Does the real interest rate parity hold for OECD countries? New evidence using panel stationarity tests with cross-section dependence and structural breaks*

Mariam Camarero† Josep Lluís Carrion-i-Silvestre

Jaume I University University of Barcelona

Cecilio Tamarit

University of Valencia

January 2008

Abstract

This paper tests for real interest parity (RIRP) among the seventeen major OECD countries over the period 1978:Q1-2006:Q1. The econometric methods applied consist of combining the use of panel data tests that are valid under cross-section dependence and presence of multiple structural breaks. This feature is important since the misspecification errors due to not accounting for structural breaks and/or cross-section dependence can lead to misleading conclusions. Our results support the fulfillment of the weak version of the RIRP for short term interest rate differentials once dependence and structural breaks are considered.

Key words: real interest rate parity, economic integration, panel data tests, structural breaks, cross-section dependence

JEL classification: C32, C33, F21, F32

*M. Camarero and C. Tamarit gratefully acknowledge the financial support from CICYT Project SEJ2005-01163 ECON (Spanish Ministry of Education and FEDER) and Generalitat Valenciana ACOMPLE07/102. Carrion-i-Silvestre gratefully acknowledges the financial support from CICYT Project SEJ2005-08646/ECON.

†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

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. Second, the degree of fulfilment of the real interest rate parity (RIRP hereafter) 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). Third, the RIRP is also an important assumption in several monetary models of exchange rate determination (i.e., Frenkel, 1976).

The literature that tests for the real interest rate parity 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 have 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 have reduced the deviations from uncovered interest parity and from purchasing power parity, the sum of which is the deviation from RIRP. Therefore, the study of real interest rate differentials across countries either under the Bretton-Woods regime or under the present floating exchange rate system that replaced it deserves further attention (Goldberg et al., 2004).

The hypothesis of RIRP has been primarily analyzed in the literature through the assessment of the stochastic properties of the real interest rate differentials. Thus, most of the studies have relied on the application of unit root and stationarity tests to entail whether the shocks affecting the real interest rate differentials have permanent or temporary effects. However, the use of the standard unit root and stationarity statistics can lead to misleading conclusions when the presence of structural breaks is not accounted for – see Perron (1989) and Lee, Huang and Shin (1997). This feature is of special relevance for the analysis of real interest rate differentials since recent literature has found evidence of the presence of structural breaks in the series of real interest rates and inflation for OECD countries. More specifically, Rapach and Wohar (2005) found structural breaks in the two variables considered separately around similar dates, and they claim that infrequent shifts in the mean of real interest rates and inflation are a stylized fact of international macroeconomic data.

The presence of structural breaks affecting either the real interest rates and/or inflation is to be expected associated to episodes such as, among others, (i) large political changes (for example, in party control of either a branch of Congress or the presidency for the US), as in Caporale and Grier (2000) – (ii) regime changes in the process governing the inflation rate – see Rapach and Wohar (2005) – (iii) supply (technology) shocks due to the oil embargo, (iv) expansionary fiscal shocks (Evans, 1985), and (v) in the EU countries, the creation of the European Monetary System and the nominal convergence process implied. Therefore, the presence of structural breaks should be considered when testing for RIRP fulfillment.

To the best of our knowledge, the empirical literature has not considered testing for the real interest parity using panel data allowing for both multiple structural breaks and cross-section dependence. As mentioned above, existing evidence points to the presence of multiple structural breaks on the real interest rate and the inflation rates at the international level, which implies that structural breaks have to be accounted for when analyzing the stochastic properties of the series. Therefore, the aim of this paper is to test for the RIRP among the major OECD countries over the period 1978:Q1-2006:Q1 using panel data statistics. Our main contribution to previous literature on the RIRP is both in terms of economic results and econometric methodology. Thus, when testing for a measure of economic integration, such as the RIRP, conventional panel unit root and stationarity tests may not be adequate as they do not account for cross-section dependence. In addition, some authors have highlighted the importance of structural breaks in influencing the outcome of RIRP tests in panel data analysis.[1] In order to overcome these flaws, we propose a testing strategy aiming at accounting for both dependence and multiple and heterogeneous structural breaks in panels. These techniques are especially well suited when dealing with groups of countries heavily integrated or when using macroeconomic variables with cross-country links. We obtain conclusive evidence in favor of the weak version of the RIRP for the whole group of countries. Finally, the scope of the paper covers different definitions of RIRP attending to inflation expectations – ex-ante and ex-post – that provides a comprehensive analysis of the topic.

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

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

$t = 1, \ldots, T$, where r is the real interest rate, and the asterisk denotes foreign variables. We impose the cointegration vector (1,-1) and then test for the stationarity of the real interest rate differential or rid_t. Since rid_t is assumed to be $iid\,(0, \sigma_v^2)$, 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 rid dissipate and the series return to their long-run zero mean level. This objective can be accomplished by performing unit root and stationarity tests on the series of rid. Consider that rid_t

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

follows a general stochastic process:

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

Following Ferreira and León-Ledesma (2007), the former equation can be represented as a pth-order autoregressive process,

$$\Delta rid_t = a_0 + \delta rid_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta rid_{t-j} + \varepsilon_t. \tag{3}$$

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

$$\delta > 0 \tag{4}$$

$$\delta = 0 \tag{5}$$

$$\delta < 0 \text{ and } a_0 = 0 \tag{6}$$

$$\delta < 0 \text{ and } a_0 \neq 0. \tag{7}$$

Inequality (4) accounts for the case in which the parameter δ is statistically greater than zero. The path of rid_t in this case would be explosive and the series would not converge to any mean in the long-run. In (5) the series contain a unit root and rid_t follows a random walk with shocks affecting the variable on a permanent basis. Both cases, random walks and permanent or explosive rid_t are inconsistent with the RIRP hypothesis.

Conversely, if either (6) or (7) hold, (2) is a stationary process, which means that deviations from the mean are temporary and the estimated root provides information on whether rid_t is short-lived or persistent. In (6) the process converges to a zero mean and a *strong* definition of RIRP holds while in (7) the process converges to a non-zero mean and the *weak* version of RIRP prevails. There are a number of reasons – such as the existence of transaction costs, non-traded goods, non-zero country specific risk premia or different national tax rates – that explain the non-fulfillment of the strong version of the RIRP. Furthermore, the definition of the weak version is consistent with the purchasing power parity concept as defined by Balassa (1964) and Samuelson (1964). Note that, according to Balassa (1964) and Samuelson (1964), the difference in the inflation rates might evolve around a constant different from zero due to the differences in productivity of the economies that are compared. On the contrary, the purchasing power parity hypothesis defined in Cassel (1918) implies that the difference in inflation rates has to evolve around zero if the hypothesis has to be satisfied. Therefore, in the next sections we will concentrate on testing for the weak version of the RIRP, which in fact includes the strong version of the RIRP as a special case.

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. Therefore, we will focus on the literature that directly verifies the RIRP hypothesis making use of different econometric methods. The early studies (Mishkin, 1984 or Cumby and Obstfeld, 1984) were direct tests of the real interest rate equality that performed classical OLS regression analysis and obtained 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., 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 (Holmes and Maghrebi, 2003, and Ferreira and León-Ledesma, 2007) finding more supportive evidence for weak RIRP for various OECD and Asian countries.

Within the group of studies that directly test 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 those that apply classical univariate unit root tests (basically ADF- type) with non-conclusive results. This outcome can be explained by two commonly accepted flaws associated 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 unattended structural breaks. 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 have more power 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] 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.[3]

Bearing these considerations in mind, a newer strand of the empirical literature tries to increase the power of the unit root tests using the recent statistics developed for panel data. The main advantage of the panel tests is that by adding the cross-section dimension, they increase the amount of information available for each time period. In this context, Wu and Chen (1998) and Holmes (2002) have found more promising results using Levin et al. (2002), Maddala and Wu (1999), and Im et al. (2003) panel unit root tests. Notwithstanding this favorable evidence, it is widely recognized that these tests have some flaws in terms of lack of power[4] and size distortion in the presence of correlation among contemporaneous cross-sectional error terms (O'Connell, 1998). Furthermore, an additional source of problems appear when applying these panel data statistics if structural breaks are unatteded.

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

[3] Fountas and Wu (1999), and Goldberg et al. (2004) apply unit root tests allowing for structural breaks, finding rejection of the null in most cases.

[4] Especially in the Levin et al. (2002) test, due to the restrictiveness of the alternative hypothesis. Although this test has higher statistical power than the conventional single-equation unit root test, it requires the coefficient (ρ) of the lagged dependent variables to be homogeneous across all cross-section units of the panel. This implies that all the series should revert to their respective unconditional mean over time at the same rate. This flaw has been overcome by the Im et al. (2003) test, which allows for a higher degree of heterogeneity across cross-sectional units.

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 in 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.

4 Empirical methodology and results

In this paper we investigate the stochastic properties of the real interest differential over the period 1978:Q1 to 2006:Q1 – i.e., post Bretton Woods era. We have chosen this period due to its relevance for the financial integration process both at a global and at an European level, which covers relevant features such as the beginning of the European Monetary System and the launching of the euro. Obviously, a change in the behavior of the RIRP may be expected in the euro-area from 1999 for two reasons. First, the exchange rate regime moved (for more than half of the countries in the sample) from an adjustable peg system to a fixed one. Second, the European Central Bank took over the monetary policy of the euro-area. Thus, nominal short-run interest rates are set centrally and differences in the real rates can only come from distinct country-risk premia and inflation rates.

The sample includes quarterly data of money market interest rates and consumer prices for 17 OECD countries: Australia, Austria, Belgium, Canada, Denmark, France, Germany, Ireland, Italy, Japan, Netherlands, 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 outstanding role within the industrialized economies. The data have been obtained from the International Financial Statistics database of the IMF.

We have chosen onshore short-term asset rates for the analysis because these rates reflect market forces better than deposits ones.[5] In order to account for macroeconomic

[5] 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., 2003).

policy measures, domestic interest rates are the most important. Using offshore rates would prevent capital controls from exerting influence (if any) on the assessment of the RIRP. Although general results overwhelmingly lead to rejection of the RIRP, it is generally accepted that the results 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 (Fuijii and Chinn, 2002). Therefore, our study focuses on the fulfillment of the RIRP with the more demanding short run dataset.[6] The short-run rates are T-bill rates when available for the whole period (Canada, UK and US) and call money rates otherwise.

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. There are two alternative ways to estimate *ex-ante* real interest rates. In the first one, practitioners use survey data to measure expected inflation (i.e., Tanzi, 1985), while in the second, they simulate data using different methods.[7] Alternatively, most of the researchers use *ex-post* real interest rates to test for RIRP. They 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 forecasting ability. If RIRP holds and inflation forecast errors are random, then the observed real interest differential should be random as well.

In order to asses the sensitivity of the results to the (ex-ante or ex-post) nature of the variables, in our study we use both quarterly *ex-post* (RIRPEXPO) and *ex-ante* (RIRPEXA) estimates of real rates of return on short-term securities. For the RIRPEXA we have used the Hodrick and Prescott (1997) filter to proxy price expectations over a time horizon as this filter exhibits the ideal statistical properties for this purpose (Hodrick and Prescott, 1997),[8] while for the RIRPEXPO we have computed the actual CPI annual

[6] All in all, we have additional evidence using the methodology proposed in the paper to 10-year bonds yielding to similar results. These results are available upon request to the authors.

[7] Evans (1985) uses 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.

[8] However, some authors have claimed that all the methods based on simulating the data can be biased. During inflation episodes realized real rates tend to be less than the real rate calculated using the inflation

variation.

Concerning the empirical methodology, we have applied panel stationarity tests following a two-step testing strategy that addresses the problems related to the issues of multiple structural breaks and cross-section dependence.

First, we have tested for the RIRP allowing for multiple structural changes in a panel setting which, to the best of our knowledge, has not been applied yet in this literature. Previous evidence has revealed that there might be some events that affect real interest rates in a permanent way. It is well known that non accounting for structural breaks biases both unit root and stationarity tests towards concluding in favor of non-stationarity in variance.[9] 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. Second, we consider the existence of cross-section dependence amongst the individuals in the panel. 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 and stationarity 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. Therefore, 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, in addition to testing the null hypothesis of cross-section independence, it provides guidance about the best way to model cross-section dependence.

The application of this statistic reveals that cross-section dependence is present in the panel data sets that are studied. Then, our analysis considers two different ways to accommodate cross-section dependence. First, following the approach by Carrion-i-Silvestre et al. (2005) we compute the bootstrap critical values of the panel data stationarity test

forecast, and conversely, when inflation falls, the realized real rate lies well above the predicted real rate (Darin and Hetzel, 1995).

[9] See Perron (1989) for univariate statistics, or Carrion-i-Silvestre, del Barrio and López-Bazo (2001) for panel data statistics.

statistic, which allows us to consider a wide form of cross-section dependence. Second, we compute the panel stationarity statistic proposed by Harris et al. (2005), which models the presence of cross-section dependence through the estimation of approximate common factor models as in Bai and Ng (2004). In both cases, the analysis considers the existence of the estimated structural breaks. It is worth mentioning that the approach that is adopted here is general enough to consider the non-break situation as a particular case embedded in the testing procedure.

Finally, note that 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 Testing for the presence of multiple structural breaks

The discussion above based on previous evidence reported in the literature suggests that the real interest rates series might suffer the effects of structural breaks. Unattended structural breaks may affect the statistical inference, as the tests would be biased to conclude in favor of non-stationarity. The first stage of our analysis consists of assessing the presence of structural breaks using the following specification:

$$rid_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{i,k,t} + \varepsilon_{i,t}, \tag{8}$$

$t = 1, \ldots, T$, $i = 1, \ldots, N$, with $DU_{i,k,t} = 1$ for $t > T^i_{b,k}$ and 0 elsewhere – $T^i_{b,k}$ denotes the kth break point for the ith individual, $k = 1, \ldots, m_i$ – and where $\{\varepsilon_{i,t}\}$ are assumed to be a stationary process satisfying the strong-mixing conditions given in Phillips (1987) and Phillips and Perron (1988). This specification permits a high degree of heterogeneity assuming that the structural breaks may have different effects on each individual time series. For this purpose, the break points are located at different dates for each individual, and the individuals may have different number of structural breaks. Under these conditions, the estimation of the number and position of the structural breaks, if any, can be carried out using the sequential testing procedure proposed 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 $m_i = 5\ \forall i$. The number of structural breaks is estimated using critical values at the 5% level of significance. It is worth mentioning that the application of the Bai-Perron methodology to estimate the number and position of the structural breaks requires the variables under analysis to be stationary in variance, which is consistent with the null hypothesis that we have specified, i.e., that the RIRP hypothesis holds. Furthermore, the test statistic that is used is consistent against the alternative hypothesis of non-stationarity in variance, even when structural breaks are present in the analysis – see Lee, Huang and Shin (1997), Kurozumi (2002) and, Carrion-i-Silvestre (2003), among others.

Table 1 reports the estimated number and position of the structural breaks for each individual in the two panel data sets. We can see that, except for the RIRPEXPO of Belgium and Canada, the procedure detects at least one structural break for each individual, which indicates that previous analyses in the literature that do not account for the presence of structural breaks may have obtained misleading conclusions. Except for the Australian RIRPEXPO interest rate, it should be stressed that the estimated number of structural breaks does not attain the maximum that has been defined.[10]

Figures 1 and 2 present the pictures of the RIRPEXPO and RIRPEXA for all the countries involved in our analysis, along with the estimated deterministic component. The countries have been divided according to their condition of EU members during the analyzed period. This presentation allows us to establish a comparison of the break dates and the direction of the changes that have been estimated. The break points that have been estimated here are related to some important monetary policy changes in the analyzed period. In order to ease interpretation, we have computed the 95% confidence intervals for the estimated break points. This allows us to get a better picture when

[10]In order to check if the maximum number of breaks was correct for all the countries in the sample, we have increased the maximum number of structural breaks for Australia to up to eight. However, the procedure selected just five structural breaks for this country.

identifying short time periods where break points are located accounting for the fact that the same event might have affected different individuals, although not at the same precise moment. We have also included a summary of the breaks in Table 5. The purpose of this table is just to highlight some economic and political events that can be associated with the changes detected in the real interest differential with the US. Although in some cases there may be more than one explanation for some of them, we are interested in detecting the changes that are common to a group of countries in the sample.

According to the results reported in Table 2, the sample period can be truncated into up to four breakpoints (with one exception). Table 5 tries to simplify the information concerning the estimated break dates. 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). At the same time, the post-Bretton Woods era knew a first removal of capital controls in the OECD. 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 monetary integration in the EU. This process meant the progressive abolition of any remaining capital controls among the European countries by 1990. However, the strong appreciation of the dollar during the eighties that ended after the Plaza Agreements may also be behind this group of structural changes. A third break is placed, in many cases, around 1990-1993, which coincides with German unification in July 1990. This fact generated a large asymmetric shock that boosted the EMS crisis in September 1992 and the exit of Italy and the UK from the exchange rate mechanism of the EMS. Moreover, in August 1993, took place a formal widening of exchange rate bands of the EMS to $\pm 15\%$ followed to the adherence of the prospective euro members to the Maastricht conditions on nominal convergence. In addition, three of the more inflationary countries in the EMS reduced drastically their interest rates to adapt the economies to the Maastricht criteria. This may explain the structural changes of Portugal, Ireland and Spain at the end of the nineties. There are other monetary policy factors of importance for the rest of the OECD countries in the sample. For example, Canada decided to apply an inflation targeting strategy in 1991, whereas Japan suffered from severe deflation from 1995 to 2001. In addition, Australia was hit by the Asian crisis in 1997-98 and the international convulsion after September 11th 2001 may also explain some of the structural breaks found at the beginning of the present decade.

Once the break points have been dated, we proceed to analyze the order of integration of the rid_t time series. The estimation of the model in (8) with the break points that have been obtained above can be used to compute individual KPSS statistics given by

$$\hat{\eta}_i(\lambda_i) = \hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}_{i,t}^2, \tag{9}$$

where $\hat{S}_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ is the partial sum process that is obtained using the estimated OLS residuals of (8), $\hat{\omega}_i^2$ denotes a consistent estimate of the long-run variance of the error term $\varepsilon_{i,t}$, which, based on the evidence reported in Carrion-i-Silvestre and Sansó (2006), has been estimated following the procedure described by Sul et al. (2005), using the Quadratic spectral kernel. In (9), λ_i is defined as the vector $\lambda_i = (\lambda_{i,1}, ..., \lambda_{i,m_i})' = \left(T_{b,1}^i/T, ..., T_{b,m_{i,j}}^i/T\right)'$, which indicates the relative position of the dates of the breaks on the entire time period T for each individual. Thus, the computation of the individual KPSS statistic permits to get a first analysis of the stochastic properties of the real interest rates. Table 1 offers the computation of the individual KPSS along with the corresponding simulated critical values at the 5 and 10% level of significance. Looking at these results we can conclude that there is mild evidence against the null hypothesis of variance stationarity, as the null is rejected at the 5% level for Australia, Austria and Portugal, and at 10% for Ireland for the ex-post variable. For the ex-ante one, the null hypothesis is rejected for Austria and Italy at the 5 and 10% levels of significance respectively.

This individual based inference can be improved if we combine the individual statistics through the definition of panel data statistics. Thus, the literature on non-stationary panel data statistics argues that a better characterization of the stochastic properties of the time series can be obtained if we increase the amount of information when performing the inference. However, some cautions have to be taken into account when computing these panel-data-based statistics, since some of them rely on the critical assumption of cross-section independence. This assumption is investigated in the next section for our panel data set.

4.2 The issue of cross-section independence

The independence assumption imposed in the so-called first generation panel data statistics has been widely criticized in the 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 et al. (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. We can allow for the presence of the structural breaks when testing the null hypothesis of non correlation amongst individuals in the panel. We will then estimate an autoregressive model to isolate cross-section dependence from the autocorrelation that might be driving the individual time series. 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.

In brief, the procedure works as follows. First, we get rid of the autocorrelation pattern in the individual time series through the estimation of an AR model. 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 ($\bar{p}_j = |\hat{p}_j|$) 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 by minimizing 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 some trimming is required. Following Ng (2006) the trimming is set 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) in Ng (2006), which under the null hypothesis converges to the standard normal distribution.

The results in Table 3 show that the null hypothesis of non cross-section correlation is rejected for the whole and L samples of spacings, while it is not rejected for the S sample at the 5% level of significance, regardless of the data set that is used. 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. In this case, the factor models suggested by Bai and Ng (2004) is a suitable approximation to account for this type of cross-section dependence in panels. Therefore, the evidence that is obtained in this section indicates that cross-section dependence has to be considered when computing the panel data statistics if misleading conclusions are to be avoided.

4.3 Panel data tests with cross-section dependence and structural breaks

The specification estimated above permits the computation of two different panel data stationarity statistics. First, we have applied the approach suggested in Carrion-i-Silvestre et al. (2005) to test the null hypothesis of panel variance stationarity allowing for multiple level shifts. Thus, note that the specification given in (8) is one of the two models considered by these authors. The OLS estimated residuals from (8) are used to obtain the individual KPSS statistics computed in the previous sections, which in turn can be combined to define the panel stationarity test statistic:

$$LM(\lambda) = N^{-1} \sum_{i=1}^{N} \hat{\eta}_i (\lambda_i),$$

with $\hat{\eta}_i (\lambda_i)$ defined in (9). Note that $\hat{\eta}_i (\lambda_i)$ has been defined such that the long-run variance is heterogenous across individuals. However, it would be possible to use an homogeneous estimate of the long run variance, i.e., $\hat{\omega}^2 = N^{-1} \sum_{i=1}^{N} \hat{\omega}_i^2$. Using these elements we can define the panel data statistic $Z(\lambda) = \sqrt{N} \left(LM(\lambda) - \bar{\xi} \right) \Big/ \bar{\varsigma}$, where $\bar{\xi} = N^{-1} \sum_{i=1}^{N} \xi_i$ and $\bar{\varsigma}^2 = N^{-1} \sum_{i=1}^{N} \varsigma_i^2$, with ξ_i and ς_i^2 being the individual mean and variance of $\eta_i(\lambda_i)$ respectively. Note that these two possibilities for the definition of the long-run variance estimate gives rise to two different statistics, i.e., the $Z(\lambda)$ when the long-run variance homogeneity is imposed and the $Z(\lambda)$ for heterogeneous long-run variance.

Under the null hypothesis of variance stationarity and assuming cross-section independence, the $Z(\lambda)$ panel data statistics are shown to converge to the standard normal distribution. However, this limiting result is not obtained when individuals are cross-section dependent, as it is in our case. In this situation, we can compute the bootstrap distribution of the $Z(\lambda)$ statistics to account for the presence of a general form of cross-section dependence. The computation of the bootstrap distribution follows the lines given in Maddala and Wu (1999). To be specific, we have defined the $(T \times N)$-matrix of the OLS estimated residuals from (8) $\hat{\varepsilon} = (\hat{\varepsilon}_1, \ldots, \hat{\varepsilon}_N)$, and have resampled with replacement the rows of the $\hat{\varepsilon}$ matrix so that the first matrix of resampled residuals $\hat{\varepsilon}^{*(1)}$ is obtained, where the superscript " $*(1)$ " indicates the first resampling. Conditional on the estimated parameters and structural breaks, we have computed the bootstrap variables

$$rid_{i,t}^{*(1)} = \hat{\alpha}_i + \sum_{k=1}^{\hat{m}_i} \hat{\theta}_{i,k} DU_{i,k,t} + \varepsilon_{i,t}^{*(1)},$$

for each i, where $\hat{\alpha}_i$ and $\hat{\theta}_{i,k}$ are the OLS estimates of the parameters in (8). This is repeated 2,000 times so that we define $rid_{i,t}^{*(1)}, \ldots, rid_{i,t}^{*(2,000)}$ series for each individual, which can be used to approximate the empirical distribution of the $Z(\lambda)$ statistics.

Table 4 presents the $Z(\lambda)$ statistics as well as the bootstrap critical values. According to these statistics, the null hypothesis of variance stationarity cannot be rejected regardless of the assumption made about the long-run variance estimation.

Although we have already obtained favorable results on the fulfillment of the RIRP, our previous evidence of pervasive dependence found using Ng (2006) tests, recommends using common factors as a better approximation to this type of dependence. Thus, we have also computed the panel data stationarity statistic by Harris et al. (2005), which controls for the presence of cross-section dependence through the estimation of common factor models defined in Bai and Ng (2004). According to Ng (2006), when cross-correlation is pervasive, this approach has better properties. Moreover, it controls for cross-section dependence given by cross-cointegration relationships, where individuals in the panel might be cointegrated – see Banerjee et al. (2004), and Gegenbach et al. (2004). For this purpose, we decompose the estimated OLS residuals $\hat{\varepsilon}_{i,t}$ obtained from

(8) in two different components, i.e., the idiosyncratic component $\left(\xi_{i,t}\right)$ and the common factor component that is given by the $(r \times 1)$ vector F_t of factors. The idiosyncratic disturbance terms and the common factors are estimated using principal components on the first difference of $\hat{\varepsilon}$. The estimated factors $\hat{f}_{1,t}, \ldots, \hat{f}_{r,t}$ are the r eigenvectors that correspond to the r largest eigenvalues of the $(T-1 \times T-1)$ matrix $\Delta\hat{\varepsilon}\Delta\hat{\varepsilon}'$ – see Harris et al. (2005) for further details on the estimation of the common factors and the idiosyncratic disturbance terms. The number of common factors r can be consistently estimated using the panel BIC information criterion in Bai and Ng (2002) – here we specify a maximum of $r_{\max} = 6$ common factors.

The test statistic by Harris et al. (2005) is given by $S_k^F = \left(\hat{C}_k + \hat{c}\right) / \hat{\omega}\left\{\hat{a}_{k,t}\right\}$, with $\hat{C}_k = T^{-1/2} \sum_{t=k+1}^{T} \hat{a}_{k,t}$ the autocovariance of order k, $\hat{a}_{k,t} = \sum_{i=1}^{N+\hat{r}} \hat{z}_{i,t}\hat{z}_{i,t-k}$, and $\hat{z}_{i,t}$ as the ith element of the $(N+\hat{r}) \times 1$ vector $\left(\hat{F}_{1,t}, \ldots, \hat{F}_{\hat{r},t}, \hat{\xi}_{1,t}, \ldots, \hat{\xi}_{N,t}\right)'$, which contains the estimated common factors $\left(\hat{F}\right)$ and the idiosyncratic disturbance $\left(\hat{\xi}_i\right)$, with $\hat{c} = (T-k)^{-1/2} \sum_{i=1}^{N} \hat{c}_i$, being $\hat{c}_i$ a correction term defined in Harris et al. (2005) and, $\hat{\omega}^2\left\{a_t\right\}$ is a consistent estimate of the long-run variance of $\left\{a_t\right\}$. Under the null hypothesis of joint variance stationarity of the common and idiosyncratic components the statistic $S_k^F \rightarrow^d N(0,1)$. In this paper we follow Harris et al. (2005) and use $k = \left[(3T)^{1/2}\right]$.

The results that have been obtained with the $Z(\lambda)$ statistics are confirmed when we compute the Harris et al. (2005) statistic allowing for multiple level shifts and common factors using the break points estimated above, since the S_k^F statistic in Table 4 does not reject the null hypothesis of variance stationarity at the 5% level of significance for the short-run real interest rates.

To sum up, our results show that there is strong evidence of the weak version of real interest rate parity, once structural breaks and cross-section dependence are allowed for, regardless of whether ex-ante or ex-post real interest rates are used. This conclusion builds upon, first, the individual analysis that is carried out in previous section where the null hypothesis of stationarity was only rejected in a few cases and, second, the application of panel data statistics that are robust to the presence of cross-section dependence.

5 Conclusions

Although many studies have reexamined the real interest rate parity condition, few have been able to obtain its fulfilment empirically, especially for short-term interest rates. In this paper we present new evidence in support of long-run reversion in short-term 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 other 16 major industrial economies from 1978:Q1 to 2006:Q1. Our analysis is based on the use of panel data stationarity test statistics that accommodate the presence of cross-section dependence and 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 over time. We focus on two 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 a long-run equilibrium.

The results crucially depend on the allowance of both 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. The statistical procedures that have been applied in the paper reveal that these features are present in our setting. Thus, once we consider both of these characteristics we conclude in favor of 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 time series unit root and stationarity 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 failure in the RIRP hypothesis.

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Table 1: Estimated structural breaks and individual KPSS statistics for the RIRPEXPO and RIRPEXA data sets

Panel A: RIRPEXPO Individual information

	Tests	m_i	$T^i_{b,1}$	$T^i_{b,2}$	$T^i_{b,3}$	$T^i_{b,4}$	$T^i_{b,5}$	Critical values 10%	5%
Australia	0.058	5	84:1	88:4	93:2	98:1	02:1	0.044	0.050
Austria	0.073	4	81:4	89:3	94:1	01:1		0:058	0:068
Belgium	0.052	0	-	-	-	-		0:349	0:457
Canada	0.142	0	-	-	-	-		0:352	0:468
Denmark	0.045	3	89:4	94:1	00:4	-		0:087	0:108
France	0.039	4	84:4	89:4	94:1	00:4		0:055	0:063
Germany	0.037	4	81:4	89:4	94:1	00:4		0.058	0.068
Ireland	0.075	4	84:3	94:1	98:1	02:1		0.066	0.079
Italy	0.078	3	84:3	97:2	01:3	-		0.095	0.118
Japan	0.053	3	90:1	94:3	01:1	-		0.088	0.110
Netherlands	0.043	3	89:3	94:1	02:1	-		0.089	0.109
Norway	0.038	4	84:3	90:1	94:1	01:1		0.055	0.064
Portugal	0.093	4	85:1	90:4	96:1	01:4		0.054	0.063
Spain	0.050	4	86:4	93:4	98:1	02:1		0.063	0.075
Switzerland	0.051	4	85:4	89:4	94:4	00:4		0.056	0.066
United Kingdom	0.037	4	84:3	89:3	93:4	01:1		0.056	0.065

Panel B: RIRPEXA Individual information

	Tests	m_i	$T^i_{b,1}$	$T^i_{b,2}$	$T^i_{b,3}$	$T^i_{b,4}$	$T^i_{b,5}$	Critical values 10%	5%
Australia	0.027	4	85:1	93:2	97:2	01:3		0.061	0.072
Austria	0.070	4	81:4	89:3	94:1	01:1		0.058	0.068
Belgium	0.070	3	90:2	94:2	01:1	-		0.091	0.114
Canada	0.058	4	85:4	89:4	95:3	02:1		0.059	0.068
Denmark	0.040	3	89:3	94:3	01:1	-		0.085	0.105
France	0.059	3	89:3	95:3	01:1	-		0.083	0.103
Germany	0.079	3	90:1	94:2	00:4	-		0.089	0.11
Ireland	0.050	4	85:4	90:2	94:2	98:3		0.061	0.071
Italy	0.059	4	82:2	90:3	96:3	01:1		0.058	0.067
Japan	0.063	3	90:1	94:2	01:2	-		0.089	0.109
Netherlands	0.074	3	89:3	94:2	01:1	-		0.085	0.103
Norway	0.042	3	86:1	94:1	01:1	-		0.071	0.083
Portugal	0.022	3	83:1	89:4	95:2	-		0.079	0.095
Spain	0.048	4	86:1	90:1	94:1	98:1		0.063	0.073
Switzerland	0.051	3	89:3	94:3	00:4	-		0.083	0.103
United Kingdom	0.024	4	84:4	89:3	93:3	01:2		0.058	0.068

Table 2: 95% Confidence interval for the estimated break points

	RIRPEXPO				
	$T^i_{b,1}$	$T^i_{b,2}$	$T^i_{b,3}$	$T^i_{b,4}$	$T^i_{b,5}$
Australia	(83:3, 85:2)	(88:2, 89:3)	(92:3, 93:3)	(96:1, 99:1)	(01:3, 02:3)
Austria	(80:4, 83:2)	(89:1, 90:1)	(93:3, 94:2)	(00:2, 01:2)	-
Belgium	-	-	-	-	-
Canada	-	-	-	-	-
Denmark	(88:4, 90:3)	(93:3, 94:3)	(00:1, 03:1)	-	-
France	(84:2, 86:1)	(89:1, 90:1)	(93:3, 94:3)	(99:4, 03:4)	-
Germany	(80:4, 84:1)	(89:1, 90:1)	(93:3, 94:3)	(99:4, 01:2)	-
Ireland	(83:1, 85:4)	(93:4, 94:2)	(97:2, 98:2)	(01:1, 03:1)	-
Italy	(83:4, 85:1)	(97:1, 98:2)	(01:2, 02:3)	-	-
Japan	(89:2, 92:3)	(94:1, 94:4)	(00:3, 01:3)	-	-
Netherlands	(89:1, 91:4)	(93:3, 94:2)	(01:2, 02:3)	-	-
Norway	(84:1, 86:4)	(87:3, 90:2)	(93:3, 95:1)	(99:1, 02:2)	-
Portugal	(84:3, 86:3)	(90:1, 91:2)	(95:3, 96:3)	(00:4, 06:1)	-
Spain	(86:2, 88:4)	(93:2, 94:4)	(96:4, 98:4)	-	-
Switzerland	(85:2, 88:2)	(89:1, 90:1)	(94:2, 95:2)	(00:2, 01:1)	-
United Kingdom	(84:1, 85:4)	(88:3, 90:2)	(93:2, 94:2)	(00:3, 01:4)	-
	RIRPEXA				
	$T^i_{b,1}$	$T^i_{b,2}$	$T^i_{b,3}$	$T^i_{b,4}$	$T^i_{b,5}$
Australia	(84.3, 85.4)	(92.4, 97.1)	(96.4, 97.3)	(01.1, 01.4)	-
Austria	(81.2, 85.3)	(88.4, 89.4)	(93.3, 94.3)	(00.2, 01.2)	-
Belgium	(89.3, 90.4)	(93.4, 94.3)	(99.4, 01.3)	-	-
Canada	(84.2, 89.4)	(88.3, 90.3)	(95.1, 96.1)	(00.4, 03.2)	-
Denmark	(88.1, 90.3)	(94.1, 95.2)	(00.2, 02.1)	-	-
France	(88.3, 90.1)	(95.1, 95.4)	(00.1, 01.2)	-	-
Germany	(80.2, 86.1)	(89.2, 90.2)	(93.4, 94.3)	(00.1, 01.1)	-
Ireland	(84.4, 87.3)	(86.1, 90.4)	(94.1, 94.3)	(96.4, 98.4)	-
Italy	(81.4, 83.1)	(88.3, 90.4)	(96.1, 97.2)	(00.2, 02.4)	-
Japan	(89.2, 92.1)	(93.4, 94.3)	(00.4, 01.3)	-	-
Netherlands	(88.4, 90.3)	(93.4, 94.3)	(00.3, 01.2)	-	-
Norway	(84.4, 87.1)	(93.3, 95.4)	(99.1, 01.4)	-	-
Portugal	(82.3, 84.2)	(89.2, 90.1)	(94.4, 95.4)	-	-
Spain	(85.2, 90.3)	(89.1, 91.4)	(93.3, 94.3)	(96.1, 99.2)	-
Switzerland	(89.1, 90.2)	(94.1, 95.1)	(00.1, 01.1)	-	-
United Kingdom	(84.2, 86.1)	(88.4, 90.3)	(93.1, 93.4)	(00.3, 02.1)	-

Table 3: Spacing Variance Ratio statistic for the RIRPEXPO and RIRPEXA panels with level shifts

	Whole sample		Small group			Large group	
	$svr\ (\eta)$	p-val	$svr\ (\eta)$	p-val	$\hat{\eta}$	$svr\ (\eta)$	p-val
RIRPEXPO	3.564	0.000	-0.987	0.838	12	1.903	0.029
RIRPEXA	1.783	0.037	-0.467	0.680	14	4.116	0.000

Table 4: Panel data stationarity tests with multiple structural breaks and cross-section dependence

	RIRPEXPO		RIRPEXA	
		Bootstrap		Bootstrap
	Test	5% crit.val.	Test	5% crit.val.
$Z(\lambda)$ Homogeneous	0.438	3.700	0.833	9.858
$Z(\lambda)$ Heterogeneous	0.361	3.270	1.631	6.943
	Test	p-value	Test	p-value
S_k^F	1.097	0.136	1.384	0.083

Table 5: Summary of the estimated structural changes

Structural breaks and events	EU countries	Non-EU countries
Early 80's:		
Post oil shocks	AUS, GER, ITA	
Removal K controls		
Mid 80's:		
Plaza Agreement		AUSL, CAN, NORW, SWIZ
Basilea-Nyborg (EMS)	FR, IRE, ITA, POR, SPA, UK	
Reduction in fiscal deficits		AUSL
Early 90's:		
German unification	AUS, DK, FR. GER, BEL, NET, POR, UK	
Mid 90's:		
EMS crisis	AUS, DK, FR, GER, BEL, NET, POR, SPA, UK	NORW
Japanese deflation		JAP
Maastricht criteria recession	POR, IRE, SPA	
End 90's - 2001:		
Asian crisis		AUSL
Launching of the euro	AUS, GER, DK, FR, GER, IRE, ITA, BEL NET, POR, UK	SWIZ
September 11th 2001		AUSL, CAN, JAP, NORW

Figure 1. Ex-post RIRP (RIRPEXPO)

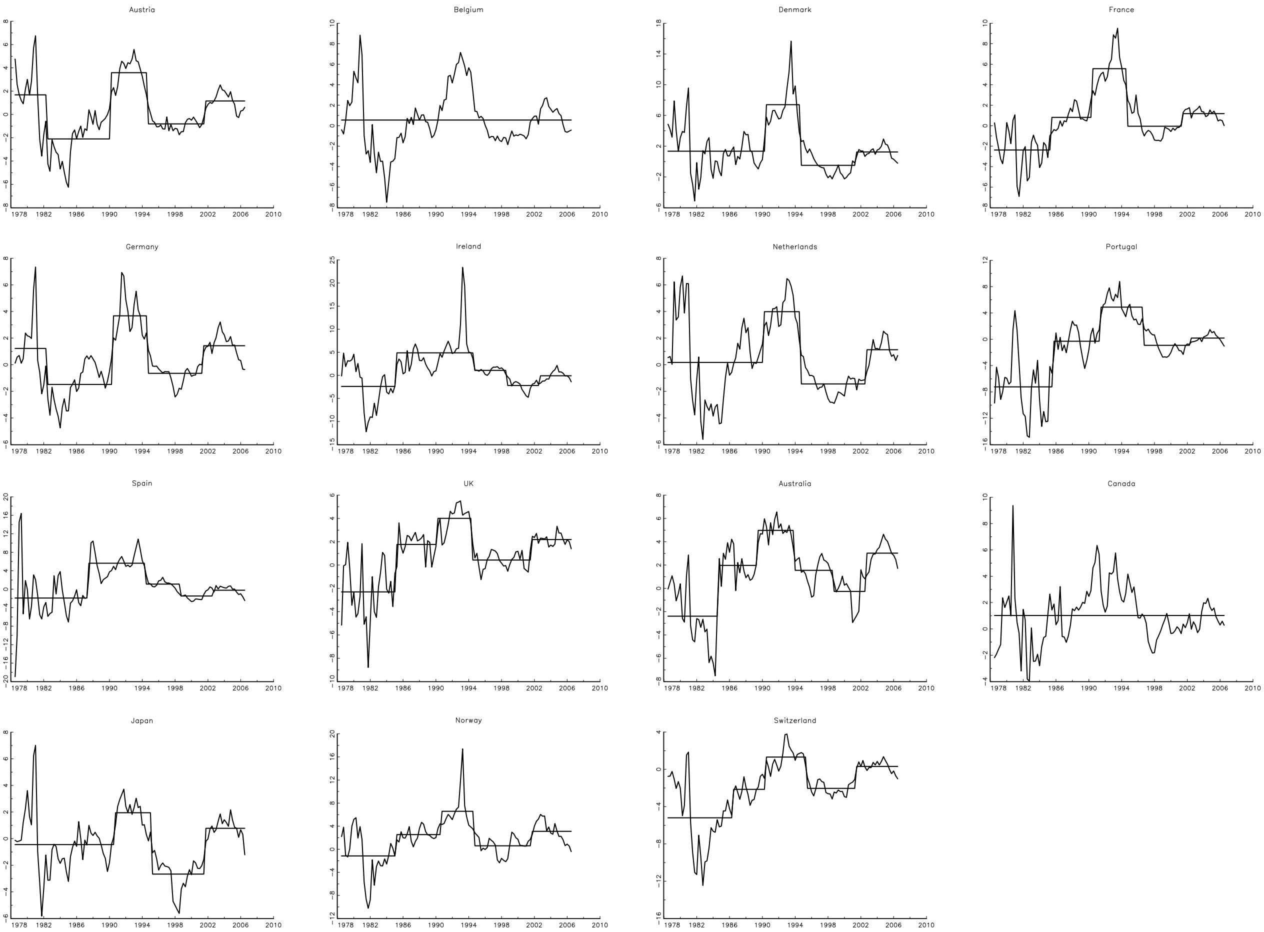

Figure 2. Ex-ante RIRP (RIRPEXA)