Does the Real Interest Parity Hypothesis Hold? Evidence for Developed and Emerging Markets

Alex Luiz Ferreira
Miguel A. León-Ledesma

Department of Economics, University of Kent

August 2003

Abstract: Evidence is presented on the Real Interest Parity Hypothesis for a set of emerging and developed countries. This is done by carrying out a set of unit-root tests on the real interest differentials with respect to Germany and the US. Our results support the hypothesis of a rapid reversion towards a zero differential for developed countries and towards a positive one for emerging markets. An important result is that this adjustment tends to be highly asymmetric and markedly different for developed and emerging countries. Our evidence reveals a high degree of market integration for developed countries and highlights the importance of risk premia for emerging markets.

Keywords: Real Interest Rate Differentials, Market Integration, Unit Roots, Asymmetric adjustment.

JEL Classification Numbers: F32, F21, C22.

Acknowledgements: We would like to thank, without implicating, Yunus Aksoy, Alan Carruth and Ole Rummel for helpful comments on earlier drafts of this paper. The paper also benefited from comments of participants at the EcoMod2003 Conference in Istanbul, July 3-5, 2003. Ferreira acknowledges financial support from the Conselho Nacional de Desenvolvemento Cientifico e Tecnológico (CNPq) of Brazil.

Address for correspondence: Miguel A. León-Ledesma, Department of Economics, Keynes College, University of Kent, Canterbury, Kent, CT2 7NP, UK. Phone: +00 + 44 (0)1227 823026. Fax: +00 +44 (0)1227 827850.
Email: M.A.Leon-Ledesma@kent.ac.uk

Does the Real Interest Parity Hypothesis Hold? Evidence for Developed and Emerging Markets

1. Introduction

The Real Interest Rate Parity Hypothesis (RIPH) states that if agents make their forecasts using rational expectations, and arbitrage forces are free to act in the goods and assets markets, then real interest rates between countries will equalise. Several studies have tested this hypothesis since the pioneer papers of Mishkin (1984) and Cumby and Obstfeld (1984). However, the empirical literature does not offer a conclusive answer regarding the existence of real interest rate differentials (*rids*). For instance, the evidence found by Gagnon and Unferth (1995), Ong et al. (1999), Evans et al. (1994), Chinn and Frankel (1995), Alexakis et al. (1997), Cavaglia (1992), Phylaktis (1999), Awad and Goodwin (1998), Frankel and Okongwu (1995), Fujii and Chinn (2000) and Jorion (1996) is mixed. These authors tend to conclude that *rids* are relatively short-lived and mean-reverting but different from zero in the long-run.

The importance of this hypothesis stems from the fact that empirical evidence can be interpreted as a measure of international integration in goods and assets markets. This is particularly emphasised in Chinn and Frankel (1995), Phylaktis (1999), Alexakis et al. (1997), Awad and Goodwin (1998), Obstfeld and Taylor (2002) and Mancuso et al (2002). This is because the RIPH is based on the existence of frictionless markets. It follows that a test of the real interest rate parity is a test of the degree of market integration.

The paper presents further evidence on the RIPH for a sample of small open-economies in relation to the USA and Germany. We aim to unveil whether some of our set of small open economies have been experiencing *ex post* real interest rate differentials (*rid*(s) hereafter) in relation to larger ones. We do so by carrying out a set of unit root tests that will characterise the dynamic behaviour of *rids*. There are few papers in the economic literature investigating the real interest rate parity hypothesis through unit root tests on *rids*. The main examples are Meese and Rogoff (1988), Edison and Pauls (1993) and Obstfeld and Taylor (2002). Our study complements these authors in three main directions. First, we make use of more powerful unit root tests and take structural changes into account. Second, in line with recent theoretical and empirical models of capital flows, we simultaneously test for the existence of asymmetries and unit roots in the behaviour of *rids*.[1] Third, we focus both on developed and emerging market economies during a period with a high degree of financial and goods market liberalisation, which is in accordance with the assumptions underlying the real interest rate parity hypothesis. This will allow us to compare the behaviour of *rids* in developed and emerging markets. Our findings show that *rids* are in general quickly mean reverting, with a positive mean for emerging markets and zero or close to zero for developed ones. We also show that *rids* show strong features of asymmetry, but the behaviour for emerging and developed markets is substantially different.

The paper is organised as follows. In the Section 2 we give some theoretical background and describe the methodology involved in the tests; in Section 3 we describe the data; Section 4 presents the results of unit root tests; Section 5 presents the results of the asymmetry tests and Section 6 concludes.

2. Theoretical background and methodology

As far as agents make their forecasts using rational expectations [represented in equation (3) below], arbitrage forces in the goods and assets markets ensure that the

---

[1] See, for instance, Kraay (2003), Pakko (2000) and the review of Stiglitz (1999). Asymmetries could also arise in the adjustment of prices due to goods market frictions arising from transaction costs as in Obstfeld and Rogoff (2000).

real interest rates parity hypothesis hold. Arbitrage forces are formalised by the uncovered interest rate parity (UIRP) and the relative purchasing power parity (PPP) conditions stated in equations (1) and (2), respectively:

$$i_t - i_t^* = ds_t^e \tag{1}$$

$$ds_t = \pi_t - \pi_t^* \tag{2}$$

$$ds_t^e = ds_t + \varepsilon_t \tag{3}$$

Where i is the domestic interest rate and i^* is the exogenously determined foreign interest rate that matures at time t. The exchange rate is the domestic price of the foreign currency and is represented by S_t; the expected rate of depreciation of the exchange rate is $ds_t^e = \frac{S_t^e}{S_{t-1}} - 1$, with the superscript e denoting expected values. The rates of domestic and foreign inflation are π_t and π_t^* respectively; and d is the first difference of the logarithm. ε_t is a disturbance term that exhibits the classical properties, i.e. ε_t is iid $\mathrm{N}(0, \sigma_\varepsilon^2)$ and σ_ε^2 represents its variance. Lower case variables, except interest rates, here and elsewhere represent natural logarithms.

If PPP holds, we can substitute equation (2) into (3) and the result into (1), which yields:

$$i_t - i_t^* = \pi_t - \pi_t^* + \varepsilon_t\,. \tag{4}$$

Equation (4) can also be rewritten as

$$(i_t - \pi_t) - (i_t^* - \pi^*) = rid_t = \varepsilon_t \tag{5}$$

Since ε_t are iid $\mathrm{N}(0, \sigma_\varepsilon^2)$, the expected value of the *rid* is zero.

Now consider that rid_t follows a more general stochastic process:

$$rid_t = a_0 + a_1 rid_{t-1} + \varepsilon_t \tag{6}$$

Assuming that rid_0 is a deterministic initial condition, the solution to the difference equation above is:

$$rid_t = \frac{a_0(1-a_1^t)}{(1-a_1)} + a_1^t rid_0 + \sum_{i=0}^{t-1} a_1^i \varepsilon_{t-i} \,. \tag{7}$$

If $|a_1| < 1$ and allowing t to increase to infinity, the limit towards which the *rid* converges in the long run is equal to:

$$\lim_{t\to\infty} rid_t = \frac{a_0}{1-a_1} \tag{8}$$

Taking expectations of equation (8) and considering the RIPH, we have

$$E(rid_t) = \frac{a_0}{1-a_1} = 0 \iff a_0 = 0 \tag{9}$$

It follows from equation (5) that if UIRP, PPP and rational expectations hold, the *rid* is equal to the unforeseeable disturbance term related to the forecast of exchange rate depreciation[2]. From equation (9) we observe that a *rid* does not exist in the long run because its unconditional mean, or expected value, is equal to zero. The problem is to verify whether shocks to the series of *rids* dissipate and the series returns to its long-run zero mean level. This objective can be accomplished by performing unit root tests on the series of *rids*.

We can represent the model of equation (6) as a *p*th-order autoregressive process,

$$\Delta rid_t = a_0 + \psi rid_{t-1} + \sum_{i=2}^{p} \beta_i \Delta rid_{t-i+1} + \varepsilon_t \,, \tag{10}$$

where,

$$\psi = \sum_{i=1}^{q} a_i - 1 \,. \tag{11}$$

[2] In fact the *ex post rid* can also be derived from the Fisher (1930) equation. In that case the *ex post* rid equals the *ex ante rid* plus a disturbance term related to the inflation forecast, given that rational expectations is assumed [Mishkin (1984)].

The following possibilities arise from the estimation of this ADF-type equation (10):

$$\psi > 0 \tag{12}$$

$$\psi = 0 \tag{13}$$

$$\psi < 0 \text{ and } a_0 = 0 \tag{14}$$

$$\psi < 0 \text{ and } a_0 \neq 0 \tag{15}$$

Inequality (12) represents 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 (13) the series contains a unit root and *rids* follow a random walk with shocks affecting the variable on a permanent basis. In cases (14) and (15) the estimated parameter (ψ) is such that $\sum_{i=1}^{q} a_i < 1$. Deviations from the mean are temporary and the estimated root provides information on whether the *rid* is short-lived or persistent. In (14) the *rid* follows a stationary process and converges to a zero mean. The RIPH holds and the speed of adjustment of the *rid* to its equilibrium level is a measure of the degree of persistence. In (15) *rids* converge to a mean that is different from zero. In summary, short-lived *rids* are consistent with the RIPH because the series rapidly reverts to zero. Persistent *rids* that converge to a constant mean that is equal to zero are also consistent with the RIPH, since shocks eventually dissipate. The existence of a mean different from zero may arise theoretically from a country specific risk premium. However, random walks, permanent or explosive *rids* are inconsistent with the real interest rate parity hypothesis.

Three usual problems with standard unit root tests, such as the ADF, arise. First, it is well known that the power of these tests tends to be very low, leading to over-acceptance of the null of a unit root. The low power problem is magnified for small samples because a stationary series could be drifting away from its long-run equilibrium level in the short-run. Another serious problem of unit root tests is not considering the existence of structural breaks in the series. When there are structural changes, the standard tests are biased towards the non-rejection of a unit root [Perron (1989)]. Finally, since the work of Neftci (1984), it has been increasingly recognised that macroeconomic time series show strong asymmetry over the business cycle. If asymmetry is present in *rids*, linear unit-root tests will suffer from a loss of power.[3]

Several tests have been put forward to alleviate these problems. Kwiatkowski et al (1992) use the LM statistic to test the null hypothesis of stationarity (KPSS test). The time-series in their model is written as the sum of a deterministic trend, a random walk and a stationary error. The null corresponds to the hypothesis that the variance of the random walk equals zero, in other words, the variance of the error is constant. When the series has an unknown mean or linear trend, the tests suggested by Elliot et al (1996) (ERS test hereafter) and Elliot (1999) are recommended. These tests use information contained in the variance of the series to construct a test statistic (DF-GLS and ADF-GLS) that has more asymptotic power than the standard ones. The initial condition is assumed to be zero in the ERS test while it is drawn from its unconditional distribution in Elliot (1999). Regarding the existence of structural breaks, Perron (1997) developed a procedure to test for unit-roots that endogenously searches for structural breaks in the series using two methods. In the first method, the break date is chosen to be the one in which the t-statistic for testing the null hypothesis of a unit root is smallest among all possible break points. In the second method, the break point corresponds to a maximum of the absolute value of the t-statistic on the parameter associated with the change in the intercept. We will make use of both when testing the RIPH assuming symmetric behaviour of the *rid* to

[3] See Enders and Granger (1998). They also show evidence of asymmetry in the adjustment of the term structure of interest rates.

positive and negative shocks. Later on we relax this assumption and apply the tests proposed by Caner and Hansen (2001) that allows for testing unit roots and asymmetry using threshold autoregressive methods. This is because some theoretical and empirical models of credit markets with imperfect information point out to possible asymmetric behaviour of *rids* as changes in interest rates may influence subjective risk perceptions by creditors. If the series are asymmetric, the power of unit-root tests will improve. The pattern of asymmetry showed by *rids* is also a relevant issue in itself, especially when we compare different countries.

While there is a substantial number of papers testing unit roots in nominal interest rates, inflation and even real interest rates, few studies are concerned with *rids* especially when the objective is to test the real interest rate parity hypothesis. Meese and Rogoff (1988) performed unit root tests in the series of *rids* of the US, UK, Japan and Germany over the period 1974M2 to 1986M3. They could not reject the hypothesis that there is a unit root in the series of long term real interest rate differentials, but not in short-term differentials. In fact, they found that both nominal and real short-term interest rate differentials appear to be stationary in levels. Along the same line of Meese and Rogoff (1988), Edison and Pauls (1993) performed ADF-tests on *rids* using quarterly observations from 1974 to 1990 for the G-10 countries. They could not reject the unit root hypothesis.

Obstfeld and Taylor (2002) questioned why Meese and Rogoff (1988) Edison and Pauls (1993) and McDonald and Nagayasu (2000) could not reject the unit root hypothesis given the increasing globalisation of capital markets "…if capital is perfectly mobile, this dooms to failure any attempts to manipulate local asset prices to make them deviate from global prices, including the most critical macroeconomic asset price, the interest rate" (pp.17). In their view, the failure to reject the null stems from the fact that these authors focused attention on the recent float, had shorter samples, and used tests of low power such as the ADF test. The sample used by Obstfeld and Taylor (2002) includes *rids* of three countries relative to the USA (UK, France and Germany) from 1870 to 2000. Their results, using standard ADF and Elliott (1999) tests, show that the hypothesis of a unit root can be rejected at the 1% level in all periods except for the recent float: "The most striking impression conveyed by the figure is that differentials have varied widely over time, but have stayed relatively close to a zero mean. That is the series appear to have been stationary over the very long run, and even in shorter sub periods." (pp.26). By splitting the sample of the recent float in two sub periods (1974-1986 and 1986-2000) they found that the evidence against a unit root is stronger over the second sub period.

3. Data

The countries chosen for our tests can be split into three groups. The first one comprises some small open-economies of emerging markets: Argentina, Brazil, Chile, Mexico and Turkey. The second group is composed of small open-economies of developed countries: France, Italy, Spain and the UK. Finally, the third group is the one with the countries used as the reference large economies for the calculation of *rids*: Germany and USA. The period of the tests corresponds to the interval that spans from 1995M3 to 2002M5, with the exceptions of Argentina, for which we have calculated *rids* until 2002M3, and Chile and the UK, with *rids* calculated until 2002M4. The shorter period of the former countries is due to data availability. This heterogeneous sample of countries allows inter-group comparisons and the detection of similar patterns between them.

Our sample period starts in the mid 90s because harmonised data for the construction of the *rid*s for some of our countries did not exist before this period.[4] An advantage of using this period is that after the mid-90s most of the countries had liberalised capital markets and had advanced substantially in their trade liberalisation process. As shown previously, the RIPH is based on the assumptions of frictionless goods ands assets' markets. If there are restrictions to trade in these markets, arbitrage would be constrained and different outcomes from those predicted by the RIPH may arise. The process of trade and financial liberalisation happened during different periods for the countries in our sample. Trade liberalisation in developing economies was carried out in the late 1980's and early 1990's[5]. Financial liberalisation happened almost simultaneously[6]. Hence, we focus on the second half of the 1990's using data available until the most recent period.

The *ex post* real interest rate is defined as

$$r_t = i_t - \pi_t \ , \tag{16}$$

where i_t is the nominal interest rate earned on a one-period bond or deposit that matures at time t, i.e., it is the nominal return from holding the one-period bond from t-1 to t; π is the actual (or *ex post*) inflation from t-1 to t. Data on interest rates was obtained from IMF's International Financial Statistics (IFS). Among the several categories of interest rates available in the IFS database, we considered the Treasury Bill Rate as being the most appropriate for the tests. In practice, there is no unique variable that international arbitrageurs use to compare their prospective returns at home and abroad. However, the Treasury Bill Rate is available in domestic markets to

[4] We decided to test the RIPH for the same period for all countries to allow for comparison of the results.
[5] See UNCTAD (1999).
[6] Edwards (2001), among others, acknowledged the difficulty in measuring the "true" degree of capital mobility and thus the starting period of the financial liberalisation. In spite of this difficulty, however, it is recognised that there has been a marked increase in the flows of capital across countries especially during the nineties.

international arbitrageurs and has a fixed maturity. For these reasons, we have chosen to use the Treasury Bill Rates for Brazil, Mexico, Italy, Spain, UK, USA and Germany. We use deposit rates for Argentina, Chile and Turkey because the availability of data on Treasury Bill Rate was limited for these countries. This is the only other short-term interest rate available with a specified maturity. As regards the choice of maturity, as stated by the liquidity premium theory, investors tend to prefer bonds with short-term maturities rather than bonds with longer-term maturities, since the former bear less interest-rate risk. Forecast errors of exchange rate changes are also more likely to increase as time increases. Because we are interested in verifying the degree to which real interest rates are different across countries, we have decided to use short term rates instead of long-term ones in order to avoid a greater influence of risk premium and forecast errors in the composition of *rids*.

Hence, in order to calculate *rids* we transformed the annualised monthly interest rate into a compounded quarterly rate; the real interest rate was then calculated by subtracting the quarter-on-quarter inflation rate from the compounded nominal interest rate of three months. The inflation rate is the rate of growth of the Consumer Price Index (CPI).[7] Our choice of interest rate and inflation is in accordance with the data used by the majority of the authors testing the UIRP. For instance, Mishkin (1984), Knot and de Haan (1998), Nakagawa (2002), Phylaktis (1999), Alexakis et. al. (1997) used interest rates that included either the 3-month Treasury Bill or the 3-month deposit rate. The great majority of authors also used the CPI as the appropriate deflator.

Figure 1 plots the different *rid*s with respect to Germany and the US. With the exception of Chile, *rids* were high in all developing countries at the beginning of the sample period and behaved differently afterwards. The *rids* of Argentina, for example, were stable until mid-1998 when they experienced a substantial increase that accelerated with the 2001 crisis. The *rids* of Brazil initially diminished but started to increase again until 1999 when they fell and stayed relatively constant. A possible explanation for this apparent structural break in Brazil is that the change in the exchange rate arrangement in 1999M1 released monetary policy from the objective of attracting capital flows to sustain the hard peg[8]. The *rids* of Turkey were volatile around a positive mean during the whole sample period. The *rids* of Mexico showed a "negative trend" until 1999 and a positive mean afterwards. It is difficult to see any pattern in the *rid* of Chile, so we prefer to describe it as being volatile with a positive mean. The shortage of international liquidity triggered by the Mexican crisis of 1994-1995 may explain the high common level of *rids* in the initial period of the sample.

The graphs of the *rids* of developed countries tell different stories. The pattern of the *rid* is very similar for Spain and Italy. These countries experienced positive *rids* in a first period that finished by mid-1997 and negative *rids* during the second period. For France and the UK it is difficult to see a clear pattern. They are relatively volatile and seem to fluctuate around a zero mean. Much of the evolution of *rids* for these countries can be explained in terms of the closing gaps in nominal rates due to the convergence criteria imposed for the launch of a common currency. The speed of convergence increased considerably after the establishment of the irrevocable parities in 1999M1 (with the exception of the UK). In fact, convergence to a lower level of nominal interest rates, given a higher inflation rate, may explain the negative mean of the *rids* of Spain and Italy in the period that started after mid-1997.

---

[7] The results using the Producer Price Index (PPI) were remarkably similar and are not reported here to save space.

[8] See Frankel et al (2002) for an analysis of the empirical regularities concerning the sensitivity of domestic interest rates to international ones under different currency regimes. They also verify in the paper whether floating exchange rate regimes allow independent monetary policy.

4. Unit root tests

The results of ADF tests are reported in Table 1. We found the optimal augmentation lags by using a sequential general-to-specific criteria. The results show that we can reject the hypothesis of a unit root only for Brazil, Mexico, Turkey and UK(Ger).[9] It must be stressed, however, that our test statistics were very sensitive to the number of lags, which means that inaccuracy in the lag selection may have led to biased conclusions. Increasing the number of lags of Brazil from 3 to 5, or in the case of Mexico from 1 to 5, for example, imply the non rejection of the null hypothesis of a unit root. Nonetheless, as briefly discussed in section II, failure to reject the unit root is likely to be explained by the low power of ADF tests. Hence, we performed the already mentioned more powerful tests. As we can see in Table 1, the results using ERS (1996) were slightly different. Using the same number of lags chosen for the ADF tests, we could reject the unit root hypothesis not only for Brazil, Mexico and Turkey but also for the *rid* of Chile-US. The findings of the tests using the method proposed by Elliot (1999) were very similar to that of ERS (1996), the only difference is that we could also reject the null of a unit root for the *rid* of UK(Ger) which we were already able to reject using ADF tests. The KPSS test allowed us to accept the hypothesis of stationarity for most countries of the sample. Apart from the countries mentioned before, we could not reject the hypothesis of level-stationary for the *rids* of Argentina, France(US) and Chile(Ger) but not for Brazil. We could also not reject the null of stationary for the *rid* of the UK and US(Ger).

Although these methods provide more powerful alternatives to the ADF test, they do not take into account structural breaks. The plots of *rid*s in Figure 1 reveal that many of these series may contain a break in their mean. This is especially so for

---

[9] When we refer to the *rid* of a country 1 with respect to country 2 we will use the notation country1(country2). So the *rid* of, for instance, Turkey with respect to Germany would be Turkey(Ger). When we mention only "Turkey" we are referring to both *rids*.

Argentina, Mexico, Brazil, Italy and Spain.[10] For this reason we applied Perron's (1997) tests assuming that the series contain an innovational outlier with a change in the intercept. This model can be represented as

$$rid_t = a_0 + \theta DU_t + \lambda D(T_b)_t + a_1 rid_{t-1} + \sum_{i=1}^{p} \beta_i \Delta rid_{t-i} + \varepsilon_t\,, \qquad (17)$$

where T_b denotes the break date; $DU_t = 1(t > T_b)$ and $D(T_b)_t = 1(t = T_b + 1)$. The test is performed using the t-statistic for the null hypothesis that a_1=1. The results of this test are reported in Table 2.

We were able to reject the unit root for Brazil, Chile, Mexico, Turkey, Italy(Ger), Spain(Ger) and UK(Ger) using the date break suggested by the first method. The unit root hypothesis was also rejected for Mexico(US), Turkey, Italy and Spain(Ger) using the date break of the second method. Nevertheless, we could not find evidence of non-stationarity for the *rids* of Spain(US) in any of the tests. We suspect that systematic forecast errors due to excessive credibility in the convergence to low inflation levels could explain the unit root behaviour of the *rid* of Spain(US).

The date breaks retrieved by the tests suggest that the Asian crisis (starting in mid 1997) impacted on the *rids* of Chile and, less likely, the rids of France(US), Italy and Spain. Another explanation for the break dates of the latter countries is that *rids* were affected by the Stability and Growth Pact signed by the European Council in 1997M6 and the prospect of the establishment of the European Central Bank (ECB), that officially took place in 1998M6. The Mexican crisis (1994M12) appears to have impacted on the *rids* of Brazil, as can be seen in the date break found by the first method. The Russian crisis (mid 1998) may have had an effect on the *rids* of Mexico and Chile(US). The Brazilian crisis (1999M1) is captured by the date break of the *rids* of that country retrieved by the second method. The Brazilian crisis probably

[10] This is further confirmed by plots of the recursive Chow tests of the AR parameter, not reported here but available on request.

impacted the *rids* of Argentina as can be seen in the date break suggested by the second method. The free float of the Peso in Argentina 2002M1 is reflected in the date break of the first method. The results also indicate that the Turkish crisis, which culminated in 2001M2 with the free floating of the Lira, may have its origins at the beginning of 1999.

According to our results, the irrevocable parities announced in 1999M1 for the Euro area and the introduction of the Euro as a medium of exchange in 2002M1, have not been reflected in the form of a structural break during the sample period. Our results also suggest that the Asian financial crisis and/or the establishment of the European System of Central Banks may have affected the *rids* of developed countries in a structural manner.

The tests carried out also allowed us to calculate half-lives of deviations from equilibrium and the equilibrium itself as given by equation (8). It must be stressed that *rids* converge to an equilibrium level only if there is not a unit root in the series. As previously stated, the low power of the traditional tests implies that a unit root may not exist even if we are not able to reject the null. Hence, we decided to calculate the half-life and equilibrium level of *rids* for all countries including Spain(US). The results are reported in Table 3.

According to the estimated roots obtained with standard ADF tests, some countries of our sample have highly persistent *rids*. The half-life of the *rid* of Argentina, for example, varies from 3.4 to 13.1 months. In the case of Italy and Spain, the half-life varies from 5 months to 10.3 months. The most persistent *rids*, according to our results, are those of Argentina and Italy. On the other hand, the tests using the Perron (1997) methods suggest a smaller degree of persistence for the *rids* of all countries. Half-lives vary between 0 and 3.2 months, with the exception of

Argentina(US). Thus, when possible structural changes are taken into account, *rids* of almost the whole sample are short-lived.

Estimated equilibrium levels for the *rid*s from the ADF and Perron (1997) equations are reported in Table 4.[11] Equilibrium levels of *rids* are significantly different from zero if both the intercept and estimated root are significant. Inspection of Table 4 shows that the *rids* of Brazil, Chile, Mexico, Turkey, UK(Ger) and US(Ger) converge to a mean value that is statistically different from zero. These values were higher for Turkey, Brazil, Argentina, Mexico and Chile in descending order. These results point out to risk premium as a likely explanation of permanent higher levels of real interest rates. When we allowed for structural changes using the date breaks retrieved by the first method of Perron (1997), we found that the *rids* of Argentina, Brazil, Chile, Mexico, Turkey, Italy(US) and Spain converge to equilibrium values that are statistically significant.

5. Asymmetry and unit roots

The previous unit-root tests assume that *rids* follow a linear representation or linear path around a breaking trend. However, recent developments in the theory of imperfect capital markets/imperfect information suggest that the behaviour of *rids* may be asymmetric because risk perceptions may vary with changes in interest rates themselves. The argument is summarised in Stiglitz (1999) and a similar argument put forward in Pakko (2000).[12] Given the existence of asymmetric information in international credit markets, lenders will look at increases in interest rates as a signal that determines their subjective probability of bankruptcy (or default). As Stiglitz (1999) explains "[…] the probability of bankruptcy may depend on the interest rate

[11] We just report equilibrium levels obtained using Perron (1997) for break search method 1, as method 2 gave similar results.

charged, so that beyond a point, increases in the interest rate charged actually lead to lower expected returns" (p. 64) which relates to the idea that "The dominant effect of large, unanticipated increases in interest rates is thus induced bankruptcies and an increase in non-performing loans" (p. 65). The consequence of these arguments for the RIPH is that the country risk premium may depend on changes of the interest rate and, hence, *rids* would converge to different equilibrium differentials if previous changes in *rids* surpass a certain threshold. This multiple equilibria idea would induce asymmetries in the time series behavior of *rids*.[13]

Evidence on this *rid* nonlinearity is presented, for instance, in Mancuso et al (2002). If asymmetries are present in the adjustment of *rids*, unit-root tests may lose power and suffer size distortions unless they are incorporated in the tests (Enders and Granger, 1998). Our approach allows us to simultaneously test for asymmetry and unit roots in the *rids* series, revealing interesting features about the RIPH. If *rids* behave asymmetrically, we can use the following TAR (Threshold Autoregression) representation (Caner and Hansen, 2001):

$$\Delta y_t = \theta_1' y_{t-1} 1_{\{z_{t-1} < \lambda\}} + \theta_2' y_{t-1} 1_{\{z_{t-1} \geq \lambda\}} + \sum_{j=1}^{p} \gamma_j \Delta rid_{t-j} + \zeta_t \ , \qquad (18)$$

where $y_{t-1} = (1 \quad rid_{t-1})$, $1_{\{.\}}$ is the indicator function that takes the value of 1 if z_{t-1} is higher or lower than a threshold λ, and 0 otherwise. The variable z_t is any stationary variable that would determine the change of regime. For our purposes, we set $z_t = rid_t$

[12] Some of these features are commonly introduced in models of speculative attacks with asymmetric information. The idea in the context of the 1997 South East Asian crisis is discussed in Radelet and Sachs (1998).

[13] Recent papers by Nakagawa (2002) and Obstfeld and Rogoff (2000) amongst others present evidence that suggests that convergence towards PPP may be non-linear. This is usually associated with theoretical models in which market segmentation arising from various transaction costs introduce non-linearities in the adjustment of real exchange rates (RER) as in Obstfeld and Rogoff (2000). These kinds of non-linearities may also induce asymmetry in the speed of adjustment of *rids* to positive and negative shocks.

– rid_{t-m}. That is, we assume that *rids* may have a different behaviour depending on whether past *changes* in r*ids* have been higher or lower than a certain threshold λ. This is a momentum-TAR model or M-TAR as in Enders and Granger (1998). The lag length *m* for the changes in *rids* will be data determined as will be the search for the optimal threshold λ. Finally, the parameter vectors θ_1 and θ_2 can be partitioned as

$$\theta_1 = \begin{pmatrix} \mu_1 \\ \rho_1 \end{pmatrix}, \qquad \theta_2 = \begin{pmatrix} \mu_2 \\ \rho_2 \end{pmatrix}$$

The choice of the threshold λ could be simply made on an a priori basis, such as setting $\lambda = 0$ or equal to the sample mean of Δrid_t. However, this would be a biased estimate of the threshold if asymmetric adjustment exists and a subjective measure. In order to search for the optimal threshold, we follow Chan (1993) and find λ as the value of Δrid_t that minimises the residual sum of squares of the OLS estimation of (18).[14]

In order to test for the existence of asymmetry in the adjustment under both regimes we test the null hypothesis $H_o : \theta_1 = \theta_2$ on the OLS estimation of (18), making use of the Wald statistic (W) proposed in Caner and Hansen (2001). The RIPH would imply rejecting H_o: $\rho_1 = \rho_2 = 0$, and we also make use of two Wald statistics (R1 and R2). Finally, we also chose *m* to minimise the residual sum of squares. Given that the Wald test of asymmetry is a monotonic function of the residual variance, we choose *m* as the value which maximizes the Wald test of asymmetry.

The procedure we follow to test simultaneously for asymmetry and unit roots implies first estimating a baseline model for the linear ADF regression to determine the lag augmentation of the DF regression using general-to-specific techniques as in previous sections. We then select the threshold by minimising the residual sum of squares of (6) as mentioned earlier and fit the M-TAR model by OLS for every value of m. We choose the m that minimises the residual sum of squares for all values of m.[15]

Given that the asymptotic null distribution of the asymmetry test (W) is non-standard, Caner and Hansen (2001) recommend the use of bootstrap methods to obtain p-values. In a Monte Carlo experiment they show that the power and size of the test does not crucially depend on whether we impose a unit-root. Hence, we obtained p-values by carrying out 1,000 iterations of the unconstrained asymmetry test, i.e. not imposing the existence of a unit root. Finally, the unit root hypothesis involves testing for H_o: $\rho_1 = \rho_2 = 0$. There are two possible alternatives: H_1: $\rho_1 < 0$ *and* $\rho_2 < 0$ and

$$H_2 : \begin{cases} \rho_1 < 0 \quad and \quad \rho_2 = 0 \\ or \\ \rho_1 = 0 \quad and \quad \rho_2 < 0 \end{cases}$$

The first alternative corresponds to the stationary case, whilst the second implies stationarity in only one of the regimes, which implies overall non-stationarity but a different behaviour from the classic unit-root. Caner and Hansen (2001) develop asymptotic theory for the distribution of this unit-root test. However, for finite samples they recommend the use of bootstrapping. As the distribution of the test statistic will depend on whether or not a threshold effect exists, p-values obtained through the bootstrap are not unique. We hence obtained the bootstrapped p-values from 1,000 iterations under the hypothesis that the threshold is not identified (R1) and under the hypothesis that it is identified (R2). These two tests have substantially more power than the ADF test as threshold effects become more important. In order to discriminate between the two alternatives in H_2, Caner and Hansen (2001) recommend looking at the t-ratios of ρ_1 and ρ_2.

---

[14] In practice we eliminated the highest and lowest 10% values of Δrid_t.

[15] Usually, for monthly data we take $m = 1,\ldots, 12$ and for quarterly $m = 1,\ldots, 4$.

The results are provided in Table 5 where we report the estimated threshold (λ), the lag of the change in *rids* for the determination of the threshold (m), and the estimates of the parameters of (18) in both regimes. Asymmetry appears to be a prevalent feature in the data. We can reject the null of no asymmetry for at least one pair of *rids* for all countries except Mexico and the UK. Observing the values and t-ratios of the intercepts and autoregressive terms we can see that this asymmetry is associated with both differences in intercepts in both regimes and asymmetric adjustment speeds, although the former appears more frequently. As for unit roots, the results confirm that taking asymmetry into account is important, as we can reject the null of non-stationarity for all the countries by at least one of the R tests except for Argentina, Italy and Spain.[16] In several cases, such as US(Ger) and UK, *rid*s appear to be stationary when decreasing and non-stationary when increasing above the threshold. The other way around occurs for Brazil, Chile and Turkey.

Other important features appear when observing the different behaviour of emerging and developed markets in our sample. For France, UK and US(Ger) the intercept is either statistically insignificantly different from zero or close to it in both regimes. The speed of adjustment for these countries, as already mentioned, is higher when decreasing and lower when increasing. This pattern in the speed of adjustment is reversed for Brazil, Chile, Mexico and Turkey. Furthermore, for these countries the intercept tends to be close to zero when the *rids* are growing below the threshold and significantly higher than zero when it is rising. This positive intercept would imply large equilibrium *rids* especially for Turkey and Brazil. In relation to the theoretical models of imperfect information in credit markets, these results may seem to indicate that large increases in interest differentials may be negatively interpreted by the market, which in turn imposes a higher risk premium. The fact that this pattern does not seem to arise for developed markets also supports this idea as, during this period,

[16] For Italy and Spain structural breaks may be driving most of the results as seen previously.

none of these countries has suffered large swings of their interest rates that may have induced this change in risk perception effect.

6. Conclusions

We have presented evidence on the Real Interest Parity Hypothesis for a set of developed and emerging markets for the period that spans from the mid-90s until the middle of 2002. Our results show that, despite the short time span, we were able to find mean reversion in *rids*. The speed of mean reversion is high, indicating that real differentials tend to be short lived. This is especially so if we allow for the likely possibility of structural breaks in the series. We were able to reject the unit root hypothesis or to accept the null of stationarity for all countries, excluding Spain relative to the US. This evidence supports the hypothesis of a high degree of market integration, which is consistent with financial liberalisation and the emergence of global capital markets. The pattern of adjustment is asymmetric, that is, whenever *rids* grow above or below a certain threshold they tend to behave differently. For emerging markets adjustment is quicker when *rids* grow fast, while for countries such as France and the UK adjustment is quicker when *rids* grow below the threshold.

Nonetheless, we found evidence supporting the existence of a positive long-run mean in the *rids* of, especially, emerging markets. The long run mean of emerging markets economies tends to be higher than for developed ones, for which it is zero or close to zero. Our results also suggest that foreign financial crisis may have generated structural changes in *rids*. Finally, we found evidence that equilibrium *rids* for emerging markets are high in periods of rapid growth of the *rid*. All these features point out to the existence of large risk premia for emerging markets, but not for developed markets.

In general, our results support recent evidence on the RIPH for developed countries despite the short sample of our study. It also complements this literature with evidence from emerging markets. For these countries, the RIPH with a risk premium component seems to be a more realistic specification. We also find that asymmetries induced by either risk perception changes or transaction costs seem to be an important feature when explaining real interest rates differentials.

References

Alexakis, P., Apergis, N. and Xanthakis, E. (1997). "Integration of International Capital Markets: Further Evidence from EMS and Non-EMS Membership," *Journal of International Financial Markets, Institutions and Money*, (7)3, 277-287.

Awad, M. A. and Goodwin, B.K. (1998). "Dynamic Linkages Among Real Interest Rates In International Capital Markets," *Journal of International Money and Finance*, (17)6, 881-907.

Caner, M. and Hansen, B.E. (2001). "Threshold Autoregressions With a Unit Root," *Econometrica*, 69, 1555-1596.

Cavaglia, S. (1992). "The Persistence of Real Interest Differentials," *Journal of Monetary Economics*, (29)3, 429-443.

Chan, K.S. (1993). "Consistency and Limiting Distribution of the Least Squares Estimator of a Threshold Autoregressive Model," *The Annals of Statistics*, 21, 520-533.

Chinn, M.D. and Frankel, J. A. (1995). "Who Drives Real Interest Rates Around the Pacific Rim: the USA or Japan?," *Journal of International Money and Finance*, (14)6, 801-821.

Cumby, R. and Obstfeld, M. (1984). "International Interest Rate and Price Level Linkages Under Flexible Exchange Rates: a Review of Recent Evidence," in Bilson, J. and Marston, R. C., eds., *Exchange Rate Theory and Practice*, University of Chicago Press, Chicago.

Edison, H. J., and Pauls, B. D. (1993). "A Re-Assessment of the Relationship Between Real Exchange Rates and Real Interest Rates: 1974-1990," *Journal of Monetary Economics*, 31, 165-187.

Edwards, S. (2001). "Capital Mobility and Economic Performance: Are Emerging Economies Different?", *NBER Working Paper*, nº 8076.

Elliott, G. (1999). "Efficient Tests for a Unit Root when the Initial Observation is Drawn from its Unconditional Distribution," *International Economic Review*, 40, 767-783.

Elliott, G., Rothenberg, T. J. and Stock, J.H. (1996). "Efficient Tests for an Autoregressive Unit Root," *Econometrica*, 64, 813-836.

Enders, W. and Granger, C.W.J. (1998). "Unit Root Tests and Asymmetric Adjustment with an Example Using the Term Structure of Interest Rates," *Journal of Business and Economic Statistics*, 16, 304-311.

Evans, L.T., Keef, S.P. and Okunev, J. (1994). "Modelling Real Interest Rates," *Journal of Banking and Finance*, (18)1, 153-165.

Fisher, I. (1930). *The theory of Interest*. New York: McMillan, 1930.

Frankel, J. A. and Okongwu, C. (1995). "Liberalized Portfolio Capital Inflows in Emerging Markets: Sterilization, Expectations, and the Incompleteness of Interest Rate Convergence," *NBER working paper*. nº 5156.

Frankel, J. A., Schmukler, S. L. and Servén L. (2002). "Global Transmission of Interest Rates: Monetary Independence and Currency Regime," *NBER Working Paper,* nº 8828.

Fujii, E. and Chinn, M.D. (2000). "Fin de Siecle Real Interest Parity," *NBER Working Paper*, nº. 7880.

Gagnon J.E. and Unferth, M. D. (1995). "Is There a World Real Interest Rate?," *Journal of International Money and Finance*, (14)6, 845-855.

Jorion, P. (1996). "Does Real Interest Parity Hold at Longer Maturities?," *Journal of International Economics*", (40)1-2, 105-126.

Knot, K. and de Haan, J. (1995). "Interest Rate Differentials and Exchange Rate Policies in Austria, The Netherlands, and Belgium," *Journal of Banking and Finance*, 363-386.

Kraay, A. (2003). "Do High Interest Rates Defend Currencies During Speculative Attacks?" *Journal of International Economics*, 59, 297-321.

Kwiatkowski, D., Phillips, P. C.B., Schmidt, P. and Shin, Y. (1992). "Testing the Null Hypothesis of Stationary against the Alternative of a Unit Root," *Journal of Econometrics*, 54, 159-178.

MacDonald, R. and Nagayasu, J. (2000). "The Long-Run Relationship Between Real Exchange rates and Real Interest rates Differentials: A Panel Study." *IMF Staff Papers,* 47, 116-128.

Mancuso, A. J., Goodwin, B. K. and Grennes, T. J. (2002). "Non-Linear Aspects of Capital Market Integration and Real Interest Rate Equalization," *International Review of Economics and Finance*, forthcoming.

Meese, R. and Rogoff, K. (1988). "Was It Real? The Exchange Rate-Interest Differential Relation over the Modern Floating-Rate Period," *The Journal of Finance*, 43, 933-948.

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.

Nakagawa, H. (2002). "Real Exchange Rates and Real Interest Rate Differentials: Implications of Nonlinear Adjustment in Real Exchange Rates," *Journal of Monetary Economics*, 49, 629-649.

Neftci, S. N. (1984). "Are Economic Time Series Asymmetric Over the Business Cycles?" *Journal of Political Economy*, 85, 281-291.

Obstfeld, M. and Rogoff, K. (2000). "The Six Major Puzzles in International Finance: Is there a Common Cause?" *NBER Macroeconomics Annual*, 15, 339-390.

Obstfeld, M. and Talyor, A. M. (2002). "Globalization and Capital Markets," *NBER working paper*, nº 8846.

Ong, L. L., Clements, K. W. and Izan, H.Y. (1999). "The World Real Interest Rate: Stochastic Index Number Perspectives," *Journal of International Money and Finance*, (18)2, 225-249.

Pakko, M.R. (2000). "Do High Interest Rates Stem Capital Outflows?" *Economics Letters*, 67, 187-192.

Perron, P. (1989). "The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis," *Econometrica*, 57, 6, 1361-1401.

Perron, P. (1997). "Further Evidence on Breaking Trend Functions in Macroeconomic Variables," *Journal of Econometrics*, 80, 2, 355-385.

Phylaktis, K. (1999). "Capital Market Integration in The Pacific Basin Region: an Impulse Response Analysis," *Journal of International Money and Finance* (18)2, 267-287.

Radelet, S. and Sachs, J.D. (1998). "The East Asian Financial Crisis: Diagnosis, Remedies, Prospects," *Brookings Papers on Economic Activity*, 1, 1-74.

Stiglitz, J.E. (1999). "Interest Rates, Risk, and Imperfect Markets: Puzzles and Policies," *Oxford Review of Economic Policy*, 15(2), 59-76.

UNCTAD (1999). "Payments Deficits, Liberalization and Growth in Developing Countries." In: *Trade and Development Report.* UNCTAD, Geneva.

Figure 1. Real interest rate differentials

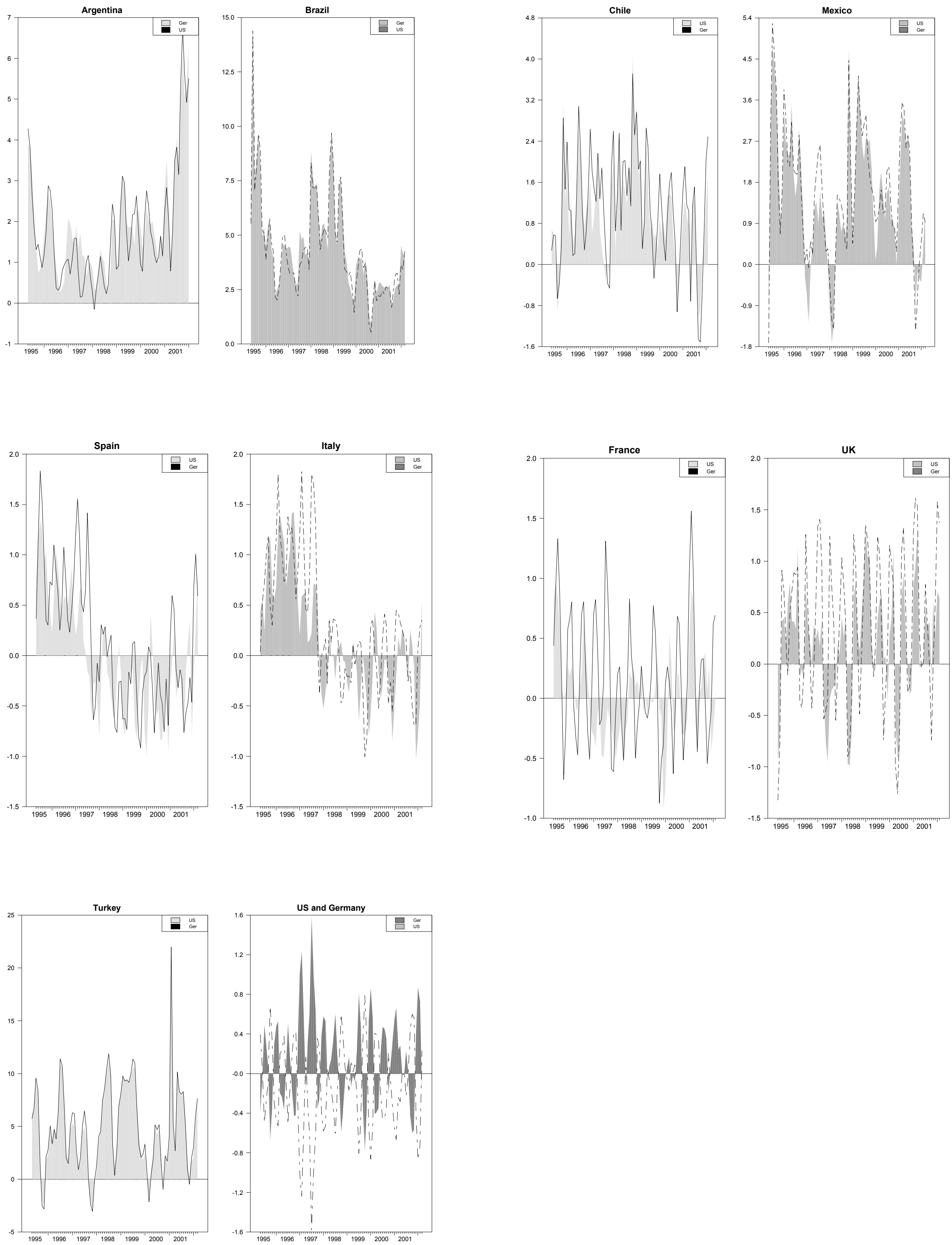

Table 1. Unit Root Tests

Country - Reference	Nº of Lags	ADF	KPSS	ERS (DF-GLS)	Elliot (1999) (DF-GLS$_\mu$)
Argentina (US)	10	-0.421	0.389*	-0.855	-0.878
Argentina (Ger)	10	0.418	0.411*	-0.738	-0.554
Brazil (USA)	3	-3.015*	0.702	-2.489*	-2.937*
Brazil (Ger)	3	-3.119*	0.751	-2.852*	-3.077*
Chile (US)	12	-2.054	0.179*	-1.971*	-2.086
Chile (Ger)	6	-2.024	0.260*	-1.587	-2.041
Mexico(US)	1	-4.593*	0.239*	-2.901*	-4.580*
Mexico(Ger)	1	-4.890*	0.222*	-2.690*	-4.821*
Turkey(US)	3	-4.230*	0.078*	-3.948*	-4.201*
Turkey(Ger)	7	-3.827*	0.067*	-3.734*	-3.849*
France (US)	10	-2.259	0.186*	-0.424	-1.783
France (Ger)	9	-2.490	0.199*	-0.845	-1.785
Italy(US)	9	-1.485	0.688	-0.945	-1.312
Italy(Ger)	9	-0.993	0.653	-1.357	-1.065
Spain(US)	9	-1.759	0.643	-0.349	-1.144
Spain (Ger)	12	-1.533	0.527	-0.990	-1.233
UK (US)	10	-1.723	0.153*	-0.434	-1.624
UK (Ger)	12	-3.203*	0.222*	-0.199	-2.796*
US (Ger)	7	-2.752	0.082*	-0.795	-1.803

Notes:
* indicates rejection of the null of a unit root at the 5% confidence level for the ERS (1996) and Elliott (1999) tests and acceptance of the null for the KPSS test.

Table 2. Perron (1997) tests

	Break Search Method I			Break Search Method II		
Country – Reference	**Lags**	**Break Date**	**T-ratio**	**Lags**	**Break Date**	**T-ratio**
Argentina (US)	9	2001:07	-1.529	9	1999:04	-0.055
Argentina (Ger)	9	2001:07	-1.238	9	1999:04	0.738
Brazil (USA)	0	1995:09	-5.184*	5	1999:04	-3.683
Brazil (Ger)	0	1995:08	-5.224*	5	1999:04	-3.707
Chile (US)	0	1998:10	-5.726*	12	1997:10	-2.046
Chile (Ger)	2	1997:10	-4.837*	10	1998:02	-1.056
Mexico(US)	1	1998:09	-4.613*	1	1998:09	-4.613*
Mexico(Ger)	1	1998:09	-4.901*	7	1998:12	-3.084
Turkey(US)	3	1999:07	-4.299*	3	1999:08	-4.068*
Turkey(Ger)	1	1999:07	-4.984*	1	1999:08	-4.758*
France (US)	10	1996:10	-2.637	10	1996:11	-2.493
France (Ger)	10	2000:11	-2.780	10	2000:11	-2.780
Italy(US)	3	1997:08	-3.315	11	1996:11	-4.128*
Italy(Ger)	9	1997:11	-4.729*	9	1997:08	-4.747*
Spain(US)	9	1997:05	-3.375	9	1997:04	-3.439
Spain (Ger)	9	1997:06	-4.843*	9	1997:07	-4.390*
UK (US)	9	1998:03	-2.871	9	1998:04	-2.444
UK (Ger)	12	2001:09	-3.824*	10	2000:11	-3.311
US (Ger)	7	1996:10	-3.114	7	1996:11	-2.897

Notes:
* Indicates rejection of the null of a unit root at the 5% confidence level.

Table 3. Half-Lives

	ADF		Structural Break			
			Method 1		*Method 2*	
Country – Reference	**Estimated Root**	**Half Life (months)**	**Estimated Root**	**Half Life (months)**	**Estimated Root**	**Half Life (months)**
Argentina (US)	0.95	13.1	0.81	3.2	0.99	72.8
Argentina (Ger)	1.06	--	0.80	3.1	1.14	--
Brazil (USA)	0.79	3.0	0.55	1.2	0.64	1.6
Brazil (Ger)	0.77	2.7	0.56	1.2	0.62	1.5
Chile (US)	0.54	1.1	0.48	0.9	0.54	1.1
Chile (Ger)	0.61	1.4	0.30	0.6	0.74	2.3
Mexico(US)	0.66	1.6	0.65	1.6	0.65	1.6
Mexico(Ger)	0.62	1.5	0.62	1.4	0.56	1.2
Turkey(US)	0.48	1.0	0.46	0.9	0.48	1.0
Turkey(Ger)	0.34	0.6	0.49	1.0	0.51	1.0
France (US)	0.58	1.3	0.43	0.8	0.46	0.9
France (Ger)	0.51	1.0	0.33	0.6	0.33	0.6
Italy(US)	0.91	7.3	0.60	1.4	0.63	1.5
Italy(Ger)	0.94	10.3	0.19	0.4	0.46	0.9
Spain(US)	0.87	5.0	0.40	0.8	0.46	0.9
Spain (Ger)	0.88	5.6	0.51	1.0	0.45	0.9
UK (US)	0.68	1.8	0.52	1.1	0.55	1.1
UK (Ger)	0.10	0.3	-0.03	0.2	0.21	0.5
US (Ger)	0.58	1.3	0.50	1.0	0.53	1.1

Notes: Half-lives were calculated according to the formula $-\left(\frac{\ln(2)}{\ln(1-\psi)}\right)$, where ψ is the estimated Autoregressive coefficient in the ADF equation.

Table 4. Equilibrium Level of *rids*

	ADF equation			Perron (1997) Method 1			
				Period I		*Period II*	
Country – Reference	Intercept	Estimated Root	Long Run Equilibrium Value	Intercept	Long run Equilibrium Value	Intercept	Long run Equilibrium Value
Argentina (US)	0.12	0.95	2.33	0.29*	1.48*	2.01	10.42
Argentina (Ger)	-0.05	1.06	0.82	0.33*	1.67*	1.96	9.90
Brazil (USA)	0.80*	0.79*	3.82*	4.94*	11.10*	1.75*	3.92*
Brazil (Ger)	0.88*	0.77*	3.87*	5.02*	11.54*	1.79*	4.12*
Chile (US)	0.45**	0.54*	0.98**	0.55*	1.05*	0.39	0.74
Chile (Ger)	0.44**	0.61*	1.12**	0.95*	1.37*	0.79	1.13
Mexico(US)	0.47*	0.66*	1.37*	0.39*	1.12*	0.57	1.64
Mexico(Ger)	0.58*	0.62*	1.25*	0.52*	1.35*	0.66	1.73
Turkey(US)	2.30*	0.48*	4.44*	2.54*	4.70*	2.28	4.22
Turkey(Ger)	3.29*	0.34*	4.95*	2.61*	5.09*	2.15	4.19
France (US)	-0.02	0.58*	-0.05	0.11	0.20	-0.04	-0.08
France (Ger)	0.05	0.51*	0.10	0.05	0.08	0.17	0.25
Italy(US)	-0.01	0.91	-0.11	0.32*	0.79*	-0.07*	-0.17*
Italy(Ger)	-0.02	0.94	-0.31	0.84*	1.03	-0.07*	-0.08
Spain(US)	-0.04	0.87**	-0.31	0.32*	0.53*	-0.20*	-0.34*
Spain (Ger)	-0.02	0.88	-0.17	0.33*	0.68*	-0.13*	-0.26*
UK (US)	0.07	0.68	0.22	0.03	0.07	0.16	0.34
UK (Ger)	0.31*	0.10*	0.35*	0.35*	0.34	0.70*	0.68
US (Ger)	0.06**	0.58*	0.15**	0.00	0.00	0.09	0.19

Notes:
1) We used the intercept model to calculate long run equilibrium levels.
2)The null hypothesis is that the long run equilibrium level is equal to zero.
* denotes significance at 5%
** denotes significance at 10%

Table 5. M-TAR model for *RIDs*.

Country	*ARGUS*	*ARGER*	*BRAUS*	*BRAGER*	*CHIUS*	*CHGER*	*MEXUS*	*MEXGER*	*TURUS*	*TURGER*
λ	-0.381	0.247	-0.459	-0.420	-0.237	-0.502	-0.886	-0.872	-0.926	1.290
m	2	5	2	2	1	3	1	1	1	2
μ_1	-0.013 (-0.055)	-0.177 (-0.524)	1.679 (1.829)	2.468 (2.572)	1.428 (2.610)	1.358 (2.389)	0.687 (2.421)	0.643 (2.237)	3.027 (2.997)	5.956 (3.196)
μ_2	-0.584 (-1.867)	-0.187 (-0.515)	0.696 (0.770)	-0.178 (-0.189)	0.207 (0.325)	0.714 (1.414)	0.194 (0.293)	0.543 (1.956)	-2.427 (-1.337)	0.355 (0.200)
ρ_1	-0.031 (-0.213)	0.002 (0.011)	-0.472 (-3.558)	-0.397 (-3.718)	-0.720 (-2.727)	-0.799 (-3.168)	-0.341 (-4.127)	-0.377 (-4.532)	-0.506 (-4.127)	-0.760 (-3.946)
ρ_2	-0.355 (2.260)	-0.454 (-1.739)	-0.123 (-0.786)	0.030 (0.187)	-0.304 (-1.001)	-0.131 (-0.486)	-0.217 (-1.658)	-0.639 (-2.211)	-0.402 (-1.180)	-0.274 (-1.452)
W *p-value*	**17.64** ***0.040***	6.128 *0.343*	0.865 *0.903*	**9.417** ***0.023***	5.367 *0.440*	**8.210** ***0.010***	1.56 *0.707*	2.43 *0.487*	**21.470** ***0.000***	**27.912** ***0.000***
R1 *p-value*	2.954 *0.227*	0.532 *0.587*	**9.112** ***0.010***	**10.256** ***0.003***	**8.695** ***0.050***	**8.137** ***0.023***	**22.603** ***0.000***	**26.154** ***0.000***	**21.16** ***0.000***	**15.866** ***0.006***
R2 *p-value*	4.017 *0.360*	1.234 *0.733*	**9.139** ***0.047***	**13.859** ***0.000***	8.437 *0.133*	**7.675** ***0.087***	**19.783** ***0.000***	**25.425** ***0.000***	**20.575** ***0.003***	**15.960** ***0.037***
Lag	10	10	3	3	12	6	1	1	3	7

Table 5. Continued

Country	*FRAUS*	*FRAGER*	*ITAUS*	*ITAGER*	*SPAUS*	*SPAGER*	*UKUS*	*UKGER*	*USGER*
λ	0.264	0.399	-0.087	0.347	0.030	-0.220	-1.110	0.801	0.506
m	6	5	4	5	2	6	3	1	1
μ_1	0.627 (2.381)	0.004 (0.020)	0.026 (0.153)	-0.453 (-1.416)	0.100 (0.615)	0.044 (0.264))	-0.025 (-0.191)	-0.818 (-0.917)	0.510 (2.433)
μ_2	-0.221 (-1.659)	0.177 (1.482)	-0.048 (-0.279)	0.290 (1.455)	-0.448 (-3.029)	-0.230 (-1.145)	-0.441 (-0.964)	0.328 (2.774)	0.019 (0.262)
ρ_1	-0.278 (-1.198)	-0.566 (-2.699)	-0.124 (-0.945)	-0.005 (-0.031)	-0.126 (-0.834)	-0.092 (-0.688))	-0.388 (-1.482)	-0.313 (-0.444)	-0.379 (-1.798)
ρ_2	-0.382 (-1.757)	-0.472 (-2.139)	-0.041 (-0.273)	-0.204 (-1.502)	0.074 (0.515)	-0.044 (-0.242)	-0.795 (-2.202)	-0.903 (-3.235)	-0.549 (-3.671)
W *p-value*	5.941 *0.357*	**16.984** ***0.040***	**15.679** ***0.016***	8.631 *0.183*	**13.504** ***0.057***	4.774 *0.433*	0.727 *0.947*	5.597 *0.437*	**20.420** ***0.006***
R1 *p-value*	**7.710** ***0.060***	**7.977** ***0.043***	0.823 *0.523*	0.823 *0.416*	0.135 *0.783*	1.08 *0.423*	**5.432** ***0.090***	**16.754** ***0.000***	**10.105** ***0.010***
R2 *p-value*	7.105 *0.177*	3.871 *0.353*	0.968 *0.790*	1.892 *0.583*	0.804 *0.773*	0.734 *0.846*	6.027 *0.217*	**10.782** ***0.093***	**16.098** ***0.010***
Lag	10	10	9	9	9	12	11	12	7

Notes: T-ratios in parentheses. Bold indicates rejection of the null of symmetry or unit roots at the 10% level. P-values for the asymmetry and unit-root tests were obtained by the bootstrap method of Caner and Hansen (2001). W is the Wald test for asymmetry and R1 and R2 are Wald tests for the null of a unit root assuming an unidentified and an identified threshold respectively.