Testing for real interest rate parity using panel stationarity tests with dependence: a note*

Mariam Camarero† Jaume I University

Josep Lluis Carrion-i-Silvestre University of Barcelona

Cecilio Tamarit University of Valencia

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Abstract

This paper tests for real interest parity (RIRP) among the nineteen major OECD countries over the period 1978:Q1-2006:Q1 using both short and long-run definitions of interest rates. Once the independence hypothesis is rejected among these series, we test for RIRP using panel data unit root and stationarity tests based on common factor models that allow for pervasive forms of dependence. Our results indicate that there is no evidence in favor of the weak version of the RIRP since one of the common factors that have been estimated is non-stationary.

Key words: Real interest rate parity, economic integration, panel data unit root and stationarity tests, cross-section dependence

JEL classification: F32, F21, C32, C33.

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†Corresponding author: Department of Economics, Jaume I University, Campus Riu Sec, E-12071 Castellon (Spain). Phone: 34+964728595. Fax: 34+964728591. e-mail: camarero@eco.uji.es; http://www3.uji.es/~camarero

1 Introduction

The empirical literature that tests for the real interest rate parity (RIRP hereafter) is abundant and extends back to the pioneer papers of Mishkin (1984) and Cumby and Obstfeld (1984). The flurry of papers that have analyzed this topic has given mixed results but, in general, the short-run RIRP is overwhelmingly statistically rejected (Chinn and Frankel, 1995). The empirical literature has explained this result by the existence of non-traded goods and/or transaction costs (Goodwin and Grennes, 1994). However, recent financial and real sector integration is expected to reduce the deviations from uncovered interest parity and from purchasing power parity, the sum of which are the deviations from RIRP. Thus, the study of real interest rate differentials across countries either under the Bretton-Woods regime or under the present of floating exchanges that replaced it deserves further attention (Goldberg et al., 2004).

The aim of this empirical note is to test for RIRP among the major OECD countries over the period 1978:Q1-2006:Q1 using panel data unit root and stationarity tests. The main contribution of this study to the existing literature on RIRP is in terms of the econometric methodology. We pool data on real interest differentials between the United States and other major OECD countries as a panel, and then use panel data based statistics as a way to increase the power of the statistical inference controlling for cross-section dependence.

Starting with Levin et al. (2002), much work has also been done on testing for unit roots in panels, including the IPS test developed by Im et al. (2003) or the test proposed by Hadri (2000). However, it is worth to note that when there is cross-sectional dependence in the disturbances, these tests do not longer converge to a standard normal. Therefore, one of the major concerns about the application of panel data based statistics is the assumption of cross-section independence. This assumption is rarely found in practice, especially in a globalised economy where the shocks overpass country-borders. This is of special interest in our study, due to the inclusion of several EU countries in the panel data set, which are partially ruled by common governmental institutions. These facts question the validity of the independence assumption, which is tested here using the

Ng (2006) statistic. In order to overcome this criticism, we have applied some statistics that allow controlling for the presence of different kinds of cross-section dependence. In particular, we apply approximate common factor models, as suggested by Bai and Ng (2004b). The application of the panel techniques that we use allows us to disentangle the sources of non-stationarity and carry out the analysis both at a global and at an individual basis.

The remainder of the paper is organized as follows. Section 2 briefly describes the theoretical background. Section 3 presents the data, the test statistics and the econometric results. Finally, Section 4 concludes.

2 Theoretical issues

A standard derivation of the RIRP condition can be found in Moosa and Bhatti (1996). Starting with the Fisher equation for two countries and after using some algebra, we arrive to an expression for the RIRP in a univariate framework such as:

$$r_t - r_t^* = v_t = rid_t, \tag{1}$$

$t = 1, \ldots, T$, where r_t is the real interest rate, and the asterisk denotes foreign variable. In order to test for RIRP, we impose the cointegration vector (1,-1) on r_t and r_t^* and then test for the stationarity of the error term $\{v_t\}$. Since $\{v_t\}$ is assumed to be $iid\,(0, \sigma_v^2)$, the expected value of rid_t is zero. This procedure is effectively testing for mean reversion in the real interest differential, which implies verifying whether shocks to the series of rid_t dissipate and the series return to their long-run zero mean level. This goal can be accomplished by performing unit root and stationarity tests on the series of rid_t.

Now consider that rid_t follows a more general stochastic process. As in Ferreira and León-Ledesma (2007), the former equation can be represented as a pth-order autoregressive process, so that we get an expression suitable for unit root testing:

$$\Delta rid_t = a_0 + \delta rid_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta rid_{t-j} + \varepsilon_t. \tag{2}$$

The following possibilities arise from the estimation of the former ADF-type equation:

$$\delta = 0 \tag{3}$$

$$\delta < 0 \text{ and } a_0 = 0 \tag{4}$$

$$\delta < 0 \text{ and } a_0 \neq 0. \tag{5}$$

In (3) the series contain a unit root and rid_t follows a random walk with shocks affecting the variable on a permanent basis. This case is inconsistent with the RIRP hypothesis.

Conversely, if either (4) or (5) hold, rid_t is a stationary process, which means that deviations from the mean are temporary and the estimated root provides information on whether the rid_t is short-lived or persistent. In (4) the process converges to a zero mean and a *strong* definition of RIRP holds, while in (5) the process converges to a non-zero mean and the *weak* version of RIRP prevails. The equality of real interest rates can be derived from the uncovered interest parity (UIP) condition together with the purchasing power parity (PPP) hypothesis. The strict equality of real interest rates (strong version of RIRP) requires the fulfillment of PPP, as well as perfect substitutability of financial assets located in different countries. There is an abundant empirical evidence against the validity of PPP. If PPP does not hold, then real interest rates cannot be equalized. It is worth noticing that strong RIRP can be violated due to the existence of transaction costs, non-traded goods, non-zero country specific risk premia or different national tax rates, among others.

Finally, it would be interesting to analyse whether a long-run relationship between r_t and r_t^* exists, and then to estimate the cointegrating vector instead of imposing it to be (1, -1). However, in this case the analysis departs from the RIRP hypothesis, as testing for the RIRP requires equality of real rates – note that cointegration of real interest rates is a weaker requirement than the equality of real rates. Cointegration of the real rates means that the variances of two or a group of real interest rates are bounded, and in the long-run these rates approach an equilibrium. In the next sections we will test for the weak version of the RIRP imposing the cointegrating vector to be (1, -1).

3 Empirical methodology and results

The empirical literature on RIRP is quite abundant and diverse. Although some authors have been able to find supportive evidence for weak RIRP in OECD countries using panel data (Fuijii and Chinn, 2002) or non-linearities (Mancuso et al., 2003, Holmes and Maghrebi, 2004, Ferreira and León-Ledesma, 2007), the traditional time series unit root tests have not been able to provide satisfactory results and additional empirical refinement is needed. More specifically, we can find two different clusters of research based on the type of unit root test that are used. A first one would include those that apply classical univariate unit root tests (basically ADF- type) with non-conclusive results.[1]

This outcome can be explained by a commonly accepted flaw associated with standard unit root tests: the power of these tests tends to be very low when the root is close to one, especially in small samples (Shiller and Perron, 1985). Therefore, we can conclude that the traditional time series unit root tests did not provide satisfactory results and additional empirical refinements can be a useful line of research.

In an attempt to solve the above-mentioned problems, Moosa and Bhatti (1996) find that a series of alternative univariate unit root tests that are more powerful than the conventional ADF tests lead to more promising results. Some other authors try to find more accurate evidence by enlarging the sample period considered.[2] Other empirical studies have tried to increase the power of the unit root tests using recent tests developed for panel data. The main advantage of the panel tests is that they add the cross-section dimension and increase the amount of information for each time period. In this context, Wu and Chen (1998), Holmes (2002) and Baharumshah et al. (2005) have found more promising results using panel unit root tests. Notwithstanding, it is widely recognized that these

[1] See for instance Meese and Rogoff (1988) and Edison and Pauls (1993).

[2] Lothian (2000) uses annual data on real interest rate differentials over the long period 1791-1992 with mixed results.

tests have some flaws in terms of lack of power and size distortion in the presence of correlation among contemporaneous cross-sectional error terms (O'Connell, 1998). As mentioned above, cross-section independence is hardly found in practice, especially when using macroeconomic time series that derive from globalized financial markets. However, the interest rates linkages that exist, both by construction of the variables and as a result of an economic integration process among a group of countries, are usually neglected in this type of analysis. In this particular case, using panel methods may increase the power of the tests, but the commonly used assumption of cross-section independency would not adequately capture the actual cross-relations present in the data.

In this paper we present an alternative testing procedure to deal with the problem of cross-section dependence. We first suggest to compute the test statistic by Ng (2006) to assess whether time series in the panel are cross-section independent. In addition, the Ng (2006) test statistic provides some guide about the best way to model cross-section dependence. Since panel data unit root tests are known to be biased towards concluding in favor of variance stationarity when individuals are cross-section dependent – see O'Connell (1999) and Banerjee et al. (2004, 2005) – we proceed in a second stage to compute statistics that account for such dependence when required.

There are several alternative proposals formulated in the literature to overcome the cross-section dependency problem. First, Levin et al. (2002) suggest to compute the test removing the cross-section mean. Although simple, this implies assuming, quite restrictively, that cross-section dependence is driven by one common factor with the same effect for all individuals in the panel data set. Second, Maddala and Wu (1999) propose obtaining the bootstrap distribution to accommodate general forms of cross-section dependence. Third, Breuer et al. (2002) also propose a panel unit root test that allows for contemporaneous correlation among the errors. Separate null and alternative hypotheses are tested for each panel member using the information captured through the variance-covariance matrix in a system estimated within a SUR framework and the critical values are obtained by bootstrap methods.

More recently, Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) and

Bai and Ng (2004b) have suggested other proposals that are especially relevant. They assume that the process is driven by a group of common factors, so that it is possible to distinguish between the idiosyncratic component and the common component. Although there are differences among the methods proposed, their driving idea is similar.

Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) focus on the extraction of the common factors that generate the cross correlations in the panel to assess the non-stationarity of the series, while in Bai and Ng (2004b) the non-stationarity of the series can come either from the common factors, the idiosyncratic component or from both. Moreover, Pesaran (2007) and Phillips and Sul (2003) only consider the existence of one common factor, while in Moon and Perron (2004) and Bai and Ng (2004b) there can be multiple common factors. Finally, Bai and Ng (2004b) consider also the possibility of cointegration relationships among the series of the panel. Banerjee et al. (2004) stated that there is a tendency to over-reject the null of stationarity when cointegration is present. As the existence of cointegrating relations between interest rate series is a very plausible hypothesis in economic integrated areas, the proposal in Bai and Ng (2004b) is the best approach in our case. Moreover, Monte Carlo comparisons developed by Gengenbach et al. (2004) and Jang and Shin (2005) show that, for all the specifications considered in their simulation experiments, the test in Bai and Ng (2004b) has more power than those by Moon and Perron (2004) and Pesaran (2005), and better empirical size than that of Phillips and Sul (2003). Consequently, our analysis is based on the Bai and Ng's (2004) approach. Finally, we apply the Harris, Leybourne and McCabe (2005) test for the null hypothesis of stationarity, which also uses common factors to account for dependence.

To the best of our knowledge, cross-section dependence has only been taken into account for RIRP testing in three recent papers. The SURADF test of Breuer et al. (2002) has been used in Chan et al. (2007) for East Asian economies and in Kim (2006) for some OECD countries distinguishing between traded and non-traded goods. Singh and Banerjee (2005) have applied the Pesaran (2005) CADF test to emerging economies. Therefore, our analysis increases the empirical evidence on this topic.

We test the null hypothesis of stationarity in the real interest differential over the period 1978:Q1 to 2006:Q1 – i.e. post Bretton Woods and EMU era. We have chosen this period due to its relevance for the financial integration process both at a global and at a regional (i.e. European) level. In fact, it covers from the beginning of the European Monetary System (EMS) up to now. The advent of the flexible exchange rate regime in 1973 and the relaxation of capital controls in some major OECD countries had opposite effects on the degree of real interest rate convergence among these countries. However, for countries that belonged to the exchange rate mechanism (ERM) of the EMS, the relaxation of the capital controls in the 90s along with the lower variability of nominal and real exchange rates, as the member countries were increasingly coordinating their monetary policies, should be expected to lead to increasing long-run real interest rate convergence. All in all, under a system of flexible exchange rates (or under an adjustable peg system like the EMS), real interest rate equalization may not be obtained because of expectations about exchange rate changes and foreign exchange risk premia. However, this would be compatible with the fulfillment of the weak version of RIRP, and therefore with the stationarity of interest rate differentials, which can lead to cointegrating relationships in the cross-sections of our panel data set.

The sample includes quarterly data of money market interest rates, long-term bond yields and consumer prices for up to 18 OECD countries[3], being the US the numeraire. The data have been taken from the International Financial Statistics database of the IMF. We have chosen both short-term and long-term asset rates for the analysis because these rates reflect market forces better than deposits ones.[4] The short-run rates are T-bill rates when available for the whole period (Canada, UK and US) and call money rates otherwise. Unfortunately, data unavailability excludes from the analysis the short-run real interest rates from Luxembourg and New Zealand. The long-run rates are 10-year bond yields. It is generally accepted that results on RIRP depend crucially on the maturities considered.

[3]Namely, Australia, Austria, Belgium, Canada, Denmark, France, Germany, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Switzerland and the UK.

[4]While deposit rates are much more widely available, they are often subject to administrative controls and in many cases display little movement over prolonged periods, which renders them uninformative (Frankel et al., 2004).

The empirical literature is far more supportive of the RIRP at five to ten-year horizons while the RIRP hypothesis is decisively rejected with short-run data (Fuijii and Chinn, 2002). Therefore, our study compares the results using short-term horizon instruments with the long-term ones.

In addition, we allow for two different definitions of real interest rates, depending on whether they are ex-ante (RIRPEXA) or ex-post (RIRPEXPO). For the ex-ante real interest rate we have used the Hodrick-Prescott filter to extract the trend and cycle of inflation and obtain its expectation. For the ex-post real interest rate we have used the actual CPI annual variation.

3.1 Panel unit root and stationarity tests and dependence

We present empirical evidence in two stages. First, we test the assumption of independence using the Ng (2006) statistic and find evidence that points to the presence of cross-correlation amongst the time series. It is worth mentioning that strong-correlation is expected, as cross-section correlation arises almost by construction given that the RIRP series are defined using the same base country. However, instead of assuming in the analysis that the time series are cross-section dependent, we have preferred to check it to get rid of unproved priors. Then, we perform the panel data statistical analysis accounting for cross-section dependence.

3.1.1 Testing the null hypothesis of cross-section independence

In this subsection we test the null hypothesis of independence against the alternative hypothesis of correlation using the approach suggested by Ng (2006). Besides, this framework allows us to gain some insight on the kind of cross-section dependence in terms of how pervasive and strong is the cross-section correlation – see Ng (2006).

In brief, the procedure works as follows. First, we get rid of the autocorrelation pattern in individual time series through the estimation of an AR model – we use the *MBIC* criterion suggested by Ng and Perron (2001) to select the order of autoregressive correction with $p_{\max} = \left[12 * (T/100)^{1/4}\right]$, where $[\cdot]$ denotes the integer part, as the maximum order of the autoregressive. This allows us to isolate the cross-section regression from serial correlation. Taking the estimated residuals from the AR regression equations as individual series, we compute the absolute value of Pearson's correlation coefficients ($\bar{p}_j = |\hat{p}_j|$) sorted in ascending order for all possible pairs of individuals, $j = 1, 2, \ldots, n$, where $n = N(N-1)/2$, with N denoting the number of individuals. As a result, we obtain the sequence of ordered statistics given by $\left\{\bar{p}_{[1:n]}, \bar{p}_{[2:n]}, \ldots, \bar{p}_{[n:n]}\right\}$. Under the null hypothesis that $p_j = 0$ and assuming that the individual time series are normally distributed, $\bar{p}_j$ is half-normally distributed. Furthermore, let us define $\bar{\phi}_j$ as $\Phi\left(\sqrt{T}\bar{p}_{[j:n]}\right)$, where Φ denotes the cdf of the standard Normal distribution, so that $\bar{\phi} = \left(\bar{\phi}_1, \ldots, \bar{\phi}_n\right)$. Finally, let us define the spacings as $\Delta\bar{\phi}_j = \bar{\phi}_j - \bar{\phi}_{j-1}$, $j = 1, \ldots, n$.

Second, Ng (2006) proposes splitting the sample of (ordered) spacings at arbitrary $\vartheta \in (0, 1)$, so that we can define the group of small (S) correlation coefficients and the group of large (L) correlation coefficients – we have followed Ng (2006) and set the required trimming at 0.10. Once the sample has been split, we can proceed to test the null hypothesis of non correlation in both sub samples. Obviously, the rejection of the null hypothesis for the small correlations sample will imply also rejection for the large correlations sample as the statistics are sorted in ascending order. Therefore, the null hypothesis can be tested for the small, large and the whole sample using the Spacing Variance Ratio $(SVR(\eta))$ in Ng (2006), where $\eta = [\vartheta n]$ is the number of statistics in the small correlations group. Under the null hypothesis that a subset of correlations are jointly zero, $svr(\eta) = \sqrt{\eta}SVR(\eta)/\sqrt{\omega_q^2} \rightarrow^d N(0, 1)$, $\omega_q^2 = 2(2q-1)(q-1)/(3q)$, as $\eta \rightarrow \infty$.

As can be seen from Table 1, we can split the whole sample of spacings in two groups, where the break point is estimated at $\hat{\eta} = 12$ or $\hat{\eta} = 15$, depending on the maturity and the definition of real interest rates. The analysis indicates that except for the short-run RIRPEXPO, the $svr(\eta)$ statistic rejects the null hypothesis of non correlation when focusing on the whole sample of correlations. When we split the whole sample in two groups, we observe that the null hypothesis of non correlation is strongly rejected for the L group, whereas, with the exception of the long-run RIRPEXPO interest rate, it is not rejected for the S group. This leads us to conclude that some form of cross-section correlation is present amongst time series, so that it has to be accounted for when assessing the stochastic properties of the real interest rates.

In addition, the fact that the break point is estimated at the beginning – $\hat{\eta} = 12$ or $\hat{\eta} = 15$ – implies that the proportion of correlation coefficients that form the S group is small compared to the correlation coefficients in the L group. This indicates that pervasive cross-correlation is present amongst the time series in the panel data sets. In this case, approximate factor models as suggested in Bai and Ng (2004b) reveal as a good option to account for cross-section dependence in panels.

3.1.2 Panel data unit root and stationarity tests with cross-section dependence

From the different approaches in the panel literature to deal with cross-section dependence the one we consider is based on the approximate common factor models of Bai and Ng (2004b). This is a suitable approach when cross-correlation is pervasive, as the analysis with Ng (2006) has revealed. Furthermore, this approach controls for cross-section dependence given by cross-cointegration relationships, where time series in the panel might be cross-cointegrated – see Banerjee et al. (2004).

The Bai and Ng (2004b) approach decomposes the $rid_{i,t}$ time series as follows:

$$rid_{i,t} = D_{i,t} + F_t'\pi_i + e_{i,t},$$

$t = 1, \ldots, T$, $i = 1, \ldots, N$, where $D_{i,t}$ denotes the deterministic part of the model – either a constant or a linear time trend – F_t is a $(r \times 1)$-vector that accounts for the common factors that are present in the panel, and $e_{i,t}$ is the idiosyncratic disturbance term, which is assumed to be cross-section independent. Unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model. The estimation of the number of common factors is obtained using the panel BIC information criterion in Bai and Ng (2002), with a maximum of six common factors.

Table 2 reports the results of applying this method, which admit two interpretations. At first sight, for both the short-term and long-term interest rates, the ADF statistic computed from the idiosyncratic disturbance terms rejects the null hypothesis of unit root, while the procedure detects one non-stationary common factor in all cases – r_0 and r_1 denote the number of stationary and non-stationary common factors, respectively, so that $r = r_0 + r_1$. This result is not surprising if we bear in mind that the $rid_{i,t}$ time series are constructed using the same base country, i.e. the US in our case. Therefore, what we are capturing with this common factor is the US real interest rate that is common to all the $rid_{i,t}$ time series, which turns out to be non-stationary. Given that the common factor that has been detected is non-stationary, we conclude that there is not evidence in favor of RIRP neither in the short-run nor in the long-run, regardless the definition of inflation that is used.

The picture derived from the Bai and Ng (2004b) test is especially interesting since it allows us to discriminate between the different sources of non-stationarity in the series. The results obtained indicate that idiosyncratic shocks are stationary, hence, they do not affect the interest rates in the long-run. However, we have found that interest rates are led by one common stochastic trend. Thus, the non-stationarity lays on one common factor. This feature has been interpreted in Gengenbach et al. (2004) as a sign of the presence of cointegration in the cross-section among the series of the panel. Note that the existence of one common stochastic trend implies that the $rid_{i,t}$ time series cointegrate. This indicates that there is a high degree of integration among the international markets where, although the real interest rates of each country are non-stationary, they share the same stochastic trend.

It is possible to complement the analysis by testing the null hypothesis of stationarity with cross-section dependence using the $\hat{S}_k^F$ statistic of Harris et al. (2005). Their statistic is given by $\hat{S}_k^F = \left(\hat{C}_k + \hat{c}\right) \Big/ \hat{\omega}\left\{\hat{a}_{k,t}\right\}$, with $\hat{C}_k = T^{-1/2} \sum_{t=k+1}^{T} \hat{a}_{k,t}$ the autocovariance of order k, $\hat{a}_{k,t} = \sum_{i=1}^{N+r} \hat{z}_{i,t}\hat{z}_{i,t-k}$, and $\hat{z}_{i,t}$ as the ith element of the $(N + \hat{r}) \times 1$ vector $\left(\hat{F}_{1,t}, \ldots, \hat{F}_{\hat{r},t}, \hat{e}_{1,t}, \ldots, \hat{e}_{N,t}\right)'$, which contains the estimated common factors and the idiosyncratic disturbance terms obtained as described above. $\hat{c} = (T-k)^{-1/2} \sum_{i=1}^{N} \hat{c}_i$, being

$\hat{c}_i$ a correction term defined in Harris et al. (2005) and, $\hat{\omega}^2\{a_t\}$ is a consistent estimate of the long-run variance of $\{a_t\}$. Under the null hypothesis of joint stationarity in both common factors and idiosyncratic disturbance terms, $\hat{S}_k^F \to^d N(0,1)$.

Table 2 reports the $\hat{S}_k^F$ statistics. These results reveal that the null hypothesis of stationarity is not rejected at the 5% level of significance only for the short-run ex-post interest rate, whereas it is rejected for the other cases. Note that for the evidence drawn from the $\hat{S}_k^F$ statistic to be coherent with the results obtained using Bai and Ng's procedure, we should reject the null hypothesis of stationarity, since we have found non-stationary common factors. As the $\hat{S}_k^F$ statistic takes both the idiosyncratic and common factor components into account, the presence of non-stationary common factors should lead to the rejection of the null hypothesis of the $\hat{S}_k^F$ statistic. One explanation for the non-rejection of the null hypothesis of joint stationarity for the short-run RIRPEXPO interest rates might be that the size (or proportion) of the non-stationary component is smaller than the stationary one. In this case, stationarity test statistics will have low power – see Bai and Ng (2004a).

To sum up, our analysis has revealed that the RIRP hypothesis is not satisfied due to the non-stationarity of the common factor detected, even though the idiosyncratic disturbance terms are found to be stationary. According to White and Woodbury (1980) and Holmes (2005), the existence of common factors reflects the high degree of financial markets integration achieved in the OECD countries, as we have found that there is only one stochastic trend that governs the different interest rates. This is not surprising due to the process of ongoing financial integration that started in the 80s in the OECD. In fact, during the 80s and 90s there was an increasing opening up of the financial markets in OECD countries together with an important innovation process (new markets and instruments) that helped financial integration. For almost three decades the OECD countries have taken steps to promote economic efficiency by liberalizing their domestic financial systems and removing restrictions on capital flows. Financial liberalization efforts in these countries followed almost the same pattern and took place primarily in two stages. In the first stage, foreign exchange controls, as well as the ceilings on deposits and lending rates were progressively removed, though at different times. The second stage of the liberalization process witnessed the opening up of the capital accounts during the late 80s. Once national markets have been deregulated, real interest differentials between two national markets should be close to those found in the Eurocurrency markets. In that case, average real interest rates will differ only if nominal returns in one currency are consistently higher than in another (i.e., UIP does not hold) or if the relative prices consistently diverge between the two countries (i.e., PPP does not hold). Relative financing costs will no longer depend on the peculiar features of national loan markets shielded from international competition.

4 Conclusions

Many studies have reexamined the real interest rate parity condition and found rather hard to establish its fulfillment empirically. In this paper we present new evidence showing that RIRP is not satisfied either in the short-run or in the long-run for a group of OECD countries. We examine the behavior of cross-country real interest rate differentials for the US and eighteen other major industrial economies from 1978:Q1 to 2006:Q1. Our analysis is based on the use of panel data unit root and stationarity test statistics that accommodate the presence of strong and pervasive cross-section dependence that has been found. Taking into account dependence is important to overcome potential biases in statistical inference that will lead to conclude in favor of the stationarity hypothesis. By exploiting the cross-section information, these tests have higher power relative to the classical unit root and stationarity tests.

Cross-section dependence has been modelled through the use of common factor models. This has allowed us to shed light on the source of non-stationarity. Thus, we have detected the presence of one non-stationary common factor, i.e. one common stochastic trend, while the idiosyncratic disturbance terms are found to be stationary. This means that the non-stationarity of the real interest rates differentials is due to the presence of one common stochastic trend, which can be interpreted as a sign of high market integration. In our case, the non-stationary common factor represents the US real interest rate as, by definition, the differential of real interest rates has been computed relative to this benchmark variable. The presence of this non-stationary common factor implies that RIRP is not satisfied, although it shows that real interest rate differentials share one non-stationary common factor. We can interpret this result as a consequence of the increasing opening up of the financial markets in OECD countries during the 80s and the 90s, together with an important innovation process in the form of new markets and instruments that helped financial integration. This outcome is not surprising in a highly integrated area as the OECD, where there is an increasing synchronization of business cycles. In this context, tests that do not impose independency across the panel are more adequate in empirical research on economic integration.

How can we interpret the result that the real interest differentials between the US and other OECD countries are non-stationary? We must recall that there is no direct economic mechanism that ensures the equality of real interest rates. So RIRP relies on its two underlying components: (nominal) uncovered interest parity and PPP assuming that the two deviations almost exactly cancel out. If deviations from UIP and PPP are driven by a common factor, e.g. exchange rate forecast errors, then this cancelling out is to be expected. On the other hand, if each deviation is driven by independent factors, deviations from UIP driven by risk premia and deviations from PPP driven by real trade factors – such as secular changes in competitiveness – then this cancelling out could only have occurred by chance. Unfortunately, distinguishing between these two possibilities lies beyond the scope of this paper and is left for future research.

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Table 1: Spacing Variance Ratio statistic for the RIRPINF and RIRPHP panels

	Short-run real interest rates						
	Whole sample		Small group			Large group	
	$svr(\eta)$	p-val	$svr(\eta)$	p-val	$\hat{\eta}$	$svr(\eta)$	p-val
RIRPEXPO	-2.485	0.994	-0.776	0.781	12	4.449	0.000
RIRPEXA	2.794	0.003	0.834	0.202	12	8.612	0.000

	Long-run real interest rates						
	Whole sample		Small group			Large group	
	$svr(\eta)$	p-val	$svr(\eta)$	p-val	$\hat{\eta}$	$svr(\eta)$	p-val
RIRPEXPO	3.025	0.001	0.209	0.417	15	8.372	0.000
RIRPEXA	4.669	0.000	2.027	0.021	15	6.851	0.000

Table 2: Panel data statistics based on approximate common factor models

Panel A: Short-run real interest rates

Bai and Ng (2004b) statistics

	RIRPEXPO		RIRPEXA	
	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-5.150	0.000	-5.233	0.000
	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-7.646	1	-7.576	1
MQ test (non-parametric)	-10.777	1	-11.520	1

Harris et al. (2005) statistic

	RIRPEXPO		RIRPEXA	
	Test	p-value	Test	p-value
$\hat{S}_k^F$	0.545	0.293	2.157	0.016

Panel B: Long-run real interest rates

Bai and Ng (2004b) statistics

	RIRPEXPO		RIRPEXA	
	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-4.822	0.000	-2.246	0.012
	Test	$\hat{r}_1$	Test	$\hat{r}_1$
MQ test (parametric)	-8.166	1	-9.980	1
MQ test (non-parametric)	-9.719	1	-11.102	1

Harris et al. (2005) statistic

	RIRPEXPO		RIRPEXA	
	Test	p-value	Test	p-value
$\hat{S}_k^F$	2.144	0.016	2.011	0.022