of ε_{it}. Similarly to McCoskey and Kao´s test, the $\hat{\varepsilon}_{it}$ residuals can be obtained by the DOLS or the FMOLS methods. In any case the statistic is asymptotically free of nuisance parameters, so that it verifies asymptotically (and sequentially, that is, when $T\to\infty$ followed by $N\to\infty$) that $\left(\sqrt{N}(PC-\mu)\right)/\sigma\Rightarrow N(0,1)$, where μ and σ^2 are the first and second order moments of the *PC* statistic, which have been tabulated by Westerlund (2005a).

Amongst the non-cointegration tests, Westerlund (2005b) proposes two non-parametric tests that do not require any hypothesis about the time dynamics of the errors ε_{it} (although they assume cross-independency of the errors). These two variance ratio tests are the within group –panel type– VR_P, with $VR_P=\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{E}_{it}^{2}\left(\sum_{i=1}^{N}\hat{R}_i\right)^{-1}$ and the between group –group average– VR_G, with $VR_G=\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{E}_{it}^{2}\hat{R}_i^{-1}$, where $\hat{E}_{it}=\sum_{j=1}^{t}\hat{\varepsilon}_{it}$ and $\hat{R}_i=\sum_{t=1}^{T}\hat{\varepsilon}_{it}^{2}$. Westerlund shows that both statistics have an asymptotic normal standard distribution.

Finally, Westerlund (2005c) proposes two cointegration tests for panel data that do not impose cross-independence among the units of the panel. To model the cross-dependencies, a factorial approach like that of Bai and Ng (2004a,b) is considered, but it is applied to the long-run regressions' errors ε_{it}. In particular, the tests proposed in our case are given by the following system of equations:

$$
\begin{gathered}
s_{i,t}=\beta_{0,i}+\beta_{1,i}p_t^{USA}+\beta_{2,i}p_{i,t}+\varepsilon_{i,t}\\
\varepsilon_{it}=\lambda_i^{'}F_t+e_{it}\\
(I-\gamma L)F_t=C(L)w_t\\
(I-\rho_i L)e_{it}=D_i(L)v_t
\end{gathered}
$$

where *F* is a *K*-vector of common factors, e_i are vectors of idiosyncratic errors and *C(L)* and *D(L)* are, respectively, matrices and polynomials in the lag operator *L*. Once all the elements of this system have been estimated, the two statistics are constructed following the Durbin-Hausman principle, obtaining a panel test $DH_P=\hat{\sigma}^2\hat{\gamma}_0^{-2}(\tilde{\rho}-\hat{\rho})^2E_{22}$ and a group average test $DH_G=\sum_{i=1}^{N}\hat{\sigma}_i^2\hat{\gamma}_{io}^2(\tilde{\rho}_i-\hat{\rho}_i)^2E_{i22}$, which, when normalised, have normal standard limit distributions.

Table 6 shows the estimations of the 3 groups of tests put forward by Westerlund. Interestingly, they all indicate that there is heterogeneous cointegration between nominal exchange rates and domestic and foreign (US) prices, which provides further evidence in favour of PPP in the weak sense.

Table 6

WESTERLUND'S TESTS FOR PANEL COINTEGRATION OF THE PROCESS (s, p^{USA}, p), 1973:1-2004:4

	Statistic	Prob.
Null: No unit root in residuals (cointegration)		
CUSUM	0.607	0.27
Null: Unit root in residuals (no cointegration)		
VR_G	-4.073***	0.00
VR_P	-2.984***	0.00
DH_G	1.846**	0.03
DH_P	2.608***	0.01

Notes: 1) All of the statistics have been standardized by the means and variances given in Westerlund (2005a,b,c) so that all reported values are distributed as $N(0,1)$ under the *null hypothesis of cointegration or no cointegration*; 2) An * (**) [***] indicates rejection of the null hypothesis at the 10% (5%) [1%] significance level based on the appropriate critical values (1.28, 1.64 and 2.33, respectively); 3) The residuals for the *CUSUM* test have been estimated using the fully modified least squares estimator (FM-OLS) [the CUSUM test based on DOLS estimates was 0.996, with p-value of 0.16]; 4) The residuals for the *VR* tests have been estimated using the least squares estimator; 4) For the *DH* tests the number of factors have been estimated using the $IC_1(K)$ criterion with the maximum number of factors set equal to five.

The tests of Pedroni, McCoskey and Kao and Westerlund all share a weakness due to the fact that they all assume that there is only one cointegrating vector among the variables (s, p^{USA}, p). To avoid this problem, Larsson *et al.* (2001) have developed a panel statistic based in the multivariate approach of Johansen (1988, 1991), which allows for the presence of multiple cointegration relationships amongst the variables[17]. Larsson-Lyhagen-Löthgren's test is given by the expression $\overline{LR}_{NT}[H(r)/H(p)]=\frac{1}{N}\sum_{i=1}^{N}[LR_{iT}[H(r)/H(p)]]$, where LR_{iT} represents the trace statistic proposed by Johansen to test the hypothesis $H_0\text{:rank}(\Pi_i)=r_i \leq r$ against the alternative $H_1\text{:rank}(\Pi_i)=p$ for each country, and where *p* is the number of variables in the model (3 in our case). The null hypothesis of the test is that all countries of the panel have a maximum common rank of *r* cointegrating vectors, even though it is allowed that each country has

[17] Recently, Breitung (2005) has extended the approach of Larsson *et al.* (2001) to more general cases. In particular, deterministic components in the underlying VAR model are allowed (this is why in our empirical analysis we use the tabulated critical values given in Breitung's article), and a new estimator for the cointegrating vector(s) is proposed, which can be modified in case of contemporaneous cross-correlation.

its own r_i number of stable equilibrium relationships. Larsson *et al.* demonstrate that the asymptotic distributions of the standardised version of $\overline{LR}_{NT}$ is given by $\Psi_{\overline{LR}}[H(r)/H(p)]=\frac{\sqrt{N}\left(\overline{LR}_{NT}[H(r)/H(p)]-E(Z_k)\right)}{\sqrt{Var(Z_k)}}\Rightarrow N(0,1)$, where $E(Z_k)$ and $Var(Z_k)$ are, respectively, the mean and variance of the asymptotic distribution to which the trace statistic $LR_{iT}[H(r)/H(p)]$ converges, with $k = p\text{-}r$.

The estimations of the individual statistics LR_{iT}, and of the average and standardised Larsson-Lyhagen-Löthgren ones are reported in table 7[18]. At the individual level, the trace tests reject in all cases (at least at the 95% confidence level) the null hypothesis of absence of cointegration (r=0), and this conclusion is further reinforced by the panel statistic $\Psi_{\overline{LR}}[H(0)/H(3)]$. On the other hand, some of the individual trace statistics and the panel statistics $\Psi_{\overline{LR}}$ reject the hypothesis of one and even two cointegrating vectors. It is worth pointing out that a similar result has been obtained in other studies, like those of Coakley and Fuertes (2000), Cerrato and Sarantis (2002) or Caporale and Cerrato (2004). Nonetheless, this last result should be taken with caution as the Larsson-Lyhagen-Löthgren test is based on the individual Johansen statistics and it is well-known that they tend to overestimate the number of cointegrating vectors and, further, they are very sensitive to the inclusion of different deterministic components, to the error distribution, to the number of lags chosen and to the size of the time series used (Maddala and Kim, 1998)[19].

Table 7

INDIVIDUAL (JOHANSEN) AND PANEL (LARSSON-LYHAGEN-LÖTHGREN) TRACE COINTEGRATION TESTS FOR THE PROCESS $\left(s,p^{USA},p\right)$, 1973:1-2004:4

	Lag (k_i)	LR(r=0)/Prob[1]	LR(r=1)/Prob[1]	LR(r=2)/Prob[1]
United Kingdom	5	60.06 (0.00)	24.75 (0.00)	5.77 (0.02)
Austria	5	45.26 (0.00)	20.08 (0.01)	6.17 (0.02)
Belgium	5	57.44 (0.00)	21.56 (0.01)	4.85 (0.03)
Denmark	2	64.29 (0.00)	20.39 (0.01)	0.85 (0.36)
France	2	46.55 (0.00)	18.98 (0.01)	0.13 (0.71)

(Sigue)

[18] The optimal lag length of the VAR model for each country has been chosen according to the Akaike (AIC) and Schwarz (SBC) information criteria, and for the individual trace tests we have used the probabilities of McKinnon *et al.* (1999).

[19] In particular, it is shown in Maddala and Kim (1998, p. 173-220) that the Johansen statistics for testing the second and subsequent cointegrating vectors suffer from size distortions and tend to show multiple cointegrating vectors when the proportion between the number of time observations and the number of parameters is relatively low.

(Continuación)

	Lag (k_i)	LR(r=0)/Prob[1]	LR(r=1)/Prob[1]	LR(r=2)/Prob[1]
Germany	5	30.88 (0.04)	15.83 (0.04)	4.76 (0.03)
Italy	4	36.09 (0.01)	8.54 (0.41)	0.30 (0.59)
Netherlands	5	47.53 (0.00)	18.51 (0.02)	6.43 (0.01)
Norway	4	45.59 (0.00)	13.18 (0.11)	3.67 (0.06)
Sweden	4	32.07 (0.03)	13.39 (0.10)	0.88 (0.35)
Switzerland	5	42.82 (0.00)	23.20 (0.00)	9.32 (0.00)
Canada	2	34.54 (0.01)	3.51 (0.94)	0.39 (0.53)
Japan	4	72.78 (0.00)	15.82 (0.05)	4.94 (0.03)
Finland	5	60.10 (0.00)	18.73 (0.02)	7.12 (0.01)
Greece	5	32.34 (0.02)	12.60 (0.13)	3.80 (0.05)
Iceland	4	41.64 (0.00)	12.87 (0.12)	0.12 (0.73)
Ireland	4	49.98 (0.00)	14.41 (0.07)	2.56 (0.11)
Portugal	2	63.27 (0.00)	21.81 (0.01)	5.14 (0.02)
Spain	5	35.45 (0.01)	10.15 (0.27)	3.71 (0.05)
Australia	2	64.13 (0.00)	18.10 (0.02)	3.46 (0.06)
New Zealand	4	32.87 (0.02)	10.32 (0.26)	0.62 (0.43)
$\overline{LR}_{NT}[H(r)/H(3)]$		47.41	16.03	3.57
$E[Z_k]$		19.35	8.27	0.98
$Var[Z_k]$		31.84	14.28	1.91
$\Psi_{\overline{LR}}[H(r)/H(3)]$ *panel test*		22.79	9.42	8.59

Notes: 1) The panel test $\Psi_{\overline{LR}}[H(r)/H(3)]$ is $\sqrt{N}$ times the standardized version of the $\overline{LR}_{NT}[H(r)/H(3)]$ statistic (using the mean and variance given in Breitung, 2005, Table B.1/Case 3) so the reported value is distributed as $N(0,1)$ under the *null hypothesis of no cointegration*; 2) Prob denotes McKinnon-Haug-Michelis (1999) *p*-values.

The evidence from all the range of individual and panel cointegration tests applied in this section clearly points towards the presence of a long-run equilibrium relationship between nominal exchange rates and domestic and foreign prices according to the equation $s_{i,t}=\beta_{0,i}+\beta_{1,i}p_t^{USA}+\beta_{2,i}p_{i,t}+\varepsilon_{i,t}$. Therefore, the next stage in our analysis is the estimation of the parameters of each of these relationships, and of the parameters of an average function for the complete panel of 21 exchange rates.

We have taken into account two issues when deciding the estimator to be used. A first issue, already mentioned before, refers to the problems that arise from the Johansen method, whose estimators are generally not very robust to changes in the initial VAR model used[20]. Second, the study of Maddala and Kim (1998) reveals that amongst the alternative methods proposed by the literature for the estimation of cointegration equations, the DOLS/DGLS methods of Saikkonen (1991) and Stock and Watson (1993) offer the best results in finite samples with respect to other estimators asymptotically more efficient.

Thus, our estimation of the long-run relations for the nominal exchange rates is based on these estimators and the results for our sample of 21 countries are reported in table 8. It can be seen that there is a high variability in the significance and in the coefficients' estimates, both for the foreign price $\left(p^{USA}\right)$ and for the domestic price (*p*). Thus, the foreign price is significant in the UK, Belgium, the Netherlands, Japan, Iceland, Ireland and Portugal (a third of the total panel), with the correct expected sign in all cases except for Japan. The domestic price, on the other hand, is significant in 14 out of the 21 countries (the UK, Austria, Belgium, Denmark, France, Italy, the Netherlands, Finland, Greece, Iceland, Ireland, Portugal, Australia and New Zealand), presenting the correct expected sign in all these cases. In the last column we present the Wald statistics to test the validity for each country of the joint symmetry and proportionality restrictions, $\beta_{1,i}=-1$ and $\beta_{2,i}=1$. We reject the null for 5 countries (Austria, Denmark, France, Canada and Australia) at the 5% level and for 2 countries (United Kingdom and Japan) at the 1% level, and thus, for the remaining 14 countries we do not reject the null and so we cannot reject the strong version of PPP for them. Hence, the evidence on the rejection of the symmetry and proportionality conditions at the individual country-level is in line with similar results obtained by other studies (see, for example, Cheung and Lai (1993b) or Cerrato and Sarantis (2002)).

The last row of table 8 presents a pooled estimation for the complete panel to establish a basis for comparison, using a fixed effects model for each country but assuming homogeneity in the slopes of the exchange rate equation, that is, a specification of the type $s_{i,t}=\beta_{0,i}+\beta_1 p_t^{USA}+\beta_2 p_{i,t}+\varepsilon_{i,t}$[21]. We have used a DGLS type

[20] Further, some works have pointed out the weaknesses of the statistical properties of the maximum likelihood estimators for small simples (Phillips, 1994; Hansen *et al.*, 1998; Brüggemann y Lütkepohl, 2004).

[21] It might seem very restrictive to impose homogeneity of the effect of the domestic and foreign prices amongst all the members of the panel, but for our data, both the individual likelihood ratio tests for the domestic and foreign prices ($\chi^2_{LR}=14.44$ and $\chi^2_{LR}=20.60$) and the joint hypothesis of homogeneity of the price effect among countries ($\chi^2_{LR}=29.91$) do not reject the corresponding null hypothesis. However, the generality of the model is still limited as it restricts the dynamics of the variables –leads and lags– and the autoregressive process for the errors.

estimator, similar to the one used for the individual regressions, adding four leads and lags of the explanatory variables $(p^{USA}$ and $p)$ and an AR(2) model for the errors of the model. It can be seen that in this specification the two parameters are highly significant and their values are statistically indistinguishable ($\chi^2_{WALD}=0.52$, Prob = 0.77) from the theoretical values needed for the PPP to hold in its strong version ($\beta_1=-1$ and $\beta_2=1$)[22].

Table 8

LONG-RUN EQUILIBRIUM NOMINAL EXCHANGE RATES FUNCTIONS (STOCK-WATSON DGLS ESTIMATES)

	p^*	p	**AR (p)**	**LL (q)**	χ^2_{WALD}
United Kingdom	-1.26* (-1.89)	0.96** (1.91)	2	3	13.11*** (0.00)
Austria	0.63 (0.54)	3.68** (2.26)	1	2	7.25** (0.03)
Belgium	-2.20** (-2.35)	2.87** (2.35)	1	3	2.74 (0.25)
Denmark	-0.09 (-0.07)	3.15*** (2.65)	1	2	6.80** (0.03)
France	-1.21 (-1.19)	3.25*** (3.45)	1	0	7.21** (0.03)
Germany	-1.10 (-1.17)	1.18 (0.78)	1	3	0.02 (0.99)
Italy	-1.85 (-1.40)	2.65*** (3.06)	1	3	3.88 (0.14)
Netherlands	-1.65** (-2.41)	2.33** (2.25)	1	2	2.29 (0.32)
Norway	-0.38 (-0.72)	0.62 (1.45)	1	0	2.46 (0.29)
Sweden	-0.31 (-0.52)	0.71 (1.53)	2	0	3.35 (0.19)
Switzerland	-0.50 (-0.82)	0.13 (0.13)	1	2	0.83 (0.66)

(Sigue)

[22] Our estimates under the homogeneity hypothesis are very similar to those obtained by Breitung (2005) using the FM-OLS method and the new two-stage estimator proposed in his work. Further, his Wald statistics does not reject in both cases the null hypothesis of $\beta_1=-1$ and $\beta_2=1$.

(Continuación)

	p^*	p	AR (p)	LL (q)	χ^2_{WALD}
Canada	-0.18 (-0.38)	-0.22 (-0.47)	1	0	7.05** (0.03)
Japan	2.57** (2.45)	0.75 (0.90)	1	1	12.09*** (0.00)
Finland	-0.86 (-1.47)	1.16** (2.01)	2	0	1.59 (0.45)
Greece	0.50 (0.69)	0.52** (2.45)	1	1	5.08* (0.08)
Iceland	-0.88** (-2.04)	1.00*** (9.72)	1	3	0.75 (0.69)
Ireland	-1.52*** (-2.66)	1.37*** (3.14)	1	2	0.85 (0.65)
Portugal	-1.29* (-1.67)	1.18*** (4.14)	1	2	0.46 (0.79)
Spain	-0.82 (-0.81)	0.87 (1.42)	2	3	0.05 (0.97)
Australia	-0.47 (-0.85)	0.80* (1.86)	1	1	5.92** (0.05)
New Zealand	-0.39 (-0.66)	0.60* (1.64)	1	1	1.17 (0.56)
Pool 21 OECD (homogeneous)	-1.00*** (-12.95)	1.02*** (22.17)	2	4	0.52 (0.77)

Notes: 1) The numbers within parentheses (below coefficients) are *t* values; 2) *AR*(*p*) denotes the order of the autoregressive model used in the estimation (we have used a step-down procedure starting from *p*=2); 3) *LL*(*q*) denotes the order of the leads-lags terms used in the estimation (we have used a step-down procedure starting from *Q*=4); 4) χ^2_{WALD} denotes the Wald test of joint symmetry-proportionality restriction and the number within parentheses (below Wald statistics) are *p*-values; 5) An * (**) [***] indicates statistical significance at the 10% (5%) [1%] level; 5) The model that we have used to pool time-series and cross-sectional data involves the assumptions of varying intercepts and homogeneous slope coefficients.

5. CONCLUSIONS

This article has carried out a detailed empirical study of long-run PPP in the post-Bretton Woods period. In doing so, we have reviewed the current status quo of the empirical analysis of PPP. We have analysed the statistical properties of the real exchange rates, which is equivalent to testing for PPP in its strong version. We have also examined the relationship between nominal exchange rates and domestic and foreign prices for each country, which implies the analysis of PPP in its weak version.

Overall, the results obtained through the analysis implemented in this article point in favour of the validity of PPP. Thus, when analysing real exchange rates, through individual and panel unit root tests, through non-linear stationarity models and through high persistence ARFIMA models, evidence indicates that a considerable number of the 21 real exchange rates examined are stationary. Further, the group of cointegration tests proposed by Pedroni (1999, 2004), McCoskey-Kao (1998), Westerlund (2005a,b,c) and Larsson-Lyhagen-Löthgren (2001), clearly indicate the presence of a long-run equilibrium relationship between nominal exchange rates and domestic and foreign prices for each country, which gives support to the validity of the weak version of PPP. There is also evidence in favour of the strong version of PPP when tested through cointegration techniques.

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SÍNTESIS

PRINCIPALES IMPLICACIONES DE POLÍTICA ECONÓMICA

Aunque la hipótesis de la paridad del poder adquisitivo (PPA) es una de las proposiciones más analizadas en la literatura económica reciente, sin embargo aún no existe un consenso claro sobre su validez empírica. En particular, aún sigue abierto el debate sobre la validez a largo plazo de la PPA en el período que siguió al final del sistema de Bretton Woods y la consecuente aparición generalizada de tipos de cambio más o menos flotantes.

El objetivo básico de esta investigación consiste en analizar los tipos de cambio de un conjunto de países industrializados para determinar si se encuentra evidencia favorable sobre el cumplimiento de la PPA. En primer lugar se analizan los resultados respecto a la presencia de raíces unitarias en los tipos de cambio reales, lo cual iría contra la PPA en sentido estricto. En segundo lugar, se examinan con detalle las propiedades de cointegración de las relaciones entre los tipos de cambio nominales y los precios externos e interior de cada país, al objeto de examinar si se cumplen versiones más débiles de la PPA.

En general, nuestros resultados pueden considerarse globalmente favorables a la hipótesis de la PPA. Así, cuando se analizan los tipos de cambio reales, la evidencia apunta hacia la estacionariedad de dichas series para una parte importante de los 21 países examinados. Por otra parte, todos los contrastes de cointegración realizados señalan claramente la existencia de una relación de equilibrio a largo plazo entre los tipos de cambio nominales y los precios interno y externo de cada país, lo que implica el cumplimiento de la versión débil de la PPA.

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