Testing for Long Run Relative Purchasing Power Parity in Europe

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Abstract

This paper tests for long run relative PPP using recently developed nonstationary panel regression estimators that can accommodate cross sectional dependence and both permanent and temporary shocks. The PPP null in our framework is a unit elasticity of nominal exchange rates with respect to relative prices. Using US dollar and deutschemark denominated exchange rates over the 1977:1-20001:12 period for 15 European countries we cannot reject the hypothesis that the long run relative price elasticity is unity. We conclude that long run relative PPP holds in our European sample.

Keywords: Real exchange rate; nonstationary panel estimators; permanent shocks; cross sectional dependence.

JEL Classifications: C32; F31

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1. Introduction

Purchasing power parity (PPP) states that there is a proportional relationship between the prices - proxied by a representative basket of goods - in one country relative to that of another when expressed in the same numéraire currency. Although there are a number of PPP specifications, the concept that has been the focus of recent empirical studies is long run PPP that permits short run deviations. The concept has been the subject of much debate both in the theoretical and econometric literature. As Dornbusch and Krugman (1976) comment, most macroeconomists have a deep-seated belief that a variant of PPP is justified in some sense. Since it forms a cornerstone of many macroeconomic models of trade and of exchange rate determination, failure to support this parity empirically would somewhat undermine the basis for such models.[1]

Support for PPP has fluctuated wildly over the post-Bretton Woods period. Following the advent of floating exchange rates in the early 1970s, the concept of continuous PPP was tested and found to be flawed. This failing was attributed to the excess volatility in the spot rate *vis-à-vis* changes in domestic and foreign prices. This view was modified somewhat by Dornbush's (1976) overshooting model which allows for goods prices to be sticky and therefore permitted deviations from PPP over short horizons. However, the post-Bretton Woods period is still relatively short and so the power of time series tests has been augmented by adding cross-sectional data.

While initially this was seen as the solution to the power problem, it was soon realised panel tests of PPP are subject to the cross-section dependence (CSD) or contemporaneous correlation problem. Authors such as Higgins and Zakrajsek (1999), O'Connell (1998), Papell (2003), Taylor (2003), and Wu and Wu (2001) question the strong support for the PPP evidence found in the literature for the post-Bretton Woods period. They show that problems with the econometric testing techniques used lead to unreliable results. One notable problem is that of CSD. The typical results from tests that accommodate CSD are found support PPP less strongly or indeed to reject it as do the studies by O'Connell (1998) and Wu and Wu (2001).

The first contribution this paper makes to the PPP literature is to adopt an econometric methodology that is robust to CSD. This issue has been relatively neglected despite the seminal contributions of Abuaf and Jorion (1990) and O'Connell (1998). The approach popularised by Abuaf and Jorion (1990) is to apply the seemingly unrelated regression (SUR) framework in panel estimation of individual slope coefficients or in panel unit root tests. In PPP studies, the base country price index is part of relative prices for all countries and so the omitted common factor is correlated with the regressor. Thus the implication is that the SUR technique is not generally appropriate for PPP studies. We deal with this problem by employing the common correlated estimators (CCE) approach developed by Pesaran (2003a). This augments the PPP regression by cross sectional averages of both nominal exchange rates and relative prices.

The second contribution is that we adopt a panel regression approach that can accommodate both permanent and temporary shocks. In this respect it follows the approach of Coakley, Flood, Fuertes and Taylor (2004). The motivation for this framework is that both nominal exchange rates, and to a lesser degree relative prices, seem persistent processes in finite samples. Moreover, real exchange rates are generally recognised as persistent processes and therefore one cannot exclude a role for permanent shocks. The slope coefficient in our panel regression approach can be interpreted as the long run elasticity of nominal exchange rates with respect to relative

[1] For recent reviews of the literature see Rogoff (1996), Sarno and Taylor (2002), and Taylor (2003).

prices. The null hypothesis is a long run slope coefficient of unity or long run relative PPP (RPPP).

In contrast to our approach, most of the existing literature adopts the unit root or cointegration approach to testing PPP. This literature has two potential problems. On one hand, it is predicated on the presence of temporary (monetary) shocks only. On the other, and it implicitly imposes the symmetry and proportionality restrictions in defining real exchange rates. It tests long run PPP through examining the time series properties of the real exchange rate. If real exchange rates are found to have a unit root, this implies a violation of PPP. Otherwise they are stationary or mean reverting and this is interpreted as support for long run PPP.

The remainder of this paper is organised as follows. Section 2 shows the relationship between long run absolute and relative PPP. Section 3 presents a brief outline of the Mean Group (MG) panel method. Section 4 contains the details of the dataset and empirical results while a final section concludes.

2. Concepts of Purchasing Power Parity

Purchasing power parity is a logical extension of the law of one price (LOP) that states that the price of good *i* in the domestic country should equal the price in another, when expressed in the same numéraire currency:

$$p_t\,(i) = p_t^*(i) + s_t \qquad (1)$$

where $p_t\,(i)$ is the price of good *i* in the domestic country at time t, $p_t^*(i)$ the equivalent foreign country price, s_t the spot exchange rate (the domestic price of foreign currency) and all variables are in logarithmic form.

Others include Breuer (1994) and Bleaney and Mizen (1995).

Making the assumption that the LOP holds across all goods, then it must hold for a convex combination or basket of goods. It therefore follows that the price of a basket of goods in one country should, when compared in a like currency, equal that of another, assuming that the baskets are identical. Traditionally proxies such as the consumer price index (CPI) and wholesale price index (WPI) are used to represent these baskets. This extension of the law of one price to price indices yields the concept of absolute PPP:

$$p_t = p_t^* + s_t$$

$$s_t = p_t - p_t^* \tag{2}$$

The second concept, RPPP, can be illustrated by taking the differential of the logged nominal exchange rate and the other variables in (2)

$$\begin{aligned} ds_t / s_t &= \left(dp_t / p_t - dp_t^* / p_t^*\right) \\ (ds_t / s_t) / \left(dp_t / p_t - dp_t^* / p_t^*\right) &= 1 \end{aligned} \tag{3}$$

RPPP can be interpreted as implying that small inflation differentials are reflected exactly or one-for-one in exchange rate depreciation. More formally it implies a unit elasticity of nominal exchange rates with respect to relative prices.

In the empirical literature, one can contrast two main approaches to testing for long run PPP:

(a) On one hand, there is the unit root (Engle and Granger, 1987) and cointegration (Johansen 1988) approach.

(b) On the other one can use a nonstationary panel regression approach that avoids the spurious correlation problem.

2.1 Unit root studies

Since most of the literature looks at the PPP relationship as a long-run equilibrium condition, it can readily be tested by applying unit root and/or cointegration frameworks. The general approach is that these studies look at the real exchange rate, via the addition of an error term to equation (2):

$$
\begin{aligned}
s_t &= p_t - p_t^* + u_t \\
q_t = u_t &= s_t + p_t^* - p_t
\end{aligned}
\tag{4}
$$

where q_t is the real exchange rate. They test the null hypothesis of a unit root to determine whether q_t is a nonstationary process. If mean reversion is not found for the real exchange rate or the error term is I(1), then this indicates a rejection of PPP in the long-run. Nonrejection of long run PPP would imply that the exchange rate is proportional to relative prices and is not subject to permanent shocks. The latter implication is that the error term is stationary.

Generally, these studies found difficulty in rejecting the null hypothesis of a unit root, which constitutes a violation of PPP. However, amidst the flurry of papers applying the methods of Engle and Granger, a discussion arose regarding the power of these tests to reject the null hypothesis of nonstationary when applied to the post Bretton Woods floating period alone. Frankel (1986) argued that the deviations from PPP could be persistent and that it might therefore require long-horizon datasets to reliably reject the null hypothesis. The source of this persistence is still under scrutiny. A recent attempt to explain this behaviour by Ng (2003) uses a semi-structural VAR to identify sticky price shocks in two countries. She finds that the US has been the main source of real exchange rate deviation post-Bretton Woods. Interestingly she notes that the real exchange rate adjusts reasonably fast to the US sticky price shock and that the persistence of the real exchange rate cannot therefore be explained by this phenomenon alone. Persistence is found to increase only when this effect is combined with other shocks.

Given that there is persistence, the use of long-horizon datasets increases the reliability of the unit root test insofar as it gives more opportunity to detect a slow rate of reversion. What it does not do is offer a rationale about why and through what mechanisms this persistence occurs. Several studies have examined the use of long-horizon datasets address this power problem. Using data from 1869-1984 Frankel is able to reject the null hypothesis, finding an estimated rate of decay of 14% per year. This value is not too dissimilar to Lothian and Taylor (1996) who demonstrate the low power of the unit root test by generating an AR(1) model of the real exchange rate using a 200 year sample. Their first-order autocorrelation coefficient implied a speed of mean reversion of just over 11% per year. Sarno and Taylor (2002) use the Lothian and Taylor results to run a Monte Carlo experiment and find support for the latter's view that using the UK/US exchange rate, there is only a 50/50 chance of rejecting the unit root hypothesis with 100 years of data.[2]

Using these long horizon datasets can be problematic. The data are subject to survival bias as it is simply not available for some countries over the entire sample span (Froot and Rogoff, 1995). Furthermore, when working with such a large span, one will encounter movements in the real exchange rate that may be attributable to real factors such as technical innovation and regime change (Hegwood and Papell, 1998). Data over such a long span, incorporating such shifts therefore need to be analysed also in the context of structural breaks.[3] Further criticism is voiced by Engel (2000). The results of his Carlo experiments suggest that studies that use data with

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[2] Further, they find that the smallest span permissible to achieve a probability of rejecting the null of least 50% would be would be approximately 75 years. This value is obtained by taking the lowest value of the 95% confidence interval employed.
[3] See for example Lothian and Taylor (1996).

long horizons may have reached the wrong conclusion (by rejecting the null hypothesis) since tests for long run PPP suffer from size biases. Cheung and Lai (1997) suggest that it may not be necessary to use long-horizon datasets to increase the power of the test, rather that one should utilise more powerful Dickey-Fuller test. Employing two forms of Dickey-Fuller test as modified by Park and Fuller (1995) and Elliott, Rothenberg and Stock (1996), they find mean reversion in the post-Bretton Woods period.[4]

An alternative to long horizon datasets is to use panel data. The advantage of this method is that the power of unit root tests can be increased by adding series, and therefore it overcomes the need for long time horizons and their attendant problems. Two often cited examples of this type of study are Abuaf and Jorion (1990) and Papell (1998). Abuaf and Jorion (1990) test for the presence of a unit root in the real exchange rate using a form of multivariate GLS. The null hypothesis of joint nonstationarity is tested across a series of 10 countries from 1973-1987. This study rejects this hypothesis, and this therefore is seen as supporting long run PPP.[5]

Papell (1998) follows this methodology, using 20 industrialised countries over a 22 year post Bretton Woods period, with both monthly and quarterly data. Papell notes that Abuaf and Jorion do not incorporate serial correlation in the disturbances when calculating the critical values for the panel unit root test. Here, the hypothesis of unit root is rejected for monthly data, but not for quarterly. Further, as with the study here, this type of analysis tends to offer support for long run PPP.

Such multivariate tests, whilst appealing in that they allow us to believe in PPP in some sense, are not without criticism. Notably Taylor and Sarno (1998) find

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[4] Pantula, Gonzalez-Farias and Fuller (1994) show that these tests have approximately the same power.

[5] Also tested is a long-horizon dataset that shows shocks to the real exchange rate cancel out over time. A half life of 3 years is observed here, and of 3-5 years in the post Bretton Woods sample. This is in keeping with Frankel and Rose (1996) who find a half life of approximately 4 years

that these tests reject the joint null of nonstationarity when just one of the composite processes are stationary. They suggest another test where the null is only violated when all the processes in question are stationary. In this study they find strong evidence of mean reversion in the real exchange rate.

2.2 Panel regression approach

The regression test approach to PPP marks a departure from the unit root or cointegration approach. Much of the early work using time-series studies was of this type and suffered from the nonsense regression problem. This was because at that time there was a lack of empirical tools to distinguish between the short- and long-run effects. The results from these earlier studies predominantly reject PPP. Frenkel (1981) rejected PPP for industrialised countries, and suggested that this rejection was in some way related to short run factors pushing the slope coefficient from its hypothesised value. One of the often-cited exceptions from this early literature is Frenkel (1978). He ran regressions for a number of hyperinflationary economies and did indeed find the slope of the price differential to be close to 1.[6]

This somewhat older literature focused on the expression of equation (2) in regression form to test for PPP. This paper presents the results of a new test for long run RPPP by extending the simple regression equations employed by the early literature. The problem of spurious regression is avoided via the application of recent innovations in the nonstationary or I(1) panel regression literature. This panel method allows for consistent estimation of a long run slope coefficient even in the presence of I(1) errors or when the real exchange rate is subject to permanent shocks.

[6] The dataset covered 1921 to 1925, with the study supporting both forms of PPP, but RPPP most of all.

To see how this panel regression framework is of use in the context of PPP, consider again equation (2) showing that the nominal exchange rate is proportional to the price differential. This method of testing long run PPP is rather restrictive in two respects. On one hand, it imposes the symmetry and proportionality restrictions on the real exchange rate without testing them. On the other hand it is not easy empirically to make an argument for the error term always being stationary, $u_t \sim$ I(0). One possible solution to the potential problem of non-stationary errors is to take the first difference.

$$\Delta s_t = \left(\Delta p_t - \Delta p_t^*\right) + \Delta u_t \tag{5}$$

so that Δu_t is I(0) by definition. This specification was used in Flood and Taylor (1996) and, employing long differences, was seen as a test of long-run relative PPP. However, in actually testing equation (5), one is left with only the information on the short run or high frequency dynamics. As with all first difference equations, the long run or low frequency information has been lost.

While long run PPP can be tested within a unit root framework, long run RPPP is more readily tested within a nonstationary panel regression framework. This approach permits the consistent estimation of a long run slope coefficient irrespective of whether the error term is I(0) or I(1).

$$s_{it} = \alpha_i + \beta_i\left(p_{it} - p_{it}^*\right) + u_{it} \tag{6}$$

This facilitates testing the null hypothesis of a unit relative price elasticity of the nominal exchange. The finding of a value that was not significantly different from 1 would imply an acceptance of long-run RPPP irrespective of whether $u_t \sim$ I(0) or u_t ~I(1).

3. Panel estimation framework

3.1 Nonstationary panels

It is well known that, in a time series analysis of I(1) variables, the absence of cointegration leads to the statistical problem of spurious regression. However, recent theoretical contributions by Phillips and Moon (1999) and Kao (1999) establish that large *N* and large *T* panel datasets offer the prospect of overcoming the nonsense regression problem of pure time series. More particularly, they demonstrate that in panels one can consistently estimate a long-run average parameter or mean effect even if there is no time-series cointegration at an individual level. The later pertains to situations where the error term and the variables are nonstationary. The intuition is that the averaging or pooling over independent countries lessens the `noise' in the relationship - the covariance between the I(1) error and the I(1) regressor - that induces the nonsense regression problem and leads to a stronger overall 'signal' than in pure time-series approaches.

Coakley, Fuertes and Smith (2001) examine the applicability of these results to the field of applied econometrics by employing Monte Carlo simulations. They investigate the small sample properties of the fixed effects (FE), pooled OLS and mean group (MG) panel estimators with I(1) errors. It is shown that the bias in the estimates declines at $\sqrt{N}$ using a static regression with I(1) errors for all three panel methods. They find that the standard *t*-tests for the MG estimator are correctly sized for the case of both I(1) or I(0) errors while those for the FE and POLS estimators are subject to potentially severe size distortions.

The particular technique we employ in this nonstationary framework is the MG estimator of Pesaran and Smith (1995). This is due to its ability to deal with both I(0) and I(1) errors and because we can rely on its standard errors for inference purposes in finite samples. It allows one to estimate consistently the long run association between non-cointegrating I(1) variables and so avoid the problem of spurious regression in panels. The MG estimator has the added advantage of accommodating country heterogeneity by means of country specific intercepts and slopes.[7]

While the underlying theory is complex, the application of the MG panel method is reasonably straightforward. Firstly, for each country selected for the panel, a single OLS regression is run, equation (5) from above:

$$s_{it} = \alpha_i + \beta_i \left(p_{it} - p_{it}^*\right) + u_{it} \tag{5}$$

where $i = 1,2,3,\ldots,N$ and $t = 1,2,3,\ldots T$, where N is the number of countries, and T the number of observations. From equation (5) one obtains the individual estimates of the slope β_i for each country by OLS. The MG estimator and its standard error are calculated as follows:

$$\hat{\beta}^{MG} = \overline{\beta} = \frac{\sum_{k=1}^{N} \hat{\beta}_k}{N} \tag{6}$$

$$se(\hat{\beta}^{MG}) = \frac{\sigma(\hat{\beta}_k)}{\sqrt{N}} = \sqrt{\sum_{k=1}^{N} \frac{(\hat{\beta}_k - \overline{\beta}\,)^2}{(N-1)}} \Big/ \sqrt{N} \tag{7}$$

Then inference can be undertaken in the usual manner. The above standard errors for the MG estimates (in contrast with the usual pooled standard errors) remain correct in the presence of the group-wise heteroskedasticity typical of panels and of autocorrelated disturbances which include the I(1) case also. The MG-based *t*-statistic is asymptotically distributed as a standard normal and exactly as a Student *t* with *N-1* degrees of freedom if the underlying estimated β_i sample is normal.

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[7] This heterogeneity is useful as factors ranging from economic fundamentals to demographics will

3.2 Cross sectional dependence

Thus far, the discussion has assumed cross sectional independence. However this is rather unrealistic in the PPP case for two reasons. On one hand, such dependence will be induced through the use of a common numeraire currency. On the other, the use of a common foreign price index will have a similar effect. This raises the question of whether our panel estimators (or a modified version thereof) are robust to cross sectional dependence. It turns out that the MG estimator can be modified to address this problem. In particular, two additional MG estimator versions are deployed that are based on obtaining the individual β_i estimates whilst treating with CSD.

One is a seemingly unrelated regression (SUR) system in which the individual β_i coefficients are estimated by a two-step FGLS procedure. The standard errors are calculated as in the simple MG estimator, with the resulting estimator named SUR-MG. The other approach builds on recent contributions that suggest augmenting the regression of interest by the cross-section means of the variables in order to capture the unobserved common macroeconomic variables or shocks that may induce the cross-section dependence (Pesaran, 2003a, 2003b). Accordingly, we obtain β_i by OLS in the RPPP regressions augmented by cross sectional averages of both nominal exchange rates and relative prices.

$$s_{it} = \alpha_i + \beta_i d_{it} + \gamma_i \bar{s}_t + \delta_i \bar{d}_t + u_{it} \tag{8}$$

where $d_{it} = p_{it} - p_{it}^*$ is the price differential, $\bar{d}_t = d_{it}/N$ is the cross sectionally averaged price differential and $\bar{s}_t$ is defined analogously.

The resulting cross-section augmented MG estimator is called CMG hereafter. Pesaran shows analytically that this estimator is consistent in a rather general setup. This includes the cases where the common factor can have I(0) or I(1) properties, can

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vary extensively across countries.

be correlated with the regressor and where it can heterogeneous effects on different panel members.[8]

Coakley, Fuertes and Spagnolo (2004) have supplemented the earlier Coakley et al. (2001) Monte Carlo simulations with others to examine the properties of the MG, SUR-MG and CMG estimators in the presence of cross sectional dependence. They consider DGPs with I(0) errors, I(1) errors and a mixture of both and the panel dimensions they employ are N=12 and T=84. Their results suggest that all three MG estimators are unbiased and that their standard errors are essentially correct. When explicitly accounting for CS dependence, the SUR-MG and CMG estimators result in efficiency gains versus the baseline MG estimator.

4. Data and results

4.1 Data and unit root tests

There are two datasets comprising of 2 different sets of currencies for 15 European economies. The first has nominal exchage rates quoted in terms of US dollars and consists of Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, and the UK. The second has spot rates in terms of German Marks, where Luxembourg replaces Germany in the panel estimation (Germany being the base currency). This second set contains all 15 European Union states plus Switzerland. The data are of monthly frequency and span the period 1977:01 to 2001:12 (yielding 300 observations). CPI data was gathered for all 17 countries over the same period. (15+US+Luxembourg).

Each dataset's dimensions exactly match those in the Coakley et al. (2001) Monte Carlo simulations. The implication is that one can be reasonably confident of

[8] One could deploy 2-way FE to capture a common factor. However, the FE approach assumes that the

the small sample properties when applying MG panel estimator to our data. The data collected are exclusively from Europe, as it is more likely that PPP will hold in a group of geographically contiguous economies such as those constituting Europe (Froot and Rogoff, 1994). This is because heterogeneity in their respective consumer (producer) price indices[9] is likely to be minimised and because barriers to trade are less likely.

Two unit root tests are applied to the variables in levels: the Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) (1988) tests. The results are given in Table 1.[10]

[Table 1 around here]

All 15 US dollar exchange rate series are found to be indistinguishable from I(1) processes compared to 10 for the DM series. The ADF price differential results are rather similar with a majority of series failing to reject the null..

Given the mainly I(1) nature of the variables, a next logical step is to test for a unit root in the residuals from the regressions in levels. Here the Augmented Engle-Granger (AEG) test for nonstationarity is used. These results are shown in Table 2.

[Table 2 around here]

These indicate that all the residuals for the US dollar series are indistinguishable from I(1) processes and so are 11 of the 15 DM series. In other words, exchange rates appear to be subject to permanent as well as temporary shocks. Such strong evidence of nonstationarity in the residuals negates a valid econometric interpretation of the slope coefficient from individual time series regressions. The results from all three

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latter has the same impact on all units.

[9] Of course this does not preclude heterogeneity in other aspects such as output and population.

[10] Results for the data in mean-differenced form are qualitatively similar and are available from the authors on request.

tests are consistent with our a priori expectations and are in keeping with those in the extant literature.

4.2 Panel regression results

Thes above unit root tests provide the basis for the nonstationary panel regression approach we employ for testing long run RPPP. Recall that the MG estimator is robust to I(1) or I(0) errors or indeed a mixture of both. Table 3 presents the MG panel regression results.

[Table 3 around here]

It shows the MG, SUR-MG, and CMG estimates for the slope coefficient, standard errors, and *t*-statistics for the null that the slope is 0 and 1 for both panels. In economic terms, null of 1 is the RPPP hypothesis that the long run relative price elasticity of exchange rates is unity. Table 3 also includes the *p*-values for the Kolmogorov-Smirnov (KS) normality test on the $\hat{\beta}_k$ series for each of the three MG estimates. We are unable to reject the null that the components of the MG panel estimates are normally distributed for all three specifications and so our inference procedures are valid in this context.[11]

For both the US dollar and DM series, all three panel estimators yield long run slope estimates that are significantly different from zero and insignificantly different from 1 at the 5% significance level for both one and two tailed tests.[12] Interestingly the point estimates between specifications vary considerably. For the US the baseline MG slope estimate is very close to 1 at 0.97 while those that accommodate cross sectional dependence are rather less so. The SUR-MG method yields a slope of 0.90

[11] Note that for the KS test a statistically significant result would reject normality

while the CMG approach produces the smallest slope estimate of 0.58. The DM series results are similar. The baseline MG slope estimate is 0.7687, SUR-MG yields a slope of 0.7122, and the CMG a slope of 0.5980. Although the CMG point estimates for both series are low, just below 0.60, their standard errors are around 0.2-0.25. The latter explains why the null cannot be rejected in either case – not even for a one-tailed test. Consequently, we conclude that long run RPPP holds in Europe.

Finally Table 3 gives the average of the absolute off-diagonal pairwise correlation of the terms in the regression residual correlation matrix, labelled Ω. The CMG average for the US dollar series is about half the value of the others but still not close to zero. We observe a much less dramatic but nevertheless substantial difference for the DM series.

The results found here complement those of Coakley et al. (2004) insofar as both find support for long run RPPP. The former study uses a larger panel comprising of OECD and developing countries separately over a period of 29 years. The panels used by Coakley et al. (2004) are comprised of series in which the underlying countries have more heterogeneous baskets of goods than those in our study. The point estimators in Coakley et al. are in many cases closer to the desired value of one and the null that the slope is one generally cannot be rejected with a smaller standard error than in this study.

[12] For both datasets under MG and SUR-MG specifications the null hypothesis that the slope coefficient is 1 is also not rejected for $\alpha = 0.1$ and the null hypothesis that the coefficient is 0 is rejected for $\alpha = 0.01$ for both one and two tailed tests.

5. Conclusions

This study investigates the long-run relative PPP hypothesis using a non-stationary panel regression approach for a panel of 15 European economies European region. This approach avoids the problem of spurious regression since the cross section dimension N is sufficiently large to overcome. Our mean group panel estimators take account of country heterogeneity and we deal with cross sectional dependence by using both the SUR-MG estimator and the novel CCE estimator proposed by Pesaran (2003a).

Our results indicate that all the residuals for the individual US dollar series are indistinguishable from I(1) processes and so are 11 of the 15 individual DM series for the 1977:1-2001:12 period. This vindicates the use of our nonstationary panel approach. This yields support for long-run relative PPP or that relative prices are reflected one-for-one in the nominal exchange rate depreciation. While this result holds for both US dollar and DM exchange rate series, the remaining residual cross sectional dependence seems less serious for the DM series.

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Table 1. ADF and PP test results

	US dollar series		DM series		Price differential	
	ADF	PP	ADF	PP	ADF	PP
Austria	-1.7639 (1)	-1.6639 (5)	-2.01998 (2)	-3.7066 (7)	-3.6775 (12)	-4.4435 (8)
Belgium	-1.5228 (1)	-1.4244 (6)	-2.65767 (0)	-2.5115 (7)	-3.4881 (12)	-2.1997 (11)
Denmark	-1.6356 (1)	-1.4975 (5)	-5.16326 (0)	-5.18758 (6)	-2.1469 (0)	-2.1377 (5)
Finland	-1.5988 (1)	-1.3930 (6)	-2.10897 (0)	-2.09558 (5)	-1.3044 (2)	-1.42736 (9)
France	-1.4727 (1)	-1.4537 (7)	-3.49077 (0)	-3.31883 (12)	-1.4712 (3)	-1.6786 (12)
Germany	-1.7561 (1)	-1.6530 (5)	-	-	-5.0227 (1)	-4.9003 (8)
Ireland	-1.4692 (1)	-1.3589 (6)	-2.19983 (0)	-2.21633 (4)	-2.4181 (9)	-2.5561 (12)
Italy	-1.2306 (1)	-1.1456 (7)	-2.5422 (0)	-2.53342 (1)	-4.9939 (1)	-5.1894 (9)
Luxembourg	-	-	-2.53296 (0)	-2.53296 (0)	-2.4257 (12)	-2.18998 (12)
Netherlands	-1.6916 (1)	-1.5657 (5)	-3.96879 (0)	-4.4735 (12)	-2.8831(12)	-4.7841 (4)
Norway	-1.5797 (1)	-1.1991 (1)	-2.31458 (0)	-2.29472 (2)	-1.7183 (12)	-1.2723 (9)
Portugal	-2.0043 (1)	-2.8763 (6)	-5.72137 (0)	-5.60794 (6)	-4.1263 (1)	-5.1580 (7)
Spain	-1.4999 (1)	-1.3995 (6)	-2.84612 (0)	-2.8301 (4)	-1.7024 (12)	-6.0234 (7)
Sweden	-1.3238 (1)	-1.1045 (5)	-2.15021 (0)	-2.08359 (2)	-1.7163 (0)	-1.7874 (4)
Switzerland	-2.5476 (1)	-2.3034 (2)	-3.64576 (1)	-2.97896 (8)	-5.1914 (0)	-4.9391 (7)
UK	-2.1051 (1)	-1.8322 (4)	-1.49248 (0)	-1.55309 (4)	-1.9410 (12)	-2.8843 (12)
95% Critical Value: -2.8710						

ADF(): denotes the number of lags as selected by the Schwarz Bayesian information criterion
PP(): denotes the optimal truncation lag for the Newey-West correction

Table 2. Augmented Engle-Granger test results

	US	DM
Austria	-1.5756 (1)	-2.409846 (2)
Belgium	-1.5048 (1)	-2.420447 (0)
Denmark	-1.5977 (1)	-2.747664 (0)
Finland	-1.5137 (1)	-1.4082 (0)
France	-1.2993 (1)	-2.485741 (0)
Germany	-1.5957 (1)	-
Ireland	-1.9175 (4)	-1.74306 (0)
Italy	-1.6810 (1)	-1.554747 (0)
Luxembourg	-	-3.405195 (0)
Netherlands	-1.7470 (1)	-3.83705 (0)
Norway	-1.7565 (1)	-1.83397 (0)
Portugal	-1.5625 (1)	-3.38353 (0)
Spain	-1.4141 (1)	-2.152242 (2)
Sweden	-1.0051 (1)	-2.070733 (0)
Switzerland	-2.1149 (1)	-3.644057 (1)
UK	-2.4850 (1)	-1.802574 (0)
Critical Value	-3.3587	

(): Denotes number of lags selected by the Schwarz-Bayesian information criterion

Table 3. Mean group estimates

Dollar Panel	MG	SUR-MG	CMG
$\hat{\beta}^{MG}$	0.9685	0.9025	0.5798
$se(\hat{\beta}^{MG})$	0.0965	0.0927	0.2219
t-statistic (β=0)	10.028	9.7329	2.6130
t-statistic (β=1)	-0.325	-1.051	-1.893
KS p-values	0.267	0.578	0.236
Abs(Corr), Ω	0.8581	0.8628	0.4202

DM Panel	MG	SUR-MG	CMG
$\hat{\beta}^{MG}$	0.7687	0.7122	0.5980
$se(\hat{\beta}^{MG})$	0.2363	0.2216	0.2429
t-statistic (β=0)	3.252991	3.214362	2.461853
t-statistic (β=1)	-0.97875	-1.29863	-1.65494
KS p-values	0.447	0.326	0.816
Abs(Corr), Ω	0.3345	0.3600	0.2348