New evidence of the real interest rate parity for OECD countries using panel unit root tests in a "*SURE*" framework.*

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Abstract

This paper tests for real interest parity (RIP) among the six major OECD countries over the period 1957-2003. The econometric methods applied consist of combining the use of univariate and multivariate unit root tests in an attempt to gain in power and size properties. In addition to univariate ADF tests where the lags are selected using a Modified AIC criterion, we apply a sequential procedure based in two multivariate ADF-type panel unit root tests in a SURE framework. This methodology offers substantial advantages over the univariate ADF tests and panel tests that assume independence between the cross-sections. Our results strongly support the fulfillment of RIP for the whole period considered.

Key words: Real interest rate pariy, economic integration, panel unit root tests, SURE.

JEL classification: F32, F21, C32, C33.

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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 a open economy, real interest rates are an important channel for transmission of macroeconomic policies. Secondly, the degree of fulfillment of the real interest parity (RIP) can be used as a criterion to measure market integration because RIP 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, RIP 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 RIP 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 RIP 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 are expected to reduce the deviations from uncovered interest parity and from purchasing power parity, the sum of which is the deviations from RIP. Thus, the study of real interest rate differentials across countries under the Bretton-Woods regime and the present of floating exchanges that replaced it deserves further attention (Goldberg et al., 2002).

The aim of this paper is to test for RIP among the major OECD countries over the period 1958-2003 using unit root tests. The econometric methods applied consist of combining the use of univariate and multivariate unit root tests with good power and size properties. Thus, the main contribution made by this study to the literature on RIP is in terms of the econometric methodology. First, we apply ADF unit root tests where the lag truncation is selected using a method that avoids over-rejecting the null hypothesis, that is, to assure a good size of the tests. Second, we apply multivariate ADF-type unit root tests that have better power properties than their univariate counterparts. Third, these multivariate tests use the information content in the variance-covariance matrix, so that the unrealistic assumption of cross-section independence made in the majority of panel tests can be avoided. Fourth, the tests allow for an important degree of heterogeneity: the lag order of the augmented test can vary among the individuals using in each case the one selected previously in the univariate analysis and the autoregressive parameter β_0 can also differ for every cross-section. Finally, the sequential use of two complementary tests permits us to cluster the countries in two groups: those who fulfill the RIP and those that do not.

The remainder of the paper is organized as follows. Section 2 briefly presents the theoretical background. The third section reviews the previous relevant literature while Section 4 describes the techniques and tests used in the paper. Section 5 presents the data and the econometric results. Finally, Section 6 concludes.

2 Theoretical issues.

The extent of deviation from RIP is a measure of the lack of goods and financial market integration. This can be seen more clearly by deriving the RIP 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* RIP 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

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

$$ds_t^e = \pi_{t+1}^e - \pi_{t+1}^{*e} \quad (5)$$

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

$$r_{t+1}^e = r_{t+1}^{*e} \quad (6)$$

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

$$r_{t+1} = r_{t+1}^e + u_{t+1} \quad (7)$$

and

$$r_{t+1}^* = r_{t+1}^{*e} + u_{t+1}^* \quad (8)$$

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

$$r_{t+1}^* - r_{t+1}^* = v_{t+1} \quad (9)$$

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

$$r_t - r_t^* = v_t = rid_t \quad (10)$$

This last equation (10) can be used to test RIP in a univariate framework which by imposing the cointegration vector (1,-1) a priori and testing for the stationarity of the error term v_t. Since v_t are iid N(0, σ_ε^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:

$$rid_t = a_0 + a_1 rid_{t-1} + \varepsilon_t \quad (11)$$

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

$$\Delta rid_t = a_0 + \Psi rid_{t-1} + \sum_{j=1}^{k} \beta_j \Delta rid_{t-j+1} + \varepsilon_t \qquad (12)$$

where,

$$\Psi = \sum_{j=1}^{k} a_{j-1.} \qquad (13)$$

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

$$\Psi > 0 \qquad (14)$$

$$\Psi = 0 \qquad (15)$$

$$\Psi < 0 \text{ and } a_0 = 0 \qquad (16)$$

$$\Psi < 0 \text{ and } a_0 \neq 0 \qquad (17)$$

Inequality (14) accounts for the case in which the parameter Ψ 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 (15) 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 RIP hypothesis

On the contrary, if either (16) or (17) 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 (16) the process converges to a zero mean and a *strong* definition of RIP holds while in (17) the process converges to a non zero mean and a *weak* version of RIP prevails. It is worth to note that strong RIP 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 RIP 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 made of studies on RIP with implications for other hypotheses or theories and to some extent can be used as indirect evidence of RIP. This literature includes research on the analysis of the international monetary policy transmission and the currency dominion hypothesis (i.eg. Chinn and Frankel (1995) for the Pacific Rim case[1]), the existence of time-varying risk premia on foreign exchange series (Nieuwland et al, 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 the one devoted to verify the RIP hypothesis in itself 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 tecniques (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 et al., 2003) finding more supportive evidence for weak RIP for various OECD and Asian countries.

Within this second group of studies of direct testing for RIP, 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 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, especially in small samples. 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 leads to more promising results.

[1]Recent papers on the subject include Wu and Fountas (2000), Chinn and Frankel (2003) and Frankel et al. (2003).

Some other authors try to find a more accurate evidence enlarging the sample period considered[2]. However, 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]. 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. Thus, a very promising 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 every time period. In this context, Wu and Chen (1998) reject a unit root in the majority of the countries considered using the Levin and Lin (1992), Im, Pesaran and Shin (2003) and Maddala and Wu (1996) panel unit root tests and Holmes (2002) is also able to reject the null using the IPS tests for a different set of countries and sample periods. However, it is widely recognized that these tests have some flaws connected to the size of the tests, cross-section dependence and the restrictive null hypothesis.

In this paper we present a testing procedure that overcomes previous problems in panel unit root tests. We contribute to the empirical literature on the RIP on various respects. First, we use univariate tests with good size and power properties so that some of the problems traditionally found in the empirical literature can be avoided. Second, the multivariate tests applied do not impose, as the majority of the panel tests, cross-section independence. This posibility is crucial in this type of analysis, when the variables are closely linked in a globalized environment. Finally, the use of a sequential procedure avoids another drawback associated with panel unit root tests: the presence of a single stationary process may cause the rejection of the null hypothesis. In our case, after rejecting the null, we apply additional individual tests that can cluster the variables into the stationary and the non-stationary ones.

4 Empirical methodology.

In this paper we test the null hypothesis of unit root in the real interest differential of all possible country pairs combinations made up of 6 OECD countries, namely Canada, France, Germany, Japan, the UK, and the United States, over the period 1958Q1 to 2003Q3. Therefore, we test the fulfillment

[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. (2002) apply unit root tests that allow for structural breaks in the series finding rejection of the null in more cases.

of the real interest parity using univariate and also multivariate ADF-type tests estimated using SUR methods. We apply first, the ADF tests as a benchmark in the analysis and the basis to compute their multivariate equivalent and apply a modified AIC ($MAIC$) selection rule for k, to gain size and power.

Then, the multivariate tests are applied in two stages. First, the multivariate ADF test ($MADF$), that is a Wald test proposed by Sarno and Taylor (1998), where rejection of the null means that the RIP is fulfilled in some cases. However, as this test is a joint test, rejection does not provide information about how many panel members follow the null and how many don't, being impossible to identify which are the stationary and non-stationary cross-section elements. The second multivariate test applied here, and proposed by Breuer, McNown and Wallace (2002), on the contrary, identify which variables contain a unit root and which do not. Thus, it complements the $MADF$ test, and should be applied in a second stage of the analysis.

In addition, the choice of the individual cross-section lags in the $MADF$ test is made using the same $MAIC$ selection rule as in the univariate case.

4.1 Truncation lag selection for the unit root tests.

Testing for unit roots is a difficult task, due to problems of size and power associated with the tests. Size problems can arise from a near common factor in the moving-average and autoregressive polynomials in the time series ARMA representation. Thus, Ng and Perron (2001) argue that the selection of the lag truncation (k) plays a crucial role in the size of the unit root test. Traditional information criteria, such as the AIC and the BIC tend to select a truncation lag that is too low. This, in turn, can provoke Type I error (that is, rejecting the null hypothesis of non-stationarity when true). In particular, when there are errors with a moving-average root close to -1, a high order augmented autorregression would be necessary to avoid over-rejecting the null hypothesis of a unit root. In order to account for this type of problems, they suggest using instead a modified AIC ($MAIC$) with a penalty factor that is sample dependent.

Ng and Perron (2001) emphasize that many economic variables exhibit a large negative moving-average. The inflation rate, where omitted outliers induce a negative moving average root in the error process, would be a candidate for size distortions[4]. As the inflation rate is used to compute the real interest rate variables in this paper, we select the number of lags in the ADF

[4] See Ng and Perron (2002) for a discussion and an empirical application to the US inflation rate.

tests using the $MAIC$.

They define a stochastic process such that

$$y_t = d_t + u_t, \qquad u_t = \alpha u_{t-1} + v_t \tag{18}$$

with $E(u_0^2) < \infty$, $v_t = \delta(L)e_t = \sum_{j=0}^{\infty} \delta_j e_{t-j}$ with $\sum_{j=0}^{\infty} j \mid \delta_j \mid < \infty$ and $\{e_t\} \sim i.i.d.(0, \sigma_e^2)$.

$$\Delta y_t = d_t + \Psi y_{t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_{tk} \tag{19}$$

where $d_t = \psi' z_t$, and z_t is a set of deterministic components, $d_t = \sum_{i=0}^{p} \psi_i t^i$, with special focus on $p = 0, 1$.

To implement this correction but, at the same time, avoiding selecting a large k when not needed, the modified information criterion, $MAIC$, can be written as follows:

$$MAIC = -2\left(\frac{\ell}{T}\right) + \frac{2(k+\tau)}{T} \tag{20}$$

with

$$\tau = (\hat{\sigma}_e^2)^{-1} \hat{\Psi} \sum_{t=1}^{T} y_{t-1}^2$$

where ℓ is the likelihood function.

According to Ng and Perron (2001), the proposal of $MAIC$ is motivated by the observation that the bias in $\hat{\Psi}$ decreases non-linearly as k increases. Alternative model selection rules, such as the AIC and BIC do not take this non-linearity into account. Thus, they underpenalize models with a small k and select autoregressive approximations that are too parsimonious for models with a negative MA component.

4.2 Sarno and Taylor (1998) multivariate augmented Dickey-Fuller test (MADF)[5].

Sarno and Taylor (1998) proposed a Multivariate ADF test, where the sum of the autoregressive coefficients may vary across countries under the alternative

[5] A first application of the test appeared in 1997 as a CEPR Discussion Paper that was finally published as Taylor and Sarno (1999).

hypothesis. To apply the $MADF$ test, they consider an $N-$dimensional stochastic process defined by:

$$y_{it} = a_i + \sum_{j=1}^{k} \Psi_{ij} y_{it-j} + u_{it} \tag{21}$$

for $i = 1, ..., N$, where N denotes the number of series in the panel. The disturbances $\mathbf{u}_t = (u_{1t}...u_{Nt})'$ are assumed to be independently, normally distributed, with zero means. In contrast to the standard ADF test, that involves separately testing each of the N nulls of non-stationarity, Sarno and Taylor (1998) estimate the system (21) by the $SURE$ method, taking into account the contemporaneous correlations among the disturbances. Their joint null is:

$$H_0 : \sum_{j=1}^{k} \Psi_{ij} - 1 = 0, \qquad \forall i = 1, ..., N \tag{22}$$

and is tested by way of a Wald statistic.

The Ψ_i coefficients are allowed to differ across the panel members and the test also permits heterogeneous lags.

Process (21) can also be specified in differences:

$$\Delta y_{it} = \mu_i + \Psi_i y_{it-1} + \sum_{j=1}^{ki} \beta_{ij} \Delta y_{it-j} + u_{it} \qquad t = 1, ..., T; \quad i = 1, ..., N \tag{23}$$

when the $MADF$ test becomes a joint test of the null $\Psi_1 = \Psi_2 = ... = \Psi_N = 0$.

We should emphasize that, in this paper, the use of panel methods may increase the power of the tests, whereas the modified AIC selection rule applied for the lags choice in the MADF tests may improve the size of the tests.

The critical values of the test are specific for each sample and the number of cross-sections and have to be computed by simulation.

4.3 Breuer et al. (2002) multivariate test (SURADF).

Breuer et al. (2002) also propose a panel unit root test that allows for heterogeneous serial correlation across the panel, contemporaneous correlation among the errors, and different autoregressive parameters for each panel member under the alternative. In contrast to the $MADF$ test, separate null and alternative hypotheses are tested for each panel member within a $SURE$ framework.

Similarly to the $MADF$ test, this test has nonstandard distributions and the critical values must be obtained by simulation. The simulation procedures provide critical values for testing the null hypothesis that $\Psi_i = 0$, in an equation such as (23) for each individual member of the panel. The critical values are specific to the estimated covariance matrix for the system considered and the sample size and number of panel members. The procedure allows identification of how many and which members of the panel contain a unit root and which do not.

5 Data construction and empirical results.

The data used in the analysis have been obtained from the International Financial Statistics of the IMF in its CD version. The sample includes quarterly data of money market interest rates and consumer prices for six OECD countries (Canada, France, Germany, Japan, the United Kingdom, and the United States) over the the period 1957Q1 to 2003Q3, that is the longest time period available in this database. 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. We have chosen short-term domestic money market rates for the analysis because these rates reflect market forces better than deposits ones[6]. T-bill rates are used when available for the full period (Canada, UK and US) and call money rates otherwise. The price level data used to construct the real interest series is the consumer price index (CPI).

It is generally accepted that results on RIP depend crucially on the maturities considered. At five to ten-year horizon the empirical evidence becomes far more supportive while the RIP hypothesis is decisively rejected with short horizon data (Fujii and Chinn, 2000). Therefore, our study focus on short term horizon instruments in order to ascertain if our methodology helps to find more supportive and robust evidence on RIP.

Although RIP is an *ex-ante* concept involving expected rather than actual inflation, since expected inflation rates are unobservable, most of the empirical studies use *ex-post* variables. In order to asses the sensitivity of the results to the (ex-ante or ex-post) nature of the variables, we both use

[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 et al., 2002).

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 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]. 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 RIP holds and if inflation forecast errors are random, then the observed real interest differential should be random as well. In this case, we can test RIP by determining whether real interest differentials are systematically related to variables in the current information set.

The results of applying the univariate unit root tests to the RIP are presented in tables 1 and 2 for *ex-post* and *ex-ante* inflation, respectively. In the first column of the tables we show the ADF test[9] results, where the truncation lag has been chosen according to $MAIC$[10], as described above. When inflation is measured ex-post, we find that the RIP is fulfilled for a relatively large group of cases, the rejections mainly associated with the UK (Germany-UK, Japan-UK and US-UK). In addition, the evidence is also weak for France-Germany and Germany-Japan, as the rejection of non-stationarity only occurs at 10%. In table 2, when we use the ex-ante inflation measure, the results for the ADF test are mixed: the RIP is not futfilled in five cases (Canada-UK, Germany-Japan, Germany-UK, Japan-UK and UK-US), also related to the participation of the UK and Japan in the parities. Tight rejections (at 10%) are those of Canada-US, France-US and Germany-US.

Thus, the application of the univariate unit root tests gives mixed results

[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.

[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]Specified with a constant term.

[10]We also computed the ADF tests when the lags were selected using the AIC statistic. In this case, non-stationarity was rejected in the majority of the parities, and the lags selected were much lower. These results confirm the size problems associated with that selection criteria, as argued by Ng and Perron (2001), and are available upon request.

on the fulfillment of the RIP. However, the use of univariate unit root tests may not be appropriate to study a measure of financial integration and specially when the fifteen series analyzed involve just six countries. Thus, the linkages that exist, both by construction and as a result of economic integration among this group of variables, are neglected in this type of analysis. In this particular case, using panel methods may increase the power of the tests, but the commonly used assumption of cross-section independency would not adequately capture the actual cross-relations present in the data. As an illustration, we have included the correlation matrix of the fifteen real interest rate differentials in tables A1 and A2 in the appendix. The actual correlation between many of the real interest differentials is well above 0.5 and this is not only the case of parities involving the same countries, but also between parities that are not linked by construction. See, for example, in table 4, a correlation between Canada-Japan and France-UK of 0.43.

Thus, in this paper we apply a multivariate methodology that accounts for possible cross-section dependence among the variables involved in the analysis. This information is captured through the variance-covariance matrix in a system estimated by the *SUR* method. At the same time, the degree of heterogeneity allowed in the univariate statistics is also present in these tests, as the value of the autoregressive parameter and the number of lags can differ[11]. The *MADF* test proposed by Sarno and Taylor (1998) has been applied to the RIP in the two cases considered: ex-ante and ex-post inflation. The null hypothesis of non-stationarity is easily rejected, for the two inflation measures (see bottom of tables 3 and 4). However, this rejection does not imply that all the real interest differentials are stationarity, so that, in a second stage, we individually test which parities contain a unit root and which do not. The results of the *SURADF* test proposed by Breuer et al. (2002) are also presented in tables 3 and 4, together with the critical values that have been obtained by simulation. Non-stationarity is now rejected at 1% for all the cases and for the two types of inflation, with the only exception of the parity Canada-UK, where rejection occurs at 5%.

Additionally, after the RIP has been accepted for the system considered, the analysis of the significance of the intercept can provide information about the version of RIP that is fulfilled in each case. In the SURADF tests, non-significant intercepts have been excluded, as suggested by Breuer et al. (2002). According to the definitions given above, a non-significant intercept (see equation (16)), implies that the *strong version* of RIP holds. The *weak version* would hold, in contrast, for a non-null intercept (see equation (17)).

[11] Note that the lags used to implement this test are those selected in the univariate analysis using *MAIC*.

In our particular case, the results differ depending on the adopted inflation measure. For ex-post inflation, the strong version holds for all the country-pairs. The only exceptions are linked to Canada (Canada-France, Canada-Germany, Canada-UK and Canada-US). When using the filtered version of inflation, the strong version only holds in six cases: France-Japan, Germany-UK, Germany-US, Japan-UK, Japan-US and UK-US.

6 Conclusions.

In this paper we have applied robust multivariate SURE unit root tests to a group of 6 OECD countries to test the real interest rate parity hypothesis (RIP) over the period 1957-2003. We contribute to previous empirical literature on the RIP in various respects. First, we use univariate tests with good size properties in finite samples, thanks to the lag selection method, so that some of the problems traditionally found in the empirical literature can be avoided. Second, we also apply multivariate tests that do not impose, as the majority of the panel tests, cross-section independence. This possibility is crucial in this type of analysis, when the variables are closely linked in a globalized world. Third, also in the multivariate tests, the individual cross-section lags are selected using the modified AIC rule. Finally, the use of a sequential procedure avoids another drawback associated with panel unit root tests: the null can be rejected if one finds just one stationary series. In our case, after rejecting the null, we can cluster the variables into the stationary and the non-stationary ones.

From this multivariate strategy we can conclude that the evidence is favorable to the stationarity of the rid in all the OECD country-pairs considered. Therefore, RIP hold in all the cases studied either in its strong or weak version. This is in clear contrast to the results of the univariate tests, that do not account for the cross-section interdependence, where RIP could not be accepted in some cases. This outcome is not surprising in a highly integrated area as the OECD, where there is an increasing synchronization of business cycles. In this context, tests that do not impose independency across the panel would have more power to reject the null and, therefore, are more adequate in empirical research on economic integration.

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Table 1
ADF tests
Real interest rate differential computed using ex-post inflation rate
Quarterly data (1958:2-2003:3)

Parity / Test	*ADF*	*lags*
Can-Fran	-4.816^{***}	1
Can-Ger	-3.635^{***}	4
Can-Japan	-3.174^{**}	11
Can-UK	-4.906^{***}	1
Can-US	-3.036^{**}	4
Fran-Ger	-2.741^{*}	9
Fran-Japan	-3.257^{**}	12
Fran-UK	-3.134^{**}	6
Fran-US	-3.908^{***}	2
Ger-Japan	-2.695^{*}	9
Ger-UK	2.559	4
Ger-US	-3.316^{***}	1
Japan-UK	-2.285	13
Japan-US	-5.504^{***}	1
UK-US	-2.469	6

(1) ADF test. The asterisks $(^{*}), (^{**})$ and $(^{***})$ denote rejection of the hypothesis of non-stationarity at 10, 5 and 1% respectively. The critical values are -3.577 (1%), -2.925 (5%) and -2.60 (10%) and have been taken from MacKinnon (1996). The test has been computed using the lags selected by the modified AIC ($MAIC$), as suggested by Ng and Perron (2001). The maximum lag is set at 13, the result of applying $\mathrm{k}_{\max} = \mathrm{int}\left(12(T/100)\right)^{1/4}$.

Table 2
ADF tests
Real interest rate differential computed using ex-ante inflation rate
Quarterly data (1958:2-2003:3)

Parity / Test	*ADF*	*lags*
Can-Fran	-3.556***	1
Can-Ger	-3.635***	4
Can-Japan	-3.226**	9
Can-UK	-1.420	12
Can-US	-2.703*	2
Fran-Ger	-3.137**	7
Fran-Japan	-4.141***	5
Fran-UK	-3.435**	1
Fran-US	-2.801*	2
Ger-Japan	-2.550	9
Ger-UK	-1.513	10
Ger-US	-2.662*	8
Japan-UK	-1.992	9
Japan-US	-3.127**	7
UK-US	-1.309	10

(1) ADF test. The asterisks (*), (**) and (***) denote rejection of the hypothesis of non-stationarity at 10, 5 and 1% respectively. The critical values are -3.577 (1%), -2.925 (5%) and -2.60 (10%) and have been taken from MacKinnon (1996). The test has been computed using the lags selected by the modified AIC ($MAIC$), as suggested by Ng and Perron (2001). The maximum lag is set at 13, the result of applying $\mathrm{k}_{\max} = \mathrm{int}\left(12(T/100)\right)^{1/4}$.

Table 3

Taylor and Sarno (1998) *MADF* test and Breuer et al. (1999) *SURADF* test

Quarterly data (1958:2-2003:3)

Ex-post inflation

Lag truncation selection according to $MAIC$

Parity	*SURADF*	10%	5%	1%
Can-Fran	-10.331***	-3.515	-3.822	-4.237
Can-Ger	-9.484***	-3.502	-3.801	-4.339
Can-Japan	-5.228***	-3.431	-3.741	-4.304
Can-UK	-9.815***	-3.474	-3.819	-4.503
Can-US	-9.002***	-3.429	-3.714	-4.348
Fran-Ger	-7.313***	-3.209	-3.516	-3.965
Fran-Japan	-7.986***	-3.308	-3.576	-4.101
Fran-UK	-9.163***	-3.379	-3.677	-4.222
Fran-US	-11.022***	-3.409	-3.647	-4.099
Ger-Japan	-6.027***	-3.444	-3.832	-4.423
Ger-UK	-10.203***	-3.411	-3.748	-4.302
Ger-US	-10.719***	-3.465	-3.760	-4.304
Japan-UK	-5.908***	-3.232	-3.623	-4.165
Japan-US	-9.125***	-3.289	-3.680	-4.253
UK-US	-9.509***	-3.400	-3.647	-4.163
Panel test	*MADF*			
	201.012***	29.889	33.372	38.631

Note: Critical values computed by simulation, choosing the individual cross-section lags according to the results of the $MAIC$ selection rule in the univariate tests (see table 1)

Table 4
Taylor and Sarno (1998) *MADF* test and Breuer et al. (1999) *SURADF* test
Quarterly data (1958:2-2003:3)
Ex-ante inflation
Lag truncation selection according to MAIC

Parity	*SURADF*	10%	5%	1%
Can-Fran	-5.750***	-3.515	-3.822	-4.237
Can-Ger	-5.840***	-3.502	-3.801	-4.339
Can-Japan	-4.984***	-3.431	-3.741	-4.304
Can-UK	-3.964**	-3.474	-3.819	-4.503
Can-US	-5.822***	-3.429	-3.714	-4.348
Fran-Ger	-5.400***	-3.209	-3.516	-3.965
Fran-Japan	-8.044***	-3.308	-3.576	-4.101
Fran-UK	-7.089***	-3.379	-3.677	-4.222
Fran-US	-6.278***	-3.409	-3.647	-4.099
Ger-Japan	-6.167***	-3.444	-3.832	-4.423
Ger-UK	-6.080***	-3.411	-3.748	-4.302
Ger-US	-5.013***	-3.465	-3.760	-4.304
Japan-UK	-6.312***	-3.232	-3.623	-4.165
Japan-US	-5.754***	-3.289	-3.680	-4.253
UK-US	-5.588***	-3.400	-3.647	-4.163
Panel test	*MADF*			
	78.581***	29.889	33.372	38.631

Note: Critical values computed by simulation, choosing the individual cross-section lags according to the results of the *MAIC* selection rule in the univariate tests (see table 2)

A Correlations

Table A.1.
Correlations
RIR differential computed using ex-post inflation
Quarterly data (1958:2-2003:3)

Differential	Can-Fra	Can-Ger	Can-Jap	Can-UK	Can-US	Fra-Ger	Fra-Jap	Fra-UK	Fra-US	Ger-Jap	Ger-UK	Ger-US	Jap-UK	Jap-UK	UK-US
Can-Fran	1.00	—	—	—	—	—	—	—	—	—	—	—	—	—	—
Can-Ger	0.59	1.00	—	—	—	—	—	—	—	—	—	—	—	—	—
Can-Japan	0.42	0.40	1.00	—	—	—	—	—	—	—	—	—	—	—	—
Can-UK	0.44	0.40	0.27	1.00	—	—	—	—	—	—	—	—	—	—	—
Can-US	0.45	0.49	0.40	0.45	1.00	—	—	—	—	—	—	—	—	—	—
Fran-Ger	-0.42	0.47	-0.006	0.06	0.07	1.00	—	—	—	—	—	—	—	—	—
Fran-Japan	-0.31	-0.02	0.72	-0.05	0.07	0.32	1.00	—	—	—	—	—	—	—	—
Fran-UK	-0.30	0.06	-0.04	0.71	0.12	0.40	0.18	1.00	—	—	—	—	—	—	—
Fran-US	-0.67	-0.20	-0.11	-0.10	0.34	0.50	0.39	0.42	1.00	—	—	—	—	—	—
Ger-Japan	-0.01	-0.35	0.71	-0.10	0.02	-0.37	0.75	-0.09	0.04	1.00	—	—	—	—	—
Ger-UK	0.001	-0.28	-0.04	0.68	0.08	-0.32	-0.04	0.73	0.06	0.18	1.00	—	—	—	—
Ger-US	-0.27	-0.68	-0.11	-0.16	0.28	-0.47	0.09	0.03	0.52	0.41	0.38	1.00	—	—	—
Japan-UK	0.01	0.07	-0.06	0.59	0.03	0.06	-0.64	0.62	0.01	-0.67	0.60	-0.04	1.00	—	—
Japan-US	-0.18	-0.13	-0.82	-0.01	0.18	0.04	-0.72	0.12	0.33	-0.74	0.09	0.29	0.67	1.00	—
UK-US	-0.21	-0.23	-0.04	-0.82	0.13	-0.02	0.11	-0.71	0.33	0.13	-0.71	0.39	-0.64	0.13	1.00

Table A.2.
Correlations
RIR differential computed using ex-ante inflation
Quarterly data (1958:2-2003:3)

Differential	Can-Fra	Can-Ger	Can-Jap	Can-UK	Can-US	Fra-Ger	Fra-Jap	Fra-UK	Fra-US	Ger-Jap	Ger-UK	Ger-US	Jap-UK	Jap-UK	UK-US
Can-Fran	1.00	—	—	—	—	—	—	—	—	—	—	—	—	—	—
Can-Ger	0.51	1.00	—	—	—	—	—	—	—	—	—	—	—	—	—
Can-Japan	-0.15	0.28	1.00	—	—	—	—	—	—	—	—	—	—	—	—
Can-UK	0.20	0.08	0.40	1.00	—	—	—	—	—	—	—	—	—	—	—
Can-US	0.11	0.43	0.31	0.19	1.00	—	—	—	—	—	—	—	—	—	—
Fran-Ger	-0.59	0.38	0.43	-0.14	0.28	1.00	—	—	—	—	—	—	—	—	—
Fran-Japan	-0.71	-0.10	0.79	0.16	0.15	0.66	1.00	—	—	—	—	—	—	—	—
Fran-UK	-0.66	-0.35	0.43	0.60	0.05	0.37	0.71	1.00	—	—	—	—	—	—	—
Fran-US	-0.82	-0.20	0.32	-0.06	0.46	0.69	0.72	0.61	1.00	—	—	—	—	—	—
Ger-Japan	-0.50	-0.43	0.73	0.32	-0.01	0.13	0.82	0.66	0.44	1.00	—	—	—	—	—
Ger-UK	-0.20	-0.64	0.10	0.70	-0.16	-0.38	0.20	0.71	0.09	0.55	1.00	—	—	—	—
Ger-US	-0.45	-0.71	-0.05	0.06	0.31	-0.18	0.23	0.41	0.57	0.45	0.55	1.00	—	—	—
Japan-UK	0.32	-0.21	-0.66	0.40	-0.16	-0.54	-0.66	0.05	-0.37	-0.47	0.46	0.10	1.00	—	—
Japan-US	0.22	-0.04	-0.84	-0.30	0.23	-0.28	-0.73	-0.41	-0.06	-0.75	-0.20	0.23	0.59	1.00	—
UK-US	-0.11	0.19	-0.17	-0.79	0.44	0.30	-0.05	-0.51	0.35	-0.30	-0.74	0.13	-0.47	0.42	1.00