The overvaluation of PPP in Europe?

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February 2004

Abstract

This paper tests for long run PPP using a nonstationary panel regression framework that can accommodate both permanent and temporary shocks. It also uses the common correlated estimator of Pesaran (2003a) to take account of cross sectional dependence. The PPP null in our framework is a unit elasticity of nominal exchange rates with respect to relative prices. Using US dollar and German mark spot rates and the consumer price index for 15 European economies 1977:1-2001:12, we cannot reject the hypothesis that the long run relative price elasticity of exchange rates is unity. While this result supports long run PPP in our European sample, it has to be viewed with caution since some residual cross sectional dependence remains.

Keywords: Real exchange rate; common correlated estimator; permanent shocks; cross sectional dependence.

JEL Classifications: C32; F31

1. Introduction

Purchasing power parity (PPP) states that there is a proportional relationship between prices - proxied by a representative basket of goods - in one country relative to those of another when expressed in the same numéraire currency. Although there are different concepts of PPP, 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 exchange rate determination, failure to support this parity empirically would somewhat undermine the basis for such models.[1]

The literature has gone through cycles of both supporting and rejecting PPP. Initially, when sufficient time series data became available after the advent of the floating exchange rates, 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 that panel tests of PPP are subject to the cross-section dependence or contemporaneous correlation problem by Higgins and Zakrajsek (1999), O'Connell (1998), Papell (2003) and Taylor (2003), and Wu and Wu (2001). They question the strong support for the PPP evidence found

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in the literature for the post-Bretton Woods period. They show that problems with the econometric testing techniques used lead to unreliable results. The typical results from tests that accommodate cross sectional dependence 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 allow for cross sectional dependence. This issue has been relatively neglected despite the seminal contribution of O'Connell (1998). One 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 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. The implication is that the SUR technique may not be appropriate for PPP studies.[2] We deal with this problem by employing the common correlated estimator (CCE) approach developed by Pesaran (2003a). This augments the PPP regression by cross sectional averages of both nominal exchange rates and relative prices to account for one omitted common factor.

The second contribution is it adopts a panel regression approach that can accommodate both permanent and temporary shocks. This approach is based on the insights of Kao (1999) and Phillips and Moon (1999) and (2000). They show that it is possible to overcome the spurious regression problem of pure time series by panel regressions. Coakley, Fuertes and Smith (2001) and Coakley, Fuertes and Spagnolo (2004) provide Monte Carlo evidence that such an approach can provide unbiased and

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[1] For recent reviews of the literature see Rogoff (1996), Sarno and Taylor (2002), and Taylor (2003). Others include Breuer (1994) and Bleaney and Mizen (1995).
[2] Nonetheless we employ the SUR estimator for comparative purposes since it is widely used in the PPP context.

consistent estimates of a long run coefficient in small samples. The motivation for this nonstationary regression framework is that both nominal exchange rates and relative prices seem persistent processes in finite samples. Moreover, real exchange rates are generally recognised as persistent processes. Accordingly 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 prices. The null hypothesis is a long run slope coefficient of unity or long run RPPP.

Finally this paper implements tests for long run relative PPP (RPPP) whereas the existing literature has tended to focus on tests for long run absolute PPP (APPP). Note that much of the existing literature adopts the unit root or cointegration approach to testing PPP. This literature is predicated on the presence of temporary (monetary) shocks only and it implicitly imposes the strong version of PPP. It tests long run APPP through examining the time series properties of the real exchange rate. If the real exchange rate is found to have a unit root, this implies a violation of APPP. Nonrejection of APPP implies RPPP, but not vice versa. As a result, the issue of whether the real exchange rate is mean reverting can be seen as less relevant in testing long run RPPP.

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 on 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.

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:

$$p_t = p_t^* + s_t$$

$$s_t = p_t - p_t^* \qquad (2)$$

There is a second concept relative PPP (RPPP). This can be illustrated by taking the differential of the logged nominal exchange rate and the other variables in (2)

$$ds_t / s_t = \left(dp_t / p_t - dp_t^* / p_t^* \right) \qquad (3)$$
$$(ds_t / s_t) / \left(dp_t / p_t - dp_t^* / p_t^* \right) = 1$$

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 panel regression approach.

2.1   Unit root studies

Generally, these studies found difficulty in rejecting the null hypothesis of a unit root [in the real exchange rate.] By implication, this finding constituted 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 that the real exchange rate was 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 to 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.[3]

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.[4] Further criticism is voiced by Engel (2000). The results of his Carlo experiments suggest that studies that use data with

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[3] 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.
[4] 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 (1994) 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.[5]

An alternative to long horizon datasets is to use panel data. The advantage of this method is that the power can be increased by adding series, and therefore it overcomes the need for long time horizons and their attendant problems. Two 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.[6]

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

Moving away from the unit root/cointegration literature, there is the regression test

approach to PPP. Much of the early work time-series studies were 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.[7]

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.

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[7] 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. 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):

$$s_t = p_t - p_t^* + u_t$$

$$q_t = u_t = s_t + p_t^* - p_t$$

(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 APPP in the long-run. Nonrejection of APPP 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.

This method of testing long run APPP 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.

$$\Delta s_t = \left(\Delta p_t - \Delta p_t^*\right) + \Delta u_t$$

(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. Note that if long-run APPP holds, equation (5) means that long-run relative PPP must hold also. However, the converse is not true.

While APPP 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).

$$s_{it} = \boldsymbol{a}_i + \boldsymbol{b}_i \left(p_{it} - p_{it}^* \right) + u_{it} \tag{6}$$

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 regression of I(1) variables, the absence of cointegration leads to the statistical problem of spurious correlation. 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

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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 √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 results in Coakley, Fuertes and Spagnolo (2004) extend the above to include cross sectional dependence.

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.[8]

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:

$$s_{it} = \boldsymbol{a}_i + \boldsymbol{b}_i\left(p_{it} - p_{it}^*\right) + u_{it} \tag{5}$$

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[8] This heterogeneity is useful as factors ranging from economic fundamentals to demographics will vary extensively across countries.

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 $ß_i$ for each country by OLS. The MG estimator and its standard error are calculated as follows:

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

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

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 $ß_i$ sample is normal.

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 $ß_i$ estimates by different approaches.

One is a seemingly unrelated regression (SUR) system in which the individual $ß_i$ coefficients are estimated by a two-step FGLS procedure. The standard errors are calculated as in the simple MG estimator. The resulting estimator is called 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 $ß_i$ by OLS in the RPPP regressions augmented by cross sectional averages of both nominal exchange rates and relative prices.

$$s_{it} = a_i + b_i d_{it} + g_i \overline{s}_t + d_i \overline{d}_t + u_{it} \tag{8}$$

where $d_{it} = p_{it} - p_{it}^*$ is the price differential, $\overline{d}_t = d_{it}/N$ is the cross sectionally averaged price differential and $\overline{s}_t$ 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 be correlated with the regressor and where it can heterogeneous effects on different panel members. [9]

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.

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[9] One could deploy 2-way FE to capture a common factor. However, the FE approach assumes that the latter has the same impact on all units.

4. Data and results

4.1 Data and unit root tests

We use data sets comprising of two different sets of currencies for 15 European economies. One has spot rates quoted in US dollars and consists of Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, and the UK. The other has spot rates quoted in German marks, where Luxembourg replaces Germany in the panel estimation (as Germany is the base country). This second sample 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). We use the consumer price index (CPI) as a proxy for prices.

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 the small sample properties when applying MG panel estimator to our data. The data collected are exclusively from Europe for the following reason. Froot and Rogoff (1994) hypothesised that PPP is more likely to hold in a group of geographically contiguous economies such as those constituting Europe. This is because heterogeneity in their respective consumer (producer) price indices[10] 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.[11]

[Table 1 around here]

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

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

All 15 spot of the US exchange rate series are found to be indistinguishable from I(1) processes, compared to 10 from the DM series. The results from the price differentials reject the unit root null rather more often (7 of the 16 cases). [12]

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

[Table 2 around here]

This indicates that all the residuals for the US dollar series are I(1) processes, and 11 of the 15 for the DM series. 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 tests are consistent with our a priori expectations and are in keeping with those in the extant literature.

4.2   Panel regression results

The 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 coefficient is 0 and 1 for both panels. The latter null 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-

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[12] Here the claim of a series being indistinguishable from an I(1) process constitutes either the ADF or

Smirnov (KS) normality test on the $\hat{b}_k$ 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 across both panels, and so our inference procedures are valid in this context.[13]

For both the US dollar and German mark 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 This is true for both one and two tailed tests.[14] Interestingly the point estimates between specifications vary considerably. For the US dollar series 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 while the CMG approach produces the smallest slope estimate of 0.58. We interpret the DM in a similar way, with the baseline MG slope estimate of 0.77, and the SUR-MG yielding a slope of 0.71, and the CMG a slope of 0.60. Although the CMG point estimates that accommodate cross sectional; dependence for both countries are low, around 0.60, their standard errors are 0.20-0.25. This explains why the null cannot be rejected in either case – not even for the one-tailed test. Consequently, we can 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. It is seen for the US that the CMG average is about half the value of the others but still some distance from zero. It is clear that some residual cross sectional dependence remains for the dollar series since there are probably several omitted common factors in this case.

PP statistic accepting the null hypothesis.

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

[14] 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.

For the DM series we see a less dramatic but nevertheless substantial reduction in the average correlation. The CMG average absolute correlation is closer to zero (0.23) as compared with the SUR-MG average of 0.36 which is worse that the simple MG average. Intuitively, this is as expected since we would expect cross-section dependence to be less of a problem with a German mark series for Europe.

The results reported above suggest that cross-section dependence matters for tests of long run PPP. We note the contrast between the MG point estimates that are typically closer to the desired value of unity and the appreciably smaller point estimates from the CMG estimator allowing for one common factor. Thus far it is unclear how to incorporate more than one omitted common factor in the Pesaran (2003a) framework. If one could incorporate more than one common factor, then this would further reduce the average correlation. However the probable outcome then may be that strong relative PPP may not then be supported although our conjecture is that weak PPP would be supported It is in this sense that we refer to the overvaluation of PPP in Europe in the current literature.

5. Conclusions

This study investigates the long-run relative PPP hypothesis using a non-stationary panel regression approach can accommodate both permanent and temporary shocks alike. This estimator avoids the problem of spurious regression when T and N are sufficiently large. Our approach also accounts for both country heterogeneity and cross sectional dependence. In this panel regression framework relative PPP is interpreted as the hypothesis that relative price changes are reflected one-for-one in the nominal exchange rate changes in the long run. We find support for long-run relative PPP for a sample of 16 European economies for the period 1977-2001.

However, this conclusion only holds if we believe that cross sectional dependence has been dealt with well enough to avoid distorting the empirical results. Whilst we are able to avoid the problem of I(1) spurious regression in our panel, we are unable completely to resolve the cross sectional dependence problem. In future work it would be interesting to include more than one omitted common factor either within Pesaran's (2003a) CCE framework or using the principal components approach developed by Coakley, Fuertes and Smith (2004)

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Table 1. ADF and PP test results

	US Spot Rate		German Spot Rate		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 panel estimates

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

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