DOCUMENT DE TREBALL

CREAP2006-14

NEW EVIDENCE OF THE REAL INTEREST RATE PARITY FOR OECD COUNTRIES USING PANEL UNIT ROOT TESTS WITH BREAKS

Mariam Camarero; Josep Lluis Carrion-i-Silvestre
and Cecilio Tamarit

CENTRE
DE REFERÈNCIA
EN ECONOMIA APLICADA

NEW EVIDENCE OF THE REAL INTEREST RATE PARITY FOR OECD COUNTRIES USING PANEL UNIT ROOT TESTS WITH BREAKS

Mariam Camarero*; Josep Lluis Carrion-i-Silvestre†
and Cecilio Tamarit‡

Abstract: This paper tests for real interest parity (RIRP) among the nineteen major OECD countries over the period 1978:Q2-1998:Q4. The econometric methods applied consist of combining the use of several unit root or stationarity tests designed for panels valid under cross-section dependence and presence of multiple structural breaks. Our results strongly support the fulfillment of the weak version of the RIRP for the studied period once dependence and structural breaks are accounted for.

Key words: Real interest rate parity, economic integration, panel data unit root tests, structural breaks, cross-section dependence.

JEL Codes: C32, C33, F21, F32

Resum: En aquest article es duu a terme el contrast de la hipòtesi de paritat entre tipus d'interès reals (RIRP) entre dinou dels majors països de la OCDE al llarg del període 1978:Q2-1998:Q4. La metodologia economètrica aplicada es basa en la combinació de l'ús de diversos contrastos d'arrel unitària i d'estacionarietat dissenyats per un entorn de dades de panell que són vàlids sota la presència de dependència entre els individus i presència de múltiples canvis estructurals. Els nostres resultats donen fort suport al compliment de la versió dèbil de la RIRP en el període analitzat un cop la dependència i els canvis estructurals són tinguts en compte.

Paraules clau: Paritat del tipus d'interès real, integració econòmica, contrastos d'integració de dades de panell, canvis estructurals, dependència entre individus.

Classificació JEL: C32, C33, F21, F32

* Department of Economics, Jaume I University, Campus Riu Sec, 12071 Castelló, Phone: +34964728595. Fax: +34964728591. E-mail: camarero@eco.uji.es

† Research Group AQR-IREA, Department of Econometrics, Statistics and Spanish Economy, University of Barcelona, Av. Diagonal, 690, 08034 Barcelona, Phone: +34934021826; Fax: +34934021821, E-mail: carrion@ub.edu

‡ Department of Applied Economics II, University of Valencia., 46071 Valencia, Phone: +34963828349; Fax: +34963828354, E-mail: cecilio.tamarit@uv.es

1 Introduction

With the widespread removal of regulations and closer integration of international financial markets, global movements of interest rates have become increasingly linked. Therefore, the analysis of the extent to which real interest rates are equalized across countries is a matter of increasing interest to researchers for various reasons. First, in an open economy, real interest rates are an important channel for transmission of macroeconomic policies. Secondly, the degree of fulfilment of the real interest rate parity (RIRP) can be used as a criterion to measure market integration because RIRP requires efficiency in the goods market (via ex-ante purchasing power parity) and efficiency in the assets markets (via ex-ante uncovered interest parity). Thirdly, RIRP is also important because it is an assumption in several monetary models of exchange rate determination (i.e. Frenkel (1976)).

The empirical literature testing for the RIRP hypothesis is abundant and extends back to the pioneer papers of Mishkin (1984) and Cumby and Obstfeld (1984) giving mixed results, but, in general, short-run RIRP is overwhelmingly statistically rejected (Chinn and Frankel, 1995). Although casual observation suggests that international markets have become increasingly integrated, the formal empirical literature in economics and finance indicates that integration remains incomplete due to the existence of non-traded goods 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 is 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, Lothian and Okunev, 2002).

The aim of this paper is to test for RIRP among the major OECD countries over the period 1978:Q2-1998:Q4 using univariate and panel data unit root and stationarity tests. The econometric methods applied consist of combining the use of univariate and multivariate unit root tests with good size and power properties. Thus, the main contribution made by this study to the literature on RIRP is in terms of the econometric methodology. First, we analyze stochastic properties of RIRP's from a univariate point of view using standard unit root and stationarity tests. Second, we use panel data based statistics as a way to increase the power of the statistical inference. One of the major concerns about the application of panel data based statistics is the assumption of cross-section independence. In order to overcome this criticism, we have computed those statistics in the literature that allows controlling for the presence of different kinds of cross-section dependence. Besides, some authors have highlighted the importance of structural breaks in influencing the outcome of RIRP tests.[1] In order to overcome these flaws, we propose

[1] See, for instance, Fountas and Wu (1999), Holmes (2002) or Lai (2004).

a testing strategy aiming at accounting for both dependence and multiple and heterogeneous structural breaks in panels.

The remainder of the paper is organized as follows. Section 2 briefly presents the theoretical background. Section 3 reviews the previous relevant literature. Section 4 presents the data, test statistics and the econometric results. Finally, Section 5 concludes.

2 Theoretical issues

The extent of deviation from RIRP is a measure of the lack of goods and financial market integration. This can be seen more clearly by deriving the RIRP condition from its components. To do so, we use a standard presentation, as in Moosa and Bhatti (1996), starting with the Fisher equations for two countries, the domestic country and the foreign one. These equations can be written as follows

$$r^e_{t+1} = i_t - \pi^e_{t+1}, \tag{1}$$

and

$$r^{*e}_{t+1} = i^*_t - \pi^{*e}_{t+1}, \tag{2}$$

where r is the real interest rate, i is the nominal interest rate, π is the inflation rate, the superscript e indicates the expected value of the underlying variable, and the asterisk denotes foreign variables. If the Fisher equations (1) and (2) are jointly valid then

$$r^e_{t+1} - r^{*e}_{t+1} = (i_t - i^*_t) - (\pi^e_{t+1} - \pi^{*e}_{t+1}). \tag{3}$$

The fulfilment of *ex-ante* RIRP entails the joint hypothesis of the uncovered interest parity (UIP) and *ex-ante* instantaneous relative PPP to hold. Both conditions are stated in equations (4) and (5), respectively

$$ds^e_t = i_t - i^*_t \tag{4}$$

$$ds^e_t = \pi^e_{t+1} - \pi^{*e}_{t+1}, \tag{5}$$

where s is the spot exchange rate defined as the number of units of domestic currency per unit of the foreign one. Combining equations (3), (4) and (5), we obtain

$$r^e_{t+1} = r^{*e}_{t+1}. \tag{6}$$

Let us assume that expectation are formed rationally across countries, then the actual ex-post real interest rate realized at time $t+1$ will differ from the ex-ante real interest rate by a random term that is serially uncorrelated with a zero mean. This can be formally written as

$$r_{t+1} = r_{t+1}^{e} + u_{t+1}, \tag{7}$$

and

$$r_{t+1} = r_{t+1}^{*e} + u_{t+1}^{*}. \tag{8}$$

Substituting equations (7) and (8) into equation (6), we get

$$r_{t+1} - r_{t+1}^{*} = v_{t+1}, \tag{9}$$

where $v_{t+1} = u_{t+1} - u_{t+1}^{*}$. We can easily transform the expression (9) using the backward shift operator

$$r_t - r_t^{*} = v_t = rid_t. \tag{10}$$

Equation (10) can be used to test RIRP in a univariate framework imposing the cointegration vector (1,-1) and testing for the stationarity of the error term $\{v_t\}$. Since $\{v_t\}$ is assumed to be $iid\,(0, \sigma_{\varepsilon}^{2})$, the expected value of the *rid* is zero. This procedure is effectively testing for mean reversion in the real interest differential, that is, to verify whether shocks to the series of *rids* dissipate and the series return to their long-run zero mean level. This objective can be accomplished by performing unit root tests on the series of *rids.*

Now consider that rid_t follows a more general stochastic process:

$$rid_t = a_0 + a_1 rid_{t-1} + \varepsilon_t. \tag{11}$$

Following Ferreira and León-Ledesma (2003), the former equation can be represented as a pth-order autoregressive process,

$$\Delta rid_t = a_0 + \delta rid_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta rid_{t-j} + \varepsilon_t, \tag{12}$$

where $\delta = \sum_{j=1}^{p} a_j - 1$. The next possibilities arise from the estimation of the former ADF-type equation:

$$\delta > 0 \tag{13}$$

$$\delta = 0 \tag{14}$$

$$\delta < 0 \text{ and } a_0 = 0 \tag{15}$$

$$\delta < 0 \text{ and } a_0 \neq 0. \tag{16}$$

Inequality (13) accounts for the case in which the parameter δ is statistically greater than zero. The path of *rids* in this case would be explosive and the series would not converge to any mean in the long-run. In (14) the series contain a unit root and *rids* follow a random walk with shocks affecting the variable on a permanent basis. Both cases, random walks and permanent or explosive *rids* are inconsistent with the RIRP hypothesis

On the contrary, if either (15) or (16) hold, (11) is a stationary process, which means that deviations from the mean are temporary and the estimated root provides information on whether the *rid* is short-lived or persistent. In (15) the process converges to zero mean and a *strong* definition of RIRP holds while in (16) the process converges to non zero mean and a *weak* version of RIRP prevails. It is worth noticing that strong RIRP can be violated, among others, due to the existence of transaction costs, non-traded goods, non-zero country specific risk premia or different national tax rates.

3 Previous empirical literature

The empirical literature on RIRP is quite abundant and diverse depending on the purpose of the analysis. Consequently, an extensive review of the subject is far beyond the scope of the present section. However, it might be useful to distinguish, at least, between two different groups of studies according to their primary objective.

First, an initial group of papers could be classified as indirect evidence of RIRP as they analyse the implications of other hypotheses or theories that are connected to the fulfillment of RIRP. This literature includes research on the analysis of the international monetary policy transmission and the currency dominion hypothesis (i.e. Chinn and Frankel (1995) for the Pacific Rim case[2]), the existence of time-varying risk premia on foreign exchange series (Nieuwland, Verschoor and Wolff, 1998), the impact on UIP (McCallum, 1994), the efficiency of exchange rate market (MacDonald and Moore, 2001) or the international Fisher effect (Fraser and Taylor, 1990).

A second string of the literature is devoted to directly verify the RIRP hypothesis making use of different econometric methods. As already mentioned, the early studies (Mishkin, 1984 or Cumby and Obstfeld, 1984) were direct tests of real interest rate equality that performed classical OLS regression analysis and found evidence inconsistent with complete financial integration. Other studies have found hints of increasingly strong real interest linkages by comparing either summary statistics or regression coefficients considering different subsamples of the data (i.e. Glick and Hutchison, 1990 or Marston, 1995). More recent studies have applied cointegration and unit root techniques (Goodwin and Grennes, 1994 or Wu and Fountas, 2000), time-varying parameters (Cavaglia, 1992) panel data (Fuijii and Chinn, 2002) or non-linearities (Ferreira and León-Ledesma, 2003, Holmes and Maghrebi, 2003 and Mancuso, Goodwin and Grennes, 2003) finding more supportive evidence for weak RIRP for various OECD and Asian countries.

[2] Recent papers on the subject include Wu and Fountas (2000), Chinn and Frankel (2003) and Frankel, Schmukler and Servén (2003).

Within this second group of studies of direct testing for RIRP, an alternative empirical approach to which the present paper contributes has focused on the use of unit root tests. We can find two different clusters of research based on the type of unit root test used. A first one would include the papers that apply classical univariate unit root tests (basically ADF- type) with non-conclusive results. This outcome can be explained by two commonly accepted flaws that appear with standard unit root tests. First, 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). Second, a serious problem is that the standard tests are biased towards the non-rejection in the presence of structural breaks.

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 enlarging the sample period considered.[3] Nevertheless, as long as we extend the sample period a new set of problems arises linked to discontinuities in the series generated either by shocks or institutional changes.[4] All in all, we can conclude that the traditional time series unit root tests did not provide satisfactory results and additional empirical refinement can be a useful line of research. Bearing the above mentioned in mind, a second group of empirical studies try to increase the power of the unit root tests using the 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) and Holmes (2002) have found more promising results using Levin, Lin and Chu (2002), Maddala and Wu (1999), and Im, Pesaran and Shin (2003) panel unit root tests. Notwithstanding, it is widely recognized that these tests have some flaws in terms of lack of power[5] and size distorsion in the presence of correlation among contemporaneous cross-sectional error terms (O'Connell, 1998).

In this paper we present a testing procedure that overcomes previous problems common in panel unit root tests. We contribute to the empirical literature on the RIRP on various respects. First, we consider the presence of multiple structural changes that might be affecting the series. Additionally, we tackle the issue of cross-section dependence when computing panel data based statistics.

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

[4] Fountas and Wu (1999), and Goldberg, Lothian and Okunev (2002) apply unit root tests that allow for structural breaks in the series finding rejection of the null in more cases.

[5] Especially in the case of the Levin, Lin and Chu (2002) test due to the restrictiveness of the alternative hypothesis. Although this test has high statistical power relative to the conventional single-equation unit root test, the major criticism is that it requires the coefficient (ρ) of the lagged dependent variables to be homogeneous across all cross-section units of the panel, which suggests that all series revert to their respective unconditional mean over time at the same rate. This flaw has been overcome through the Im, Pesaran and Shin (2003) test, which allows for a greater degree of heterogeneity across cross-sectional units.

4 Empirical methodology and results

In this paper we test the null hypothesis of unit root in the real interest differential over the period 1978:Q2 to 1998:Q4 – i.e. post Bretton Woods and pre-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 up to the launching of the euro. The sample includes quarterly data of money market interest rates, long-term bond yields and consumer prices for up to 19 OECD countries: Australia, Austria, Belgium, Canada, Denmark, France, Germany, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Switzerland and UK, as well as US, which is defined as the numeraire. The countries have been selected depending on the span of data availability through various exchange rate regimes and their outstandig role within the industrialized economies. 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.[6] 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. At five to ten-year horizons the empirical evidence becomes far more supportive, while the RIRP hypothesis is decisively rejected with short horizon data (Fujii and Chinn, 2000). 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 (RIRPHP) or ex-post (RIRPINF). For the ex-ante real interest rate we have used the Hodrick-Prescott filter to extract the trend and cycle of inflation to obtain its expectation. For the ex-post real interest rate we have used the actual CPI annual variation.

Although RIRP is an *ex-ante* concept involving expected rather than actual inflation, most of the empirical studies use *ex-post* variables mainly because expected inflation rates are unobservable. In order to asses the sensitivity of the results to the (ex-ante or ex-post) nature of the variables, we both use quarterly *ex-post* and *ex-ante* estimates of real rates of return on short-term securities. There are two alternative ways to estimate real interest rates. In the first one, practitioners either use survey data to measure expected inflation (i.e. Tanzi, 1985) or simulate data using different methods.[7] In this paper we

[6] 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, Schmukler and Servén, 2003).

[7] Evans (1985) use some macro variables as *proxies*, Plosser (1987) and Barro and Sala-i-Martin (1991) generate inflationary expectations using AR models, Reichenstein and Elliot (1987) use P*-type monetary models of inflation expectations and other authors, like King and Rebelo (1993) use statistical filters to extract low frecuency components.

applied the Hodrick and Prescott (1997) filter (HP hereafter) to proxy price expectations over a time horizon as this filter exhibits the ideal statistical properties for this purpose (Hodrick and Prescott, 1997).[8] Therefore, the ex-ante real interest rate is approximated by we have used the Hodrick-Prescott filter to extract the trend and cycle of inflation to obtain its expectation. Alternatively, most of the researchers assume that expected inflation equals realized inflation (plus a white-noise error term). The use of realized inflation as an unbiased measure of expected inflation lies on the assumption of rational expectations and perfect forecastability. If RIRP holds and inflation forecast errors are random, then the observed real interest differential should be random as well. In this case, we can test RIRP by determining whether real interest differentials are systematically related to variables in the current information set.

Concerning the empirical methodology, we have applied sequentially a variety of panel stationarity and unit root tests following a three-step testing strategy that addresses the problem related to the issues of dependence.

First we apply those tests assuming – unrealistically in this particular case – cross-section independence. Cross-section independence is hardly found in practice, especially when using macroeconomic time series that derive from globalized financial markets, as it is the case with interest rates. As 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 (1998) and Banerjee, Marcellino and Osbat (2004, 2005) – the issue of cross-section dependence is of great importance. Thus, as the second stage of the analysis, we suggest computing the test statistic by Ng (2006) to assess whether the individuals in the panel are cross-section independent. This statistic is quite convenient since, despite of testing the null hypothesis of cross-section independence, it provides guidance about the best way to model cross-section dependence.

Finally, the last step in the testing strategy is to compute statistics that account for such dependence when required. For this purpose, we report results using two different approaches to allow for cross-section dependence.[9] First, Maddala and Wu (1999) propose obtaining the bootstrap distribution to accommodate general forms of cross-section dependence. Second we use approximate common factor models, as suggested by Bai and Ng (2004).

The application of this testing strategy distinguishes two different approaches de-

[8] However, some authors have claimed that these methods are always biased. During inflation episodes realized real rates tend to be less than the real rate calculated using inflation forecast, and conversely, when inflation falls, the realized real rate lies well above the predicted real rate (Darin and Hetzel, 1995).

[9] Additionally, we have computed the tests following the approach suggested in Levin, Lin and Chu (2002) that is to remove the cross-section mean. Although simple, this proposal implies the restrictive assumption that cross-section dependence is driven by one common factor with the same effect for all individuals in the panel data set. Moreover, Strauss and Yigit (2003) show that demeaning does not eliminate cross-section dependence due to contemporaneous correlation and, more importantly, when applied it induces false inference. Notwithstanding, results based on cross-section demeaning are available upon request from the authors.

pending on whether panel statistics allow for structural breaks. Previous literature has revealed that there might be some events that affect real interest rates in a permanent way. It is well known that not accounting for structural breaks biases both unit root and stationarity tests towards concluding in favor of non-stationarity in variance.[10] Thus, this feature should be of special interest in our case, since variables like interest rates have been affected by major events such as currency crises or economic integration processes during the period analyzed. In addition, proceeding in this fashion accounts for the existence of a tension between cross-section dependence and misspecification concerning the presence of structural breaks: the former introduces a bias towards stationarity in variance while the bias due to the latter goes in the opposite direction. This feature implies that the empirical analysis of the RIRP should be addressed carefully to avoid the effects of this tension.

4.1 Panel tests without structural breaks

In this section we conduct the standard analysis where individual time series are assumed not to be affected by structural breaks. We present empirical evidence in three stages. First, we compute the statistics under the assumption that the individuals are cross-section independent. Second, we test this assumption using the Ng (2006) statistic and find evidence that point to the presence of cross-correlation amongst the individuals. Finally, we perform the statistical analysis accounting for cross-section dependence.

Before presenting the results for the panel data statistics, we have computed the individual ADF statistics, which are reported in Table 1 – we have specified the deterministic component as a constant term and use the t-sig criterion in Ng and Perron (1995) 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 analysis is complemented with the computation of stationarity test proposed in Kwiatkowski, Phillips, Schmidt and Shin (1992) – hereafter, KPSS test. The selection of the constant term as the deterministic component obeys to economic theory as well as to visual inspection of the time series. Thus, if time series were to be stationary, they would be better described by stationary fluctuations around a mean different from zero. This implies that we are focusing on the weak version of the real interest rate parity.

From the results in Table 1, when we focus on the short-run real interest rates the null hypothesis of unit root is only rejected at the 5% level of significance for Austria. In contrast, the null hypothesis of variance stationarity is only rejected for Portugal at the 5% level by the KPSS statistic. Note that these results are achieved irrespective of whether we use the ex-ante or ex-post definitions of real interest rates.

[10] See Perron (1989), Montañés and Reyes (1998), and Lee, Huang and Shin (1997), among others, for univariate statistics, and Carrion-i-Silvestre, del Barrio and López-Bazo (2001) for panel data statistics

When analyzing the long-run interest rates, we are able to find some evidence in favor of the RIRP for the case of the RIRPINF variable, as the unit root is rejected for seven individuals at the 5% level of significance. The null hypothesis of unit root cannot be rejected for the RIRPHP in any case. These results indicate that there is strong evidence against the RIRP hypothesis. However, the converse is found when using the individual KPSS statistic, since the null hypothesis of stationarity is rarely rejected even at the 10% level for both variables.

These contradictions between unit root and stationarity statistics might be due to the fact that information contained in single time series is not enough to discriminate between stationarity and non-startionarity in variance. In what follows, we compute statistics that combine this individual information to gain power in the statistical inference.

4.1.1 Panel data assuming cross-section independent individuals

In this subsection we compute the Im, Pesaran and Shin (2003) – hereafter IPS unit root tests – and Maddala and Wu (1999) unit root tests – henceforth, MW test. These statistics are based on the ADF-type regression equation given by:

$$\Delta rid_{i,t} = \alpha_{mi} d_{mt} + \delta_i rid_{i,t-1} + \sum_{k=1}^{p} \gamma_k \Delta rid_{i,t-k} + \varepsilon_{i,t}, \tag{17}$$

$t = 1, \ldots, T$, $i = 1, \ldots, N$, where d_{mt} denotes the deterministic component, and where the usual specifications are a constant or a linear time trend. As mentioned above, in this paper we specify the deterministic component as a constant term, since this is the specification that is consistent with both the economic theory and with the visual inspection of time series. The null hypothesis is given by $H_0 : \delta_i = 0 \ \forall i$, whereas the alternative hypothesis is $H_1 : \delta_i < 0 \ i = 1, \ldots, N_1$; $\delta_i = 0 \ i = N_1 + 1, \ldots, N$. Therefore, the null is rejected if there is a subset (N_1) of stationary individuals. As a result, unit root hypothesis testing can be conducted allowing for a higher degree of heterogeneity as under the alternative hypothesis a common autoregressive parameter is not required. In addition, it accounts for idiosyncratic dynamics since different lag lengths for the parametric correction can be specified for each individual. Im, Pesaran and Shin (1997, 2003) propose two test statistics. The first test is the standardized group-mean Lagrange Multiplier (LM) bar test statistic – the $\Psi_{\overline{LM}}$ test – and the second one is the standardized group-mean t bar test statistic – the $\Psi_{\bar{t}}$ test – both of them used for testing $\delta_i = 0$ in (17). Under the assumption that the individuals are cross-section independent, it can be shown that both tests converge to the standard Normal distribution once they have been properly standarized.

Instead of combining the individual pseudo t-ratio that defines the $\Psi_{\bar{t}}$ test in Im, Pesaran and Shin (2003), Maddala and Wu (1999) suggest pooling the individual p-values. Under the null hypothesis and assuming cross-section independence, the test statistic given by $MW = -2\sum_{i=1}^{N} \ln(\pi_i) \sim \chi^2_{2N}$, where π_i denotes the p-value of the pseudo t-ratio for testing $\delta_i = 0$ in (17).

It is possible to complete the stochastic properties analysis that are drawn from the panel data unit root tests through the application of the LM test proposed by Hadri (2000), which specifies the null of stationarity allowing for heterogeneous and serially correlated errors. This test can be considered the panel version of the variance stationarity given by the KPSS test. Hadri (2000) proposes two models for the deterministic component depending on whether we only include an intercept or a linear time trend. The stochastic component is assumed to be decomposed into the sum of a random walk and a stationary disturbance term. He tests for the null hypothesis that all the variables ($rid_{i,t}$) are stationary – around deterministic levels or around deterministic trends – so that for the N elements of the panel the variance of the errors is such that $H_0 : \sigma^2_{u1} = ... = \sigma^2_{uN} = 0$ against the alternative hypothesis that some $\sigma^2_{ui} > 0$. This alternative allows for heterogeneous σ^2_{ui} across the cross-sections and includes the homogeneous alternative ($\sigma^2_{ui} = \sigma^2_u$ for all i) as a special case. It also allows for a subset of cross-sections to be stationary under the alternative. The test statistic is given by the average of the individual KPSS statistics, assuming either homogeneous or heterogeneous long-run variance. After suitable standardization and assuming cross-section independence, the tests are shown to converge to the standard Normal distribution. In the paper and based on simulation evidence reported in Carrion-i-Silvestre and Sansó (2006), the long-run variance is estimated using the procedure suggested by Sul, Phillips and Choi (2005).

Harris, Leybourne and McCabe (2005) have recently proposed a panel stationarity test statistic that introduces common factors to account for cross-section dependence. Their statistic tests the joint null hypothesis of stationarity in both common factors and idiosyncratic disturbance terms. Let us first deal with cross-section independence. In this case, their statistic is an extension of the one proposed by Harris, McCabe and Leybourne (2003) to the panel data framework, and it is based on the specification of the following model

$$\begin{aligned} rid_{i,t} &= x_{i,t}\beta + z_{i,t} \\ z_{i,t} &= \phi_i z_{i,t-1} + \varepsilon_{i,t}, \end{aligned} \tag{18}$$

where $x_{i,t}$ collects deterministic regressors in a general way – regressors such as constant, linear time trend or broken trends. We can obtain the OLS estimated residuals in (18) and, assuming cross-section independence, compute the statistic given by

$$\hat{S}_k = \frac{\hat{C}_k + \hat{c}}{\hat{\omega}\left\{\hat{a}_{k,t}\right\}}, \tag{19}$$

with $\hat{C}_k = T^{-1/2} \sum_{t=k+1}^{T} \hat{a}_{k,t}$ the autocovariance of order k, $\hat{a}_{k,t} = \sum_{i=1}^{N} \hat{z}_{i,t} \hat{z}_{i,t-k}$, and $\hat{z}_{i,t}$ denotes the OLS residuals in (18). $\hat{c} = (T-k)^{-1/2} \sum_{i=1}^{N} \hat{c}_i$, being $\hat{c}_i$ a correction term defined in Harris, Leybourne and McCabe (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 variance stationarity of the common and idiosyncratic components the statistic $\hat{S}_k \rightarrow^d N(0,1)$. In this paper we follow Harris, Leybourne and McCabe (2005) and use $k = \left[(3T)^{1/2}\right]$.

Panel A in Tables 2 and 3 reports the results for the IPS, MW and Hadri statistics when the individuals are assumed to be cross-section independent for the ex-post and ex-ante real interest rates respectively. Let us first focus on the short-run interest rates. Panel data unit root statistics lead to strong rejection of the null hypothesis of unit root, whereas the Hadri statistic that is computed assuming heterogeneous long-run variance does not reject the null hypothesis of variance stationarity. Although some discrepancy is obtained for the Hadri statistic with homogeneous long-run variance, for which the null hypothesis is rejected, the assumption of homogeneity might be hardly satisfied by the individuals in the data set. Evidence in favour of the RIRP hypothesis is reinforced by the $\hat{S}_k$ statistic of Harris, Leybourne and McCabe (2005) – see Table 5 for the $\hat{S}_k$ statistic. These conclusions are obtained for both the ex-ante and ex-post versions of real interest rates.

Regarding long-run real interest rates, the panel-data-based statistics conclude in favour of the real interest rate parity for ex-post interest rates, since all panel data unit root tests reject the null hypothesis of unit root, while the stationarity tests of Hadri (2000) and $\hat{S}_k$ indicate that the null hypothesis of stationarity cannot be rejected. Unfortunately, the evidence is not so conclusive for the ex-ante real interest rate. For this variable the IPS and $\hat{S}_k$ statistics point towards variance stationarity, while the MW and Hadri statistics do not.

4.1.2 Testing the null of cross-section independence

As mentioned above, the independence assumption imposed in the first generation panel data statistics has been widely criticized in recent literature, since it has been shown that non accounting for cross-section dependence amongst the individuals might bias the statistical inference in favor of variance stationarity – see Banerjee, Marcellino and Osbat (2004, 2005). Although it is now common practice to apply panel data unit root and stationarity tests that take into account cross-section dependence, few really test whether the individuals are cross-section dependent. In this subsection we test the null hypothesis of non correlation 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.

In brief, the procedure works as follows. First, we get rid of autocorrelation pattern in individual time series through the estimation of an AR model. This allows us to isolate the cross-section regression from serial correlation. Since we base the computation of panel data unit root tests on the ADF statistic, we specify the order of the autoregressive model that has been used for this statistic. Taking the estimated residuals from the ADF-type regression equations as individual series, we compute the absolute value of Pearson's correlation coefficients ($\bar{p}_j = |\hat{p}_j|$) for all possible pairs of individuals, $j = 1, 2, \ldots, n$, where $n = N(N-1)/2$, and sort them in ascending order. 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 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. The definition of the partition is carried out through minimization of the sum of squared residuals

$$Q_n(\vartheta) = \sum_{j=1}^{[\vartheta n]} \left(\Delta\bar{\phi}_j - \bar{\Delta}_S(\vartheta)\right)^2 + \sum_{j=[\vartheta n]+1}^{n} \left(\Delta\bar{\phi}_j - \bar{\Delta}_L(\vartheta)\right)^2,$$

where $\bar{\Delta}_S(\vartheta)$ and $\bar{\Delta}_L(\vartheta)$ denotes the mean of the spacings for each group respectively. A consistent estimate of the break point is obtained as $\hat{\vartheta} = \arg\min_{\vartheta \in (0,1)} Q_n(\vartheta)$, where the definition of some trimming is required. We follow Ng (2006) and set the trimming at 0.10.

Once the sample has been splitted, 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) in Ng (2006), which is defined as $SVR(\eta) = \left(\hat{\sigma}_q^2/\hat{\sigma}_1^2 - 1\right)$, where $\hat{\sigma}_q^2 = (m_q q)^{-1} \sum_{k=q+1}^{\eta} \left(\bar{\phi}_k^n - \bar{\phi}_{k-q}^n - \hat{\mu}_q\right)^2$, $\hat{\sigma}_1^2 = (\eta)^{-1} \sum_{k=1}^{\eta} \left(\bar{\phi}_k^n - \bar{\phi}_{k-1}^n - \hat{\mu}_1\right)^2$, $\hat{\mu}_q = (\eta - q)^{-1} \sum_{k=q+1}^{\eta} \left(\bar{\phi}_k^n - \bar{\phi}_{k-q}^n\right)$ and $\hat{\mu}_1 = (\eta - 1)^{-1} \sum_{k=1}^{\eta} \left(\bar{\phi}_k^n - \bar{\phi}_{k-1}^n\right)$, with $m_q = (\eta - q)$ and $\hat{\eta} = \left[\hat{\vartheta} n\right]$ being the number of statistics in the small correlations group. Ng (2006) shows that under the null hypothesis that a subset of correlations are jointly zero, $\sqrt{\eta} SVR(\eta) \to^d N\left(0, \omega_q^2\right)$, $\omega_q^2 = 2(2q-1)(q-1)/(3q)$, as $\eta \to \infty$. Using these results we can define the standardized statistic as $svr(\eta) = \sqrt{\eta} SVR(\eta) / \sqrt{\omega_q^2} \to^d N(0,1)$.

As can be seen from Table 4, we can split the whole sample of spacings in two groups, where the break point is estimated at $\hat{\eta} = 36$, $\hat{\eta} = 12$ or $\hat{\eta} = 15$, depending on the maturity and the definition of real interest rates. Except for the S group of the RIRPHP panel set, the computation of the $svr(\eta)$ statistic indicates that the null hypothesis of non correlation is strongly rejected. This leads us to conclude that some form of cross-section correlation is present amongst individuals, so that it has to be accounted for when assessing the stochastic properties of the real interest rates. The null hypothesis for the S group of the short and long-run RIRPHP panel set cannot be rejected at the 5% level of significance. In all, these results point to the presence of cross-section dependence amongst individuals in both panels of interest rates.

In addition, the fact that the break point is estimated at the beginning – $\hat{\eta} = 36$, $\hat{\eta} = 12$ or $\hat{\eta} = 15$ – implies that the proportion of correlation coefficients that form the S group – $\hat{\vartheta} = 0.3$, $\hat{\vartheta} = 0.1$ and $\hat{\vartheta} = 0.098$, respectively – is small compared to the correlation coefficients in the L group, which indicates that pervasive cross-correlation is present amongst the individuals in the panel data sets. In this case, approximate factor models as suggested in Bai and Ng (2004) reveal as a good option to account for cross-section dependence in panels.

4.1.3 Panel data with cross-section dependence

There are different approaches in the panel literature to deal with cross-section dependence. As mentioned above, the simplest one consists of removing the cross-section mean, which implicitly assumes that cross-section dependence is driven by one common factor with the same intensity for all the individuals in the panel.

Results in Panel B of Tables 2 and 3, where the variables have been cross-section demeaned, show contradictions between panel data unit root and stationarity statistics for all panel data sets: all these statistics reject their respective null hypotheses. This has been interpreted in some situations as misspecification errors of the deterministic component used to compute the statistics – see Cheung and Chinn (1997). Note that this feature is found irrespective of the real interest rate that is used.

Panel C in Tables 2 and 3 presents the empirical distribution of the statistics obtained using the bootstrap techniques described by Maddala and Wu (1999). The bootstrap distribution, which accommodates general forms of cross-section dependence, is based on 2,000 replications. In general, the contradiction persists for the short-run interest rates, since neither panel data unit root tests nor the stationarity tests reject their null hypothesis at the 5% level of significance for both ex-post and ex-ante variables. However, it should be noted that mild evidence is found in favour of RIRP hypothesis by all panel statistics considered here when working at the 10% level of significance.

Concerning the long-run interest rates, the results indicate that there is strong evidence in favour of the real interest rate parity for the long-run ex-post variable, since panel data unit root statistics reject the null hypothesis and stationarity tests do not.

The reverse situation is found for the long-run ex-ante variable, as the panel data unit root tests are not able to reject the null hypothesis of non-stationarity in variance, and the Hadri homogeneous long-run variance test rejects the null hypothesis of stationarity in variance. Finally, using the Hadri test with heterogeneous long-run variance the null of stationarity is rejected at 10% but rejection is not possible at 5%.

The third approach that we consider to control for the presence of cross-section dependence is the one based on the approximate common factor models of Bai and Ng (2004). 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 individuals in the panel might be cointegrated – see Banerjee, Marcellino and Osbat (2004), and Gegenbach, Palm and Urbain (2004). The Bai and Ng (2004) approach decomposes the observable variables in two different stochastic components. The first one is given by the common factors that drive the cross-section dependence amongst individuals, while the second component is the idiosyncratic disturbance term that is assumed to be cross-section independent. Thus the Data Generating Process for each individual is given by

$$\begin{aligned} rid_{i,t} &= D_{i,t} + F_t' \pi_i + e_{i,t}; && (20)\\ (I - L) F_t &= C(L) u_t; && (21)\\ (1 - \rho_i L) e_{i,t} &= H_i(L) \varepsilon_{i,t}, && (22) \end{aligned}$$

$t = 1, \ldots, T$, $i = 1, \ldots, N$, where $C(L) = \sum_{j=0}^{\infty} C_j L^j$ and $H_i(L) = \sum_{j=0}^{\infty} H_{i,j} L^j$. $D_{i,t}$ 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. Unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model

$$\begin{aligned} M_i \, \Delta rid_i &= M_i \, \Delta F \pi_i + M_i \, \Delta e_i \\ y_i &= f \pi_i + z_i, && (23) \end{aligned}$$

where $M_i = I_{T-1}$ for the only constant case and $M_i = I_{T-1} - (T-1)^{-1} \iota\iota'$ for the linear time trend deterministic specifications. The estimated factors $\hat{f}_{1,t}, \ldots, \hat{f}_{r,t}$ are the r eigenvectors that corresponds to the r largest eigenvalues of the $(T - 1 \times T - 1)$ matrix yy', being $y = [y_1, \ldots, y_N]$. The matrix of estimated weights, $\hat{\Pi} = (\hat{\pi}_1, \ldots, \hat{\pi}_N)'$, is given by $\hat{\Pi} = y' \hat{f}_t$. As a result, we can obtain an estimate of z_i from $\hat{z}_i = y_i - \hat{f} \hat{\pi}_i$, that, after computing its cumulated sum, produces a consistent estimation of the idiosyncratic disturbance term, $\tilde{e}_{i,t} = \sum_{j=1}^{t} \hat{z}_{i,j} = \sum_{j=1}^{t} (M_i \Delta \hat{e}_i)_j$. Similarly, we can recover the common factors as $\hat{F}_t = \sum_{j=1}^{t} \hat{f}_t$. The panel data unit root hypothesis on $\tilde{e}_{i,t}$ can be tested using the idiosyncratic ADF statistic. When the estimated number of common factors is $\hat{r} = 1$, we can test the null hypothesis of unit root on $\hat{F}_t$ using the usual ADF statistic. Finally, when $\hat{r} > 1$ we can use either the parametric or non-parametric MQ statistics suggested in Bai and Ng (2004) to estimate the number of common stochastic trends. The estimation of the number of common factors is obtained using the panel BIC information criterion in Bai and Ng (2002) – in this case we specify a maximum of $r_{\max} = 6$ common factors.

Table 5 reports the results of applying the Bai and Ng (2004) approach. Using this procedure we are not able to detect any common factor. This is in sharp contrast with the previous dependence analysis, where there was a large number of significant large correlations amongst individuals. This feature can be due to the moderate number of individuals (sixteen) that defines these panel data sets, which makes the estimation the number of common factors more difficult – the use of panel BIC information criteria provides a consistent estimation of the factor subspace for large N. For this reason, in this case it would be better an exogenous selection of the number of common factors. If we impose one common factor, the ADF statistic computed from the idiosyncratic disturbance terms rejects the null hypothesis of unit root, while the common factor is non-stationary in variance. As for the long-run interest rates, we have detected one common factor in both panel data sets using Bai and Ng (2004) approach. This common factor is characterized as non-stationary using either the parametric or non-parametric MQ statistics. Concerning the idiosyncratic component of the ex-post rate, the ADF statistic indicates that they are stationary in variance, whereas for the ex-ante variable we find evidence of non-stationarity.

It would be possible to test the null hypothesis of stationarity assuming that the individuals are cross-section dependent using the $\hat{S}_k$ statistic in Harris, Leybourne and McCabe (2005). When individuals are cross-section dependent, we can compute the statistic given in (19) defining $\hat{z}_{i,t}$ 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 drawn from the Bai and Ng (2004) procedure. The statistic that can be used to test the null hypothesis of variance stationarity has the form given by (19), although we introduce a superscript in $\hat{S}_k^F$ to denote that its computation is based on approximate common factor models. 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 5 reports the $\hat{S}_k^F$ statistics. The evidence for the ex-ante and ex-post short-run interest rates, and the long-run ex-post data sets is conclusive, since the null hypothesis of variance stationarity is not rejected by the $\hat{S}_k^F$ statistic. However, the $\hat{S}_k^F$ statistic does reject the null hypothesis of variance stationarity for long-run RIRPHP.

4.2 Panel tests with structural breaks

The results obtained up to now in the paper are based on the assumption that individual time series are not affected by structural changes. However, the visual inspection of the pictures of the real interest rates casts doubts on this underlying assumption. In addition, some contradictory results found when computing panel unit root and stationarity statistics might be understood as an indicator of misspecification error of the deterministic component. Unattended structural breaks may affect the statistical inference leading us to conclude in favor of non-stationarity in variance.

4.2.1 Panel data assuming cross-section independent individuals

In order to account for the presence of structural breaks we have applied the approach suggested by Carrion-i-Silvestre, del Barrio and López-Bazo (2005) to test the null hypothesis of panel variance stationarity allowing for multiple level shifts. This statistic extends the approach in Hadri (2000) through the specification of the following deterministic component

$$y_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{i,k,t} + \varepsilon_{i,t}, \tag{24}$$

with $DU_{i,k,t} = 1$ for $t > T_{b,k}^i$ and 0 elsewhere, where $\{\varepsilon_{i,t}\}$ are assumed to be independent across i. This specification permits a high degree of heterogeneity since the structural breaks may have different effects on each individual time series, the break points can be located at different dates for each individual, and the individuals may have different number of structural breaks. The OLS estimated residuals from (24) can be used to compute the individual KPSS statistic, which in turn can be combined to define the panel stationarity statistic – hereafter, we denote this statistic as $Z(\lambda)$, where λ is the vector of relative positions of the break points, i.e. the break fraction parameters. Under the null hypothesis of variance stationarity and assuming cross-section independence, the standardized panel data statistic is shown to converge to the standard normal distribution.

The estimation of both the number of structural breaks and their position in (24) is done using the sequential procedure by Bai and Perron (1998). When computing the statistic we have to specify a maximum number of structural breaks, which in this case has been set equal to $m_i = m = 5\ \forall i$. This maximum number of structural break is never reached for the short-run interest rates. However, the maximum number of structural breaks is selected for one individual in the case of the ex-post variable and in five situations for the ex-ante long-run interest rates. In this case, we consider that increasing the number of breaks is unrealistic due to the number of observations ($T = 83$)

that we use.[11] Tables 6 to 9 present individual KPSS statistics, with the estimated break points and the corresponding critical values – these critical values were obtained by direct simulation – for ex-ante and ex-post, short and long-run real interest rates. Figures 1 and 2 present the picture of the RIRP series along with the estimated break points. Mild evidence against the null hypothesis of variance stationarity is found for the short-run interest rates. Thus, the null is rejected at the 10% level for Australia and Netherlands for the ex-post variable. For the ex-ante one, the null hypothesis is rejected for Austria, Canada and Spain (at the 10% level of significance for the latter).

Concerning the long-run interest rates, the null hypothesis of variance stationarity is not rejected at 5% for almost all cases in the ex-post data set – the exceptions are Australia and Norway, where the null hypothesis is rejected at the 5% level, and United Kingdom, where it is rejected at 10%. Similar results are obtained for the ex-ante rate, where the null hypothesis is not rejected in most cases – the exceptions are Netherlands, where the null is rejected at the 5% level, and Austria, France and Spain, for which the mild evidence against the null hypothesis is found when working at the 10% level of significance.

The break points that have been estimated here correspond with some important features of monetary policies undertaken in the analysed period. In order to ease interpretation, we have computed the 95% confidence intervals for the estimated break points. This allow us to get better picture when identifying short time periods where break points are located acounting for the fact that the same event might be affected different individuals, although not at the same time period. According to the results reported in Tables 10 and 11, the sample period can be truncated by up to four breakpoints (with one exception). The first structural break is estimated to occur for the majority of the countries considered around early 1981. The rising inflation expectations in the pre-1981 period were due mainly to the oil shocks in the mid and late 70s and also partly to lack of monetary policy credibility. By the end of 1980, a significant reversal of inflation expectations took place after the US economy experienced a steep recession (Evans and Lewis, 1995) and the rise in the federal budget deficit (Garcia and Perron, 1996). In the case of the EU countries, the European Monetary System (EMS) inception can be an explanatory factor as well. A second break can be dated around the middle of the 80s with the launching of the new EMS (Basle-Nyborg agreement) as a mechanism to achieve the monetary integration in the EU. This process meant the progressive abolition of any remaining capital controls among the European countries by 1990. A third break is around 1990-1993, which coincides with German unification in July 1990. This fact meant a large asymmetric shock that gave birth to the EMS crisis in September 1992 and

[11] The tests were computed as well for a maximum of $m_i = m = 8$ $\forall i$ structural breaks to assess the sensitivity of the results. The null hypothesis of variance stationarity was not rejected for the homogeneous long-run variance version of the statistic at the 5% level of significance, although the opposite happened when using the statistic based on heterogeneous long-run variance.

the exit of Italy and the UK from the exchange rate mechanism of the EMS. Moreover, in August 1993, there is a formal widening of exchange rate bands of the EMS to ±15% followed to the adherence of the prospective Euro members to the Maastricht conditions on nominal convergence.

Individual information can be pooled to define the panel data statistics. Thus, assuming that the individuals are cross-section independent, the statistics in Carrion-i-Silvestre, del Barrio and López-Bazo (2005) reject the null hypothesis of variance stationarity for the short-run ex-post variable, but not for the short-run ex-ante rate and long-run ex-post and ex-ante variables – see Panel A in Tables 7 and 9. The conclusion of RIRP fulfillment is supported by the $\hat{S}_k$ statistic in Harris, Leybourne and McCabe (2005) for the short-run ex-ante (RIRPHP) and long-run ex-post (RIRPINF) data sets, but not for the short-run ex-post and long-run ex-ante ones – see results presented in Table 13 for the $\hat{S}_k$ statistic.

Evidently, while the first shock detected is clearly common to all the countries considered, other structural breaks are more EU-specific. However, the high degree of financial integration within the OECD area makes a case to analyse cross-section dependence between the different individual countries.

4.2.2 Testing the null of cross-section independence

We can allow for the presence of the structural breaks when testing the null hypothesis of non correlation amongst individuals in the panel. As above, we have estimated an autoregressive model to isolate cross-section dependence from the autocorrelation that might be driven individual time series – the order of the autoregressive model has been selected using the t-sig criterion in Ng and Perron (1995) with $p_{\max} = \left[12 * (T/100)^{1/4}\right]$ lags as the maximum order. In addition, the estimation of the autoregressive model includes dummy variables to capture the level shifts that have been detected using Bai and Perron (1998) in the previous section, which aims at isolating cross-section dependence from both autocorrelation and structural breaks in the individual time series.

The results in Table 12 show that the null hypothesis of non cross-section correlation is strongly rejected for the whole and L samples of spacings, while it is not rejected for the S sample one at the 5% level of significance, irrespective of the data set that is used. As before, the proportion of non significant correlations in the S group is small compared to the L group, which leads us to conclude that cross-section dependence is pervasive and, hence, it might be well captured by approximate common factor models.

4.2.3 Panel data with cross-section dependence

Previous results reveal that cross-section dependence is present among individuals, so it should be considered when computing the panel data statistics to avoid biases. Panel B in Tables 6 to 9 offers the value of the $Z(\lambda)$ statistic once the cross-section mean has been removed. According to this statistic, the null hypothesis of variance stationarity cannot be rejected for the two data sets irrespective of the assumption made on the long-run variance estimation and the different definitions of real interest rates. This evidence is reinforced when using bootstrap critical values – see Panel C in Tables 6 to 9 for the empirical distribution.[12]

Finally, we have computed the statistic in Harris, Leybourne and McCabe (2005) allowing for multiple level shifts and common factors – the break points are the ones estimated above. Results in Table 13 show that the $\hat{S}_k^F$ statistic do not reject the null hypothesis of variance stationarity for the short-run real interest rates – five common factors were detected – the long-run RIRPINF variable – two common factors are detected – and the long-run RIRPHP variable – one common factor is detected at the 5% level of significance. Mild evidence against the null hypothesis is found when working at the 10% level of significance.

To sum up, our results show that there is strong evidence of weak version of real interest rate parity, once structural breaks and cross-section dependence are allowed for, irrespective of whether ex-ante or ex-post real interest rates are used – Table 14 summarizes the results that have been obtained along the paper.

5 Conclusions

Many studies have reexamined the real interest rate parity condition and found rather hard to establish its fufillment empirically. In this paper we present new evidence in support of long run reversion in real interest rates differentials assessing the stochastic properties of the series for a group of OECD countries. We examine the behavior of cross-country real interest rate differentials for the US and 18 other major industrial economies from 1978:Q2 to 1998:Q4. Our analysis is based on the use of both panel data unit root and stationarity test statistics that accommodate the presence of either cross-section dependence and/or structural breaks. Taking into account these features is important to overcome potential biases of statistical inference. We investigate both the extent of financial market integration and whether and how it may have changed through time. We focus on three issues: first, whether real interest rate differentials, if not literally zero, are at least small in absolute value and hence consistent with financial integration in the presence of cross-country differences in risk; second, whether these differentials are mean reverting, and therefore, indicative of long-run equilibrium and, thirdly, whether and how their behavior has differed across exchange-rate regimes.

The results show that they crucially depend on the allowance of both structural breaks

[12] Similar conclusions were obtained when using up to eight structural breaks for the long-run real interest rates – none of the individuals achieved the maximum number of structural breaks.

and cross-section dependence when computing the statistics. 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. However, over short but still significant periods, real interest differentials have fluctuated greatly due to capital controls and to temporary responses to shocks and policy measures. Statistical procedures that have been applied along the paper reveal that these features are present in our setting. Thus, once we consider both of these characteristics we conclude in favour of ex-ante and ex-post RIRP fulfilment. The results of various panel-based unit root and stationarity tests used in this study are consistent and robust to alternative ways of estimating real interest rates (ex-ante or ex-post).

By exploiting the cross-sectional information and increasing the data span, these tests have higher power relative to the classical unit root tests. The failure of previous empirical studies to confirm mean reversion of real interest rates differentials may therefore reflect the choice of the estimation method used rather than any inherent deficiency in the RIRP. We find further that cross-country differentials are invariant to regime changes. Fluctuations in differentials occur periodically over the sample period, but while somewhat persistent, in the end prove transitory. For all 18 possible country pairs *rid* are mean reverting and RIRP holds in its weak form.

References

[1] Bai, J. and S. Ng (2002) "Determining the Number of Factors in Approximate Factor Models", *Econometrica*, 70, 191-221.

[2] Bai, J. and S. Ng (2004) "A PANIC Attack on Unit Roots and Cointegration", *Econometrica*, 72, 4, 1127-1177.

[3] Bai, J., and P. Perron (1998) "Estimating and Testing Linear Models with Multiple Structural Changes", *Econometrica*, 66, 47–78.

[4] Banerjee, A., Marcellino, M. and C. Osbat (2004) "Some Cautions on the Use of Panel Methods for Integrated Series of Macro-Economic Data", *Econometrics Journal*, 7, 2, 322-340.

[5] Banerjee, A., Marcellino, M. and C. Osbat (2005) "Testing for PPP Should We Use Panel Methods?". *Empirical Economics*, 30, 77-91.

[6] Barro, R.J. and X. Sala-i-Martín (1991) "World real interest rates", *NBER Macroeconomics Annual 1990*, 5, 15-61.

[7] Carrion-i-Silvestre, J. L., del Barrio, T. and López-Bazo, E. (2001) "Level Shifts in a Panel Data Based Unit Root Test. An Application to the Rate of Unemployment". *Proceedings of the 2001 North American Econometric Society.*

[8] Carrion-i-Silvestre, J. L., del Barrio, T. and López-Bazo, E. (2005). "Breaking the Panels. An Application to the GDP per Capita", *Econometrics Journal*, 8, 159-175.

[9] Carrion-i-Silvestre, J. L., and Sansó, A. (2006) "A Guide to the Computation of Stationarity Tests", *Empirical Economics*, forthcoming.

[10] Cavaglia, S. (1992) "The persistence of real interest differentials. A Kalman filter approach", *Journal of Monetary Economics*, 29, 429-443.

[11] Cheung, Y.-W. and M. D. Chinn (1997) "Further investigation of the uncertain unit root in GNP", *Journal of Business and Economic Statistics*, 15, 68-73.

[12] Chinn, M. D. and J. A. Frankel (1995) "Who drives real interest around the Pacific Rim the USA or Japan?", *Journal of International Money and Finance,* 14, 801-821.

[13] Chinn, M. D. and J.A. Frankel (2003) "The euro area and the world interest rates", mimeo.

[14] Cumby, R. and M. Obstfeld (1984) "International interest rate and price level linkages under flexible exchange rates a review of recent evidence", in Bilson, J. and

R.C. Marston, eds. *Exchange rate theory and practice,* University of Chicago Press, Chicago.

[15] Darin, R. and R. L. Hetzel (1995) "An empirical measure of the real rate of interest", *Federal Reserve Bank of Richmond Economic Quarterly*, 81, 17-47.

[16] Evans, P. (1985) "Do large deficits produce high interest rates?", *American Economic Review*, 75, 68-87.

[17] Evans, M. D. D. and K. K. Lewis (1995) "Do expected shifts in inflation affect estimated of the long-run Fisher relation?", *Journal of Finance*, 50, 225-253.

[18] Ferreira, A. L. and M. A. León-Ledesma (2003) "Does the real interest parity hypothesis hold? Evidence for developed and emerging markets", *Journal of International Money and Finance*, forthcoming.

[19] Fountas, S. and J. Wu (1999) "Testing for real interest rate convergence in European countries", *Scottish Journal of Political Economy*, 46, 158-174.

[20] Frankel, J.A., Schmukler, S.L. and L. Servén (2003) "Global transmission of interest rates monetary independence and currency regime", *NBER Working Paper* no. 8828.

[21] Fraser, P. and M.P. Taylor (1990) "Some efficient tests of international real interest parity", *Applied Economics*, 22, 1083-1092.

[22] Frenkel, J. (1976) "A monetary approach to the exchange rate doctrinal aspects of empirical evidence", *Scandinavian Journal of Economics*, 78, 200-224.

[23] Fuijii, E. and M.D. Chinn (2002) "Fin de siècle real interest parity", *Journal of International Financial Markets, Institutions and Money,* 11, 289-308.

[24] Garcia, R. and P. Perron (1996) "An analysis of the real interest rate under regime shifts", *Review of Economics and Statistics*, 78, 111-125.

[25] Gegenbach, C., F. C. Palm and J. P. Urbain (2004) "Panel unit root tests in the presence of cross-sectional dependencies comparison and implications for modelling", *Research memorandum* RM/04/039, Meteor, Universiteit Maastricht.

[26] Glick, R and M. Hutchison (1990) "Financial liberalization in the Pacific Basin implications for real interest rate linkages", *Journal of the Japanese and International Economies*, 4, 36-48.

[27] Goldberg, L. G. , J.R. Lothian and J. Okunev (2002) "Has international financial integration increased?", mimeo.

[28] Goodwin, B. and T. Grennes (1994) "Real interest rate equalisation and the integration of international financial markets", *Journal of International Money and Finance*, 13, 107-124.

[29] Hadri, K. (2000) "Testing for Unit Roots in Heterogeneous Panel Data", *Econometrics Journal*, 3, 148-161.

[30] Harris, D., S. Leybourne and B. McCabe (2005) "Panel stationarity tests for purchasing power parity with cross-sectional dependence", *Journal of Business and Economics Statistics*, 23, 395-409.

[31] Hodrick, R. and E. Prescott (1997) "Post-war US business cycles An empirical investigation", *Journal of Money, Credit and Banking*, 29(1), 1-16.

[32] Holmes, M.J. (2002) "Does long-run real interest parity hold among EU countries? Some new panel data evidence", *The Quarterly Review of Economics and Finance,* 42, 733-746.

[33] Holmes, M.J. and N. Maghrebi (2003) "Asian real interest rates, nonlinear dynamics, and international parity", *International Review of Economics and Finance*, forthcoming.

[34] Im, K., M.H. Pesaran and Y. Shin (1997): "Testing for unit roots in heterogeneous panels", *Working paper*, Department of Applied Economics, University of Cambridge.

[35] Im, K., M.H. Pesaran and Y. Shin (2003) "Testing for unit roots in heterogeneous panels", *Journal of Econometrics,* 115, 53-74.

[36] King, R.G. and S. T. Rebelo (1993) "Low frequency filtering and real business cycles", *Journal of Economic Dynamics and Control*, 17, 207-231.

[37] Kwiatkowski, D., P. C. B. Phillips, P. J. Schmidt, and Y. Shin (1992) "Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root How Sure are We that Economic Time Series Have a Unit Root", *Journal of Econometrics*, 54, 159–178.

[38] Lai, K. S. (2004) "On structural shifts and stationarity of the ex ante real interest rate", *International Review of Economics and Finance*, 13, 217-228.

[39] Lee, J., C. J. Huang and Y. Shin (1997) "On Stationary Tests in the Presence of Structural Breaks", *Economics Letters*, 55, 165–172.

[40] Levin, A., C.F. Lin and J. Chu (2002) "Unit root tests in panel data asymptotic and finite-sample properties", *Journal of Econometrics*, 108, 1-24.

[41] Lothian, J.R. (2000) "Capital market integration and exchange-rate regimes in historical perspective" in I. Hasan and W.C. Hunter (eds.) *Research in Banking and Finance*, 1, 141-176.

[42] MacDonald, R. and M. Moore (2001) "The spot-fordward relationship revisited An ERM perspective", *Journal of Financial Markets, Institutions and Money*, 11, 29-52.

[43] Maddala, G. S. and Wu, S. (1999) "A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test", *Oxford Bulletin of Economics and Statistics*, Special Issue, 61, 631-652.

[44] Mancuso, A.J., B. K. Goodwin and T.J. Grennes (2003) "Nonlinear aspects of capital market integration and real interest rate equalization", *International Review of Economics and Finance*, 12, 283-303.

[45] Marston, R.C. (1995) *International financial integration a study of interest differentials between the major industrial countries.* Cambridge and New York Cambridge University Press.

[46] McCallum, B.T. (1994) "A reconsideration of the uncovered interest parity relatioship", *Journal of Monetary Economics*, 33, 105-132.

[47] Mishkin, F.S. (1984) "Are real interest rates equal across countries?An empirical investigation of international parity conditions", *The Journal of Finance,* 39, 5, 1345-1357.

[48] Montañés, A. and M. Reyes (1998) "Effect of a Shift in the Trend Function on Dickey-Fuller Unit Root Tests", *Econometric Theory*, 14, 355–363.

[49] Moosa, I and R. Bhatti (1996) "Some evidence on mean reversion in *ex ante* real interest rates", *Scottish Journal of Political Economy*, 43, 177-191.

[50] Ng, S. (2006) "Testing cross-section correlation in panel data using spacings", *Journal of Business and Economics Statistics*, 24, 12-23.

[51] Ng, S. and P. Perron (1995) "Unit root tests in ARMA models with data dependent methods for the selection of the truncation lag", *Journal of the American Statistical Association* 90, 268-281.

[52] Nieuwland, F, Verschoor, W. and C. Wolff (1998) "EMS exchange rate expectations and time-varying risk premia", *Economics Letters*, 60, 351-355.

[53] O'Connell, P. (1998) "The overvaluation of purchasing power parity", *Journal of International Economics*, 44, 1-19.

[54] Perron, P. (1989) "The Great Crash, the Oil Price Shock and the Unit Root Hypothesis", *Econometrica*, 57, 1361–1401.

[55] Plosser, C.I. (1987) "Fiscal policy and the term structure", *Journal of Monetary Economics* 20, 343-367.

[56] Reichenstein, W. and J.W. Elliot (1987) "A comparison of models of long-term inflationary expectations", *Journal of Monetary Economics,* 19, 405-425.

[57] Sephton, P. S (1995) "Response surfaces estimates of the KPSS stationary test", *Economics Letters*, 47, 255–261.

[58] Shiller, R. J. and P. Perron (1985) "Testing the random walk hypothesis power versus frecuency of observation", *Economics Letters,* vol. 18, 4, 381-386.

[59] Strauss, J. and T. Yigit (2003): "Shortfalls of panel unit root testing", *Economics Letters*, 81, 309-313.

[60] Sul D., P. C. B. Phillips and C. Y. Choi (2005) "Prewhitening Bias in HAC Estimation", *Oxford Bulletin of Economics and Statistics*, 67, 517-546.

[61] Wu, J. L. and S.L. Chen (1998) "A re-examination of real interest rate parity", *Canadian Journal of Economics*, 837-851.

[62] Wu, J. L. and S. Fountas (2000) "Real interest rate parity under regime shifts and implications for monetary policy", *The Manchester School*, 68, 685-700.

Table 1: Individual ADF and KPSS statistics

Panel A: Short-run real interest rates

	RIRPINF				RIRPHP			
	ADF	p	p-val	KPSS	ADF	p	p-val	KPSS
Australia	-2.340	0	0.184	0.186	-1.514	8	0.534	0.178
Austria	-2.995	9	0.049	0.076	-4.130	9	0.002	0.074
Belgium	-2.386	5	0.170	0.053	-2.593	10	0.115	0.067
Canada	-2.437	7	0.155	0.180	-1.587	4	0.500	0.128
Denmark	-2.526	11	0.131	0.065	-2.191	10	0.235	0.075
France	-2.509	11	0.136	0.260	-1.538	2	0.523	0.276
Germany	-2.473	5	0.145	0.054	-1.719	11	0.438	0.067
Ireland	-2.250	2	0.214	0.136	-2.105	2	0.268	0.134
Italy	-1.608	2	0.490	0.290	-2.202	5	0.231	0.297
Japan	-2.755	5	0.083	0.063	-2.377	2	0.172	0.073
Netherlands	-2.312	11	0.193	0.047	-1.901	10	0.355	0.051
Norway	-1.548	6	0.518	0.156	-2.977	0	0.051	0.133
Portugal	-1.129	9	0.697	0.736	-1.741	5	0.428	1.459
Spain	-1.671	5	0.461	0.233	-1.924	5	0.344	0.241
Switzerland	-1.747	5	0.426	0.221	-2.202	9	0.231	0.242
United Kingdom	-1.720	3	0.438	0.262	-1.593	7	0.498	0.231

Panel B: Long-run real interest rates

	RIRPINF				RIRPHP			
	ADF	p	p-val	KPSS	ADF	p	p-val	KPSS
Australia	-2.575	7	0.119	0.128	-1.305	8	0.627	0.177
Austria	-3.315	9	0.023	0.092	-2.720	3	0.089	0.125
Belgium	-4.233	5	0.002	0.110	-2.804	10	0.075	0.157
Canada	-2.655	5	0.102	0.204	-1.584	10	0.502	0.559
Denmark	-2.139	0	0.255	0.203	-2.028	1	0.299	0.271
France	-2.040	0	0.295	0.161	-1.541	10	0.522	0.349
Germany	-3.203	3	0.030	0.071	-2.923	3	0.058	0.108
Ireland	-2.611	8	0.111	0.172	-1.799	2	0.401	0.221
Italy	-1.412	4	0.581	0.306	-1.218	7	0.662	0.334
Japan	-3.211	10	0.029	0.351	-2.710	6	0.091	0.362
Luxembourg	-3.127	6	0.036	0.091	-2.799	1	0.076	0.102
Netherlands	-3.601	5	0.011	0.135	-2.849	3	0.068	0.177
New Zealand	-1.229	8	0.658	0.223	-1.442	3	0.567	0.447
Norway	-2.346	4	0.182	0.182	-1.759	11	0.420	0.181
Portugal	-2.115	10	0.264	0.269	-1.245	2	0.652	1.462
Spain	-3.158	1	0.033	0.384	-1.447	5	0.565	0.475
Switzerland	-2.284	10	0.202	0.157	-2.318	3	0.191	0.262
United Kingdom	-2.628	7	0.108	0.088	-1.672	8	0.460	0.135

The column labelled as p denotes the order of the autoregressive correction, which has been selected using the t-sig criterion in Ng and Perron (1995) with $p_{max} = [12 * (T/100)^{1/4}]$ as the maximum order. P-values are obtained by direct simulation. Critical values for the KPSS statistic drawn from the response surface in Sephton (1995) are 0.349 (10% level of significance), 0.467 (5% level of significance) and 0.732 (1% level of significance).

Table 2: Panel data unit root and stationarity tests without structural breaks for the RIRPINF panel data set

Short-run real interest rates

	Panel A: Assuming cross-section independence			**Panel B:** Removing cross-section mean	
	Test	p-val		Test	p-val
$\Psi_{\bar{t}}$	-2.996	0.001	$\Psi_{\bar{t}}$	-6.850	0.000
$\Psi_{\overline{LM}}$	3.251	0.001	$\Psi_{\overline{LM}}$	9.361	0.000
MW	48.348	0.032	MW	100.938	0.000
Hadri (Hom.)	2.944	0.002	Hadri (Hom.)	7.666	0.000
Hadri (Het.)	0.590	0.278	Hadri (Het.)	2.215	0.013

Panel C: Bootstrap distribution (allowing for cross-section dependence)

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
$\Psi_{\bar{t}}$	-4.490	-3.823	-3.357	-2.779	1.545	2.432	3.357	4.315
$\Psi_{\overline{LM}}$	-2.677	-2.119	-1.691	-1.068	3.413	4.161	5.008	5.897
MW	11.369	14.195	16.171	19.301	48.221	54.109	60.334	66.208
Hadri (Hom.)	-2.826	-2.624	-2.420	-2.109	3.345	5.053	6.683	9.023
Hadri (Het.)	-2.739	-2.542	-2.310	-2.044	3.532	5.456	6.924	9.352

Long-run real interest rates

	Panel A: Assuming cross-section independence			**Panel B:** Removing cross-section mean	
	Test	p-val		Test	p-val
$\Psi_{\bar{t}}$	-5.615	0.000	$\Psi_{\bar{t}}$	-6.013	0.000
$\Psi_{\overline{LM}}$	7.153	0.000	$\Psi_{\overline{LM}}$	8.854	0.000
MW	90.946	0.000	MW	89.551	0.000
Hadri (Hom.)	1.233	0.109	Hadri (Hom.)	5.422	0.000
Hadri (Het.)	0.514	0.304	Hadri (Het.)	5.044	0.000

Panel C: Bootstrap distribution (allowing for cross-section dependence)

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
$\Psi_{\bar{t}}$	-3.365	-2.885	-2.470	-1.927	2.542	3.574	4.360	5.214
$\Psi_{\overline{LM}}$	-3.129	-2.679	-2.224	-1.630	2.238	2.916	3.518	4.153
MW	11.123	14.643	17.545	20.910	52.768	58.967	64.687	70.970
Hadri (Hom.)	-2.822	-2.617	-2.406	-2.118	3.452	5.199	6.923	9.066
Hadri (Het.)	-2.720	-2.501	-2.294	-2.003	3.448	5.155	6.948	9.010

Table 3: Panel data unit root and stationarity tests without structural breaks for the RIRPHP panel data set

Short-run real interest rates

	Panel A: Assuming cross-section independence			**Panel B:** Removing cross-section mean	
	Test	p-val		Test	p-val
$\Psi_{\bar{t}}$	-2.975	0.001	$\Psi_{\bar{t}}$	-6.112	0.000
$\Psi_{\overline{LM}}$	3.137	0.001	$\Psi_{\overline{LM}}$	8.116	0.000
MW	50.174	0.021	MW	104.992	0.000
Hadri (Hom.)	4.265	0.000	Hadri (Hom.)	12.707	0.000
Hadri (Het.)	1.776	0.038	Hadri (Het.)	4.906	0.000

Panel C: Bootstrap distribution (allowing for cross-section dependence)

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
$\Psi_{\bar{t}}$	-4.268	-3.428	-2.956	-2.438	1.716	2.670	3.270	4.294
$\Psi_{\overline{LM}}$	-2.549	-2.097	-1.618	-1.083	3.034	3.732	4.463	5.445
MW	12.968	15.195	17.649	20.621	49.817	56.576	62.271	69.496
Hadri (Hom.)	-2.926	-2.712	-2.466	-2.177	3.528	5.277	6.588	9.107
Hadri (Het.)	-2.81	-2.523	-2.325	-2.071	3.474	5.369	6.784	9.197

Long-run real interest rates

	Panel A: Assuming cross-section independence			**Panel B:** Removing cross-section mean	
	Test	p-val		Test	p-val
$\Psi_{\bar{t}}$	-2.487	0.006	$\Psi_{\bar{t}}$	-4.746	0.000
$\Psi_{\overline{LM}}$	2.688	0.004	$\Psi_{\overline{LM}}$	6.327	0.000
MW	49.482	0.067	MW	70.134	0.001
Hadri (Hom.)	6.421	0.000	Hadri (Hom.)	16.150	0.000
Hadri (Het.)	4.591	0.000	Hadri (Het.)	16.765	0.000

Panel C: Bootstrap distribution (allowing for cross-section dependence)

	1%	2.5%	5%	10%	90%	95%	97.5%	99%
$\Psi_{\bar{t}}$	-4.971	-4.224	-3.751	-3.079	2.175	3.447	4.712	5.804
$\Psi_{\overline{LM}}$	-3.023	-2.408	-1.921	-1.355	3.635	4.563	5.302	6.340
MW	8.280	11.658	15.702	19.523	57.293	64.013	69.138	77.115
Hadri (Hom.)	-3.041	-2.839	-2.630	-2.329	3.591	5.391	7.119	9.776
Hadri (Het.)	-2.983	-2.775	-2.566	-2.244	3.672	5.449	7.304	9.550

Table 4: 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
RIRPINF	5.140	0.000	2.156	0.016	36	3.943	0.000
RIRPHP	6.147	0.000	-1.182	0.881	12	4.957	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
RIRPINF	9.392	0.000	2.994	0.001	15	5.996	0.000
RIRPHP	3.013	0.000	0.949	0.171	15	6.740	0.000

Table 5: Panel data statistics based on approximate common factor models

Panel A: Short-run real interest rates

Bai and Ng (2004) statistics

	RIRPINF		RIRPHP	
	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-3.536	0.000	-3.064	0.001
	Test	$r = 1$	Test	$r = 1$
MQ test (parametric)	-7.552	1	-7.598	1
MQ test (non-parametric)	-7.640	1	-8.356	1

Harris et al. (2005) statistics

	RIRPINF		RIRPHP	
	Test	p-value	Test	p-value
$\hat{S}_k$	0.672	0.251	0.459	0.323
$\hat{S}_k^F$	1.332	0.091	0.706	0.240

Panel B: Long-run real interest rates

Bai and Ng (2004) statistics

	RIRPINF		RIRPHP	
	Test	p-value	Test	p-value
Idiosyncratic ADF statistic	-2.641	0.004	-1.213	0.113
	Test	$\hat{r}$ $(r_{max} = 6)$	Test	$\hat{r}$ $(r_{max} = 6)$
MQ test (parametric)	-4.668	1	-4.833	1
MQ test (non-parametric)	-6.431	1	-6.879	1

Harris et al. (2005) statistics

	RIRPINF		RIRPHP	
	Test	p-value	Test	p-value
$\hat{S}_k$	-0.192	0.576	1.110	0.134
$\hat{S}_k^F$	1.085	0.139	2.287	0.011

Critical values for the MQ tests are -20.151 (1% level of significance) -13.730 (5% level of significance) and -11.022 (10% level of significance). See Table I in Bai and Ng (2004)

2006

CREAP2006-01
Matas, A. (GEAP)**; Raymond, J.Ll.** (GEAP)
"Economic development and changes in car ownership patterns"
(Juny 2006)

CREAP2006-02
Trillas, F. (IEB)**; Montolio, D.** (IEB)**; Duch, N.** (IEB)
"Productive efficiency and regulatory reform: The case of Vehicle Inspection Services"
(Setembre 2006)

CREAP2006-03
Bel, G. (GPRE)**; Fageda, X.** (GPRE)
"Factors explaining local privatization: A meta-regression analysis"
(Octubre 2006)

CREAP2006-04
Fernàndez-Villadangos, L. (GPRE)
"Are two-part tariffs efficient when consumers plan ahead?: An empirical study"
(Octubre 2006)

CREAP2006-05
Artís, M. (AQR)**; Ramos, R.** (AQR)**; Suriñach, J.** (AQR)
"Job losses, outsourcing and relocation: Empirical evidence using microdata"
(Octubre 2006)

CREAP2006-06
Alcañiz, M. (RISC)**; Costa, A.; Guillén, M.** (RISC)**; Luna, C.; Rovira, C.**
"Calculation of the variance in surveys of the economic climate"
(Novembre 2006)

CREAP2006-07
Albalate, D. (GPRE)
"Lowering blood alcohol content levels to save lives: The European Experience"
(Desembre 2006)

CREAP2006-08
Garrido, A. (IEB)**; Arqué, P.** (IEB)
"The choice of banking firm: Are the interest rate a significant criteria?"
(Desembre 2006)

2006

CREAP2006-09
Segarra, A. (GRIT)**; Teruel-Carrizosa, M.** (GRIT)
"Productivity growth and competition in spanish manufacturing firms:
What has happened in recent years?"
(Desembre 2006)

CREAP2006-10
Andonova, V.; Díaz-Serrano, Luis. (CREB)
"Political institutions and the development of telecommunications"
(Desembre 2006)

CREAP2006-11
Raymond, J.L.(GEAP)**; Roig, J.L..** (GEAP)
"Capital humano: un análisis comparativo Catalunya-España"
(Desembre 2006)

CREAP2006-12
Rodríguez, M.(CREB)**; Stoyanova, A.** (CREB)
"Changes in the demand for private medical insurance following a shift in tax incentives"
(Desembre 2006)

CREAP2006-13
Royuela, V. (AQR-IREA)**; Lambiri, D.**; **Biagi**, **B.**
"Economía urbana y calidad de vida. Una revisión del estado del conocimiento en España"
(Desembre 2006)

CREAP2006-14
Camarero, M.; Carrion-i-Silvestre, J.LL. (AQR-IREA)**.**;**Tamarit**, **C.**
"New evidence of the real interest rate parity for OECD countries using panel unit root tests with breaks"
(Desembre 2006)

CENTRE
DE REFERÈNCIA
EN ECONOMIA APLICADA

creap@pcb.ub.es