UNIVERSITY OF MACEDONIA

INTERDEPARTMENTAL PROGRAMME OF POSTGRADUATE STUDIES IN ECONOMICS

DISSERTATION

"AN EMPIRICAL ANALYSIS OF REAL INTEREST RATE PARITY FOR INDUSTRIALIZED COUNTRIES"

MARIA C. SARPOTA

SUPERVISOR:

S. FOUNTAS

PROFESSOR IN THE DEPARTMENT OF ECONOMICS, UNIVERSITY OF MACEDONIA

THESSALONIKI 2009

CONTENTS

LIST OF TABLES

LIST OF FIGURES

ABSTRACT

This study examines whether the real interest rate parity hypothesis holds, using two different methods for computing the real interest rate. We present empirical evidence on the RIP hypothesis for thirteen industrialized countries against the US in the 1967-2008 period. This is done by employing not only classical regression analysis and standard cointegration tests, but also cointegration tests that determine the regime shift endogenously. Our results provide strong evidence in favour of bilateral real interest rate convergence between the US and several countries in our sample, in particular for long-term real interest rates. The evidence suggests that deviations from RIP have a half-life of approximately 6-9 quarters. We also provide an application of approximation of the ESTAR model, which allows for possible nonlinearities in international real interest rate dynamics.

KEY WORDS: Real interest rate parity; Cointegration tests; Regime shift; ESTAR model

INTRODUCTION

"How internationally mobile is the world's supply of capital? Does capital flow among industrial countries to equalise the yield to investors? Alternatively, does the saving that originates in a country remain to be invested there? Or does the truth lie somewhere between these two extremes? The answers to these questions are not only important for understanding the international capital market but are critical for analysing a wide range of issues…" [Feldstein and Horioka (1980), p.314]

The questions stated on the quote above, posed by Feldstein and Horioka (1980), still raise intense debate and resilient disagreement. It is peculiar that the liberalization of capital and goods markets carried out in the last decades and the increasing speed of capital movement have not sealed the enigma put forward by Feldstein and Horioka (1980) more than twenty years ago. On the contrary, according to Obstfeld and Rogoff (2000) this is still "one of the most robust and intractable puzzles in international finance".

The extent to which real interest rates are equalized across countries has occupied researchers for a number of reasons. In an open economy, real interest rates play a key role in influencing real activity through saving and investment behaviour. The real interest rates reflect the costs of borrowing and the returns from lending, adjusting for the inflation that is expected to occur over the period of time until maturity. Movements in the real interest rates are an important channel by which monetary shocks are transmitted to real economic activities. Although many studies exist on the pattern of the real interest rate, results are mixed and far from convincing. The behaviour of the real interest rate is important for policy implication. To become more specific, if the real interest rate is constant over time, the effects of monetary policy or fiscal policy on an economy are limited. Moreover, the movements in real interest rates can be used as a guide to conduct monetary policy such that factors that cause persistent changes in the real rate may call for a different policy than factors that cause temporary changes in the real rate.

Confirmation or rejection of real interest parity (RIP) provides an indication of whether countries are financially integrated or autonomous. This is particularly relevant in an era of high or perfect capital mobility in the European Union (EU). However, since RIP requires that *ex ante* purchasing power parity (PPP) holds, it can be viewed as a more general indicator of macroeconomic integration or convergence.

RIP is also important because it is an assumption in several monetary models of exchange rate determination such as Frenkel (1976) and Mussa (1986), which imply that RIP holds in the long-run. The purpose of this study is to search for long-run RIP among thirteen major industrialized countries. This study finds that RIP does not hold among eight of thirteen industrialized countries.

In the last two decades the financial systems of industrialized countries have gone through some profound changes. Capital markets have considerably developed and lots of financial innovations have emerged. Furthermore, we have witnessed a substantial shift toward institutionalized management of savings. National and international boundaries that limited the geographic scope of trade in financial services have been eroded. The activities performed by banks have altered to keep pace with this transformation. The main driving forces behind these developments were the significant demographic changes, the wave of financial liberalisation, the information technology revolution that characterized the past two decades, as well as the launch of the European Monetary Union (EMU).

We examine the existence of Real Interest Rate Parity conditions, as these are defined by the general theory of Purchasing Power Parity theory, among several industrialized countries. The study contains five main sectors. In the next section we present a literature review of the previous studies in the subject of Real Interest Rate Parity (RIP) and the general theory of this hypothesis. In section 2 we discuss the data used to generate the real interest rates and present our results for various unit root tests. Section 3 presents the linear tests of RIP when two different methods of estimating real interest rates are used and their empirical results. Section 4 deals with the effects of the choice of interest rate methodology on nonlinear tests of RIP and discusses the approximation of STAR models used in the analysis. Concluding remarks are given in the final section.

CHAPTER 1: THE THEORY OF THE REAL INTEREST PARITY

1.1 REVIEW OF LITERATURE

The test of equality of interest rates across countries dominates the literature trying to measure the degree of financial integration among nations. Such tests are based on the premise that, in a highly integrated financial market, the law of one price should hold. The law of one price implies that similar assets should yield the same return irrespective of the country of domicile and the currency in which they are denominated, if there is free movement of capital. Hence, the empirical literature on interest parity tries to use the extent of equality of interest rates to measure the degree or intensity of financial integration (see for instance, Cheung, Chinn and Fujii, 2003).

As far as agents make forecasts using rational expectations, arbitrage in goods and assets markets ensures that the real interest rate parity hypothesis holds. Arbitrage is formalized by the uncovered interest rate parity (UIP) and the relative purchasing power parity (PPP) conditions under the assumption of perfect markets. When the assumption of risk neutrality does not hold, then speculation is the driving force behind commodity and asset prices. It seems that Roll (1979) was the first author who noticed that PPP, UIP and rational expectations altogether implied real interest parity (RIP) hypothesis. The hypothesis is based on the assumptions that homogeneous goods are costless traded across countries and that arbitrageurs face a risk-free bond economy with perfect asset substitutability and perfect capital mobility. The simple monetary model of exchange rate determination is a theory fully compatible with RIP hypothesis. The speed of adjustment to equilibrium in this model is so high that equality of real rates holds at all times.

The real interest rate is one of the most widely studied variables in economics. In particular, numerous international economists have been concerned with whether or not real interest rates between countries are fundamentally connected. The early studies employed classical regression to test the existence of a linkage between countries' real interest rates and rejected the validity of the theory. For example, Cumby and Miskin (1986) test the international equality of real rates by regressing the *ex ante* real rate in a foreign country on the US real rate. They reject the RIP hypothesis; however, there is evidence that real rates in all countries show comovement with the US real rate, but this linkage is not complete (less than unity estimated coefficients). Mishkin (1984) investigates and rejects empirically three of the international parity conditions: RIP, *ex ante* PPP, and UIP. Mark (1985) also fails to find much support for the equality of real rates across countries, particularly for the US and the European countries. Similarly, Fraser and Taylor (1990) use a bivariate vector autoregression to test real interest parity and find a strong rejection of the RIP for real interest rate differentials between major OECD countries.

An alternative empirical test of interest parity has emerged in the context of cointegration and stationarity tests. This approach allows real interest rates across countries to move in a stochastic pattern and evaluate the long-run tendency toward parity relationship. According to Goodwin and Grennes (1994), situations under which rates were found to diverge from their long-run equilibrium relationship give evidence of a breakdown in the parity or equality relationship. The alternative tests were applied to real interest rates calculated from Eurocurrency and domestic money market rates for the US and nine other important countries. Their analysis has argued that the overwhelming lack of support for real interest rate equalization obtained by conventional tests may have resulted from biases raised by ignoring transactions costs. Specifically, non-synchronous variation of individual rates in response to localized financial conditions within the band created by transactions costs may have led to incorrect rejection of interest equalization or interest parity, although the markets in question were fully efficient and integrated. In addition, the presence of unit-roots in the real interest rate series utilized to evaluate interest equalization may have led to incorrect statistical inferences in conventional tests. The empirical results revealed much stronger support for the theoretical parity relationship than is commonly found in the literature. However, this support remained incomplete in that a breakdown in the parity relationship was revealed for a small number of the markets. In all, the results were reasonably consistent with the notion of a long-run equilibrium relationship between real interest rates in the US and rates in the nine other countries. In other words, their results provided strong evidence in favor of market efficiency and integration among the ten financial markets considered and suggest a much stronger link among the ten financial markets than is implied by the existing empirical literature.

Gagnon and Unferth (1995) utilize panel data techniques to evaluate real interest rate equality. Their results are favorable for the existence of a world real interest rate and find that the real interest rates of the OECD countries are highly correlated with the world interest rate such that markets are tightly integrated. Moreover, Kugler and Neusser [1](1993) , Moosa and Bhatti (1996) and Al Awad and Goodwin (1998), find that real interest rates among OECD countries are strongly linked in long-run equilibrium, while Holmes and Maghrebi (2004) find evidence that supports the RIP in South East Asian economies. Recent research on RIP recognizes the structural shifts and nonlinear stochastic dynamics of interest differential and tries to capture these effects through sophisticated models and tests. For instance, Mancuso, Goodwin, and Grennes (2003) utilize a threshold time series model and nonparametric regression to allow real interest rate differentials to adjust nonlinearly to the mean. Fountas and Wu (1999) use cointegration techniques that allow for structural shifts in the cointegrating vector. Wu and Chen (1998) examine the RIP hypothesis with panel unit-root tests. These studies find evidence that supports the convergence of the real interest rates across countries.

More recent research has focused on investigating the time-series properties of real interest rate differentials (RIRDs). This is achieved through the use of unit root tests to investigate whether these differentials are mean-reverting. Meese and Rogoff (1988), for example, tested for a unit root in long-term RIRDs for the period 1974 to 1986 and could not reject the unit root hypothesis; yet, they rejected it for short-term rates. Similar results are found in Edison and Pauls (1993). These authors used data for Japan, Germany, the UK and Canada against the US dollar and were unable to reject the null hypothesis that the differentials have a unit root using the augmented Dickey-Fuller (ADF) test. However, the negative results may reflect the poor power of the ADF test rather than evidence against RIP. In other words, these tests may fail to reject the unit root hypothesis even when RIRDs exhibit slow reversals to RIP values. This low power problem is magnified for small samples, such as the recent floating experience, because a mean-reverting series could be drifting away from its long-run equilibrium level in the short-run.

Generally, 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 low, leading to

---

[1] Kugler and Neusser (1993) investigated the validity of real interest parity using *ex post* real interest data for several countries in a stationary multivariate time-series approach and provided evidence in favour of RIP.

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 potential problem of unit-root tests is ignoring the possible 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 recognized that macroeconomic time series show strong asymmetry over the business cycle. If asymmetry is present in real interest differentials, linear unit-root tests will suffer from a loss of power. 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 null corresponds to the hypothesis that the variance of the random walk component of the series equals zero or, in other words, the variance of the error is constant. When the series has an unknown mean or linear trend, the tests suggested by Elliott, Rothenberg and Stock (1996) (ERS test hereafter) 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.

To circumvent this problem of the low power of the unit root tests, Obstfeld and Taylor (2002) have sought to increase the power of their tests by increasing the length of the sample period under examination and using the generalized least squares (GLS) version of the Dickey-Fuller (DF) test due to Elliott et al., (1996) that has more power than the conventional ADF test. Their results are generally supportive of RIP for a number of currencies and a sample period dating back to the 1890s. According to Obstfeld and Taylor (2002) "the results are generally favourable to the hypothesis that long-term real interest rates are cointegrated, and thus tend not to stray arbitrarily far apart over time. This finding contrasts with conclusions reached in earlier papers, which were based on shorter samples and weaker statistical tests than those we have used."

In order to understand the causes of RIRDs we need to verify whether and why UIP, PPP and the rational expectations hypothesis fail to hold. A common approach used to examine both the existence and causes of RIRDs is to test the individual arbitrage conditions and RIP separately. This approach was originally employed by Mishkin (1984) who did not find empirical support for RIP and concluded that models based on the assumption of costless international arbitrage cannot explain the behavior of real interest rates better than that those allowing for frictions.

Price sluggishness is a typical friction that causes PPP to be violated in the short-run. In the Dornbusch (1976) model of sticky prices, for example, a real interest differential arises whenever the exchange rate overshoots. On the other hand, transaction costs violate the assumption of perfect capital mobility. For an extensive number of authors, a simple constant term added to previous equations is able to capture transaction costs. Others, such as Phylaktis (1999) and Goodwin and Grennes (1994), credited them to the autoregressive parameters of the real interest differential. Finally, there are models based on international arbitrage in which transaction costs generate non-linearities and real interest differentials [Obstfeld and Rogoff (2000)].

In summary, the violation of the two parity conditions, relative PPP and UIP and the rational expectations hypothesis are associated with the existence of real interest rate differentials. Factors such as default and risk premium, transaction costs, price sluggishness, systematic forecast errors, government spending (Allen,1990) changes in the level or growth of money supply, (un)expected productivity or output increases and several other macroeconomic fundamentals can explain the causes of RIRDs.

So, why do some empirical studies fail to replicate the theoretical RIP relationship, while others seem to support the theory? The inconclusive findings may result from the fallacy of some assumptions required for RIP to hold. For example, Chung and Crowder (2003) pointed out that all of the studies in RIP define their real interest rates by the Fisher equation, given by

$$i_t = r_t^e + \pi_{t+1}^e \quad (1.1)$$

which implies that the nominal rate of interest i_t can be thought of as the equilibrium expected real return r_t^e plus the market's assessment of the expected rate of inflation π_{t+1}^e. Whether the Fisher relation holds or not is ambiguous, even though several studies have been devoted to examining the validity of Fisher relation[2]. Alternatively, it is also possible that the real interest rate series that authors used in RIP tests contribute to this conflicting evidence. The problem arises in testing and estimating the linkage between *ex ante* real rates in different countries as a result of unobserved expected inflation, and hence the *ex ante* real interest rate. Therefore, the

[2] See, for example, Fama (1977), Huizinga and Mishkin (1984), and Crowder and Hoffman (1996).

series of *ex ante* real rates must be estimated based on observed data with underlying assumptions in making inferences about the *ex ante* real interest rate. This process allows deviations in the methodology used in constructing the *ex ante* real interest rate and may lead to different conclusions in hypothesis testing. These empirical studies differ not only in how to treat the expected inflation in the real interest rate calculation but also in what proxies to use for the nominal interest and the price variables.

Cumby and Mishkin (1984) tested the comovement of short-term real interest rates in eight countries. They used three month interest rates in the euro deposit and domestic money markets from 1973M6 and 1983M12 for Canada, Italy, the Netherlands, France, West Germany, the UK and the US. Consumer Price Index (CPI) was used as the price index. They regressed the *ex post* real interest rate on a constant term, a time trend, the nominal interest rate and three values of lagged inflation in order to estimate the *ex ante* real interest rate (the fitted values of the regression). Their general conclusion regarding these estimates is that the timing and the extent of real rate movements differ between countries. Furthermore, they also regressed *ex post* real rates of each country against the US interest rate. The hypothesis of no linkage was rejected for all countries except Switzerland while the hypothesis of one-to-one relationship was rejected for five countries. Their finding is that there is a statistical association between real rates in nearly all pairs of countries.

Cavaglia (1992) applied Kalman filtering techniques to estimate the persistence of *ex ante* real interest differentials for the period from 1973 to 1987. He found that *ex ante* real interest differentials are relatively short-lived and mean-reverting to zero, thus providing empirical support for real rate equality in the long-run steady state.

Frankel and Okongwu (1995) proposed to investigate why real interest rates of nine Latin American and East Asian countries during the period form 1987 to 1994 have not converged to US levels in spite of the large amount of capital inflows directed to those countries in that period. Frankel and Okongwu (1995) argued that if the cause of capital inflows was external-as most of the empirical papers before the Mexican crisis have suggested-the interest rate differential should have declined. However, they recognised that a positive relationship between domestic monetary tightening and capital inflows could exist either because inflows are attracted by high interest rates or because it reflects the sterilisation of the inflows. They claimed that a methodological innovation of their work is the use of a direct measure of exchange rate expectations. Frankel and Okongwu (1995) have used exchange rate expectation from survey data on the forecasts of 45 economic agents including multinational firms and forecasting companies. They decomposed the total interest rate differential in three parts: the expected depreciation, the country-risk and the exchange rate premium. For country-risk they have employed either secondary-market debt prices or the spread between the domestic dollar interest rate and the US treasury bills, depending on data availability. For many countries, expected depreciation appeared to be accounting for most of the changes in the interest rate differential. In relation to the degree of capital mobility they found that inflows are, in general, negatively related to the US interest rates and domestic monetary expansion. On the other hand, evidence on the significance of domestic interest rates, specific country effects (measured by dummy variables), country risk and even expected depreciation to explain inflows were dubious. In summary, Frankel and Okongwu (1995) did not find support for perfect capital mobility.

Jorion (1996) investigated the validity of RIP for long-term bonds across the US, UK and Germany for period from 1973M8 to 1991M12. Using a set up for the tests similar to the one employed by Mishkin (1984), results do not support the view that expected real interest rates tend to be equalized over longer maturities, however there is evidence that RIP holds. Using monthly data over the period 1982M1 to 1993M12, Alexakis et al. (1997) demonstrated that RIP is accepted for nine European countries both on a non-EMS and an EMS basis. This relationship proves to be stronger on the EMS.

Siklos and Wohar (1997) studied the relationship between interest rates and inflation rates for 10 countries during the period 1974-1995. They found evidence of a unique cointegrating relationship between nominal interest rates of European Monetary System (EMS) countries, the US and Canada, and the US, Germany, and Japan. No similar relationship was obtained between inflation rates with one exception, namely that between the US and Canada. Then they interpreted these results as convergence in inflation but not in interest rates. Hence, if interest rates represent an indicator of monetary policy, the countries considered have attempted to implement independent policies but not to an extent which produced divergent trends in inflation.

Al-Awad and Goodwin (1998) examined weekly real interest rates for G-10 countries using a variety of time-series tests. These tests give special attention to the time-series properties of nominal interest rates, *ex-ante* expected rates of inflation and real interest rates. Term structure information was used to recover a theoretically consistent measure of *ex-ante* expected inflation. In-sample and out-of-sample Granger causality tests were also examined to evaluate lead/lag relationships among real interest rates. Their results provide strong support for well-integrated markets, but not to real interest rate equality particularly in the long-run. Moreover, the results imply leadership roles for the US in international asset markets.

Fountas and Wu (1999) examined real interest rate convergence in European Countries, by using the Engle and Granger methodology and running tests that allow endogenously determined structural breaks for pairs of countries, and as a result they reported evidence in favor of long term real interest rates convergence. Furthermore, Wu and Fountas (2000) used recently developed cointegration tests that determine endogenously the regime shift to test for bilateral real interest rate convergence (real interest rate parity) in the G7 against the US in the 1974-1995 period. In contrast with previous studies that employed classical regression analysis and standard cointegration tests, their innovative approach provided strong evidence in favor of bilateral real interest rate convergence between the US and several countries in the sample, in particular for short-term interest rates.

Wu and Chen (2001) found that one stylized fact to emerge from the empirical analysis of interest rates is that the unit-root hypothesis in nominal interest rates cannot be rejected. However, using the panel data unit root test Im, Pesaran and Shin (1997) and Wu and Chen (2001) found support for the mean-reverting property of Eurocurrency rates. Thus, neither a vector-error-correction model nor a vector autoregressive model in differences is appropriate for modeling Eurocurrency rates. Instead, conventional modeling strategies with level data are appropriate. Furthermore, the finding of stationary interest rates supports uncovered interest parity, and hence the convergence hypothesis of interest rates. This in turn suggests a limited role for a monetary authority to affect domestic interest rates.

Taylor (2001) finds that long-term real interest rate differentials are stationary and real interest rates in developed countries converge in the long run. Furthermore, there is little consensus regarding the rate of adjustment toward the world interest rate, which has important implications for monetary policy. In addition, it is not clear from the existing convergence studies whether capital movements are sufficient to equalize real rates across countries.

A number of recent studies have also investigated the causes of persistent deviations from RIP. For instance, Chung and Crowder (2004) consider five industrialized nations over the period 1960–96. They find that no single violation can explain the failure of RIP in all cases. It does appear, however, that the Fisher relationship is the least likely to violate the RIP equilibrium, whereas UIP appears to be the most commonly violated relationship. This result is consistent with a non-stationary risk premium in the foreign exchange market. Ferreira (2004) argues that departures from RIP can be explained by *ex post* deviations from PPP and UIP. A question that arises is whether real interest differentials are caused by frictions in the goods or assets markets. The findings of this article reveal the predominance of nominal interest differentials and real shocks in the path of real interest differentials for most countries which point to deviations from UIP as their driving source.

Venetis *et al.* (2004) found evidence of fractional integration for a number of monthly *ex post* real interest rate series using the GPH semi parametric estimator on data from fourteen European countries and the US. However, they posed empirical questions on certain time series requirements that emerged from fractional integration and they found that these did not hold pointing to “spurious” long memory and casting doubts with respect to the theoretical origins of long memory in the sample. Common stochastic trends expressed as the sum of stationary past errors did not seem appropriate as an explanation of real interest rate covariation.

Sekioua (2004a) tested for unit roots on real interest differentials but the focus of his paper is on the persistence of RIRDs. Sekioua (2004a) uses monthly interest rates and prices spanning from the beginning of the first quarter of the 20th century for the UK, Japan, and France relative to the US. Interest rates are long-term government bond yields of maturities of seven years or more. The inflation rate is calculated as the average value of the previous 12 months. The unit root is rejected for the three countries at the 1% significance level. Results are weaker when the sample is divided in sub-periods. The unit root, for example, cannot be rejected for the period of the recent float using the 5% and 1% significance level. He also calculates confidence intervals for the dominant root, which is estimated to be in the vicinity of 1. Using point estimates of the half-lives, Sekioua (2004a) found that it takes approximately 17 months for mean reversion in the UK which he concludes that it is compatible with RIP hypothesis. However, it may take more than 75 months for shocks to die out in France, pointing towards a very high degree of persistence. Point estimates for Japan indicate a half-life of about 24.3 months during the whole period. Sekioua (2004a) also found that the behavior of RIRDs across different exchange rate regimes seems to be uniform. The tests reject the unit root but confidence intervals for the dominant root seem to be high.

Sekioua (2004b) performed a cointegration analysis of RIP for France, Germany, Japan, Switzerland and the UK with respect to the US. The period of the tests is from 1974M1 to 1998M12. He used long-term government bond yields, with maturities of 10 years. Price indexes are the CPI and a price index of traded goods. The Johansen cointegration test is performed within a VAR framework. The finding is that there is at least one cointegrating vector between interest rate and inflation differentials. Deviations from the estimated cointegrating relationship adjust fully within three years.

The recent econometric literature provides evidence of asymmetries in key economic variables. For instance, Coakley and Fuertes (2002) identify the presence of asymmetric dynamics in the behavior of real interest rates. Studies by Enders and Granger (1998) and Enders and Siklos (2001) find evidence of asymmetries in nominal interest rates. Ramsey and Rothman (1996) identify asymmetries in inflation and attribute them to downward price rigidities. Cover (1992) provides more general evidence of asymmetries that corroborates the implications of price adjustment models where prices are primarily sticky in a downward direction. In fact, there are several reasons for regarding such relationships and adjustments as linear with suspicion. It is difficult to impose a linear and symmetric adjustment process in RIP linkages when the speed of realignment towards long-run equilibrium through market arbitrage may differ significantly across the financial and goods markets. Whereas adjustments in financial markets are rapid, they are as noted by Dumas (1992), rather sluggish, gradual and costly in the goods markets. This is consistent with the theoretical assumption of short-term stickiness in goods prices underlying the overshooting model by Dornbusch (1976). It is conceivable that these prices can be rigid or sticky in a downward direction.

The sign and trend of deviations are important in determining how quickly the monetary authorities are likely to respond to deviations from equilibrium. In particular, their importance is manifested under the exchange rate regime of a managed float. Here, monetary authorities may show greater aversion or tolerance towards inflation and currency appreciation, and in turn a greater willingness or reluctance to raise nominal interest rates. As argued by Goodhart (1999), monetary authorities may exhibit a greater aptitude towards raising interest rates on gradual basis to contain inflationary pressures, and lowering them rather more rapidly afterwards. There are recent attempts at modeling asymmetries in monetary policy rules with nonlinear central bank preferences and nonlinear reaction functions to output gaps. As shown by Kim et al. (2005) and others, the international evidence is suggestive of asymmetric monetary-policy rules, which are in turn conducive to asymmetric adjustments towards RIP.

Ferreira and León-Ledesma (2007) test for the real interest rate parity for both developed and emerging markets but do not adjust for a time-varying intercept to reflect the improved macro fundamentals (and hence lower real interest differentials over time) in emerging markets. Their results support the hypothesis of a rapid reversion towards a zero differential for developed countries and towards a positive one for emerging markets. Their evidence reveals a high degree of market integration for developed countries and highlights the importance of risk premia for emerging markets. They also find that asymmetries induced by either risk perception changes or transaction costs seem to be an important feature of the dynamics of real interest rate differentials.

An assessment of the equilibrium relationship between real interest rates across countries is useful in providing a measure of the degree of market frictions or integration. Assessing the economic significance and persistence of deviations from real interest parity (RIP) requires an econometric approach that is able to capture the time-series properties of real interest rates as well as the characteristics of the adjustment process towards long-run equilibrium. Much of the early empirical studies on uncovered interest parity (UIP) and purchasing power parity (PPP), on which RIP theory rests, have traditionally relied on linear dynamics. However, recent evidence suggests that when the nonlinearity of the mean-reversion process is not explicitly recognized, the power of standard stationarity tests remains low. As noted by Taylor et al. (2001), the lack of evidence for a long-run realignment of real exchange rates towards PPP across industrialized countries may be attributed in part to the low power of unit-root tests as they demonstrate that the smaller the deviation from PPP, the higher the likelihood for real exchange rates to exhibit unit-root behavior.

In a different approach Evans and Lewis (1995) allow the data to follow a non-linear process and their different approach allow the data to follow a non-linear process and their results are supportive of the parity relationship. In the context of the debate over fiscal policy rules in a monetary union, Haldane and Pradhan (1992), who have tested for *ex-ante* PPP-CIP and risk-premia effects, suggested that real interest rates are not as yet sufficiently interdependent to support those who have argued that one country's fiscal deficit will necessarily affect fully real interest rates for all other member countries. However, to the extent that a monetary union entails a major regime change, questions of this type are difficult to answer satisfactorily using as benchmark a non-monetary union regime such as the ERM.

Due to the mixed results in past research, the interest in nonlinear aspects of real interest rate convergence has grown rapidly recently. There is further evidence that transactions costs in particular can inhibit realignments towards PPP, UIP and RIP. For example, Balke and Wohar (1998) find non-linearities in the adjustment towards covered interest parity (CIP) between the UK and the US where non-linearity is likely to depend on the magnitude of deviation from CIP relative to the transactions cost bandwidth. In the case of arbitrage in the goods market, transactions costs may straddle the equilibrium value of PPP so that as one moves further away from central parity, arbitrage becomes more feasible (see, inter alia, Obstfeld and Taylor, 1997).

Nakagawa (2002) investigated nonlinear regressions of the real exchange rate on the *ex post* real interest differential and found stronger statistical evidence of a link between the real interest differential and the real exchange rate than in earlier studies that worked in a linear environment. Mancuso et al. (2003) point out that transaction costs, contractual arrangements, trading rules and sluggishness of arbitrage in good markets may contribute to nonlinear functional relationships among real interest rates across countries. Using Threshold Autoregression (TAR) models, they find strong support for nonlinear adjustments towards parity among industrialized countries' real interest rates. In particular, small deviations from parity generally evoke modest (slow) adjustments while large deviations bring about much faster adjustments. If this nonlinearity is ignored, deviations from parity tend to have unreasonably long half-lives, indicating less integration of markets.

Recent empirical work by Holmes and Maghrebi (2004) and Liew et al. (2004), among others highlight the importance of nonlinearities in influencing the outcome of international parity tests. Furthermore, studies have shown that the half-live of shocks in such model is found to be dramatically shorter than that obtained in linear models. Perhaps an important result from all these studies is that evidence in favor of the parity condition markedly strengthens when nonlinearities are accounted for in the adjustment process.

Given the paucity of nonlinear studies of the RIP relationship and the mixed evidence from linear modeling, recent researchers examine asymmetries in the adjustment mechanism towards long-run RIP based on the method of nonlinear cointegration power and are therefore more conducive towards the acceptance of the non-cointegration null. Using a bilateral cointegration test, which has enhanced power in the presence of asymmetric adjustment, our analysis of RIP across thirteen industrialized countries indicates that RIP with respect to the US is more likely to hold in the presence of incremental deviations resulting from falling US real interest rates because the adjustment mechanism towards RIP is relatively faster than under decreasing deviations.

It is obvious, from the empirical analyses described above, that from the 1980's, empirical evidence is showing a change in trend from less to more supportive tests on RIP. These results may reflect, on the one hand, the evolution over the last twenty five years towards a more integrated international financial market, and, on the other hand, the implementation of new developments in econometrics.

1.2 REAL INTEREST RATE PARITY

1.2.1 Fisher Hypothesis

Based on the Fisher equation, the nominal rate of interest can be thought of as the equilibrium expected real return plus the market's assessment of the expected rate of inflation, given by

$$i_t = r^e_t + \pi^e_{t+1} \qquad (1.2)$$

where i_t is the nominal interest rate from holding the one-period bond from t to t+1, r^e_t is the one-period real rate of interest expected for the bond maturing at time t+1; and π^e_{t+1} is the rate of inflation from t to t+1, expected by the agents in the market at time t.

The *ex ante* real interest rate, r^e_t, is defined as:

$$r^e_t = i_t - \pi^e_{t+1} \qquad (1.3)$$

Using realized inflation rate during period t+1, π_{t+1} , one may compute *ex post* real returns from the one-period bond as:

$$r^{p}_{t}=i_{t}-\pi_{t+1} \qquad (1.4)$$

Several methods of constructing *ex ante* real rates are based on the assumption of rationality of inflation expectations. By assuming rational expectations, π^{e}_{t+1} is the mathematical expectation of π_{t+1} conditional on all the relevant information available to the agents at time t. Letting φ_t be the set of all available information at the time inflation expectations are formed, then we have

$$\pi^{e}_{t+1}=E_{t}(\pi_{t+1}|\ \varphi_{t}) \qquad (1.5)$$

and hence

$$\pi_{t+1} - \pi^{e}_{t+1}=\varepsilon_{t+1} \qquad (1.6)$$

where ε_{t+1} is the inflation forecast error with zero mean and, by construction, is uncorrelated with φ_t. Thus the rational expectations hypothesis implies that the *ex ante* real rate equals the *ex post* real rate and the forecast error of inflation

$$r^{p}_{t}\equiv i_{t}-\pi_{t+1}\equiv r^{e}_{t}-(\ \pi_{t+1}-\pi^{e}_{t+1})\equiv r^{e}_{t}-\varepsilon_{t+1} \qquad (1.7)$$

Since the Fisher hypothesis in equation (1.2) implies a neutrality of expected inflation, such that an increase in inflation will not affect real interest rates in the long run, testing this relation is simply done by regressing swings in nominal interest rates against swings on inflation and we hope to obtain a unity hypothesized coefficient. The testing regression is given by

$$i_{t}=\beta_{0}+\beta_{1}\pi^{e}_{t+1}+u_{t} \qquad (1.8)$$

However, if both the nominal rate of interest and inflation are nonstationary processes, then the hypothesis holds if there exists a stationary long-run relationship between these two series.

1.2.2 Real Interest Rate Parity

The real interest rates parity (RIP) is the second strand of theory that researchers test using real interest rates. When agents form their expectations rationally and there is no barrier to trade or capital flow, real interest rates should be equalized across countries. The RIP can be viewed as a more general indicator of whether countries are integrated or autonomous, which has important policy implications such that it constrains the ability of domestic monetary authorities to intervene in foreign exchange markets. This is particularly relevant as trade impediments have been reduced in the past three decades, especially in the industrialized countries. RIP relies upon four parity conditions: the Fisher relation in each country, *ex ante* purchasing power parity (PPP) and the uncovered interest parity (UIP).

Theory states that in a perfect world, arbitrage in goods and assets market, under the assumption of rational expectations, ensures the equality of real interest rates across countries. In other words, if relative PPP, UIP and the efficient market hypothesis hold, interest rate differentials follow a zero mean-reverting process.

Over the past 30 years, ever-increasing global integration in the financial and goods market has turned out to be one of the most significant and profound developments in the world economy. As a result, linkages among national financial markets have gradually strengthened, and an integrated international capital market has started to emerge. In theory, in a one-world market, investors should be able to allocate their capital freely, thereby reducing arbitrage opportunities across countries. In such an environment of growing interdependence among markets, country-specific interest rates should exhibit a long-run convergence trend. Such complete convergence is known as the real interest rate parity (RIP) hypothesis.

The real interest rate parity hypothesis (RIP) 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 equalize. This notion is of practical importance because the violation of real interest rate equality is a necessary condition for domestic monetary authorities to influence policy variables through the real interest rate channel (Mark, 1985). However, despite the significant reduction in barriers to trade that has characterised the economies of industrialized countries in the last few decades, the evidence on the equalization of real interest rates appears to be mixed at best. This indicates that capital and goods market liberalization has yet to reach the stage where rates of return are equalized across national borders (Fujii and Chinn, 2000).

It is advantageous to first formally define RIP. Assume that uncovered interest parity (UIP) holds, such that

$$i - i^{*} = E_t(\Delta s) \qquad (1.9)$$

where i is the domestic nominal interest rate and i^* is the exogenously determined foreign nominal interest rate. The exchange rate is the domestic price of the foreign currency and is represented by s. The expected rate of depreciation of the exchange rate is $E_t(\Delta s)$, conditional on current information. Intuitively, Eq. (1.9) states that differences in the nominal interest rate reflect expected changes in the exchange rate and that these rates adjust to equalize the return on domestic and foreign assets. The UIP condition, as defined above, formalises the impossibility of no exploitable excess profits in the assets market.

Another fundamental relationship between open economies is purchasing power parity (PPP), which can be stated as

$$E_t(\pi) - E_t(\pi^*) = E_t(\Delta s) \qquad (1.10)$$

where π and π^* are the domestic and foreign inflation rates, respectively. Eq. (1.10) means that differences in expected price levels are offset by expected changes in the exchange rate, such that a unit of the domestic currency can purchase the same bundle of goods and services in either country. In other words, PPP states that once converted to a common currency, national price levels should be equal. Using equations (1.9) and (1.10), RIP can be found as

$$i - i^* = E_t(\pi) - E_t(\pi^*)$$

$$i - E_t(\pi) = i^* - E_t(\pi^*)$$

$$r = r^* \qquad (1.11)$$

where the final step utilizes the Fisher relationship, described previously. As the derivation makes clear, RIP is a joint hypothesis of UIP and PPP. Perhaps more importantly, the algebra supports the statement above that RIP is a condition that pertains to *ex ante* interest rates.

Alternatively, the form can be rewritten by using algebra and the *ex post* version of the fisher equations of the domestic and foreign countries as follows: $r_t - r_t^* = (i - i^* - \Delta s) - (p - p^* - \Delta s)$, where r denotes real interest rate, i denotes nominal interest rate, s denotes exchange rate, p denotes inflation rate and Δs is the difference operator. The first three terms on the right hand side represent deviations from the UIP (country premium) and the last three terms deviations from the PPP (exchange risk premium) (Fountas and Wu, 1999).

The real interest rate differential (RIRD) is simply the deviation from RIP expressed as:

$$r_t - r_t^* = RIRD_t \qquad (1.12)$$

If expectations are rational, then:

$$\Delta p_t + \varepsilon_t = \Delta p_t^e \qquad (1.13)$$

$$\Delta p_t^e + \varepsilon_t^* = \Delta p_t^{*e} \qquad (1.14)$$

where the forecast errors of inflation, ε_t and ε_t^* are I(0). In this case, tests for *ex-post* or *ex-ante* differentials are equivalent (Mishkin, 1992).

Implicitly, Eq. (1.11) assumes that there are no transaction costs and that any difference between a domestic and a foreign interest rate is arbitraged away. Transaction costs may alter these dynamics. If the difference between the rates is less than the cost of actually realizing the profit, then the transaction will not take place. If, however, the profit is greater than the cost of capitalizing on it, then arbitrage will occur and the rates will tend to converge. Even if two capital markets are perfectly integrated, transaction costs may delineate an area in which rates have no tendency to equalize or converge. Econometric techniques, which do not account for this behaviour, may incorrectly reject a hypothesis of integration.

The deviation from UIP is due to the country premium (e.g. capital controls, differential tax systems, political risk) and the currency premium (i.e. exchange risk premium). The developments in the international financial markets in the 1970s and 1980s would be expected to lead to changes in these premia. For example, the increasing dismantlement of capital controls and the increasing integration among national financial markets in industrial countries would be a contributing factor to the reduction of the country premium. On the other hand, the increasing volatility of exchange rates following the collapse of the Bretton Woods system would be associated with an increase in the currency premium. In addition, the increasing volatility of exchange rates would most probably be associated with increasing deviations from *ex ante* PPP as the highly variable exchange rates would deviate from the less variable price levels. In summary, the impact of exchange rate flexibility and the integration of financial markets on real interest rate convergence would be ambiguous as the first factor tends to contribute to interest rate divergence whereas the second factor tends to point towards interest rate convergence.

Broadly speaking, the tendency for interest rates to equalize can result from two causes. The first cause derives from the presence of arbitrage[3] opportunities, in which interest rate movements are viewed as being determined by 'financial flows in fluid, profit-seeking capital markets' (Barassi, Caporale and Hall, 2000). The interest rate parity theory argues that, with a high degree of international capital mobility resulting from capital account liberalization, two countries' financial assets would be substitutes for each other, and arbitrage brings one country's interest rates into parity with the interest rates of the other, plus the forward premium on the two currencies. Thus, the two interest rates may move together over time when the forward premium has stationary time series properties (Zhou, 2003). Based on this view, empirical studies often test the different interest parity conditions, namely: covered interest parity (CIP), uncovered interest parity (UIP), real interest parity (RIP) and closed interest parity (CLIP) conditions. In the context of a currency bloc, such as a common monetary area, where exchange rate risk (uncertainty about the future value of a currency) is absent, an empirical test of the parity conditions simply measures the co-movement between the two interest rates (Adam, et al., 2002).

The second cause arises from the use of interest rates as policy instruments, so that a policy objective such as exchange parity or an inflation target may determine their time paths. In this view, co-movement in interest rates is considered as a product of policy convergence. This may occur when a smaller country (financial unit) aligns its interest rates (policy) with that of a dominant economy (financial unit), because of the possibly stronger influence the latter will exert on the former. This explains why the US and Japan have been dominant in the international financial markets, with the result that many countries (financial markets) try to align their interest rates to reflect the trends in the US or Japan.

It is not expected, however, that perfect integration (full interest rate parity) can be achieved anywhere in the world, even in the most advanced economies with highly developed and liberal financial systems. Often there are several barriers such as asymmetric information, transaction costs, differences in tax systems, political and sovereign risks and the like, that impede the process of integration. Besides, in most

[3] Arbitrage refers to a combination of transactions designed to profit from an existing discrepancy amongst prices/interest rates in different markets without risk of these changing (Deardorff's Glossary of International Economics).

developing countries, banking systems are usually highly regulated and often enjoy high market power because of limited competition, thereby making market determined lending and deposit rates unobservable.

In summary, real interest rate parity is an essential assumption in most open-macroeconomic models. This assumption states that rates of interest for similar assets in two different countries must be equal once they have been adjusted by their respective expected inflation rates. The policy implication of this assumption is straightforward. In a context where goods and capitals flow freely and real interest rates are settled in the international markets, individual countries will find their scope for stabilization policies very limited. In other words, the scope of economic policies over real economic variables depends to a great extent on the degree to which international real interest rates can influence domestic monetary policy.

CHAPTER 2: DOES THE REAL INTEREST RATE PARITY HOLD? EMPIRICAL TESTS AND RESULTS

2.1 DATA

We use both short term and long term interest rates for thirteen industrialized countries: Australia, Belgium, Canada, France, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Spain, Switzerland and the UK[4]. The US is taken as the numeraire (home) country. The choice of the United States as the reference country is motivated by the fact that it is the main trading partner of the countries involved.

Data on interest rates was obtained form the International Financial Statistics (IFS), of the International Monetary Fund (IMF). Among the several categories of interest rates available in the IFS database, we considered deposit rate (as short-term interest rate) and general government bond (as long-term interest rate) as being the most appropriate for the tests. The inflation rates are constructed using the Consumer Price Index (CPI).

The data are quarterly and cover the period from the first quarter of 1967 till the fourth quarter of 2008 even though, in some cases, data concerning the most recent quarters are not available. It should be noted that the period is slightly different for each country.

We first explore if the computations of the inflation rate affect the dynamics of the obtained inflation series and hence the constructed real interest rates. There are two common approaches for calculating the annual rate of inflation in order to apply the Fisher equation. First, most researchers construct the inflation rate by obtaining the period-to-period changes in the logarithm of price and then annualize the series; that is, for the quarterly annualized inflation is defined as

$$\pi_t = \ln(P_t / P_{t-1})^4 \qquad (2.1)$$

Alternatively, the annual inflation rate can be constructed as the following:

$$\pi_t = \ln(P_t / P_{t-4}) \qquad (2.2)$$

[4] The countries have been selected depending on the span of data availability through various exchange rate regimes and their outstanding role within the industrialized economies.

for quarterly data. This year-to-year inflation rate calculation tends to yield a slightly smoother inflation process, since it avoids discreteness of reported CPI data. CPI is reported quarter-to-quarter in terms of discrete numbers with small changes from one period to another. Using a period-to-period inflation rate will magnify the effect of price changes by the exponential of 4 for quarterly data. Thus, the obtained inflation rate will fluctuate dramatically.

Most of the literature that mentions how the rate of inflation is obtained uses the period-to period approach to calculate the inflation rates; for example, Chen (2001) calculates the quarter-to-quarter annualized inflation. However, Gagnon and Unferth (1995) and Fountas and Wu (1999) use the year-to-year approach to calculate quarterly annualized rate of inflation.

Following the latter approach, we constructed the *ex post* real interest rate series by using the Fisher equation as follows:

$$R_t=I_t-(P_{t+4}-P_t)/P_t \qquad (2.3)$$

where R_t is the real interest rate at time t earned from holding the investment for four quarters. I_t is the nominal interest rate and P_t is the price index, thus $(P_{t+4}-P_t)/P_t$ is the inflation rate from time t to time t+4. In constructing the *ex ante* rate we created an expected inflation series using a four-period moving average of actual inflation rates.

Real rates could differ because of rational risk premia. It is therefore important to choose assets that are similar in terms of risk characteristics. This study focuses on government bonds[5] which are essentially free from default risk. Of course, even after controlling for default risk, theory does not necessarily predict that real interest rates should be equal across countries[6]. Differing risk premia could occur across countries, either because of heterogeneity in consumers-investors, or because of restrictions on capital movements.

However, the more positive results that accompany the use of yields on long-term debt instruments are not without cost. These instruments are more heterogeneous than the offshore deposit rates that have typically been used in the analyses of capital mobility and that we use in our study, for the sake of completeness. Moreover, it is not appropriate to characterize long-term bonds as zero discount bonds, so the

[5] This, however, does not mean that short-term deposit rates are underestimated or ignored.

[6] Benninga and Protopapadakis (1983), for instance, show that, under uncertainty, real interest rates include a risk premium, stemming from the covariance between the real value of future consumption and the real value of nominal assets.

reported interest rate data provide only approximate measures of the true returns that investors obtain. Yet, in many ways, these long-term instruments are more appropriate for testing capital mobility. First, firms do not usually make their investment decisions on the basis of short-term yields; in fact, depending upon the market structure of the economy, firms may rely on bank debt or equity. However, to the extent that firms borrow in bond markets, long-term bond yields will be the most informative series. Second, also from the investor's point of view, the long-term real rates are most relevant since they more closely measure rates of return expressed in terms of physical goods. Much of the previous literature has focused on the equality of short-term real interest rates and ignored any long-run dynamics. Since one of the assumptions that RIP rests on, PPP, is convincingly rejected in the short-run it seems more appropriate to test RIP in the long-run irrespective of its short-run validity (Kugler and Neusser, 1993). Finally, if our aim is to assess the equalization of returns in differing political jurisdictions, then on-shore, rather than off-shore, rates are once again more appropriate.

Before testing the RIP hypothesis, we need to examine the time series properties of the underlying real interest rate series that will be used later. A visual plot of the data is usually the first step in the analysis of any time-series because if a trend is observed it might indicate that the data is nonstationary. The graphs of the real interest rate differentials (RIRDs) in level form for Australia, Belgium, Canada, France, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Spain, Switzerland and the UK relative to the US are plotted in figures 1-4. These graphs indicate that RIRDs were relatively volatile during the sample 1967-2008. There are significant negative real interest rate differentials in our sample for some countries. Overall, the time series plots show similar movements of real interest rates across time. The ups and downs of *ex post* and *ex ante* real interest rate adjustments seem to occur at the same time intervals. Moreover, the substantial fluctuations of real interest rate differentials across time imply that there might be non-stationary series.

Figure 1: Short-Term *Ex Post* Real Interest Rate Differential

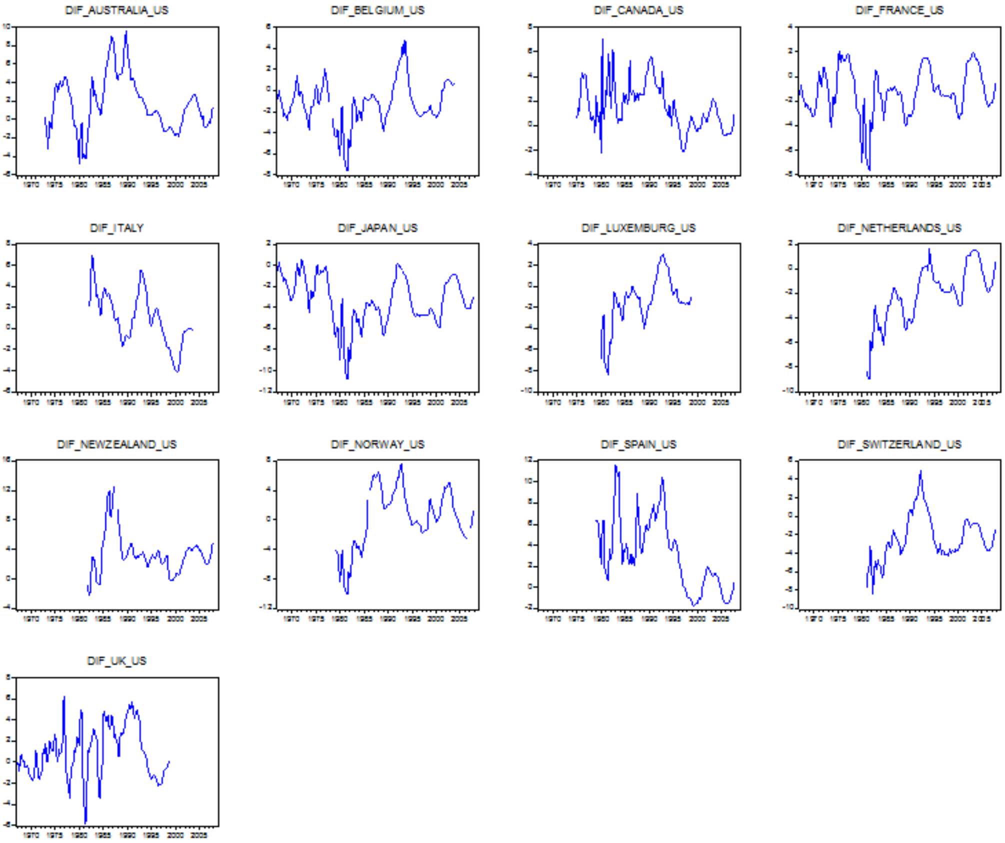


Figure 2: Short-Term *Ex Ante* Real Interest Rate Differential

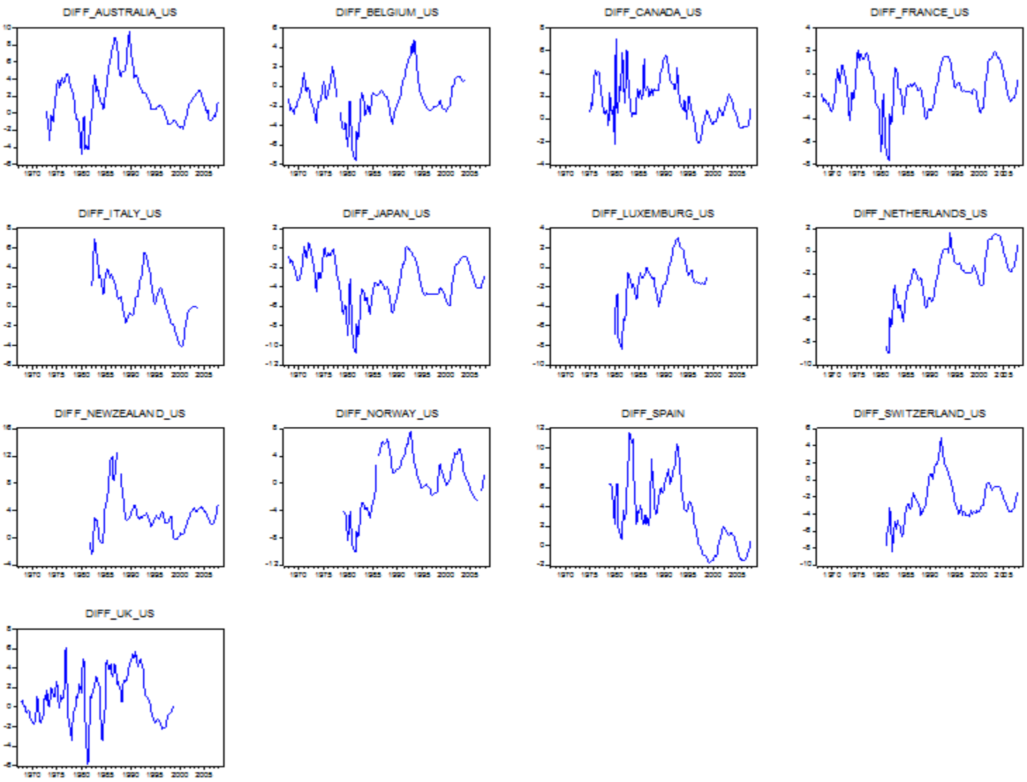


Figure 3: Long-Term *Ex Post* Real Interest Rate Differential

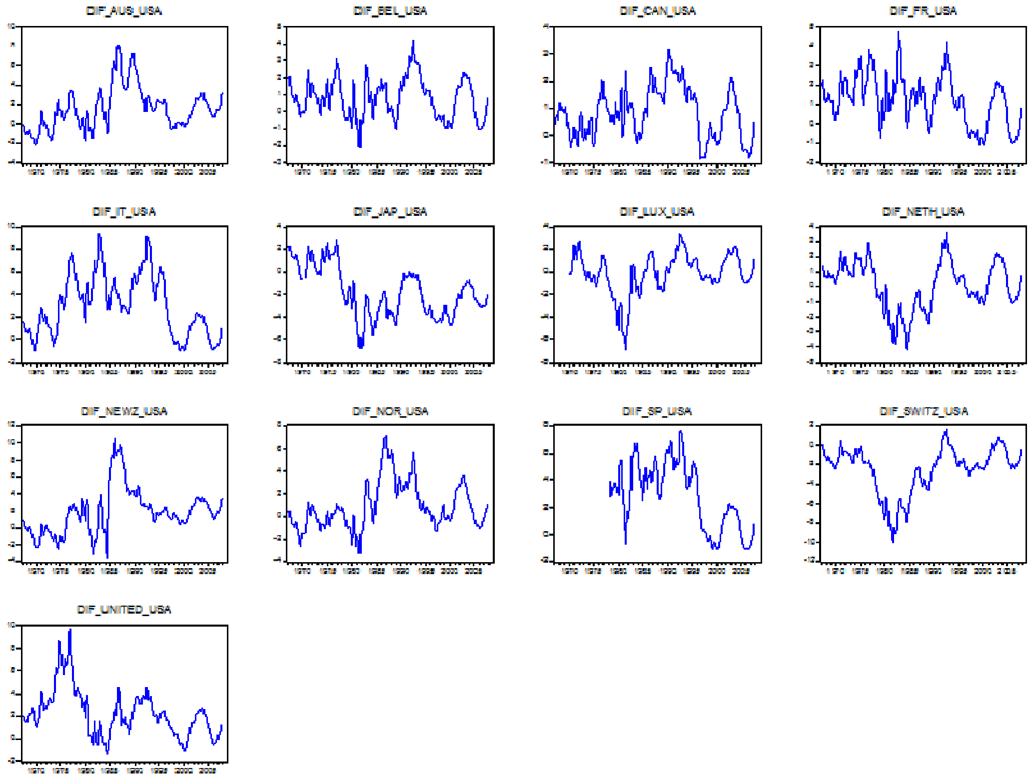
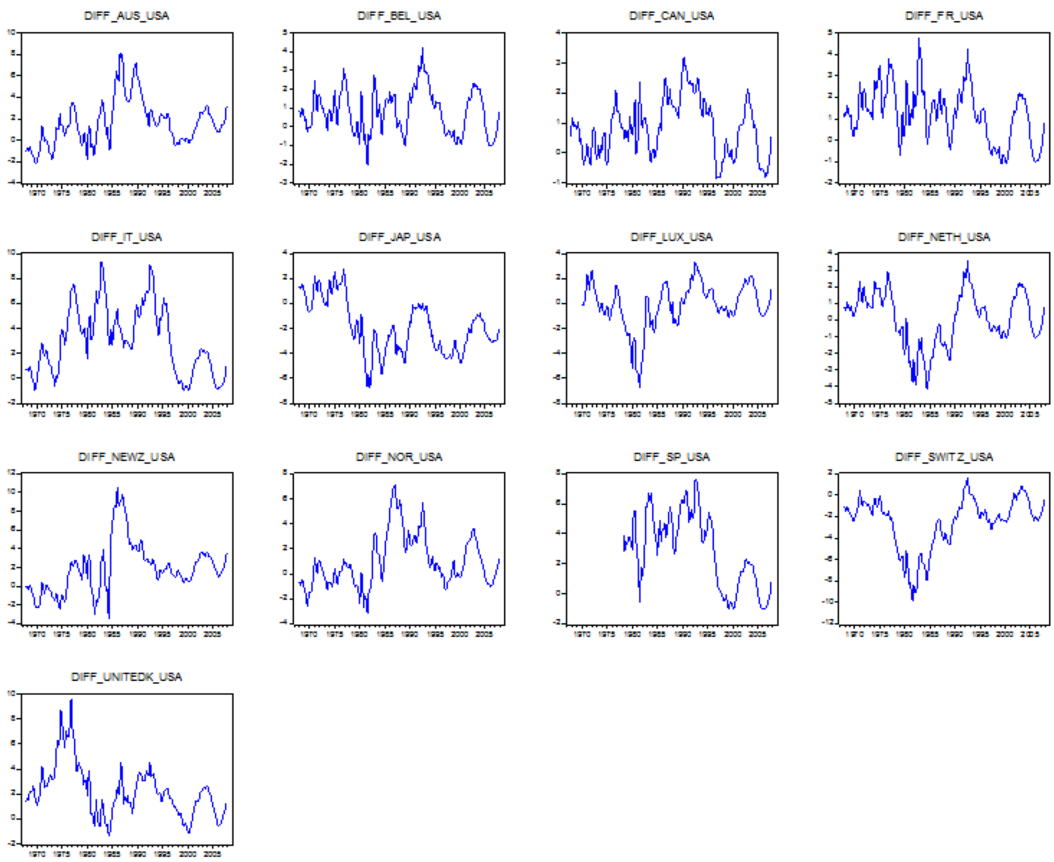


Figure 4: Long-Term Ex Ante Real Interest Rate Differential



2.1.1 Nominal Interest Rates and Inflation Rates

Tables 2.1, 2.2, 2.3 and 2.4 summarize the descriptive statistics of quarterly nominal interest rates (short-term and long-term interest rates) and inflation rates, constructed according to the above definitions for quarterly CPI data. For deposit rates, all series (except for those of Belgium, Japan and Luxembourg) are slightly positively skewed. Nearly all series (except for those of Canada, the Netherlands, New Zealand and the US) have small platykurtosis. Using the Jarque-Bera normality test, we reject normality in all the series (except for those of Italy and Luxembourg). For government bonds, most of the series seem to be slightly skewed with a long right tail. The government bonds of most of the countries (except for those of Canada, France and the US) suffer from platykurtosis. Finally, for the inflation rate, both models of calculation show similar mean and median values. All of the series of inflation rates are positively skewed but not normally distributed. The results from kurtosis are mixed: half of the series suffer from platykurtosis (k<3), and half of them suffer from leptokurtosis (k>3).

Table 2.1: Descriptive Statistics for Deposit Rates

Country	Series								
	Mean	**Median**	**Max.**	**Min.**	**St.Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia (1972Q4-2008Q4)	7.655	7.290	17.230	2.750	3.683	0.474	2.134	9.975	0.007
Belgium (1967Q1-2003Q4)	4.940	5.000	8.420	1.470	1.699	-0.048	1.946	6.816	0.033
Canada (1975Q1-2008Q4)	7.365	6.940	20.910	1.990	3.954	0.831	3.371	16.43	0.000
France (1967Q1-2008Q4)	4.632	4.500	8.500	2.000	1.749	0.592	2.346	12.81	0.002
Italy (1982Q1-2003Q4)	6.545	6.715	15.750	0.810	3.808	0.432	2.709	3.051	0.218
Japan (1967Q1-2008Q4)	2.489	2.705	6.000	0.027	1.783	-0.067	1.631	13.25	0.001
Luxembourg (1980Q1-1999Q1)	5.497	5.670	7.500	3.250	1.263	-0.266	2.071	3.677	0.159
Netherlands	3.579	3.395	6.250	2.280	0.919	1.152	4.340	33.14	0.000

(1981Q1-2008Q4)									
NewZealand (1981Q4-2008Q4)	8.604	7.800	18.750	4.220	3.359	1.146	3.846	26.61	0.000
Norway (1979Q1-2008Q3)	6.183	5.300	12.250	1.148	2.811	0.512	2.498	6.285	0.043
Spain (1979Q1-2008Q4)	8.914	8.985	19.800	2.060	5.204	0.232	1.764	8.711	0.013
Switzerland (1981Q1-2008Q4)	3.188	2.925	9.750	0.100	2.542	0.857	2.815	13.88	0.001
UK (1967Q1-1998Q4)	7.584	7.085	15.00	2.500	3.377	0.315	1.916	8.382	0.015
US (1967Q1-2008Q4)	5.807	5.370	15.090	0.390	2.810	0.858	4.305	32.53	0.000

Table 2.2: Descriptive Statistics for Government Bonds

Country	Series								
	Mean	**Median**	**Max.**	**Min.**	**St. Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia (1967:Q1-2008Q4)	8.438	7.470	16.430	3.810	3.414	0.657	2.114	17.59	0.000
Belgium (1967:Q1-2008Q4)	7.651	7.620	13.810	3.250	2.590	0.446	2.673	6.311	0.043
Canada (1967:Q1-2008Q4)	7.629	7.210	18.200	2.240	2.957	0.713	3.464	15.75	0.000
France (1967:Q1-2008Q4)	8.077	7.945	16.860	3.230	3.112	0.629	3.080	11.13	0.004
Italy (1982:Q1-2003Q4)	9.799	9.875	21.210	3.390	4.484	0.460	2.404	8.418	0.015
Japan (1967:Q1-2008Q4)	5.031	5.749	9.533	0.657	2.794	-0.122	1.537	15.39	0.000
Luxembourg (1970:Q1-2008Q4)	6.814	7.020	10.770	3.140	1.813	-0.023	2.431	2.106	0.349
Netherlands (1967:Q1-2008Q4)	6.934	6.940	12.000	3.220	1.973	0.118	2.413	2.807	0.246
NewZealand (1967:Q1-2008Q4)	8.749	7.090	18.650	5.180	3.573	1.005	2.845	28.46	0.000
Norway (1967:Q1-2008Q4)	7.773	6.700	13.690	3.130	3.079	0.574	2.046	15.59	0.000
Spain	9.790	11.01	17.810	3.180	4.595	0.003	1.531	11.06	0.004

(1978:Q2-2008Q4)									
Switzerland (1967:Q1-2008Q4)	4.388	4.370	7.330	1.950	1.231	0.184	2.477	2.861	0.239
UK (1967:Q1-2008Q4)	9.064	9.195	16.540	4.000	3.268	0.192	2.050	7.343	0.025
US (1967:Q1-2008Q4)	6.814	6.420	15.790	1.480	2.773	0.782	3.786	21.43	0.000

Table 2.3: Descriptive Statistics for Inflation Rates (for *ex post* rates)

Country	**Series**								
	Mean	**Median**	**Max.**	**Min.**	**St.Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia (1967:Q1-2008Q4)	0.059	0.049	0.177	-0.003	0.041	0.735	2.764	15.14	0.001
Belgium (1967:Q1-2008Q4)	0.041	0.031	0.161	0.005	0.030	1.499	5.392	99.95	0.000
Canada (1967:Q1-2008Q4)	0.046	0.039	0.126	0.000	0.033	0.890	2.704	22.25	0.000
France (1967:Q1-2008Q4)	0.051	0.034	0.149	0.002	0.040	0.879	2.535	22.59	0.000
Italy (1967:Q1-2008Q4)	0.075	0.052	0.256	0.000	0.061	1.091	3.001	32.56	0.000
Japan (1967:Q1-2008Q4)	0.034	0.021	0.236	-0.014	0.047	2.262	9.373	417.4	0.000
Luxembourg (1967:Q1-2008Q4)	0.039	0.032	0.113	-0.012	0.028	0.947	3.179	24.74	0.000
Netherlands (1967:Q1-2008Q4)	0.038	0.028	0.109	-0.012	0.027	0.828	2.789	19.07	0.000
NewZealand (1967:Q1-2008Q4)	0.070	0.047	0.190	-0.005	0.056	0.655	2.008	18.44	0.000
Norway (1967:Q1-2008Q4)	0.053	0.043	0.146	-0.014	0.036	0.605	2.378	12.66	0.002
Spain (1967:Q1-2008Q4)	0.079	0.058	0.273	0.015	0.058	1.104	3.473	34.83	0.000
Switzerland (1967:Q1-2008Q4)	0.029	0.020	0.108	-0.001	0.025	1.091	3.521	34.37	0.000
UK (1967:Q1-2008Q4)	0.067	0.048	0.268	0.011	0.054	1.626	5.337	109.6	0.000
US (1967:Q1-2008Q4)	0.047	0.038	0.145	0.013	0.029	1.477	4.659	78.42	0.000

Table 2.4: Descriptive Statistics for Inflation Rates (MA(4) approach)

Country	Series								
	Mean	**Median**	**Max.**	**Min.**	**St.Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia **(1967:Q4-2007Q4)**	0.060	0.048	0.168	-0.001	0.040	0.688	2.602	13.763	0.001
Belgium **(1967:Q4-2007Q4)**	0.041	0.031	0.149	0.008	0.030	1.401	4.736	72.780	0.000
Canada **(1967:Q4-2007Q4)**	0.047	0.040	0.124	0.002	0.032	0.886	2.613	22.058	0.000
France **(1967:Q4-2007Q4)**	0.051	0.033	0.142	0.003	0.040	0.816	2.335	20.845	0.000
Italy **(1967:Q4-2007Q4)**	0.075	0.051	0.221	0.011	0.060	0.977	2.601	26.709	0.000
Japan **(1967:Q4-2007Q4)**	0.034	0.019	0.232	-0.010	0.046	2.072	8.011	283.678	0.000
Luxembourg **(1967:Q4-2007Q4)**	0.040	0.032	0.109	-0.006	0.028	0.943	3.064	23.872	0.000
Netherlands **(1967:Q4-2007Q4)**	0.041	0.037	0.136	0.015	0.028	1.462	4.448	71.393	0.000
NewZealand **(1967:Q4-2007Q4)**	0.071	0.048	0.181	-0.001	0.055	0.553	1.813	17.681	0.000
Norway **(1967:Q4-2007Q4)**	0.053	0.045	0.140	0.005	0.035	0.553	2.205	12.456	0.002
Spain **(1967:Q4-2007Q4)**	0.080	0.059	0.249	0.018	0.057	1.038	3.117	29.001	0.000
Switzerland **(1967:Q4-2007Q4)**	0.003	0.021	0.104	0.000	0.024	1.068	3.355	31.450	0.000
UK **(1967:Q4-2007Q4)**	0.068	0.049	0.249	0.013	0.052	1.483	4.638	76.971	0.000
US **(1967:Q4-2007Q4)**	0.047	0.037	0.136	0.015	0.028	1.462	4.448	71.393	0.000

2.1.2 Ex Post Real Interest Rates

As mentioned above, in order to construct each of the real interest rate series, we use the nominal interest rate data and compute the inflation rates for the entire sample period of 1967Q1 to 2008Q4 for quarterly frequency.

The plots of the *ex post* real interest rates (short-term and long-term) are exhibited in Figures 5 and 6. The dynamic patterns of these constructed real interest rate series share some similarities. These estimated *ex post* real rates seem to fluctuate persistently during the sample period. As it is easily seen, real interest rates have varied widely over recent decades. Calculated in the conventional manner – that is, deflating the nominal interest rate by the rate of inflation- we see that these time series appear to be very variable throughout the period studied. This observation applies broadly whether we look at short-term interest rates or long-term interest rates. Overall, the time series plots show similar movements of real interest rates (both short-term and long-term) across time.

Figure 5: Short-Term *Ex Post* Real Interest Rates

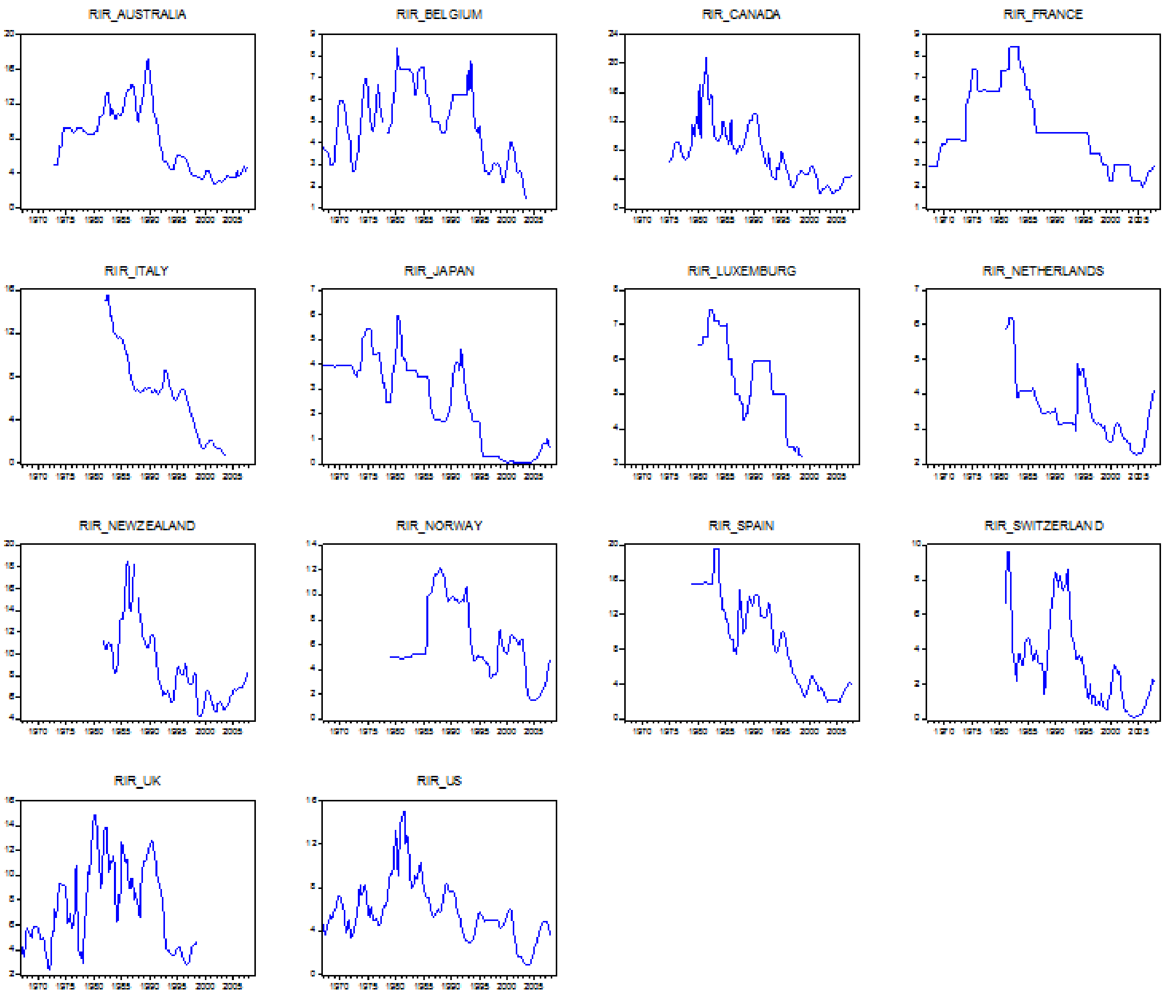


Figure 6: Long-Term Ex Post Real Interest Rates

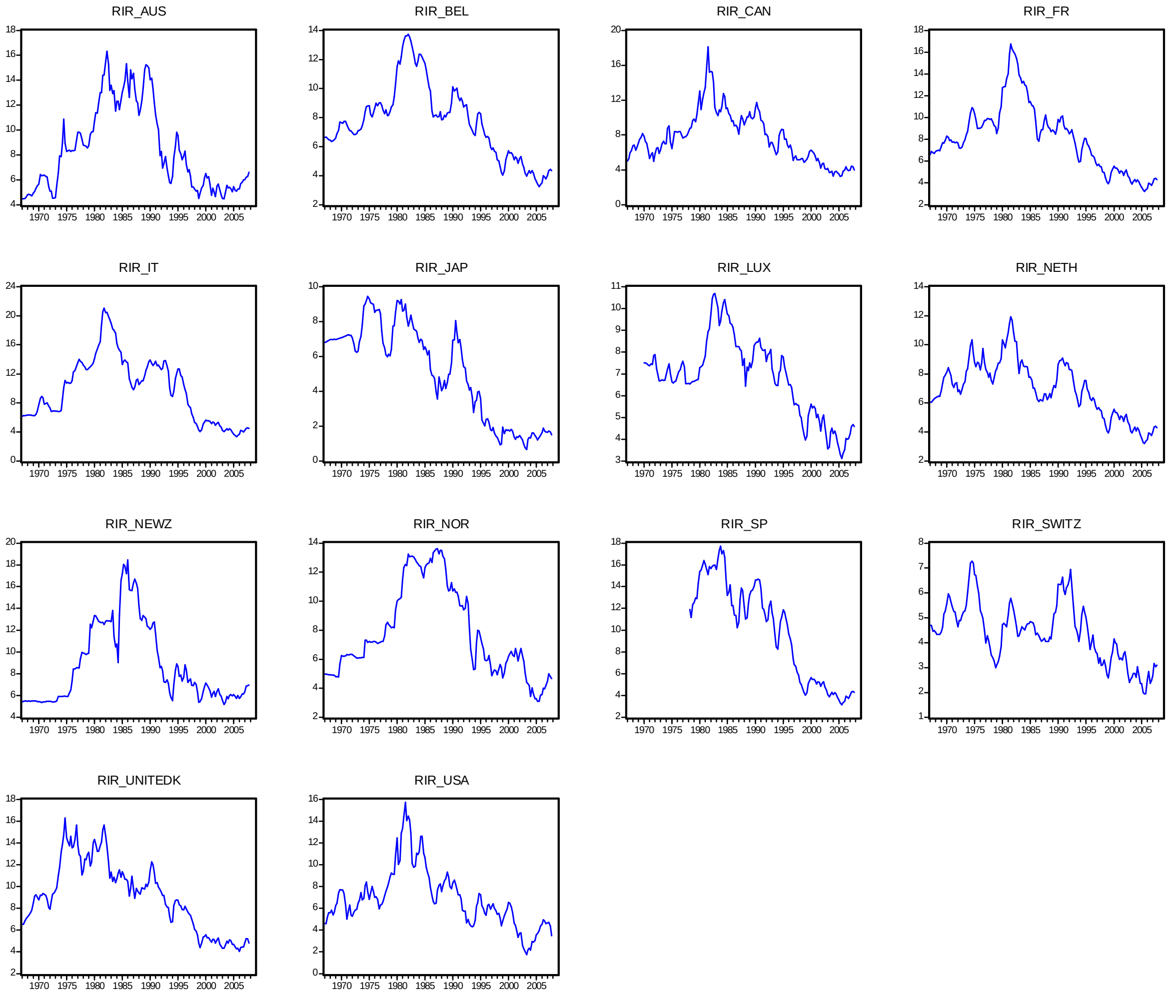


Table 2.5 lists the descriptive statistics of the short-term *ex post* real interest rates.

Table 2.5: Descriptive Statistics for the Short-Term *Ex Post* Real Interest Rates

Country	**Series**								
	Mean	**Median**	**Max.**	**Min.**	**St.Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia **(1972:Q4-2007Q4)**	7.665	7.126	17.16	2.721	3.691	0.441	2.108	9.238	0.009
Belgium **(1967:Q1-2003Q4)**	4.897	4.972	8.348	1.447	1.687	-0.049	1.943	6.861	0.032
Canada **(1975:Q1-2007Q4)**	7.455	7.008	20.803	1.945	3.914	0.812	3.368	15.26	0.000
France **(1967:Q1-2007Q4)**	4.605	4.466	8.422	1.987	1.736	0.557	2.309	11.73	0.003

Italy (1982:Q1-2003Q4)	6.494	6.662	15.610	0.788	3.781	0.423	2.693	2.964	0.227
Japan (1967:Q1-2007Q4)	2.502	2.964	5.958	0.029	1.751	-0.120	1.667	12.54	0.002
Luxembourg (1980:Q1-1999Q1)	5.464	5.628	7.419	3.221	1.246	-0.285	2.077	3.741	0.152
Netherlands (1981:Q1-2007Q4)	3.528	3.289	6.218	2.263	0.919	1.233	4.502	37.53	0.000
NewZealand (1981:Q4-2007Q4)	8.599	7.569	18.567	4.206	3.381	1.101	3.689	22.85	0.000
Norway (1979:Q1-2007Q4)	6.153	5.240	12.20	1.462	2.843	0.499	2.416	6.299	0.043
Spain (1979:Q1-2007Q4)	9.033	9.279	19.70	2.021	5.167	0.169	1.762	7.951	0.019
Switzerland (1981:Q1-2007Q4)	3.223	3.004	9.695	0.086	2.553	0.813	2.715	12.26	0.002
UK (1967:Q4-1998Q4)	7.506	6.928	14.88	2.419	3.363	0.326	1.923	8.452	0.015
US (1967:Q1-2007Q4)	5.867	5.365	15.03	0.886	2.744	0.922	4.456	37.70	0.000

Table 2.5 indicates that the mean and the median of the series above vary from 2.502 to 9.033 maximum for the countries examined. Furthermore, none of the series is distributed normally, except for those of Italy and Luxembourg with skewness close to zero. Most of the series suffer from platykurtosis, with the exception of the series of Canada, the Netherlands, New Zealand and the US which suffer from leptokurtosis (as kurtosis>3).

Table 2.6 below lists the descriptive statistics of the long-term *ex post* real interest rates.

Table 2.6: Descriptive Statistics for the Long-Term *Ex Post* Real Interest Rates

Country	Series								
	Mean	Median	Max.	Min.	St.Dev.	Skewness	Kurtosis	Jarque-Berra	Prob.
Australia (1967:Q1-2007Q4)	8.447	7.812	16.32	4.467	3.405	0.634	2.072	16.87	0.000
Belgium (1967:Q1-2007Q4)	7.689	7.656	13.72	3.235	2.558	0.425	2.696	5.563	0.062
Canada (1967:Q1-2007Q4)	7.697	7.271	18.09	3.274	2.882	0.765	3.537	17.95	0.000
France (1967:Q1-2007Q4)	8.119	7.956	16.75	3.213	3.064	0.619	3.118	10.58	0.005
Italy (1967:Q1-2007Q4)	9.850	10.10	21.04	3.368	4.431	0.433	2.408	7.520	0.023
Japan (1967:Q1-2007Q4)	5.085	6.005	9.448	0.660	2.741	-0.173	1.579	14.61	0.001
Luxembourg (1970:Q1-2007Q4)	6.819	7.007	10.69	3.112	1.794	-0.052	2.488	1.731	0.421
Netherlands (1967:Q1-2007Q4)	6.962	6.950	11.94	3.207	1.936	0.092	2.486	2.037	0.361
NewZealand (1967:Q1-2007Q4)	8.742	7.175	18.47	5.157	3.569	0.969	2.774	26.06	0.000
Norway (1967:Q1-2007Q4)	7.803	0.737	13.62	3.106	3.054	0.555	2.019	14.98	0.001
Spain (1978:Q2-2007Q4)	9.910	11.01	17.71	3.144	4.529	-0.064	1.563	10.32	0.006
Switzerland (1967:Q1-2007Q4)	4.398	4.427	7.266	1.938	1.204	0.166	2.535	2.228	0.328
UK (1967:Q1-2007Q4)	9.106	9.182	16.28	4.024	3.192	0.158	2.069	6.596	0.037
US (1967:Q1-2007Q4)	6.879	6.454	15.73	1.742	2.699	0.867	3.915	26.26	0.000

Table 2.6 indicates that all of the series (except for those of Japan, Luxembourg and Spain) are slightly positively skewed. Most of the series suffer from platykurtosis, except for those of Canada, France and the US. Using the Jarque-Bera normality test, we reject normality in all the series-except for those of Belgium, Luxembourg, the Netherlands and Switzerland- where the probability values are lower than 5% (or alternatively, JB>5.99).

2.1.3 Ex Ante Real Interest Rates

Studying real interest rates seems to be problematic in the sense that an *ex ante* real interest rate is unobservable. Thus, many studies have to develop a method to estimate the *ex ante* real interest rate and then impose some structure and maintained assumptions into models. The simplest assumption is to assume perfect rational expectations, and thus the *ex post* rate is the best prediction of the *ex ante* real rate with a zero mean error term, which is exactly what we tried to do in the previous section. However, many authors have tried to mimic how agents form their expectations about inflation rates using a wide range of models from simple AR models to elaborate general equilibrium models. Consequently, there is very little agreement among researchers on how to construct an *ex ante* real interest rate, and that lack of agreement might leads to very different results in the time series properties of the constructed real rates.

As mentioned earlier in our analysis, there are two main strands of empirical studies that focus extensively on the use of the real rate of interest: the Fisher hypothesis and the Real Interest Parity (RIP) hypothesis. The Fisher relation indicates that the nominal interest rate adjusts fully to changes in the expected rate of inflation such that there is a one-to-one relationship between them and that the expected real rate of returns remains constant with respect to changes in expected inflation. For the RIP hypothesis, given that Fisher relations, the Uncovered Interest Parity condition, and the *ex ante* version of Purchasing Power Parity are satisfied for each country, the *ex ante* real interest rates are equalized across countries. Researchers employ a wide variety of approaches of measuring the expected real returns on assets when attempting to test these hypotheses. These methodologies differ in how to treat the expected inflation in the real interest rate calculation as well as what proxies to use for the nominal interest and the price variables.

This section provides a brief presentation of the approaches researchers have taken in estimating the *ex ante* real rate of interest and expected inflation. Constructing real interest rates is a difficult task. Conceptually, one must be careful in defining how agents develop their methods of inflation forecasting. As no single method can be found to have a clear superior forecasting accuracy, we present seven methodologies of constructing the *ex ante* real interest rates used in the prior literature. Each country's *ex ante* real interest rates are constructed using: (i) the *ex* *post* real interest rate, (ii) MA(4) inflation forecast, (iii) Mishkin's[7] linear projection, (iv) rolling regression, (v) recursive least squares, (vi) regime-switching technique, and (vii) the survey of inflation forecasts. The methods (ii)-(v) can be viewed as the linear regression approaches since the estimations are based on the linear regression model under different specifications and the included variables. The regime-switching method estimates nonlinearly the pattern of the real interest rate after including possible regime shifts constructed by Markov-chain probability. Once the estimated real interest rate series are obtained, the series distributions are compared by using a normality test. Moreover, we employ three unit root tests to investigate whether the real interest rates from different approaches yield different results in stationarity. The above selected methods vary in terms of the availability of agents' information set. The *ex post* real rate assumes that agents have rational expectations such that they make random forecast errors about the future rate of inflation. Thus, the actual inflation can be used as unbiased proxy of the expected rate of inflation. Kugler an Neusser (1993) and Goodwin and Grennes (1994) are examples of papers that use the *ex post* rates for the empirical methodology. We now present a brief summary of the two of the seven approaches that we use in our study:

i. Pure Rational Expectations

By assuming rationality of inflationary expectations, researchers can use the *ex post* real rate to study the behaviour of the *ex ante* real rate. The *ex post* real interest rate, given by equation (2.3) above, will differ from the *ex ante* real interest rate by a

---

[7] The autoregressive (AR) representation expresses the value of the series as a linear relationship to its past observations. An example of RIP tests using the AR specification is Baharumshah et al. (2005), who use an AR(1) specification to estimate the expected inflation. In contrast, Mishkin (1984), Cumby and Mishkin (1986) and Huizinga and Mishkin (1984) expand the autoregressive approach by adding macroeconomic variables to an AR model of the expected inflation.

This approach, hereafter referred to as the 'Mishkin approach,' implies that the *ex ante* real rate can be obtained by linearly projecting it into a set of observable variable X_t from the available information set at time t. With the linear projection function $p(E_t(r_{t+1})|X_t)$ of $E_t(r_{t+1})$ into X_t, one can estimate the real interest rate as follows:

$$E_t(r_{t+1}) = X_t\beta + u_t$$

where $u_t=E_t(r_{t+1})-P(E_t(r_{t+1})|X_t)$ is the projection error and orthogonal to Xt. Mishkin's choice of X_t includes four lags of the inflation rate, one lag of money growth (M1), the nominal Eurodollar interest rate and a fourth-order time polynomial.

random error and this error of inflation forecast, under rational expectations, is well-behaved with zero mean and orthogonal to the available information set. Since the *ex ante* real rates are less variable than the observed *ex post* rates, by definition, if the *ex post* real interest rate is stationary, this may be interpreted as indicating that the *ex ante* real interest rate is also stationary over time. The studies of Kugler and Neusser (1993), Gagnon and Unferth (1995) and Goodwin and Grennes (1994) are the example of papers that use the *ex post* rates to conduct analysis.

ii. Time Series Forecasting Models

To quantify the unobserved component of the real interest rate, time series models can be useful in approximating the expected rate of future inflation using only the past behaviour of the realized inflation rate, which is readily available. The types of time series models that have been used in prior researches of the expected inflation are: ARMA model, Mishkin's linear projection technique, the recursive least squares method, the rolling regression, and the Markov-switching model. In this study, we will follow the method that was followed by Fountas and Wu (1999). Autoregressive representations are appealing to researchers because, for forecasting purposes, they link the present observable data to the past history of the data so that we can extrapolate to form a forecast of future observable data based on present and past observations.

The plots of the *ex ante* real interest rates (short-term and long-term) are exhibited in Figures 7 and 8.

Figure 7: Short-Term *Ex Ante* Real Interest Rates

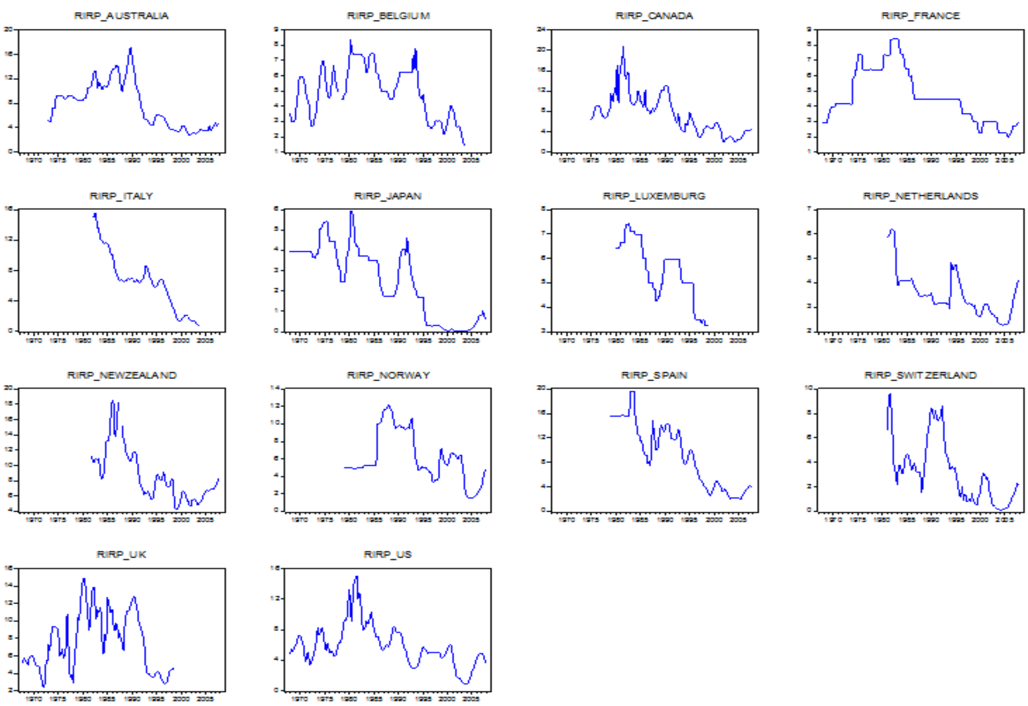
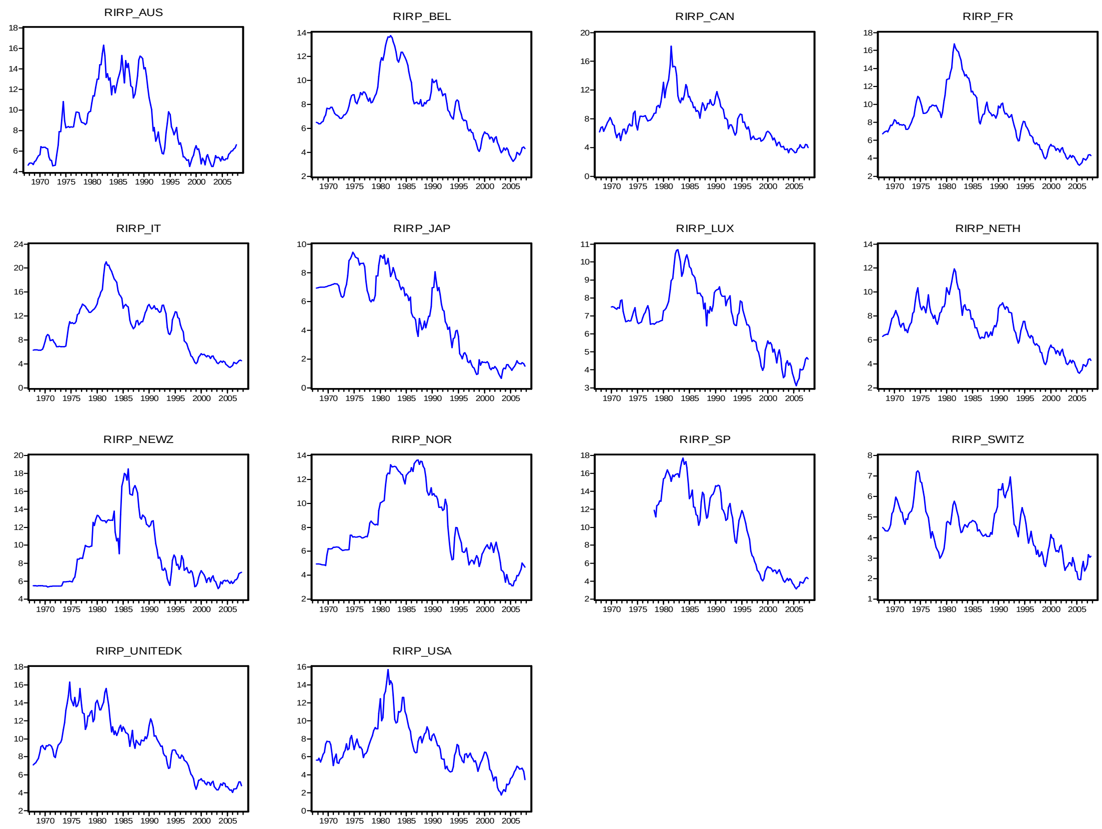


Figure 8: Long-Term *Ex Ante* Real Interest Rates



The dynamic patterns of these constructed real interest rates share some similarities. These estimated *ex ante* real rates seem to fluctuate persistently during the sample period. In other words, we see that these time series appear to be very variable throughout the period studied and this finding leaves us with the suspicion that these series might be nonstationary. This observation applies broadly whether we look at short-term interest rates or long-term interest rates. Overall, the time series plots show similar movements of real interest rates (both short-term and long-term) across time.

Tables 2.7-2.8 list the descriptive statistics of the real *ex ante* short-term and long-term interest rates.

Table 2.7: Descriptive Statistics for the Real *Ex Ante* Short-Term Interest Rate

Country	Series								
	Mean	**Median**	**Max.**	**Min.**	**St. Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia (1972:Q4-2007Q4)	7.664	7.139	17.157	2.719	3.688	0.440	2.108	9.237	0.009
Belgium (1967:Q4-2003Q4)	4.923	4.985	8.349	1.451	1.693	-0.089	1.947	6.795	0.033
Canada (1975:Q1-2007Q4)	7.454	7.012	20.795	1.959	3.912	0.812	3.366	15.225	0.000
France (1967:Q4-2007Q4)	4.635	4.467	8.409	1.983	1.735	0.526	2.289	10.824	0.004
Italy (1982:Q1-2003Q4)	6.492	6.663	15.593	0.787	3.778	0.421	2.693	2.948	0.229
Japan (1967:Q4-2007Q4)	2.474	2.719	5.941	0.032	1.753	-0.093	1.656	12.341	0.002
Luxembourg (1980:Q1-1999Q1)	5.463	5.633	7.413	3.232	1.243	-0.288	2.077	3.804	0.149
Netherlands (1981:Q1-2007Q4)	3.527	3.291	6.201	2.265	0.917	1.227	4.484	36.99	0.000
NewZealand (1981:Q4-2007Q4)	8.598	7.579	18.605	4.217	3.377	1.106	3.708	23.132	0.000
Norway (1979:Q1-2007Q4)	6.152	5.239	12.189	1.465	2.841	0.498	2.416	6.284	0.043
Spain (1979:Q1-2007Q4)	9.032	9.270	19.687	2.023	5.166	0.169	1.762	7.965	0.019
Switzerland (1981:Q1-2007Q4)	3.222	2.995	9.961	0.088	2.552	0.813	2.717	12.271	0.002
UK (1967:Q4-1998Q4)	7.595	7.319	14.86	2.428	3.349	0.286	1.912	7.867	0.019
US (1967:Q4-2007Q4)	5.899	5.399	15.016	0.890	2.756	0.891	4.387	34.19	0.000

All of the series (except for those of Belgium, Japan and Luxembourg) are slightly positively skewed. Furthermore, most of the series suffer from platykurtosis, except for those of Canada, the Netherlands, New Zealand and the US. Using the Jarque-Bera normality test, we reject normality in all the series-except for those of Italy and Luxembourg- where the probability values are lower than 5% (or alternatively, JB>5.99).

Table 2.8: Descriptive Statistics for the Real Ex Ante Long-Term Interest Rates

Country	Series								
	Mean	**Median**	**Max.**	**Min.**	**St. Dev.**	**Skewness**	**Kurtosis**	**Jarque-Berra**	**Prob.**
Australia (1967:Q4-2007Q4)	8.521	7.898	16.315	4.496	3.390	0.618	2.051	16.296	0.000
Belgium (1967:Q4-2007Q4)	7.709	7.735	13.720	3.229	2.576	0.401	2.653	5.117	0.077
Canada (1967:Q4-2007Q4)	7.739	7.363	18.085	3.275	2.889	0.735	3.499	16.187	0.000
France (1967:Q4-2007Q4)	8.145	8.033	16.728	3.212	3.084	0.591	3.059	9.409	0.009
Italy (1967:Q4-2007Q4)	9.918	10.163	21.044	3.368	4.440	0.401	2.391	6.793	0.033
Japan (1967:Q4-2007Q4)	5.052	5.809	9.415	0.659	2.754	-0.146	1.559	14.492	0.000
Luxembourg (1970:Q1-2007Q4)	6.819	7.003	10.683	3.111	1.792	-0.055	2.486	1.751	0.417
Netherlands (1967:Q4-2007Q4)	6.978	6.999	11.934	3.206	1.949	0.065	2.452	2.126	0.345
NewZealand (1967:Q4-2007Q4)	8.804	7.217	18.505	5.163	3.569	0.947	2.733	24.563	0.000
Norway (1967:Q4-2007Q4)	7.856	7.094	13.616	3.108	3.055	0.526	1.995	14.203	0.001
Spain (1978:Q2-2007Q4)	9.908	11.021	17.697	3.143	4.527	-0.064	1.563	10.324	0.006
Switzerland (1967:Q4-2007Q4)	4.394	4.383	7.251	1.938	1.213	0.168	2.486	2.532	0.282
UK (1967:Q4-2007Q4)	9.152	9.227	16.296	4.033	3.201	0.121	2.060	6.321	0.042
US (1967:Q4-2007Q4)	6.918	6.479	15.716	1.748	2.707	0.836	3.868	23.816	0.000

Table 2.8 indicates that all series (except for those of Canada, France and the US) have small platykurtosis. Using the Jarque-Bera normality test, we reject normality in all the series (except for those of Belgium, Luxembourg, the Netherlands and Switzerland). Most of the series seem to be slightly skewed with a long right tail.

2.1.4 Ex Post VS Ex Ante Real Interest Rates

As noted previously, real interest rates are non-observable and are usually proxied by the so-called *ex-post* real interest rates. As is well known, however, *ex-post* rates include two disturbing components which can render them a misleading proxy for non-observable *ex-ante* real interest rates: inflation risk premia and agents' inflation expectation errors. Like *ex-ante* real interest rates, these two variables are also non-observable. There are reasons to think that both disturbances are probably negligible. First, inflation premia can hardly be relevant if the inflation rate is not very volatile. And second, if agents are rational when forming their inflation expectations, the expectation error should be zero on average. Although it is still an open question, this view has been recently challenged in the literature, particularly in relation to the magnitude of the inflation expectation error. Thus, a series of papers have found that, due to informational or to (monetary policy) credibility problems, inflation rates can be successfully characterised by switching-regime models à la Hamilton, not only in high-inflation countries like Argentina, Israel or Mexico (see Kaminsky and Leiderman, 1996) but also in countries whose inflation rates are lower and more stable like the US (Evans and Lewis, 1995). These switching-regime models produce inflation expectation errors which have zero-mean *ex-ante* but, *ex-post* can show a non-zero mean.

In this section the focus is on determining whether the two different methods of deriving the real interest rates yield different conclusions in testing RIP theory. In particular, we aim to investigate the question of whether real interest rates are stationary or not is sensitive to the underlying approach of deriving the rates. If there is ambiguity in identifying the stationarity of the series, this could lead to problems in selecting the methodology for conducting RIP hypothesis testing. For instance, if real rates are nonstationary, cointegration techniques would be more appropriate to test for a cointegrated relationship between two or more random walk series than a simple linear regression due to *spurious* regression problems as described in Granger and Newbold (1974).

Implicitly, the past literature has assumed that the method of constructing the real rate is irrelevant to the test. Therefore, real interest rates constructed differently should have similar time series properties, and the inconclusive results of RIP may come from other theoretical sources. However, if that is not the case then differences among RIP empirical analysis may stem from deviations of the methodologies used by authors. Future investigation of issues involving use of expected real interest rates may have to be more concerned about the selection of the measuring approach.

As we found earlier, the means and the medians of the real interest rates from the two different approaches appear to be similar (as shown in Tables 2.5-2.8). Same pattern of standard deviations is also observed. The real interest rates-both short term and long term- constructed by using the year-to-year annualized inflation rate with the two methods, as shown in Figures 5-8, appear to be remarkably similar in the pattern of movements for each country. Although the derived real interest rates from the two approaches seem to follow the same pattern over the sample period, they are in fact different time series processes. As the results for *ex ante* rates are similar to those for *ex post* rates, we do not report them. We only report those for *ex post* rates in the next section. Overall, considering the stationarity of the series, all real rates seem to follow a random walk and this conclusion is robust to different choices of approach in constructing the series and the types of unit root tests[8] used. These findings for stationarity are reported separately in the following section.

Tables 2.9 and 2.10 report the results of the mean and variance equality testing for quarterly data for all countries. We employ the analysis of variance (ANOVA) to examine whether different approaches of constructing interest rates (both short-term and long–term ones) would provide the series with equal mean. For the variance equality testing, a Brown-Forsythe test is used to evaluate the null hypothesis that the variance in all series is equal against the alternative that at least one series has a different variance.

Table 2.9: Tests for Equality of Means and Variances

Real Short-Term Interest Rates from Different Approaches

Country	Mean Equality Test: Test Statistics	Variance Equality Test: Test Statistics
Australia	1.19E-06 (0.9991)	0.000184 (0.9892)
Belgium	0.016940 (0.8965)	4.47E-06 (0.9983)
Canada	3.66E-06 (0.9985)	1.05E-05 (0.9974)

[8] It should be noted however that the results from KPSS unit root tests are mixed.

France	0.025121 (0.8742)	0.001823 (0.9660)
Italy	1.89E-05 (0.9965)	3.83E-05 (0.9951)
Japan	0.019679 (0.8885)	0.005596 (0.9404)
Luxembourg	1.57E-05 (0.9968)	0.000385 (0.9844)
Netherlands	2.32E-05 (0.9962)	1.67E-05 (0.9967)
New Zealand	9.87E-06 (0.9975)	0.000101 (0.9920)
Norway	1.66E-06 (0.9990)	4.70E-05 (0.9945)
Spain	5.22E-06 (0.9982)	4.69E-06 (0.9983)
Switzerland	3.50E-06 (0.9985)	2.44E-05 (0.9961)
UK	0.044278 (0.8335)	0.002438 (0.9607)
US	0.011066 (0.9163)	0.003092 (0.9557)

Notes:
The reported test statistics are the F-statistics follow F-distribution. The parentheses display the corresponding *p*-values.

From the comparison of the tables 2.5 and 2.7 we find that the real short-term interest rates that we derived from the two methods with the year-to-year inflation rate yield a similar mean and median of the series. This finding is confirmed by the large probability (round 0.9) of accepting the mean equality between the real interest rates of each country, as shown in the second column of Table 2.9. The variance equality test clearly suggests that the variances of each country's series are equal.

Table 2.10: Tests for Equality of Means and Variances

Real Long-Term Interest Rates from Different Approaches

Country	Mean Equality Test: Test Statistics	Variance Equality Test: Test Statistics
Australia	0.038515 (0.8445)	0.003739 (0.9513)
Belgium	0.004890 (0.9443)	0.007360 (0.9317)
Canada	0.017706 (0.8942)	0.000755 (0.9781)
France	0.005565 (0.9406)	0.008297 (0.9275)
Italy	0.018907 (0.8907)	0.000462 (0.9829)
Japan	0.011735 (0.9138)	0.026267 (0.8714)
Luxembourg	6.74E-07 (0.9993)	0.000128 (0.9910)
Netherlands	0.005654 (0.9401)	0.009709 (0.9216)
New Zealand	0.023981 (0.8770)	0.002875 (0.9573)
Norway	0.024241 (0.8764)	0.001649 (0.9676)
Spain	8.03E-06 (0.9977)	1.46E-05 (0.9970)
Switzerland	0.000906 (0.9760)	0.025860 (0.8723)
UK	0.016637 (0.8974)	2.14E-06 (0.9988)
US	0.016669 (0.8974)	0.000514 (0.9819)

Notes:
The reported test statistics are the F-statistics follow F-distribution. The parentheses display the corresponding *p*-values.

From the comparison of the tables 2.6 and 2.8 we find that the real long-term interest rates that we derived from the two methods with the year-to-year inflation rate yield a similar mean and median of the series. This finding is also confirmed by the large probability (round 0.9) of accepting the mean equality between the real interest rates of each country, as shown in the second column of Table 2.10. The variance equality test also suggests that the variances of each country's series are equal.

Since these two approaches in constructing the real interest rates share some similarities in regard to the descriptive statistics, mentioned in the previous sections, our next step is to test whether these rates are correlated.

Table 2.11: Correlations: Quarterly Real Short-Term Interest Rates

	Ex post	**MA(4)**
Ex post	1.000	0.999
MA(4)		1.000

Table 2.12: Correlations: Quarterly Real Long-Term Interest Rates

	Ex post	**MA(4)**
Ex post	1.000	0.999
MA(4)		1.000

Tables 2.11 and 2.12 display the correlations of the real interest rates from the two different approaches followed. It should be noted that the results presented in the above tables are coincidentally the same for all countries and this is the reason why we do not present the results for each country separately. The findings strongly indicate that, for both short-term and long-term rates, all of the real interest rate series from the two approaches are highly correlated. Specifically, the *ex post* is highly correlated with the MA(4) approach with the correlation coefficient of 0.99. There is generally a significant positive correlation between the real interest rate in each country and the real interest rate in the other country (US).

In summary, our findings indicate that the real interest rates obtained from different approaches yield not quite different time series processes, and they appear to have the same mean and vary across time in similar patterns over the sample period. In the next section we test whether the stationarity of the real interest rates depends on the type of method used to construct the real rate of interest. In other words, we will test whether the stationarity of the real interest rate series are sensitive to the computations of the inflation rate.

2.2 UNIT ROOT TESTS

The following step is to assess the most appropriate technique to test the hypothesis of real interest rate parity (RIP) by examining whether real interest rates are stationary or not. Stationarity means that a variable, although fluctuating, tends to return to a constant mean (hence, it is called "mean reverting"). A non-stationary variable, on the other hand, would exhibit apparent changes in mean, or appear highly persistent. Such behavior in real interest rates might, superficially, appear highly unlikely, so what does the evidence on this show? In order to answer this, we conduct some statistical tests designed to reveal whether the series are stationary or not. This is not quite as straightforward as it sounds, as there is a wide range of tests in current use, reflecting the range of potential non-stationarities which have arisen in empirical testing- including breaks in series, and other changes in their means. So to investigate the question of the stationarity or otherwise in international real rates, we use a set of different tests. For presentational purposes, the full set is shown in the Appendix (see tables A1 and A2), and in Tables 2.13 and 2.14 we show only the results from the most commonly used test in literature, the Augmented Dickey-Fuller (ADF) test.

The power of these tests tends to be very low when the root is close to one, especially in small samples (Shiller and Perron, 1985). Furthermore, a serious problem is that the standard tests are biased towards the non-rejection in the presence of structural breaks. In an attempt to solve the above-mentioned problems, Moosa and Bhatti (1996) find that a series of alternative univariate unit root tests that are more powerful than the conventional ADF tests lead to more promising results. Some other authors try to find more accurate evidence enlarging the sample period considered[9]. Nevertheless, as long as we extend the sample period a new set of problems arises linked to discontinuities in the series generated either by shocks or institutional changes[10]. 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.

---

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

[10] Fountas and Wu (1999), and Goldberg, Lothian and Okunev (2003) apply unit root tests that allow for structural breaks in the series finding rejection of the null in more cases.

An important aspect of analyzing the time series process of real interest rates is to have a unit root test that is able to identify a nonstationary property. Unfortunately, as mentioned above, unit root tests are notoriously low power tests. To overcome this, we present three different unit root tests, which are: the augmented Dickey-Fuller test (ADF), the DF-GLS unit root tests and the Kwiatkowski, Phillips, Schmidt, and Shin test (KPSS). Consider the time series with serial correlation in errors described as

$$y_t = a + \rho y_{t-1} + \varepsilon_t \qquad (2.4)$$

and

$$\varepsilon_t = \phi\varepsilon_{t-1} + e_t + \theta e_{t-1} \qquad (2.5)$$

The ADF test is carried out by estimating

$$\Delta y_t = a + \alpha y_{t-1} + \sum\nolimits_{j=1}^{k} \beta_j \Delta y_{t-j} + \varepsilon_t \qquad (2.6)$$

where $\alpha = \rho - 1$ and $t = 1, ..., T$. The augmented terms Δy_t of higher order lags are included into equation (2.4) to correct the serial correlations of the disturbances ε_t. The number of k lags are selected by the Schwartz Information Criteria. The null hypothesis of a unit root ($\alpha = 0$) is tested against the alternative hypothesis of stationarity ($\alpha < 0$). The test statistic is evaluated using the conventional t-ratio for α and the critical value is obtained by MacKinnon's updated version of Dickey-Fuller critical values.

The DF-GLS[11] unit root test is developed to solve the problem of low power. Elliott et al. (1996) propose a simple modification of the ADF tests in which the data are detrended so that explanatory variables are removed from the data prior to running the test regression. The GLS detrending of the data yields substantial power gains. After obtaining the GLS detrended data, say y_t^d ,the DF-GLS test involves estimating the standard ADF test by substituting the GLS detrended y_t^d for the original y_t :

$$\Delta y_t^d = \alpha + \alpha y_{t-1}^d + \sum_{j=1}^{k} \beta_j \Delta y_{t-j}^d + \varepsilon_t \qquad (2.7)$$

[11] In order to determine whether each of the variables are I(1), we used the modified Dickey-Fuller (DF) test, based on generalized least squares (GLS) detrending series (commonly called the DF-GLS test), as proposed by Elliot, Rothenberg and Stock (1996) and, the Ng and Perron (2001) tests for unit root. While the standard Dickey-Fuller and Philip-Perron (PP) tests have been criticized for their poor size and power properties, Elliot, Rothenberg and Stock (1996) have shown that the DF-GLS test is almost uniformly most powerfully invariant.

where the lag length k in this equation is selected using a modified Akaike information criteria (MAIC), which is

$$MAIC = -2(l / T) + 2(k + \tau) / T \qquad (2.8)$$

where l is the bandwidth parameter for the kernel-based estimators of the residual spectrum at frequency zero and $\tau = \alpha^2 \sum_t \tilde{y}_{t-1}^2 / \hat{\sigma}^2$ for $\tilde{y}$ is the autoregressive spectral density estimator. Perron and Ng (1996) suggested the use of MAIC and found substantial size improvements over standard information criteria in the unit root testing.

Lastly, we use the KPSS test to test the null of stationarity against the alternative hypothesis of a random walk. The KPSS test starts with

$$y_t = \delta t + \zeta_t + \varepsilon_t \qquad (2.9)$$

where ε_t is a stationary process and ζ_t is a random walk given by

$$\zeta_t = \zeta_{t-1} + u_t, \quad u_t \sim \text{iid}(0, \sigma_u^2) \qquad (2.10)$$

The null hypothesis of stationarity is formulated as

$$\text{H}_0 : \sigma_u^2 \text{ or } \zeta_t \text{ is a constant}$$

and the alternative hypothesis is that the parameter follows a random walk. The test statistic for this hypothesis is given by

$$LM = \frac{\sum_{t=1}^{T} S_t^2}{\hat{\sigma}_e^2} \qquad (2.11)$$

where $S_t = \sum_{i=1}^{T} e_i$, $t = 1, ...T$ is a cumulative residual function for e_t are the residuals from the regression of y_t on a constant and a time trend, and $\hat{\sigma}_e^2$ is the residual variance. We use the Bartlett spectral window kernel-based estimator to obtain a consistent estimate of the variance and select the bandwidth by using the Newey-West method. The test is an upper-tailed test. Maddala and Kim (1998) do not recommend the KPSS test to be used since the KPSS test has low power such that test results can be very sensitive as shown by their Monte Carlo studies. However, we will report results from this test for the sake of completeness (see Appendix), and because it is often used in empirical studies.

Table 2.13: ADF Unit Root Test For Short-Term Real Interest Rates

Country	ADF	Critical value 5%	Probability
Level			
Australia	-2.702838 (T)	-3.442474	0.2372
Belgium	-3.052463 (T)	-3.442238	0.1221
Canada	-2.868358 (T)	-3.444756	0.1763
France	-2.132301 (T)	-3.437629	0.5236
Italy	-2.533960 (T)	-3.462912	0.3115
Japan	-3.054901 (T)	-3.437801	0.1209
Luxembourg	-1.811825 (T)	-3.470032	0.6894
Netherlands	-1.725649 (T)	-3.452358	0.7331
New Zealand	-1.149138 (T)	-3.456805	0.9145
Norway	-1.992027 (T)	-3.452358	0.5988
Spain	-3.249154 (T)	-3.449716	0.0803
Switzerland	-2.903353 (T)	-3.452764	0.1657
UK	-2.731217	-2.884477	0.0716
US	-3.234537 (T)	-3.438515	0.6392
1st difference			
Australia	-7.638668* (T)	-3.442474	0.0000
Belgium	-8.223790* (T)	-3.441777	0.0000
Canada	-15.13339* (T)	-3.444756	0.0000
France	-11.07731* (T)	-3.437801	0.0000
Italy	-6.499123* (T)	-3.462912	0.0000
Japan	-8.736694* (T)	-3.437801	0.0000
Luxembourg	-8.042551* (T)	-3.470851	0.0000
Netherlands	-8.970290* (T)	-3.452764	0.0000
New Zealand	-7.128137* (T)	-3.456805	0.0000
Norway	-4.314021* (T)	-3.452358	0.0044
Spain	-8.434944* (T)	-3.449716	0.0000
Switzerland	-8.113303* (T)	-3.452764	0.0000
UK	-9.313941*	-2.884477	0.0000
US	-5.589739* (T)	-3.438886	0.0000

Notes:
* Implies significance at 5%. (T) indicates the trend is included as indicated by the significance of the trend terms in the estimation.

The unit root tests in table 2.13 indicate that overall the real interest rates appear to be I(1). In other words, based on the conventional ADF and DG-GLS [12]unit root tests, all of the real interest rate series appear to be nonstationary, as the null hypothesis of a unit root are not significantly rejected at 0.05 level[13]. In contrast, KPSS test, where the null hypothesis is switched to be one of stationarity, indicates that real interest rates are mostly stationary. Specifically, KPSS test indicates that all real short-term rates are I(0) except for those of Australia, Belgium. France, Italy, Norway and the US (see Appendix). However, we focus our analysis on the first two unit root tests, which are the ADF and DF-GLS tests.

Table 2.14: ADF Unit Root Test For Long-Term Real Interest Rates

Country	ADF	Critical value 5%	Probability
Level			
Australia	-1.742274 (T)	-3.437629	0.7279
Belgium	-2.002527 (T)	-3.437801	0.5952
Canada	-2.217633 (T)	-3.437629	0.4763
France	-2.236527 (T)	-3.437801	0.4658
Italy	-2.009104 (T)	-3.437801	0.5916
Japan	-3.280010 (T)	-3.438154	0.0733
Luxembourg	-1.952910 (T)	-3.440059	0.6217
Netherlands	-3.296537 (T)	-3.438154	0.0705
New Zealand	-1.679204	-2.879267	0.4399
Norway	-1.634976 (T)	-3.437801	0.7749
Spain	-3.300810 (T)	-3.448681	0.0544
Switzerland	-3.294137 (T)	-3.437801	0.0709
UK	-2.921919 (T)	-3.437801	0.1584
US	-2.634493 (T)	-3.438154	0.2658
1st difference			
Australia	-11.01584* (T)	-3.437801	0.0000
Belgium	-7.750223* (T)	-3.437801	0.0000
Canada	-10.13433* (T)	-3.437977	0.0000

[12] While the standard Dickey-Fuller test has been criticized for its poor size and power property, Elliot, Rothenberg and Stock (1996) have shown that the DF-GLS test is almost uniformly most powerfully invariant.

[13] Rejection of the null hypothesis at 0.05 significant level is marked by one asterisk.

France	-7.783115* (T)	-3.437801	0.0000
Italy	-7.176661* (T)	-3.437801	0.0000
Japan	-10.52299* (T)	-3.437801	0.0000
Luxembourg	-10.03717* (T)	-3.440059	0.0000
Netherlands	-5.636647* (T)	-3.438154	0.0000
New Zealand	-10.62356*	-2.879267	0.0000
Norway	-8.892995* (T)	-3.437801	0.0000
Spain	-7.267982* (T)	-3.448681	0.0000
Switzerland	-8.305152* (T)	-3.437801	0.0000
UK	-10.73911* (T)	-3.437801	0.0000
US	-5.971386* (T)	-3.438154	0.0000

Notes:
* Implies significance at 5%.. (T) indicates the trend is included as indicated by the significance of the trend terms in the estimation.

We find that the real interest rates are I(1) in all cases under the conventional ADF unit root test. The DF-GLS test confirms the results from the unit root for the *ex post* rates (see Appendix). Note that the null hypothesis of the KPSS test is constructed differently from the other unit root tests. In the KPSS test, we test the null of stationarity against the alternative hypothesis of a unit root. The KPSS test rejected the null of stationarity in most of the cases (except for those of New Zealand, Spain and Switzerland).

Our finding that all real interest rates are I(1) deserves some discussion. Assuming that nominal interest rates and inflation rates are I(1), the Fisher equation would imply that the real interest rate is I(0) or that cointegration exists between nominal interest rates and inflation. A large part of the literature has concluded that real interest rates do follow a random walk (e.g. Rose, 1988).

CHAPTER 3: LINEAR TESTS OF REAL INTEREST RATE PARITY

3.1 INTRODUCTION

A link between the real interest rate in one country to another country's real interest rate is based on the Real Interest Rate Parity (RIP) hypothesis, which states that arbitrage should encourage a tendency toward parity of real interest rates if agents make their forecasts using rational expectations and there are no trade impediments in both good and asset markets. When the Bretton Woods era ended, the controls had been eroded by the emergence of liquid international financial markets. Since the 1970s, the industrialized countries no longer needed capital controls to preserve an exchange rate peg and hence there has been further growth of highly mobile capital flows and a deepening of international capital markets. In particular, capital controls were removed in the US, Canada and Switzerland after 1973 and the same action happened in the UK and Japan in 1979. By the early 1990s, no industrialized countries, all members of OECD, maintained capital controls of any significance. Many countries followed the same movements of capital liberalization. Economic reforms and financial innovations reduced the transaction costs and risks of foreign investment which in turn stimulated a growth of capital flows. As a result, we would expect the world economy to be increasingly integrated such that international real interest rates should be tied together.

Much research has attempted to test whether there exists a complete international linkage of real interest rates by testing for the equality of the real rates across countries. However, the empirical findings provide inconclusive answers. Generally, to test the international equality of real interest rates, authors consider a standard linear regression of the *ex ante* real rate in home country (r_t^e) on the foreign country's *ex ante* real interest rate (r_t^{e*}). Based on the assumption of rational expectations,

$$r_t^e = \alpha + \beta r_t^{e*} + \varepsilon_t \qquad (3.1)$$

where ε_t is a Gaussian error term. The evidence of real interest rate equalization relies on failing to reject the joint hypothesis of $\alpha = 0$ and $\beta = 1$. In other words, if we reject the null hypothesis, the RIP relation does not seem to hold between real interest rates in a two-country pair. Mishkin (1984) and Cumby and Mishkin (1986) followed this test equation and found that *ex ante* real interest rates across industrialized countries do not equalize. However, a potential problem with this test regression, such as spurious result, arises if the possibilities of nonstationary real interest rates are ignored. Therefore, the tests of international linkage have evolved into the test of cointegration between two random walk real interest rate series or the test of mean reversion of the real interest rate differentials. Provided that both r_t^e and r_t^{e*} are nonstationary, the RIP holds if the real interest differential, $r_t^e - r_t^{e*}$, is stationary. This implies that even though real interest rates do not appear to be equalized across countries, it is possible that they share a common trend and move similarly over time.

Several studies found evidence of mean-reverting real interest rate differentials and thus support the validity of long-run RIP hypothesis. For example, Kugler and Neusser (1993) found that the U.S. *ex ante* real interest rate has a significant predictive content for those in OECD countries, and deviations from parity conditions, although substantial in the short run, seem to die out rather quickly as time goes by. Moosa and Bhatti (1996) indicated that the failure of previous studies to reject the null hypothesis of nonstationarity in real interest rate series is due to the fact that these studies used low power unit root tests. Using more powerful tests, Moosa and Bhatti (1996) found that real interest differentials appear to be mean reverting. Furthermore, Ferreira (2003) found that RIP holds in both developed and emerging economies since the real interest differentials revert to their means rapidly once the possibility of structural breaks is allowed. Wu and Chen (1998) and Holmes and Wang (2008) supported the RIP hypothesis because real interest rate differentials seem to be stationary when employing the panel unit root test. In contrast, Chung and Crowder (2003) used several unit root tests, such as the panel unit root test, the Covariance Augmented Dickey-Fuller test, and Johansen’s test, in investigating the order of integration in the RIP relationship and concluded that RIP does not hold for G-5 countries.

3.2 EMPIRICAL RESULTS

3.2.1 Standard Regression of RIP and Unit Root Tests

We now present and discuss evidence of tests on RIP in this part of the chapter. In our analysis we will estimate the RIP regression $r_t=a+br^*_t+u_t$ (eq.3.2) where r_t and r_t^* are the dependent and reference variables, a and b denote the parameters and u_t is the error term. As reference variable (foreign) we will be using the short term and long term real interest rates of US.

We then define two forms of RIP, following the approach taken by Fountas and Wu (1999), namely the Strong form and the Weak form. The strong form holds if u_t is stationary (meaning that the real interest rates of the pair countries that we have run regressions are cointegrated) and a=0, b=1.

The weak form holds if u_t is stationary and $a\neq0$ and/or $b\neq1$. The intuition behind the weak form of RIP is that a and b may differ from the values implied by the strong RIP due to:

- The presence of transaction costs that create a neutral band with no profitable arbitrage opportunities around real interest parity
- Different national tax rates
- The existence of a constant foreign exchange risk premium
- The existence of non traded goods whose prices cannot be equalized internationally thus causing price indexes to differ across countries even if fully integrated financial markets exist.

We first assume that all the rates are stationary and test them with a standard linear regression method. Tables 3.1 and 3.2 examine the relationship between bivariate combinations of the thirteen countries' real interest rates. If the RIP holds, then the intercept and slope coefficients should equal zero and unity, respectively. The results show that the RIP hypothesis can be rejected for all combinations and hence the real interest rates do not equalize across countries when one uses standard regression methods.

Table 3.1: The Standard Regression of RIP Testing (short-term *ex post* rates)

Country	a	b	Ho: a=0,b=1
Australia-US	2.890609* (0.556503)	0.798331* (0.083750)	27.01172*
Belgium-US	2.478113* (0.264750)	0.394497* (0.039444)	183.6926*
Canada-US	0.568909 (0.353167)	1.165915* (0.053443)	52.63231*
France-US	1.946154* (0.224348)	0.453201* (0.034659)	212.9832*
Italy-US	-0.048118 (0.661580)	1.170545* (0.108512)	7.098985*
Japan-US	0.464303 (0.272175)	0.347264* (0.042047)	548.3715*
Luxembourg-US	3.705600* (0.293767)	0.254177* (0.039125)	262.4172*
Netherlands-US	2.157547* (0.118749)	0.245847* (0.018872)	1487.880*
New Zealand-US	4.389908* (0.621118)	0.794138* (0.105698)	77.57051*
Norway-US	5.590914* (0.566441)	0.094648 (0.084121)	58.23993*
Spain-US	1.600582* (0.663576)	1.253859* (0.098932)	53.31876*
Switzerland-US	0.300526 (0.424557)	0.524290* (0.067471)	96.066346*
UK-US	1.527237* (0.598604)	0.908658* (0.084845)	9.777026*

Notes:
*denotes the rejection of the null hypothesis at 5% significance level. The second and third columns represent the estimated coefficients *a* and *b* as in Eq. (3.2) with the corresponding standard errors in the parentheses. The fourth column represents the test statistics of joint significance whether the coefficients a=0 and b=1.

Having run the regressions for short term real interest rates we could not find stationary residuals by conducting ADF and DF-GLS unit root tests for any of the pair countries that were examined, except for the pairs Netherlands-US and the UK-US (the results are not reported). The implications of this finding are that we could not establish cointegration of the real interest rates in the short run for almost all countries.

Alternatively, we tested whether the real interest rate differential series contain a unit root by conducting univariate unit root test methods. A stationary real interest rate differential would be consistent with a cointegrating relationship between two real interest rate series characterised by a unity slope. For this purpose, we employed DF-GLS unit root test[14] that offer higher power and less size distortion relative to the more familiar ADF test and found that RIRDs contain a unit root for most of the countries, except for France, the Netherlands and the UK.

Table 3.2: The Standard Regression of RIP Testing (long-term *ex post* rates)

Country	a	b	Ho: a=0,b=1
Australia-US	1.734187* (0.463714)	0.975928* (0.062776)	43.16115*
Belgium-US	1.816408* (0.238298)	0.853693* (0.032260)	53.78906*
Canada-US	0.747358* (0.199747)	1.010337* (0.027041)	63.29573*
France-US	0.996812* (0.269596)	1.035454* (0.036497)	80.22288*
Italy-US	0.469196 (0.529840)	1.363733* (0.071728)	131,2774*
Japan-US	0.423499 (0.438861)	0.677635* (0.059412)	77.64589*
Luxembourg-US	3.379819* (0.254293)	0.494202* (0.033954)	112.0664*
Netherlands-US	2.978332* (0.245590)	0.579088* (0.033247)	80.56763*
New Zealand-US	2.701861* (0.573644)	0.878133* (0.077658)	40.97129*
Norway-US	1.987856* (0.436105)	0.845409* (0.059039)	20.34836*
Spain-US	1.165751* (0.535656)	1.246480* (0.069863)	95.90018*

[14] The results of this test are not reported.

Switzerland-US	3.035794* (0.231876)	0.198031* (0.031391)	757.3943*
UK-US	2.890515* (0.442636)	0.903539* (0.059923)	96.61141*

Notes:
*denotes the rejection of the null hypothesis at 5% significance level. The second and third columns represent the estimated coefficients *a* and *b* as in Eq.(3.2) with the corresponding standard errors in the parentheses. The fourth column represents the test statistics of joint significance whether the coefficients a=0 and b=1.

Having found that all of the time series of the real interest rates are I(1), we can proceed with by testing for cointegration between pairs of real interest rates with the US being the reference country, using the Engle-Granger methodology. From the results of unit root tests-which are not reported-we found that residuals are nonstationary for any of the pair countries examined except for the pair Luxembourg-US. Overall, if all real interest rate series (both short-term and long-term) are treated as stationary and used in the traditional linear regression tests for RIP condition, there are rejections of real rate equalization and convergence toward zero mean.

Tables 3.3 and 3.4 include the estimated regressions for real short-term and long-term *ex ante* rates, respectively.

Table 3.3: The Standard Regression of RIP Testing (short-term *ex ante* rates)

Country	**a**	**b**	**Ho: a=0,b=1**
Australia-US	2.883660* (0.556020)	0.799452* (0.083697)	27.05717*
Belgium-US	2.501661* (0.270264)	0.392221* (0.040017)	180.5234*
Canada-US	0.565028 (0.353341)	1.166525* (0.053484)	52.65730*
France-US	1.983287* (0.227299)	0.449612* (0.034931)	210.7672*
Italy-US	-0.050467 (0.661139)	1.170605* (0.108459)	7.088364*
Japan-US	0.374758 (0.272042)	0.355932* (0.041807)	563.0743*
Luxembourg-US	3.702792* (0.293134)	0.254541* (0.039058)	263.2335*
Netherlands-US	2.158027*	0.245694*	1495.860*

	(0.118451)	(0.018831)	
New Zealand-US	4.388294* (0.620521)	0.794235* (0.105619)	77.75058*
Norway-US	5.588576* (0.566360)	0.094982 (0.084143)	58.17021*
Spain-US	1.593641* (0.663694)	1.255009* (0.098986)	53.37920*
Switzerland-US	0.300308 (0.424799)	0.524301* (0.067533)	96.01677*
UK-US	1.650100* (0.614526)	0.095540* (0.086446)	10.20132*

Notes:
*denotes the rejection of the null hypothesis at 5% significance level. The second and third columns represent the estimated coefficients *a* and *b* as in Eq. (3.2) with the corresponding standard errors in the parentheses. The fourth column represents the test statistics of joint significance whether the coefficients a=0 and b=1.

Table 3.4: The Standard Regression of RIP Testing (long-term *ex ante* rates)

Country	a	b	Ho: a=0,b=1
Australia-US	1.834770* (0.468954)	0.966620* (0.063156)	44.42478*
Belgium-US	1.771287* (0.241870)	0.858315* (0.032574)	49.96966*
Canada-US	0.755645* (0.203742)	1.009641* (0.027439)	61.73976*
France-US	0.949557* (0.273998)	1.040167* (0.036901)	76.56830*
Italy-US	0.506421 (0.539599)	1.360500* (0.072671)	129.3484*
Japan-US	0.243469 (0.437520)	0.695093* (0.058923)	82.23794*
Luxembourg-US	3.376468* (0.253893)	0.494675* (0.033906)	112.1753*
Netherlands-US	2.958089* (0.250646)	0.581116* (0.033756)	77.21431*
New Zealand-US	2.775816* (0.582933)	0.871393* (0.078507)	40.97942*
Norway-US	2.041597*	0.840535*	20.51276*

	(0.443547)	(0.059735)	
Spain-US	1.158715* (0.535530)	1.247380* (0.069863)	96.05265*
Switzerland-US	3.001146* (0.235667)	0.201347* (0.031738)	750.7488*
UK-US	2.921045* (0.451406)	0.900681* (0.060793)	94.06361*

Notes:
*denotes the rejection of the null hypothesis at 5% significance level. The second and third columns represent the estimated coefficients *a* and *b* as in Eq. (3.2) with the corresponding standard errors in the parentheses. The fourth column represents the test statistics of joint significance whether the coefficients a=0 and b=1.

Having found that all of the time series of the *ex ante* real interest rates are I(1), we can proceed with by testing for cointegration between pairs of real interest rates with the US being the reference country, using the Engle-Granger methodology. The unit root tests-the results of which are not reported- are similar with those of *ex post* rates. They showed that the residuals are also nonstationary for all the countries, except for the Netherlands, the UK and Luxembourg. The results show that the RIP hypothesis can be rejected for nearly all combinations and hence the real interest rates do not equalize across countries when one uses standard regression methods for *ex ante* -both short-term and long-term- real rates.

In summary, our findings strongly indicate that the RIP condition does not hold among these industrialized countries. The results are consistent with Cumby and Mishkin (1986) and Jenkins and Madzharova (2007) such that the joint hypothesis of a zero intercept and a unit linear relationship between real interest rates across countries is rejected. The rejection is robust to the approaches of constructing the underlying real interest rates. However, the estimated coefficients are quite different, indicating some sensitivity to the underlying methods of construction of the real rates. While real interest rates in all of the country pairs and in both constructing approaches studied show statistically significant positive relationships (b’s)-except for those of Norway in the short run -, these linkages are not complete as predicted in the strong form (a= 0, b=1) of international real interest rate connections. These estimated connections do not differ substantially between different methods of calculating the real interest rate.

3.2.2 Cointegration

Cointegration analysis suggest that if two series, such as nominal interest rates in two different markets, are non-stationary, but there exists some linear combination of them that is a stationary process, then the two rates are cointegrated with a cointegrating parameter β. For the purpose of this analysis, following current trends in the literature, the equation (3.2) will be estimated using the Johansen (1988) cointegration method. This will be complemented by impulse response analysis. We now provide a brief description of this method.

The idea of cointegration can be related to the concept of long-run equilibrium between time series when one allows for the possibility of nonstationarity in the underlying series. If the underlying real rates were found to be unit roots, then an appropriate testing methodology of the RIP hypothesis would be to use the Johansen (1988) cointegration framework. This method tests if a linear combination of nonstationary (I(1)) variables is stationary (I(0)), then the variables are said to be cointegrated. The existence of a cointegrating vector implies that the two variables cannot move too far apart. If the real interest rates between two countries are cointegrated, for the RIP to hold, the cointegrating vector must be [1,-1]. If the cointegrating vector differs from the unit vector, the real rates do not follow each other sufficiently to equalize, but are merely comoving. Briefly, the idea of cointegration is based on a vector autoregressive (VAR) model

$$Y_t = \Phi_1 Y_{t-1} + ... + \Phi_k Y_{t-k} + U_t, \qquad (3.3) \qquad t=1,2,..,T$$

where Y_t is the n×1 vector of I(1) variables and U_t is a vector of white noise errors.

We rewrite this equation as

$$\Delta Y_t = \Pi_1 Y_{t-1} + \Pi_2 \Delta Y_{t-1} + ... + \Pi_k \Delta Y_{t-k+1} + U_t \quad (3.4)$$

where $\Pi_1 = -I + \sum_{i=1}^{k} \Phi_i$ and $\Pi_j = -\sum_{i=j}^{k} \Phi_i$ for j=2,…,k. The vector of interest is Π_1 which indicates the long-run relationship between the variables in Y_t. The rank of the Π matrix (r) conveys important information about the cointegrating behavior of the variables. If the matrix Π_1 has zero rank, then there is no cointegration among the I(1) variable. The reduced rank (r<n) of the matrix Π_1 implies that there are r cointegrating vectors among nonstationary variables. Lastly, the full rank (r=n) of Π_1 implies that all variables are stationary to begin with. Note that in the bivariate case the Π_1 has to be of a rank=1 to support the RIP hypothesis. Specifically, if we find a unit rank, then

the estimated cointegrating vector must be [1, -1] to satisfy the RIP condition. To establish the rank of the Π_1 matrix, we use the trace test and maximum eigenvalue test of Johansen (1991) and only estimate the long run cointegrating vector in the cases where a single cointegrating vector exists.

Usually Π_i has a reduced rank; that is $r \leq (n-1)$. Then we have: $\Pi_i = \alpha\beta'$ (3.5) where α is a $n \times r$ matrix and β' is a $r \times n$ matrix. Then $\beta' X_{t-1}$ are the r cointegrated variables, β' is the matrix of coefficients of the cointegrating vectors, i.e. the long-run coefficients, and α has the interpretation of the matrix of error correction terms. The rank of the matrix Π_i and the number of cointegrating relation(s) will be determined using the two commonly used likelihood ratio (LR) test statistics, as provided in Johansen (1988) i.e.: the trace statistic (λ_{trace}) and the maximum eigenvalues (λ_{max}) with their test statistics given respectively as follows:

$$\lambda_{trace} = -\mathrm{T}\sum_{i=r+1}^{n}\log(1-\hat{\lambda}_i) \quad (3.6)$$

$$\lambda_{\max} = -\mathrm{T}\log(1-\hat{\lambda}_{r+1}) \quad (3.7)$$

where λ_i is the i-th largest eigenvalue of the Π_i matrix in equation (3.5). The tests will be conducted both under the null that r = 0 and then that r = 1. Following Haug *et al.* (2000), we employed *p values* to test for cointegration and the null hypothesis is tested sequentially from low to high values of r. The testing in the sequence ends when the null is not rejected for the first time.

In addition, if cointegration between the variables were found, an error correction model (ECM) of the relationship would be estimated to examine the short run dynamics as suggested by Scholnick (1996). The ECM to be estimated is given as:

$$\Delta r_t^j = \delta_0 + \alpha_1 \Delta r_t^i + \phi \mathrm{EC}_{t-1} + \omega_t \quad (3.8)$$

where the symbol Δ represents a first difference of the relevant variable, ω_t is the white noise error term, while EC are the residuals from the cointegrating vector between the two interest rates. The coefficient, α_1 reflects the immediate or short-term pass-through, and φ is the coefficient of the error correction term, which measures the degree of adjustment to equilibrium. A statistically significant coefficient of the EC_{t-1} will suggest that economic forces (arbitrage and/or policy measures) are in operation to restore long-run equilibrium following a short-run disturbance.

From equation above, the mean lag, according to Doornik and Hendry (1994) can be obtained as:

$$ML=(1-\alpha_1)/\varphi \qquad (3.9)$$

Equation (3.9) represents the mean adjustment lag or the degree of rigidity for the symmetric error correction model. It indicates the mean adjustment lag at which a change in candidate interest rate is fully passed-through to another interest rate. A high *ML* shows high rigidity or slow adjustment of interest rates in response to changes in the candidate interest rate. The converse is the case if the *ML* is low. In the context of our analysis, a high *ML* would indicate a low degree of financial integration.

The cointegration method has become a commonly used test to examine the real interest rate co-movements in recent literature. Goodwin and Grennes (1994) applied the cointegration analysis and found strong support for a version of RIP whereby real rates vary randomly within the transaction cost band, but revert to a stable long-run equilibrium relationship. Al Awad and Goodwin (1998) and Fountas and Wu (1999)[15] also used the cointegration method to test the RIP hypothesis of the *ex ante* real interest rate and found evidence supporting the existence of interest parity. However, Fraser and Taylor (1990) tested the RIP by examining the cointegration between nominal interest rate differentials and the relative inflation in the vector autoregressive representation. This investigation led to an overwhelming rejection of real interest rate parity in a number of industrialized countries.

After establishing the first difference stationarity of the time series of the real interest rates, we can proceed with by testing for cointegration between pairs of real interest rates with US being the reference country, using the Johansen methodology. The bivariate Johansen's cointegration tests are used to examine evidence of cointegrated relationship between I(1) real interest rates across countries in Tables 3.5 and 3.6. These tables report the result of bivariate Johansen cointegration test for the short-term and long-term interest rates, respectively. Since the Johansen cointegration test relies on the assumption of Gaussian error term in a VAR system, the lag orders of bivariate VAR models must be selected a priori in order to correct for serial autocorrelation. The specification of the lag length of the VAR is tested

[15] Fountas and Wu (1999) extended the cointegration technique by allowing for structural shifts such that the timing of regime switching is not known a priori.

sequentially[16], using the five information criteria reported in Eviews, namely the Sequential Modified Likelihood Ratio (LR), the Final Prediction Error (FPE), Akaike Information Criterion (AIC), the Schwartz Information Criterion (SC) and the Hannan-Quinn Information Criterion (HQ). These criteria may produce conflicting VAR order selections. Where this occurs, preference was given to the SC if it produced economically interpretable results, which is the one that we use in this study. We use the trace and maximum eigenvalue cointegration tests that are carried out with the selected lag orders to determine the rank of the matrix Π_1 as outlined in the previous section. In these cointegration tests, we allow for linear deterministic trends in the level data and only intercepts in the cointegrating equations. Test statistics of the trace test and maximum eigenvalue tests are reported for each hypothesized number of rank in the matrix Π_1.

Tables 3.5 and 3.6 include the estimated cointegration regressions for real short-term and long-term *ex post* rates, respectively. The coefficients of the cointegrating vector are estimated and reported in the eighth column of Tables 3.5 and 3.6. For the strict form RIP to hold, the coefficients in the cointegrating vector are expected to be β =[1 ,-1]. We use the LR test to investigate whether the coefficients of cointegrating vector are significantly different from the restriction [1,-1] for the combinations that have one cointegrating vector and the probabilities of the LR test statistic are reported in the last column. We focus our attention only on the case of one cointegrating vector and find evidence for one cointegrating vector in the cases of France-US, Netherlands-US, Switzerland-US and UK-US .

Table 3.5: Johansen Cointegration Results for Bivariate System (short-term *ex post* rates)

Country	**Lag**	**Trace**		**Max. Eigen**		**# Coint. Vectors**	**Coint. Estimates**	**RIP [1 -1]**
		r=0	**r ≤1**	**r=0**	**r =1**			
Australia-US	2	17.413	6.322	11.092	6.322	0	[1,-3.457]	
Belgium-US	2	21.167	5.408	15.759	5.408	0	[1,-0.370]	
Canada-US	2	19.570	6.536	13.034	6.536	0	[1,-0.698]	

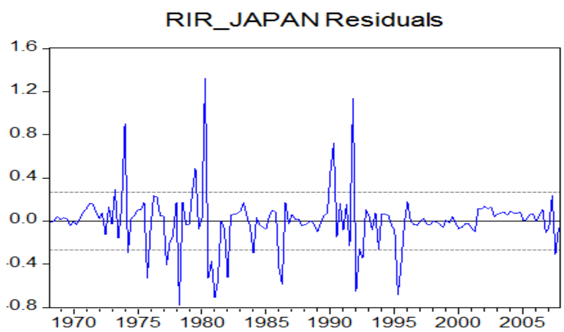
[16] Although this procedure is rarely used in the literature, the need to conform to economic theory and all the *a priori* knowledge that is associated with this theory as suggested by Seddighi *et al.* (2000), lends support for its use in this study.

France-US	1	31.385*	7.443	23.94*	7.443	1	[1,-0.655]	0.085
Italy-US	2	14.930	5.461	9.469	5.461	0	[1,0.208]	
Japan-US	4	22.722	8.706	14.016	8.706	0	[1,-0.097]	
Luxembourg-US	1	22.751	7.468	15.283	7.468	0	[1,0.854]	
Netherlands-US	2	25.893*	6.472	19.42*	6.472	1	[1,0.896]	0.042
New Zealand-US	3	8.261	0.729	7.533	0.729	0	[1,5.212]	
Norway-US	1	20.429	4.482	15.947	4.482	0	[1,2.341]	
Spain-US	2	17.382	6.084	11.298	6.084	0	[1,-0.364]	
Switzerland-US	2	30.388*	3.160	27.23*	3.160	1	[1,1,419]	0.000
UK-US	1	26.914*	6.421	20.49*	6.421	1	[1,-1.252]	0.302

Notes:
The first column represents the lag order of the cointegration test as chosen based on the SC. The maximum lag length is set at 12. *r* denotes a hypothesized number of cointegrating vectors under the null hypothesis. * denotes significance at 5% level. The trace test critical values at the 0.05 level are 15.41 (for r=0) and 3.76 (for $r\leq 1$). The maximum eigenvalue test critical values at the 0.05 level are 14.07 (for r=0) and 3.76 (for r=1). The entries in the "Coint.. Estimate" column are the estimated cointegrating vectors, normalized on the first country real interest rate in the country pair. The last column is the probability of the LR test (which is distributed as a χ_1^2) for examining whether the null hypothesis of the cointegrating vector is equal to [1 -1]′.

Having run the regressions for short term real interest rates (see section 3.2.1) we could not find stationary residuals for any of the pair countries. This means that we could not establish cointegration of the real interest rates in the short run. This might be explained if we have a closer look at the results. For example, when testing for cointegration of short term real interest rates of Japan and the US we can see from the residuals plot that structural breaks exist, thus these interest rate gaps are incorporated in the residuals and not the deterministic component of the model. These structural breaks can be attributed to different monetary policies of each country or specific macroeconomic characteristics of each economy. Similar findings of structural breaks are also reported in Fountas and Wu (1999).

Figure 9: Plot of Residuals for Japan



According to Table 3.5, there exists a cointegrating relationship between real interest rates of France and the US, of the Netherlands and the US, of Switzerland and the US and of UK and the US. Not only is there a linkage between real interest rates of France and the US, but we cannot reject the theoretical [1, -1] cointegrating vector. This finding is the same for the UK-US pair. Thus, we find evidence of strong form of RIP for these two countries, namely France and the UK. For the Netherlands and Switzerland there seems to hold the weak form of RIP, since we reject the cointegrating vector [1,-1] for these two countries.

Table 3.6 Johansen Cointegration Results for Bivariate System (long-term *ex post* rates)

Country	**Lag**	**Trace**		**Max. Eigen**		**#Coint. Vectors**	**Coint. Estimates**	**RIP [1 -1]**
		r=0	**r ≤1**	**r=0**	**r =1**			
Australia-US	1	22.339	6.098	16.239	6.099	0	[1,-1.619]	
Belgium-US	2	22.099	6.478	15.622	6.478	0	[1,-1.042]	
Canada-US	1	19.704	7.330	12.374	7.330	0	[1,-1.173]	
France-US	2	21.291	6.414	14.877	6.414	0	[1,-1.069]	
Italy-US	2	17.598	5.517	12.081	5.517	0	[1,-1.764]	
Japan-US	1	17.044	7.313	9.732	7.313	0	[1,-0.572]	
Luxembourg-US	1	28.864*	7.401	21.46*	7.401	1	[1,-0.720]	0.227
Netherlands-US	2	15.316	5.088	10.228	5.088	0	[1,-0.281]	
New Zealand-US	1	18.947	4.725	14.222	4.725	0	[1,-1.831]	
Norway-US	2	23.040	5.560	17.480	5.560	0	[1,-1.459]	
Spain-US	2	21.161	8.327	12.834	8.327	0	[1,0.379]	
Switzerland-US	2	22.301	4.513	17.789	4.513	0	[1,0.109]	
UK-US	1	17.985	7.614	10.371	7.614	0	[1,-0.019]	

Notes:
The first column represents the lag order of the cointegration test as chosen based on the SC. The maximum lag length is set at 12. *r* denotes a hypothesized number of cointegrating vectors under the null hypothesis. * denotes significance at 5% level. The trace test critical values at the 0.05 level are 15.41 (for *r*=0) and 3.76 (for *r*≤1). The maximum eigenvalue test critical values at the 0.05 level are 14.07 (for *r*=0) and 3.76 (for *r*=1). The entries in the "Coint.. Estimate" column are the estimated cointegrating vectors, normalized on the first country real interest rate in the country pair. The last column is the probability of the LR test (which is distributed as a χ_1^2) for examining whether the null hypothesis of the cointegrating vector is equal to [1 -1]′.

According to Table 3.6, both the trace test and maximum eigenvalue tests indicate one cointegrating relationship between Luxembourgish and US real interest rates. We find that the coefficients of cointegrating vector are not significantly different from the restriction [1,-1] and thus the results can support the strong form of real interest rate parity for this country.

Now we do the same test for the *ex ante* real interest rates (both short-term and long-term ones). Table 3.7 reports the results.

Table 3.7: Johansen Cointegration Results for Bivariate System (short-term *ex ante* rates)

Country	Lag	Trace		Max. Eigen		# Coint. Vectors	Coint. Estimates	RIP [1 -1]
		r=0	r ≤1	r=0	r =1			
Australia-US	2	17.355	6.309	11.045	6.309	0	[1,-0.475]	
Belgium-US	2	21.669	4.373	17.296	4.373	0	[1,-0.342]	
Canada-US	2	19.599	6.513	13.087	6.513	0	[1,-0.698]	
France-US	1	31.129*	6.447	24.68*	6.447	1	[1,-0.608]	0.042
Italy-US	2	14.913	5.465	9.448	5.465	0	[1,0.208]	
Japan-US	4	23.013	8.671	14.342	8.671	0	[1,-0.069]	
Luxembourg-US	1	22.797	7.470	15.327	7.470	0	[1,0.848]	
Netherlands-US	2	25.817	6.408	19.409	6.408	0/1	[1,0.899]	0.042
New Zealand-US	3	8.327	0.741	7.586	0.741	0	[1,5.151]	
Norway-US	1	20.358	4.440	15.918	4.440	0	[1,2.343]	
Spain-US	1	20.401	9.653	10.747	9.653	0	[1,0.076]	
Switzerland-US	2	30.416*	3.169	27.25*	3.169	1	[1,1.417]	0.000
UK-US	1	25.454	5.181	20.27*	5.181	0/1	[1,-1.29]	0.241

Notes:
The first column represents the lag order of the cointegration test as chosen based on the SC. The maximum lag length is set at 12. *r* denotes a hypothesized number of cointegrating vectors under the null hypothesis. * denotes significance at 5% level. The trace test critical values at the 0.05 level are 15.41 (for *r*=0) and 3.76 (for $r \leq 1$). The maximum eigenvalue test critical values at the 0.05 level are 14.07 (for *r*=0) and 3.76 (for *r*=1). The entries in the "Coint.. Estimate" column are the estimated cointegrating vectors, normalized on the first country real interest rate in the country pair. The last column is the probability of the LR test (which is distributed as a χ_1^2) for examining whether the null hypothesis of the cointegrating vector is equal to [1 -1]′.

For the MA(4) approach, both the trace test and maximum eigenvalue tests indicate that there is not any cointegrating relationship between nonstationary real short-term interest rate series for all countries, except for those of France and Switzerland. For these two countries the real rate series share a common trend in the adjustments toward parity conditions. However, for the Netherlands and the UK, the trace test statistics and maximum eigenvalue test statistics do not give consistent findings. Specifically, the real interest rates between Netherlands and US and the UK and US seem to have one cointegrating vector according to the maximum eigenvalue test, and none according to the trace test. Furthermore, not only is there a linkage between real interest rates of the UK and the US, but we cannot reject the theoretical [1, -1] cointegrating vector in this case. This finding is very important, since it implies the existence of the strong form of RIP for this country. For France and the US, the Netherlands and the US, as well as Switzerland and the US, we find that the coefficients of cointegrating vector are significantly different from the restriction [1,-1] and thus the results can support the weak form of real interest rate parity for these countries.

Table 3.8: Johansen Cointegration Results for Bivariate System (long-term *ex ante* rates)

Country	Lag	Trace		Max. Eigen		# Coint. Vectors	Coint. Estimates	RIP [1 -1]
		r=0	r ≤1	r=0	r =1			
Australia-US	1	21.299	5.636	15.664	5.636	0	[1,-1.593]	
Belgium-US	2	21.695	6.425	15.269	6.425	0	[1,-0.988]	
Canada-US	1	19.045	6.264	12.781	6.264	0	[1,-1.213]	
France-US	2	20.926	6.079	14.846	6.079	0	[1,-1.044]	
Italy-US	2	17.211	5.332	11.879	5.332	0	[1,-1.698]	
Japan-US	1	16.191	6.731	9.460	6.731	0	[1,-0.519]	
Luxembourg-US	1	28.759*	7.378	21.38*	7.378	1	[1,-0.718]	0.223
Netherlands-US	2	15.609	4.756	10.854	4.756	0	[1,-0.278]	
New Zealand-US	1	17.939	4.462	13.478	4.462	0	[1,-1.810]	
Norway-US	2	22.493	5.317	17.176	5.317	0	[1,-1.432]	
Spain-US	2	21.127	8.300	12.827	8.300	0	[1,0.374]	
Switzerland-US	2	21.730	4.166	17.565	4.166	0	[1,0.111]	
UK-US	1	18.111	7.331	10.779	7.331	0	[1,-0.252]	

Notes:
The first column represents the lag order of the cointegration test as chosen based on the SC. The maximum lag length is set at 12. *r* denotes a hypothesized number of cointegrating vectors under the null hypothesis. * denotes significance at 5% level. The trace test critical values at the 0.05 level are 15.41 (for *r*=0) and 3.76 (for *r*≤1). The maximum eigenvalue test critical values at the 0.05 level are 14.07 (for *r*=0) and 3.76 (for *r*=1). The entries in the "Coint.. Estimate" column are the estimated cointegrating vectors, normalized on the first country real interest rate in the country pair. The last column is the probability of the LR test (which is distributed as a χ_1^2) for examining whether the null hypothesis of the cointegrating vector is equal to [1 -1]′.

According to Table 3.8, the similar results of none cointegrating vector between these countries' real rates are also found in the long run. However, there exists a cointegrating relationship between Luxembourgish and US real interest rates. Not only are there long run connections between Luxembourg-US real interest rate pair under the MA(4) approach, but we also reject that these linkages are different from the theoretical RIP condition implied. Thus, the findings can support the strong form of RIP for this country.

The question of how long it would take for full adjustment or the attainable long run interest rate pass-through to be realized is answered in the mean adjustment lags (ML). The value of the ML indicates the exact time it takes the transmission process to be completed and shows whether the process is sluggish or fast. We estimated ML for Luxembourg from equation (3.9) and found that its value is round -14 for *ex post* rates and -9.9 for *ex ante* rates. The meaning of this finding is that a low *ML* indicates a high degree of financial integration.

Overall, this section examined the sensitivity of real interest rate linkages, when real interest rates were measured by the two approaches. The literature generally assumes the dynamic behaviours of real interest rates to be linear. Therefore, both classical regression analysis and linear cointegration tests were employed. To test the international equality of real rates, the linear regression test was based on a classical regression of home real interest rate r_t^e on foreign real rate r_t^{e*} and a joint test of the significance of intercept and slope coefficients from theoretical values of 0 and 1, respectively.

We also applied Johansen's cointegration test to investigate common stochastic trends between international real interest rates. The evidence of none long-run connection between nonstationary real interest rates seems to hold under these two different approaches for the majority of countries examined. The results show that there are overwhelming rejections of real rate equalization and convergence toward zero mean and these results are robust to both approaches examined. In other words, our findings indicate that the RIP condition does not hold between most of the countries examined. The results are consistent with Cumby and Mishkin (1986) and Mishkin (1984) such that the joint hypothesis of a zero intercept and a unit linear relationship between real interest rates across countries is rejected. However, the results obtained for the short-term real interest rates of Netherlands, Switzerland and the UK support the weak version of the RIP hypothesis for the first two countries and the strong version of the RIP for the latter. These findings are the same for both approaches. Moreover, the results for France are mixed, since those obtained for the short-term *ex post* real rates support the strong form of the RIP, while those obtained for the short-term *ex ante* real rates support the weak form of it. On the other hand, the results obtained for the long-term real interest rates of Luxembourg support the strong version of the RIP hypothesis for both approaches. These results contribute to the findings already reported by other empirical researchers. The implications of the existence of such a parity has a number of implications concerning the monetary policy of each country.

3.2.3 Cointegration Tests with Structural Breaks

Standard cointegration tests, such as Johansen cointegration test described above, suffer from a major drawback when the time period under examination includes fiscal policy changes, institutional changes and other changes in the operational mode of the monetary system. The conventional tests for cointegration are not appropriate, since they presume that the cointegrating vector is time-invariant under the alternative hypothesis. In this section we are concerned with a possibility of a more general type of cointegration, where the cointegrating vector is allowed to change at a single unknown time during the sample period. Specifically, we use the Gregory-Hansen cointegration tests that determine endogenously the regime shift in order to test for bilateral short-term and long-term real interest rate convergence in our sample period[17]. A break in the long-run relationship between pairs of real interest rates can occur for several reasons[18]; for example, our sample period 1967-2008 includes a time span of significant dismantlement of restrictions on the free movement of capital across national boundaries in the European Monetary System. Another reason could be the fact that significant changes due to the collapse of the Bretton Woods system in the 1970s or other changes in the stance of fiscal and monetary policy may take place and, thus, account for a change in the relationship between pairs of real interest rate series.

The Gregory-Hansen approach is an extension of similar tests for unit root tests with structural breaks, for example, by Zivot and Andrews (1992). Gregory and

---

[17] The sample period for some countries is 1967Q1-2007Q4, thereby covering both the Bretton Woods system of fixed exchange rates and the adoption of generalized floating exchange rates from 1973.

[18] Real shocks can affect the real interest rate and lead to structural changes in the cointegrating relationship. Supply shocks, as for example the oil price hikes in 1973 and 1979, may cause a level shift in the cointegrating relation. The same holds true for technology and preference shocks. Identification of the sources of shocks would require a structural analysis, as for example an analysis based on a structural vector-autoregression of fiscal and monetary policy transmission. This is beyond the scope of our study.

Hansen (1996) propose the cointegration tests which accommodates a single endogenous break in an underlying cointegrating relationship. The null hypothesis of no cointegration with structural breaks is tested against the alternative of cointegration by the Gregory and Hansen approach. The three models of Gregory and Hansen (1996) with assumptions about structural breaks and their specifications with two variables are as follows:

- Model 1 (Level Shift):

$$r_t = \alpha_1 + a_2 D_t + b r_t^* + u_t, \qquad t=1,\ldots,n \tag{3.10}$$

- Model 2 (Level Shift with Trend):

$$r_t = \alpha_1 + a_2 D_t + b r_t^* + ct + u_t, \quad t=1,\ldots.,n \tag{3.11}$$

- Model 3 (Regime Shift):

$$r_t = \alpha_1 + a_2 D_t + b_1 r_t^* + b_2 r_t^* D_t + u_t, \quad t=1,\ldots,n \tag{3.12}$$

where $D_t=0$, if $t \le [n\tau]$

1, if $t > [n\tau]$

and $\tau \in (0,1)$ is an unknown parameter denoting the relative timing of the change point and [] denotes integer part. The use of the dummy variable D_t is for testing for a structural break or a regime shift. In model 1, there is a level shift in the cointegrating relationship which is modeled as a change in the intercept at the time of the shift by the size of coefficient α_2. In model 2, a time trend into the level shift model is introduced. Finally, in model 3 there is a structural change in the cointegrating relationship that affects both the intercept and the slope. The null hypothesis in all three models is that u_t is nonstationary. Using cointegration tests with structural breaks, r_t and r_t^* are cointegrated if u_t is an I(0) process and α_2 (and b_2) are significantly different from zero.

To test for cointegration between these two variables (r_t, r_t^*) with structural break, Gregory and Hansen (1996) suggest the use of three tests. Their test statistics are the *smallest* values of the test statistics Z_α, Z_t and the ADF statistic. These test statistics are:

$$\begin{aligned} Z_\alpha^* &= \inf_{\tau \in \mathrm{T}} Z_\alpha(\tau), \\ Z_t^* &= \inf_{\tau \in \mathrm{T}} Z_t(\tau), \\ ADF^* &= \inf_{\tau \in \mathrm{T}} ADF(\tau) \end{aligned} \tag{3.13}$$

where $Z_\alpha(\tau)$, $Z_t(\tau)$ and ADF(τ) correspond to the choice of change point τ. In principle the set T can be any compact subset of (0,1). Gregory and Hansen (1996) compute the test statistic for each break point in the interval [0.15n],[0.85n]. In their analysis, critical values are calculated for the tests by simulation methods and a simple Monte Carlo experiment is conducted to evaluate finite-sample performance.

In most of the previous studies on real interest rate parity, an important issue that was not addressed is that the cointegration relationship may have a structural break during the sample period. Therefore, we explore in this section the stability of the real interest rates with the Gregory-Hansen techniques. The break date is found by estimating the cointegration equations for all possible break dates in the sample. We select a break date where the test statistic is the minimum. Gregory and Hansen (1996) have tabulated the critical values by modifying the MacKinnon procedure for testing cointegration in the Engle-Granger method for unknown breaks.

Table 3.9 reports the values of Gregory and Hansen (1996) statistics for the three models for short-term *ex post* rates. It should be noted that the results using *ex ante* rates are similar to those for *ex post* rates and, thus, are not reported. These results imply no evidence for cointegration between Australian and US rates, Belgium and US rates , French and US rates, Japanese and US rates, as well as Luxemburgish and US rates. However, there is evidence (at 5% level) for cointegration between US rates and rates in Canada, Italy, the Netherlands, New Zealand, Norway, Spain, Switzerland and the UK.

Table 3.9: Gregory-Hansen cointegration tests (short-term *ex post* rates)

	ADF*	Z_t^*	Z_a^*	**Brake Date**
Australia				
(1972Q4-2007Q4)				
–Model (1)	-4.35	-3.30	-17.09	1992Q2
–Model (2)	-4.69	-3.71	-23.69	
–Model (3)	-4.43	-3.57	-19.33	
Belgium				
(1967Q1-2003Q4)				
–Model (1)	-3.64	-3.74	-25.53	1982Q1
–Model (2)	-3.53	-3.72	-25.41	
–Model (3)	-3.44	-3.70	-25.17	

Canada				
(1975Q1-2007Q4)				
–Model (1)	-4.74*	-6.27*	-56.77*	1982Q4
–Model (2)	-4.75	-6.28*	-56.83*	
–Model (3)	-3.56	-5.97*	-51.49*	
France				
(1967Q1-2007Q4)				
–Model (1	-3.75	-3.03	-16.35	1982Q1
–Model (2)	-3.71	-3.02	-16.02	
–Model (3)	-3.69	-3.01	-15.56	
Italy				
(1982Q1-2003Q4)				
–Model (1)	-4.65*	-2.34	-10.66	1982Q4
–Model (2)	-4.56	-2.30	-9.37	
–Model (3)	-4.32	-2.27	-8.34	
Japan				
(1967Q1-2007Q4)				
–Model (1)	-3.87	-3.52	-23.76	1994Q2
–Model (2)	-3.78	-3.48	-23.05	
–Model (3)	-3.42	-3.13	-18.22	
Luxembourg				
(1980Q1-1999Q1)				
–Model (1)	-3.68	-2.67	-13.20	1995Q3
–Model (2)	-3.36	-2.35	-10.26	
–Model (3)	-3.21	-2.26	-6.88	
Netherlands				
(1981Q1-2007Q4)				
–Model (1)	-4.74*	-3.29	-18.59	1992Q2
–Model (2)	-5.17*	-3.74	-21.82	
–Model (3)	-4.98*	-3.39	-18.47	
New Zealand				
(1981Q4-2007Q4)				
–Model (1)	-6.59*	-3.87	-22.15	1989Q4
–Model (2)	-6.89*	-4.15	-21.04	
–Model (3)	-6.71*	-4.03	-19.29	
Norway				

(1979Q1-2007Q4)				
–Model (1)	-4.78*	-3.55	-16.69	1986Q1
–Model (2)	-4.82	-3.56	-19.88	
–Model (3)	-4.36	-2.98	-15.01	
Spain				
(1979Q1-2007Q4)	-6.42*	-4.11	-31.30	1995Q3
–Model (1)	-5.69*	-4.07	-30.52	
–Model (2)	-6.08*	-3.63	-23.51	
–Model (3)				
Switzerland				
(1981Q1-2007Q4)				
–Model (1)	-4.63*	-2.93	-16.41	1995Q3
–Model (2)	-4.69	-2.94	-16.50	
–Model (3)	-4.24	-2.66	-12.95	
UK				
(1967Q1-1998Q4)				
–Model (1)	-5.13*	-4.24	-32.72	1990Q2
–Model (2)	-5.21*	-4.26	-33.05	
–Model (3)	-5.03*	-4.15	-31.00	

Notes:
* denotes significance at 5%. The critical values are those presented in Table 1 of Gregory and Hansen (1996).

The break points are in 1992Q2 and 1982Q1 for Australia and Belgium, respectively. The break date for Australia corresponds to the peak of financial deregulation and the severe recession engulfing the Australian economy in the early 1990s, while the break date for Belgium takes place during the period of crisis in the Belgian exchange rate policy in the early 1980s. 1982Q4 represents a period where confidence in the Canadian dollar continued to erode on concerns about the commitment of Canadian authorities to an anti-inflationary policy stance, and the cancellation of a number of large energy projects[19]. The Bank also reluctantly announced in the end of 1982 that it would no longer target M1 in its fight against inflation. The break date for France coincides with a period of high French real

[19] With the dollar falling below US$0.77, the Bank of Canada allowed short-term interest rates to rise to prevent the increasing weakness of the Canadian dollar "from turning into a speculative rout" (Bank of Canada *Annual Report 1982*, 20)

interest rates accompanying the expansionary fiscal policy launched in the second half of 1981.

Moreover, during the fourth quarter of 1982, the Italian lira depreciated by 3.5% against the US dollar. 1995Q3 for Luxembourg coincides with a decreasing interest rate and monetary policy returning to a neutral or even expansionary stance. The Netherlands experienced a structural break in the short-term real interest rate at the time of the 1992 ERM crisis. New Zealand, which adopted inflation targeting in February 1990, has a break date in 1989Q4. The break date for Norway corresponds to a period of high and variable inflation and the introduction of a fixed exchange rate regime which reinstated monetary policy as an instrument of economic policy in Norway and laid the foundation for more stable economic developments. 1995Q3 represents a period where Spain successfully implemented an inflation-targeting regime. The break date for Switzerland corresponds to a period of 2.9% depreciation of the Swiss franc against the US dollar, which strengthened during the remainder of 1995, in part due to "safe haven" effects related to the financial crisis in Asia. Finally, 1990Q2 is associated with a period of high British real interest rates, as the UK applied contractionary monetary policy in preparation for joining the ERM.

Table 3.10 reports the Gregory and Hansen (1996) cointegration test results for long-term *ex post* rates. The results show that the null of a lack of cointegration is not rejected for Italy, the Netherlands and the UK. In contrast, the null hypothesis is rejected (at 5%) for Australia, Belgium, Canada, France, Japan, Luxemburg, New Zealand, Norway, Spain and Switzerland. When *ex ante* rates are used, similar results are obtained and hence are not reported.

Table 3.10: Gregory-Hansen cointegration tests (long-term *ex post* rates)

	ADF*	Z_t^*	Z_a^*	Brake Date
Australia (1967Q1-2007Q4)				
–Model (1)	-6.63*	-3.68	-26.03	1974Q1
–Model (2)	-6.60*	-4.09	-31.13	
–Model (3)	-6.64*	-3.85	-28.48	
Belgium (1967Q1-2007Q4)				
–Model (1)	-6.73*	-3.86	-27.83	1996Q3

–Model (2)	-6.71*	-3.75	-26.24	
–Model (3)	-6.82*	-4.22	-32.73	
Canada				
(1967Q1-2007Q4)				
–Model (1)	-5.73*	-3.61	-23.82	1995Q3
–Model (2)	-5.69*	-3.61	-23.68	
–Model (3)	-5.82*	-3.69	-24.73	
France				
(1967Q1-2007Q4)				
–Model (1	-6.00*	-4.23	-33.45	1995Q3
–Model (2)	-5.81*	-4.21	-33.05	
–Model (3)	-6.23*	-4.59	-38.68	
Italy				
(1967Q1-2007Q4)				
–Model (1)	-4.14	-3.31	-21.03	1975Q3
–Model (2)	-4.22	-3.59	-24.93	
–Model (3)	-4.18	-3.28	-20.72	
Japan				
(1967Q1-2007Q4)				
–Model (1)	-6.79*	-3.84	-28.23	1979Q1
–Model (2)	-6.36*	-3.32	-20.94	
–Model (3)	-6.89*	-3.59	-25.19	
Luxembourg				
(1970Q1-2007Q4)				
–Model (1)	-6.35*	-3.23	-21.22	1997Q3
–Model (2)	-6.26*	-3.31	-20.86	
–Model (3)	-6.34*	-3.31	-21.06	
Netherlands				
(1967Q1-2007Q4)				
–Model (1)	-4.31	-3.05	-17.73	1997Q2
–Model (2)	-3.78	-3.03	-17.43	
–Model (3)	-4.09	-3.11	-18.46	

New Zealand				
(1967Q1-2007Q4)				
–Model (1)	-5.18*	-3.74	-26.93	1983Q4
–Model (2)	-6.10*	-4.12	-31.88	
–Model (3)	-5.43*	-3.88	-29.06	
Norway				
(1967Q1-2007Q4)				
–Model (1)	-5.77*	-3.26	-20.59	1983Q4
–Model (2)	-6.30*	-3.92	-29.16	
–Model (3)	-5.91*	-3.51	-24.04	
Spain				
(1978Q2-2007Q4)				
–Model (1)	-5.71*	-4.64*	-34.45	1996Q3
–Model (2)	-5.59*	-4.38	-33.96	
–Model (3)	-5.82*	-4.51	-35.82	
Switzerland				
(1967Q1-2007Q4)				
–Model (1)	-5.45*	-3.42	-23.05	1996Q3
–Model (2)	-5.19*	-3.48	-23.59	
–Model (3)	-5.14*	-3.45	-23.34	
UK				
(1967Q1-2007Q4)				
–Model (1)	-3.84	-3.27	-19.16	1980Q2
–Model (2)	-3.69	-3.11	-15.99	
–Model (3)	-3.90	-3.53	-22.61	

Notes:
* denotes significance at 5%. The critical values are those presented in Table 1 of Gregory and Hansen (1996).

The break date for Australia takes place during the period of the collapse of the fixed exchange rate regime where this new political direction impact on the real exchange rate misalignment of the Australian dollar. As a part of an expansionary monetary policy, the Canadian Central Bank decreased its interest rates in May 1994, but recovery was not evident until the third quarter of 1995 which coincides with the period that the break occurs. The break date for France coincides with a period where there were positive developments in the achievement of Maasthricht budget criteria which resulted in preventing long-term interest rates from rising in France. 1975Q3 coincides with the post crisis period[20] and the depreciation of the Italian lira by a large percentage. During the first quarter of 1979, the Italian lira depreciated by 6.6% against the US dollar.

Furthermore, 1997Q3 for Luxembourg and the break date for the Netherlands coincide with decreasing interest rates. The sharp drop in the inflation rate in 1983-84 matched a sharp peak in realized real interest rates of New Zealand. 1983Q4 is associated with a period of high Norwegian real interest rates as fiscal policy turned very expansionary in the early 1980s. The break date for Spain coincides with a 1.2% depreciation of the pesetas against the US dollar. During the third quarter of 1996, the low interest rates of Switzerland acted as a corrective to the excessive rise of the Swiss franc in the foreign exchange market. Finally, the break date for UK corresponds to a period of an increasing real interest rate. This increase in the real interest rate could be attributed to an expansionary fiscal shock, as the government's deficit (as a share of GDP) increases around this time.

3.2.4 Granger Causality Tests

Granger causality tests have been used frequently to investigate short run relationships among two or more variables of interest, including real and nominal interest rates in international markets (see, for example, Swanson, 1987). A high degree of causality from one rate to another indicates that the two markets are integrated and that rate changes in one market tend to lead rate changes in the other. Alternatively, causality may be bi-directional, indicating that interest rate changes in individual markets elicit significant responses in other markets. In each case, the evidence supports integration of the markets. The lead/lag relationships revealed by Granger tests also allow an evaluation of which markets may be dominant. For example, leadership roles have often been asserted for the US in world financial markets and for Germany in the European Monetary System (EMS). Researchers have

[20] The fourfold increase in the price of oil in late 1974 proved more damaging to Italy than the other major industrialized countries.

applied Granger causality tests to evaluate these questions. However, the literature that applies these tests suffers from many limitations.

First, these tests are far from the spirit of causality suggested by Granger (1980) in which 'causality' requires evidence of improved forecasts as a result of using the causal variable. Second, they may suffer from omitted variable(s) bias. As suggested by Granger (1980), if we are looking at causality relationships between two variables X_t and Y_t, a third variable Z_t might drive both X_t and Y_t at different lags. This might produce a finding of causality between X_t and Y_t even if true causality does not exist.[21] Third, these tests ignore cointegration relationships among the variables of interest. If X_t and Y_t are cointegrated of the form $Z_t=Y_t-AY_t$, then models that do not explicitly use Z_t will be misspecified and the possible value of lagged Y_t in forecasting X_t will be missed (Granger, 1980).

We now proceed to the causality test between real interest rates of each country and the US, both in the short run and in the long run. We first present this test only for those variables that are I(1) and are cointegrated. Specifically, Wald tests for Granger-causality in bivariate cointegrated finite order VAR processes are considered. Table 3.11 reports the results.

Table 3.11: Short-Term *Ex Post R*ates: Granger Causality Test /Block Exogeneity Wald Test

Dep. variable	**Excluded**		**Coef. of ECT (eq.3.8)**	
	D(RIR_US)	**D(RIR_FRANCE)**		
D(RIR_FRANCE)	0.3169 (0.5735)		-0.043 (-3.765)	
D(RIR_US)		2.0483 (0.1524)		0.105 (3.168)
	D(RIR_US)	**D(RIR_NETHERLANDS)**		
D(RIR_NETHERLANDS)	12.8307 (0.0016)		-0.007 (-0.445)	
D(RIR_US)		1.2222 (0.5428)		0.137 (3.885)
	D(RIR_US)	**D(RIR_SWITZERLAND)**		
D(RIR_SWITZERLAND)	10.5165 (0.0052)		-0.021 (-0.932)	
D(RIR_US)		1.7562 (0.4156)		0.030 (1.398)
	D(RIR_US)	**D(RIR_UK)**		

[21] An exception to this problem is found in Katsimbris and Miller(1993). It should also be noted that in-sample tests may be influenced by correlation among variables included in the system. For example, if three variables are highly correlated, it may be difficult to assign patterns of causality in standard tests.

D(RIR_UK)	0.0314 (0.8593)			-0.201 (-4.337)	
D(RIR_US)		0.3125 (0.5761)			0.026 (0.746)

Notes:
The first two rows under each variable indicate the chi-square value with the corresponding p-value in the parentheses. The last column reports the coefficients of the error correction terms for each country with the corresponding t-statistics in the parentheses.

Block Exogeneity Wald Test was used to test the joint significance of each of the other lagged endogenous variables in each equation and also to test for the joint significance of all the other lagged endogenous variables in each equation. A chi-square test statistics of 0.31 in the D(RIR_FRANCE) equation of Table 3.11 indicate that the null hypothesis that lagged coefficients of RIR_US being equal to zero cannot be rejected. The above table indicates that there is no causal linkage between the rates of France and the US. The same results hold for the rates of UK and the US. Moreover, the coefficients of the error correction term (EC_{t-1}) in the equation with real interest rates of France and the UK as the dependent variables, respectively, are negative and statistically significant at the five percent level. This finding is important, since it means that the independent variable is indeed causally related with the dependent variable in the Granger sense through these error correction terms. A statistically significant coefficient of the EC_{t-1} also suggests that economic forces (arbitrage and/or policy measures) are in operation to restore long-run equilibrium following a short-run disturbance. A significant positive coefficient (such as those for France and the Netherlands) means that whenever the actual value of the dependent variable (the US real interest rate) falls below the value consistent with its long-term equilibrium relationship, changes in the independent variables help bring it up to the long term equilibrium value, other things being equal. It is in this sense that the error-correction term provides an additional channel of causal relationship.

In contrast, real rates of the US Granger causes real rates of the Netherlands, but real rates of the Netherlands do not Granger cause real rates of the US. Moreover, the causal relationship flows from real rates of the US to real rates of Switzerland, but not vice versa. This finding is expected to hold, since the US is the biggest economy, and, hence, a US policy change leads to a policy change in these countries. In addition to this, the coefficients of the error correction terms for the Netherlands and Switzerland in the equation with real interest rates of the Netherlands and Switzerland as the dependent variables, respectively, are negative but not statistically significant.

We find similar results for short-term *ex ante* real interest rates and thus we do not report them.

Table 3.12: Long-Term *Ex Post* Rates: Granger Causality Test /Block Exogeneity Wald Test

Dep. variable	**Excluded**		**Coef. of ECT (eq.3.8)**	
	D(RIR_US)	**D(RIR_LUXEMBOURG)**		
D(RIR_LUXEMBOURG)	0.7679 (0.3808)		-0.099 (-4.597)	
D(RIR_US)		0.6084 (0.4354)		-0.001 (-0.03)

Notes:
The first two rows under each variable indicate the chi-square value with the corresponding p-value in the parentheses. The last column reports the coefficients of the error correction terms with the corresponding t-statistics in the parentheses.

The above table indicates that there is no causal relationship between the real rates of Luxembourg and the US. Furthermore, the coefficient of the error correction term (EC_{t-1}) in the equation with real interest rates of Luxembourg as the dependent variable is negative and statistically significant at the five percent level. This finding is important, since it means that the independent variable is indeed causally related with the dependent variable in the Granger sense through this error correction term.

We now proceed to the causality test between real interest rates of each country and the US, for those variables that are not cointegrated. F-tests along with associated P-values for the Granger causality test are presented in Tables 3.13 and 3.14.

Table 3.13: Granger Causality Test (Short-Term *Ex Post* Rates)

Null Hypothesis	**F-Statistic**	**Probability**
RIR_US does not Granger Cause **RIR_AUSTRALIA**	5.04224*	0.00774
RIR_AUSTRALIA does not Granger Cause **RIR_US**	0.16837	0.84522
RIR_US does not Granger Cause **RIR_BELGIUM**	11.8026*	1.9E-05
RIR_BELGIUM does not Granger Cause **RIR_US**	0.40010	0.67103
RIR_US does not Granger Cause **RIR_CANADA**	64.2799*	6.3E-20
RIR_CANADA does not Granger Cause **RIR_US**	6.70844*	0.00171
RIR_US does not Granger Cause **RIR_ITALY**	3.56729*	0.03275
RIR_ITALY does not Granger Cause **RIR_US**	2.72819	0.07135
RIR_US does not Granger Cause **RIR_JAPAN**	4.23663*	0.01614

RIR_JAPAN does not Granger Cause **RIR_US**	1.37595	0.25563
RIR_US does not Granger Cause **RIR_LUXEMBOURG**	4.00360*	0.02258
RIR_LUXEMBOURG does not Granger Cause **RIR_US**	0.10625	0.89934
RIR_ US does not Granger Cause **RIR_NEWZEALAND**	1.41560	0.24791
RIR_NEWZEALAND does not Granger Cause **RIR_US**	0.86511	0.42433
RIR_US does not Granger Cause **RIR_NORWAY**	0.79891	0.45261
RIR_NORWAY does not Granger Cause **RIR_US**	0.15687	0.85502
RIR_US does not Granger Cause **RIR_SPAIN**	1.65786	0.19531
RIR_SPAIN does not Granger Cause **RIR_US**	1.15055	0.32028

Notes:
*An asterisk indicates statistical significance at the α=0.05.

Our results above show that there are five unidirectional causality relationships from US real interest rate to Australian, Belgium, Italian, Japanese and Luxembourgish real interest rates, respectively. There is only one bidirectional causality relationship from US short-term real interest rate to Canadian real rate. Moreover, we find no causality relationships from US rate to rates of the rest of the countries examined. It should be noted that similar results were found for the *ex ante* real interest rates and this is the reason we do not report them.

Table 3.14: Granger Causality Test (Long-Term *Ex Post* Rates)

Null Hypothesis	F-Statistic	Probability
RIR_US does not Granger Cause **RIR_AUSTRALIA**	4.21404*	0.01649
RIR_AUSTRALIA does not Granger Cause **RIR_US**	0.04909	0.95211
RIR_US does not Granger Cause **RIR_BELGIUM**	8.87224*	0.00022
RIR_BELGIUM does not Granger Cause **RIR_US**	1.80232	0.16831
RIR_US does not Granger Cause **RIR_CANADA**	1.04656	0.35358
RIR_CANADA does not Granger Cause **RIR_US**	0.00692	0.99310
RIR_US does not Granger Cause **RIR_FRANCE**	3.33264*	0.03824
RIR_FRANCE does not Granger Cause **RIR_US**	2.96766	0.05432
RIR_US does not Granger Cause **RIR_ITALY**	8.02537*	0.00048
RIR_ITALY does not Granger Cause **RIR_US**	0.50447	0.60480
RIR_US does not Granger Cause **RIR_JAPAN**	0.55700	0.57405
RIR_JAPAN does not Granger Cause **RIR_US**	2.16040	0.11869
RIR_US does not Granger Cause **RIR_NETHERLANDS**	7.91638*	0.00053

RIR_NETHERLANDS does not Granger Cause **RIR_US**	1.33869	0.26537
RIR_US does not Granger Cause **RIR_NEWZEALAND**	4.58252*	0.01163
RIR_NEWZEALAND does not Granger Cause **RIR_US**	0.34373	0.70965
RIR_US does not Granger Cause **RIR_NORWAY**	6.39567*	0.00214
RIR_NORWAY does not Granger Cause **RIR_US**	0.34994	0.70528
RIR_US does not Granger Cause **RIR_SPAIN**	1.82563	0.16588
RIR_SPAIN does not Granger Cause **RIR_US**	1.47987	0.23208
RIR_US does not Granger Cause **RIR_SWITZERLAND**	0.96260	0.38414
RIR_SWITZERLAND does not Granger Cause **RIR_US**	1.10740	0.33298
RIR_US does not Granger Cause **RIR_UK**	4.47075*	0.01293
RIR_UK does not Granger Cause **RIR_US**	1.49873	0.22659

Notes:
*An asterisk indicates statistical significance at the α=0.05.

Table 3.14 reports that there are eight unidirectional causality relationships from US real interest rate to Australian, Belgium, French, Italian, Netherlands, New Zealand, Norway and the UK real interest rates, respectively. In addition, there are not any causal linkages between the US real interest rate and the Canadian real rates. The finding of no causality relationship is the same for Japan, Spain and Switzerland. The results for long-term *ex ante* real interest rates are not reported, since they are similar to those for *ex post* rates.

3.2.5 Half-Life Measurement

For long-run RIP to hold the real interest differential should be a zero mean stationary process. The stationarity of real interest rate differentials can be verified by performing unit root tests on these differentials to determine whether they contain a unit root or not. However, if unit root is rejected, but the true value of the dominant root is close to unity, shocks will be slow to dissipate, and this stationary process may not be significantly different from a true unit root process in the economic sense (Sekioua, 2004). Consequently, the emphasis should not be on whether real interest rate differentials have a unit root, it should instead be on measuring the economic implications of their behaviour. What market participants care about is the degree of persistence in the real interest differential. One measure of persistence that has received a lot of attention in the empirical literature is the half-life. The half-life is defined as the number of periods it takes for deviations to subside permanently below 50% in response to a unit shock in the level of the real interest differential. It is computed because it essentially provides a measure of the degree of mean-reversion.

Before measuring half-life of deviations from RIP, it is important to determine what constitutes a reasonable range for this measure of persistence (*i.e.* a range consistent with RIP). Unfortunately, unlike the vast literature on PPP, there is no consensus that we can base our analysis on. Consequently, we must look at the predictions of macroeconomic models that embody the RIP hypothesis. For example, models of exchange rate determination developed by Frenkel (1976) assume real interest rate equality. Others, such as Dornbusch's (1976) overshooting model, predict that sticky goods prices would cause real interest rates to diverge across countries. If the failure of RIP is attributed to stickiness in nominal prices, then presumably we would expect substantial convergence to RIP over 12 to 24 months (4 to 8 quarters), as prices adjust to shocks. In fact, this theoretical range for the half-life estimates of price convergence is supported by Cheung *et al.* (2003) who found that these estimates are substantially short, between 12 and 24 months. Clearly, an estimate for the half-life that is less than 12 months (4 quarters) is also consistent with RIP since it implies rapid adjustment of real interest rate differentials. Therefore, our range would have an upper bound of 24 months (8 quarters), but any value less than this is obviously acceptable.

By imposing the restriction (a,b)=(0,1) in Eq. (3.2) we obtain a model for the RIRD model:

$$r_t - r_t^* = \varepsilon_t \qquad (3.14)$$

Given the specification in (3.14), RIP is said to hold in the long-run if the residuals ε_t is mean reverting. Suppose that the deviations of the RIRD series (ε_t) from its long-run value (ε_0) follows an AR(1) process, then:

$$\varepsilon_t - \varepsilon_0 = a(\varepsilon_{t-1} - \varepsilon_0) + \mu_t \qquad (3.15)$$

where μ_t is white noise. Hence, we continue our analysis by reporting in tables 3.15 and 3.16 the half-life estimates which are computed using the following equation:

t

$$\alpha^h = 0.5 \Rightarrow h_{0.5} = \ln(0.5)/\ln(\alpha) \qquad (3.16)$$

Table 3.15: Half-Lives

Country	Half-life (Quarters) (short-term *ex post* rates)	Half-life (Quarters) (short-term *ex ante* rates)
France	6.1	6.1
Netherlands	9.4	9.5
Switzerland	9.2	9.2
UK	3.4	3.4

Table 3.15 reports that the point estimates of the half life are 6.1 quarters for France, 9.4 quarters for the Netherlands, 9.2 quarters for Switzerland and 3.4 quarters for the UK. These estimates are supportive of reversion towards parity since they are all within or slightly above our benchmark which has an upper bound of 8 quarters. The lower estimates include a range of short half-lives, which can be much less than 4 quarters and, in the case of the UK, less than this benchmark. The most persistent real interest differentials, according to our results, are those of the Netherlands and Switzerland. In all, the point estimates are supportive of RIP and are consistent with the results of the unit root tests[22]. We find the same estimates of the half-life for short-term *ex ante* rates.

Table 3.16: Half-Lives

Country	Half-life (Quarters) (long-term *ex post* rates)	Half-life (Quarters) (long-term *ex ante* rates)
Luxembourg	9.0	8.9

Table 3.16 indicates that the point estimate of the half-life for Luxembourgish long-term *ex post* rates is 9 quarters, while the point estimate of the half-life for long-term *ex* ante rate is 8.9 quarters. These estimates are supportive of reversion towards parity since they are slightly above our benchmark which has an upper bound of 8 quarters.

[22] If a unit root is present, then evidently we would expect deviations never to die out and the half-life to be infinity.

3.2.6 Impulse Response Analysis

Impulse response analysis has become a common tool for investigating the interrelationship among the variables in dynamic models. Pesaran and Shin (1998) and references therein provide excellent discussions of impulse response analysis. ‘Impulse responses’ represent time path responses of variables to exogenous shocks to variables in a VAR system. In this analysis, impulse responses are utilized to evaluate the extent and nature of market integration. If two markets are integrated then an exogenous shock to real interest rate in one market should evoke an equilibrating response to real interest rate in the other market.

Consider the following undifferenced VAR of VECM :

$$X_t = B_1 X_{t-1} + ... + B_p X_{t-k} + \varepsilon_t \quad (3.17)$$

The innovation accounting from Eq. (3.17) can be used to obtain information concerning the interactions among the variables. As noted by Pesaran and Shin (1998), if the process (3.17) is stationary, forecast error impulse responses are the coefficients of moving average representation given as follows:

$$X_t = \sum_{i=0}^{\infty} \Phi_i \varepsilon_{t-i} \qquad (3.18)$$

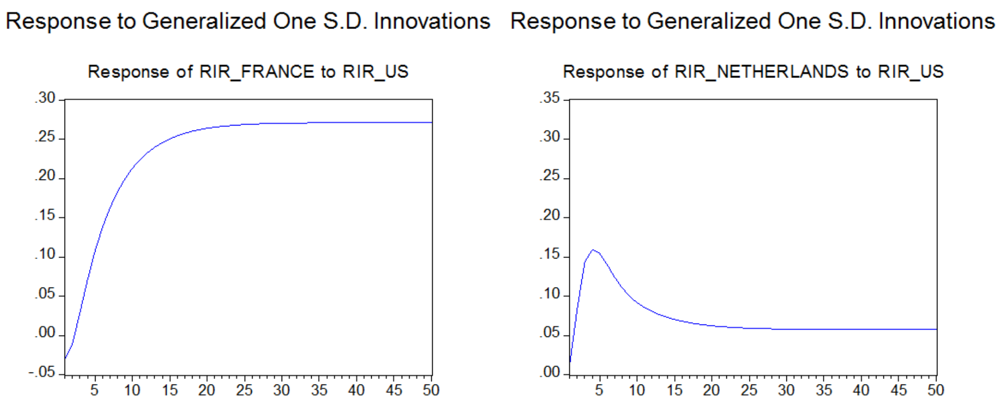
In the context of our analysis, the coefficient Φ_i can be interpreted as the response of an interest rate in any of the industrialized countries to a shock in say the interest rate in US, i period ago. In order to estimate the impulse responses in (3.18), two approaches are commonly used in the literature, namely: the generalized impulse response, as proposed by Pesaran and Shin (1998), and the Cholesky decomposition proposed by Sims (1980). The main advantage of the former approach over the Cholesky decomposition method is that it does not require orthogolization of innovations and is invariant of the ordering of the variables in the VAR (Pesaran and Shin, 1998). However, its application is based on the assumption that the shocks in the different interest rates are contemporaneously correlated; if the shocks are uncorrelated, then the two methods will coincide (Pesaran and Shin, 1998). In addition, following De Bondt (2005), our VAR model was estimated in levels. This has the advantage that it maximizes the long-term information in the data set and delivers super-consistent coefficient estimates (De Bondt, 2005). Whereas, as noted by De Bondt (2005), if inappropriate cointegration relations were imposed they could lead to biased estimates, which in turn might bias the impulse responses derived from the reduced form VARs. Finally, an appropriate VAR order was chosen using the Schwartz information criterion.

We subject our cointegrated VAR system to impulse response analysis, as follows. The generalized impulse is applied to both the short-term real interest rate (*ex post* and *ex ante*) and the cointegrating vector. We found in section 3.2.2 that there exists a cointegrating relationship between real interest rates of France and the US, of the Netherlands and the US, of Switzerland and the US and of UK and the US. Furthermore, the GI is applied to both the long-term real interest rate (*ex post* and *ex ante*) and the cointegrating vector. We found earlier in our analysis that there is one cointegrating relationship between Luxembourgish and US real interest rates and this is where we apply GI.

In a cointegrating VAR system, the impact of shocks on the individual variables is expected not to die out in the long run, or, equivalently, the variables will not return to their initial values if no further shocks occur.

Figure 10 displays the generalized impulse responses of each country to an interest rate shock in the US. The US is chosen as the source of an interest rate shock because it is viewed as the main trading partner of the countries involved and we want to determine the sensitivity of other country rates to US inspired shocks. The plot shows the dynamic response of the real interest rate to these shocks after zero periods, one period, two periods, and up to a limit of fifty periods. The time paths of the response to shocks confirm that the real interest rate might not revert to its pre-shock equilibrium. It should be noted that the plot is the same for the *ex post* and *ex ante* rates and for this reason we report only the one for *ex post* rate.

Figure 10: Generalized Impulse Responses (for short-term rates)



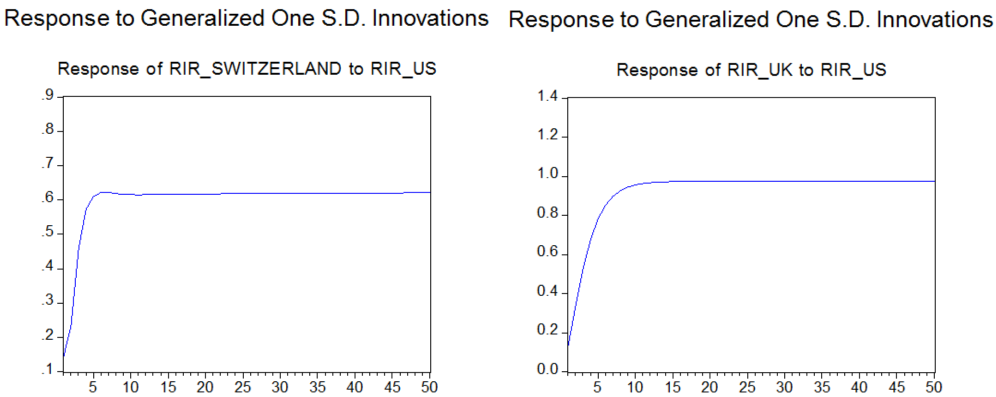
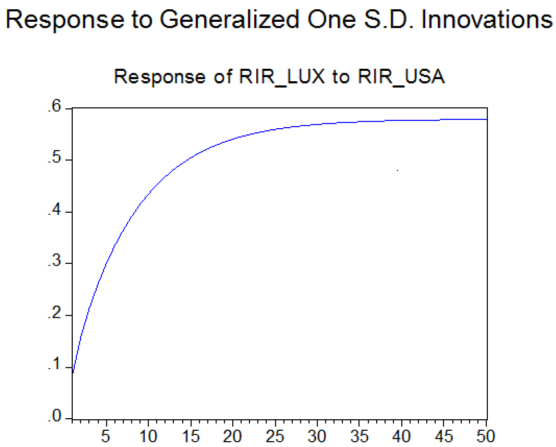


Figure 11 displays the generalized impulse responses of Luxembourg to an interest rate shock in the US. The plot shows the dynamic response of the real interest rate to this shock after zero periods, one period, two periods, and up to a limit of fifty periods and shows that the real interest rate might not revert to its pre-shock equilibrium. The plot is the same for the *ex post* and *ex ante* rates and for this reason we report only the one for *ex post* rate.

Figure 11: Generalized Impulse Responses (for long-term rates)



3.2.7 International Linkages in Real Long-Term Rates

The next step is to consider the evidence on international linkages of real long-term rates as a whole. The question posed is what evidence there is for interest rate changes in one country leading to changes elsewhere. This is a question of possible causal links between interest rates in different countries and, to investigate this, we first need to establish whether international interest rates as a whole cointegrate. The idea behind this is that economic variables, such as real interest rates, although they may individually be non-stationary, may nevertheless be related to each other over the long run. This relationship may be identified by using cointegration methods. Such methods allow us to discover if non-stationary variables are linearly related in the long run, with a stationary error. If so, they cointegrate. If this appears true in the present case of international interest rates, we can go on to determine whether there are any regular causal patterns between world interest rates revealed in this relationship. Hence, the first thing we need to consider is whether international real interest rates cointegrate. If we find they do, then additional questions can be considered. For the purpose in hand, one of the most relevant is whether any of the world's long-term real rates appear to have a causal effect on others. This caution of causality is a statistical one; we seek evidence of whether changes in one country's real interest rate regularly precede changes in another country's. Closely related to this concept is the idea that one country's interest rate may be weekly exogenous, or unaffected by movements in other country's rates.

First, we rely on the direct measure of long-term *ex post* real interest rates. Then, we test for the presence of cointegration between real long-term interest rates of the twelve countries used in Table 2.6 (Luxembourg and Spain were excluded, because of their different sample period). This exclusion has minor effect on the result, as these two countries are small in size, in comparison to the others .Second, we test the causality structure implied by these cointegrating relationships. All real long rates are found to be I(1) (see Table 2.14). The following table gives details of the tests. The selection of the lag length of the VAR is done, using the SC.

Table 3.17: *Ex Post* Real Long Rates: Johansen Cointegration Tests

H_o	λ_{trace}	5% critical value (trace)	$\lambda_{max.eigen.}$	5% critical value (max.eigen .)
r=0	429.7894*	374.9076	106.9288*	80.87025
r=1	322.8606*	322.0692	79.30144*	74.83748
r=2	253.5592	273.1889	51.49482	68.81206
r=3	202.0644	228.2979	41.28351	62.75215
r=4	160.7809	187.4701	32.37187	56.70519
r=5	128.4090	150.5585	31.18430	50.59985
r=6	97.22468	117.7082	22.87465	44.49720

r=7	74.35004	88.80380	20.27893	38.33101
r=8	54.07111	63.87610	17.64036	32.11832
r=9	36.43075	42.91525	14.85269	25.82321
r=10	21.57806	25.87211	12.99788	19.38704
r=11	8.580175	12.51798	8.580175	12.51798

Notes:
* indicates rejection of the null at 5 per cent. VAR lags:1

Table 3.17 reports the existence of two cointegrating relations, when testing for the presence of cointegration between real interest rates in the twelve countries. These findings of the cointegration suggest the presence of long-run relationships between the real interest rates of all the twelve countries concerned.

The following test we report concerns the possibility of there being causal linkages between the long rates of different countries. This uses the finding of cointegration to explore what the statistical evidence for possible linkages between the interest rates of the different countries then is. It tests for the importance of the cointegrating vectors just identified in significantly affecting the behavior of each of the country real interest rates. Where they are not significant, this is prima facie evidence that the interest rate of the country concerned is “weakly exogenous”, meaning that its interest rate does not respond to changes in the relationship captured in the cointegrating vector, and so appears to move independently as a random walk. In contrast, where the cointegrating vector is significant, it implies that the interest rate concerned is not weakly exogenous, meaning that it is determined by the other weakly exogenous rates.

Table 3.18: *Ex Post* Real Long Rates: Causality Tests

Rank	**Australia**	**Belgium**	**Canada**	**France**	**Italy**	**Japan**
r=2	16.6396 (0.1190)	41.6970* (0.0000)	8.7845 (0.6418)	25.1969* (0.0085)	58.1724* (0.0000)	11.8334 (0.3763)
	Netherlands	**New Zealand**	**Norway**	**Switzerland**	**UK**	**US**
	38.3615* (0.0001)	36.6874* (0.0001)	29.4394* (0.0019)	7.4339 (0.7629)	44.7809* (0.0000)	10.3177 (0.5021)

Notes:
Figures in brackets are probability values; * indicates rejection of the null at 5 per cent.

The above table indicates that Australian, Canadian, Japanese, Switzerland and US interest rates are weakly exogenous. These results can be interpreted as showing that changes in real long rates in these countries have effects on real long rates in other countries, but there do not appear to be effects in the reverse direction. As it is clear, we find that it is hard to argue that real interest rates are converging to a single world rate.

Now we do the same work for *ex ante* real interest rates, following the procedure used for *ex post* rates. All real long rates are found to be I(1) (results are not reported).The following table gives details of the tests. The selection of the lag length of the VAR is done, using the SC.

Table 3.19: *Ex Ante* Real Long Rates: Johansen Cointegration Tests

H_o	λ_{trace}	**5% critical value (trace)**	$\lambda_{max.eigen.}$	**5% critical value (max.eigen .)**
r=0	427.0537*	374.9076	105.7016*	80.87025
r=1	321.3520	322.0692	66.42810	74.83748
r=2	254.9239	273.1889	51.05742	68.81206
r=3	203.8665	228.2979	40.53625	62.75215
r=4	163.3303	187.4701	32.75582	56.70519
r=5	130.5745	150.5585	30.70063	50.59985
r=6	99.87382	117.7082	24.30081	44.49720
r=7	75.57301	88.80380	22.31035	38.33101
r=8	53.26266	63.87610	17.26778	32.11832
r=9	35.99488	42.91525	14.58595	25.82321
r=10	21.40893	25.87211	12.57943	19.38704
r=11	8.829504	12.51798	8.829504	12.51798

Notes:
* indicates rejection of the null at 5 per cent. VAR lags:1

The above table reports the existence of one cointegrating relation, when testing for the presence of cointegration between real interest rates in the twelve countries. These findings of the cointegration suggest the presence of long-run relationship between the real interest rates of all the twelve countries concerned. This finding is different from the one found for *ex post* rates, as for *ex post* rates we found the existence of two instead of one cointegrating relations.

We now proceed to the following test which concerns the possibility of there being causal linkages between the long rates of different countries.

Table 3.20: *Ex Ante* Real Long Rates: Causality Tests

Rank	Australia	Belgium	Canada	France	Italy	Japan
r=1	27.10641* (0.0044)	68.83157* (0.0000)	11.87126 (0.3734)	51.96001* (0.0000)	70.48057* (0.0000)	22.52116* (0.0206)
	Netherlands	**New Zealand**	**Norway**	**Switzerland**	**UK**	**US**
	34.99362* (0.0002)	72.16759* (0.0000)	43.0498* (0.0000)	34.54149* (0.0003)	22.83052* (0.0187)	13.19467 (0.2808)

Notes:
Figures in brackets are probability values; * indicates rejection of the null at 5 per cent.

The above table indicates that Canadian and US interest rates are weakly exogenous. As it is clear, we also find that it is hard to argue that real interest rates are converging to a single world rate, when *ex ante* interest rates are used.

As it is clear from this section, real yields, either calculated on an indirect *ex ante* basis or observed directly from bonds, have in fact been different between countries over different times. Moreover, it is hard to argue that real interest rates are converging to a single world rate.

CHAPTER 4: NONLINEAR TESTS OF REAL INTEREST RATE PARITY

4.1 INTRODUCTION

Although the equality of two countries' real interest rates, as described in Equation (3.2), would indicate that any difference between a domestic and a foreign real interest rate is arbitraged away, there exist several factors that may alter this parity. For instance, transaction costs may influence the dynamic of real interest rate equalization, even in a well-integrated and efficient international capital market. Goodwin and Grennes (1994) pointed out that there exists a neutral band due to transactions costs such that when the difference between real interest rates is less than transactions costs, arbitrage is not profitable and thus transactions will not take place. If, however, the real interest rate differential exceeds transactions costs, then arbitrage will quickly eliminate a disparity between these rates[23].

Another plausible explanation for deviations from the RIP is the stickiness of product prices. Previous studies find deviations from purchasing power parity (PPP) to be highly persistent in the short run. However, PPP holds better in the long run with slow reversion to the parity[24]. Note that a sufficient condition for real interest parity to hold is that both the uncovered interest parity (UIP) and the relative purchasing power parity hold. While UIP relates to financial integration driven by arbitrage between money and foreign exchange markets, the relative PPP pertains to how easily goods and services are arbitraged. Hence, the violation of purchasing power parity due to a limited strength of the forces that equilibrate goods prices will result in a breakdown of RIP. Taking the possibilities of frictions in commodity trade into account, Michael, Nobay and Peel (1997) find evidence supporting a transactions band and hence nonlinear adjustments of deviations from PPP toward the parity. Since the real exchange rate and the real interest rate differentials are associated through the theory of exchange rate determination, a nonlinear behavior of real exchange rate to restore the PPP equilibrium would imply the nonlinear adjustments of deviations from RIP.

---

[23] Balke and Wohar (1998) suggest the use of nonlinear frameworks when market frictions, namely transactions costs, exist.

[24] The half-lives of deviations from PPP are found to be approximately 3-5 years. See, for instance, Frankel and Rose (1996).

Previous literature on the RIP condition based its analysis on the assumption of a linear relationship among real interest rates such that the deviations from the parity are assumed to converge to the long-run equilibrium at a constant rate regardless of how far the process is from the mean. Although the real interest rate may be highly volatile and persistent, a long-run arbitrage equilibrium in financial markets would suggest that the individual rates do not wander arbitrarily far apart. Generally, the RIP condition is examined by the use of the cointegration approach as in Goodwin and Grennes (1994), and Fountas and Wu (1999). The cointegration tests provide a way to determine whether interest rates across countries have a common stochastic trend. One can think of home and foreign interest rates as being attracted to each other through the force of arbitrage, and that the short-term deviations from this relationship represent error corrections. They find that the differences between two real interest rates are a stationary autoregressive process that reverts back to the long-run equilibrium and thus the RIP hypothesis holds. The violation of the RIP, however, due to sluggish and costly adjustments of deviations from equilibrium in both asset and goods markets, may imply nonlinear dynamics between real interest rates (Pipatchaipoom and Norrbin, 2008). Responses to shocks that cause deviations in real interest rates may depend on the magnitude of the shocks such that larger shocks will evoke quicker adjustments to restore the parity than smaller deviations from the parity.

A failure to recognize nonlinearity in real interest rate adjustments could lead to inaccurate statistical inferences in the conventional tests for international linkages. Granger and Teräsvirta (1999) have given an example of a simple nonlinear model that can mimic a linear, long-memory series. If the true model is nonlinear, then applying the linear model estimation could result in biases. The biases are even larger when the size of a threshold band increases. Moreover, conventional unit root tests, like the Dickey-Fuller (DF) unit root test, become even more powerless to reject a unit root than they already are when nonlinearity is present. The DF unit root test tends to fail to distinguish between a nonstationary linear process and the nonlinear mean-reverting one.

4.2 EMPIRICAL RESULTS

As mentioned above, research has found evidence supporting gradual regime-switching behavior of real interest rate adjustments due to the existence of transactions costs. There may be no response of domestic rate to foreign rate changes when the deviation between the real rates is small due to transactions costs. However, when shocks to both rates are large enough to make arbitrage profitable, this may evoke quick adjustments to restore the parity. In this section we investigate nonlinear adjustments of real interest rates toward the long-run equilibrium using an approximation of the smooth transition autoregressive (STAR) models. The findings indicate that there exist nonlinearities in real interest rate adjustment.

We first apply the Ramsey Regression Equation Specification Error Test (RESET) test, which is a general specification test for the linear regression model. More specifically, it tests whether non-linear combinations of the estimated values help explain the exogenous variable. The intuition behind the test is that, if non-linear combinations of the explanatory variables have any power in explaining the exogenous variable, then the model is mis-specified. Tables 4.1 and 4.2 report the results from the Ramsey Reset test with one parameter level for short-term and long-term real interest rates, respectively. It should be noted that the results obtained for higher powers of the predicted values of the dependent variable were similar to those presented in the following tables.

Table 4.1: Ramsey Reset Test (for short-term *ex post*[25] real rates)

Country	F-Statistic	Probability
Australia	21.8519	0.0000
Belgium	0.9900	0.3214
Canada	0.0078	0.9299
France	4.7289	0.0101
Italy	2.0439	0.1565
Japan	6.5889	0.0112
Luxembourg	12.6489	0.0000
Netherlands	4.2334	0.0421
New Zealand	8.1600	0.0052

[25] Similar results were obtained for *ex ante* real interest rates and, thus, we do not report them.

Norway	17.8071	0.0001
Spain	6.5594	0.0118
Switzerland	0.0041	0.9490
UK	14.3322	0.0002

Notes:
Number of fitted terms:1

The results from the table above show that we reject the null hypothesis that the true specification is linear (which implies that the true specification is non-linear) for Australia, France, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Spain and the UK, since the F-statistics are greater than the F critical values in these cases (or, alternatively, the probability value is lower than 0.05). In contrast, we are unable to reject the null hypothesis for Belgium, Canada, Italy and Switzerland. In these cases, the results suggest that the true specification is linear.

Table 4.2: Ramsey Reset Test (for long-term *ex post*[26] real rates)

Country	F-Statistic	Probability
Australia	8.6449	0.0003
Belgium	4.1376	0.0177
Canada	0.2591	0.6115
France	0.0015	0.9687
Italy	0.3250	0.5694
Japan	18.3639	0.0000
Luxembourg	19.7966	0.0000
Netherlands	6.8315	0.0098
New Zealand	3.8248	0.0522
Norway	7.8233	0.0006
Spain	18.2313	0.0000
Switzerland	24.3137	0.0000
UK	21.1662	0.0000

Notes:
Number of fitted terms:1

Table 4.2 indicates that we reject the null hypothesis that the true specification is linear (which implies that the true specification is non-linear) for Australia, Belgium, Japan, Luxembourg, the Netherlands, Norway, Spain, Switzerland and the UK, since the F-statistics are greater than the F critical values in these cases (or,

[26] Similar results were obtained for *ex ante* real interest rates and, thus, we do not report them.

alternatively, the probability value is lower than 0.05). In contrast, we are unable to reject the null hypothesis for Canada, France, Italy and New Zealand, which means that the true specification in these cases is linear.

Furthermore, in order to show that that RIP follows a nonlinear stationary process, we tested for the existence of the long-run relationship using the Kapetanois et al. (2003) (hereafter, KSS) test. In this test, the United States is still used as the centre country. To examine whether a time series is linear or nonlinear in nature, one may use the following linearity test frameworks due to Teräsvirta, 1994:

$$y_t = \alpha_0 + \sum_{i=1}^{p} \alpha_i y_{t-i} + \sum_{i=1}^{p} (b_{1i} y_{t-i} y_{t-d} + b_{2i} y_{t-i} y_{t-d}^2 + b_{3i} y_{t-i} y_{t-d}^3) + \xi_t \qquad (4.1)$$

where a_i and b_i (i=0, 1, . . . , p) are linear and nonlinear autoregressive parameters respectively; p and d are known as optimal lag length and delay parameter respectively; and ξ_t is white noise residuals with zero mean and constant variance under the null hypothesis of linearity, in which all b's are simultaneously zero.

The null hypothesis may be tested against the alternative hypothesis of nonlinearity (at least one b is not zero) by the F-test (Teräsvirta, 1994). Following the suggestions of KSS, lag length (*p*) is determined using the significance procedure as outlined in Ng and Perron (1995).

To show that real interest rate differential is in fact stationary, KSS propose the following tests to cater for the testing of unit root in the presence of nonlinearity[27]:

$$\Delta y_t = \delta y_{t-1}^3 + error \qquad (4.2)$$

where y_t is the nonlinear time series of interest. Besides, KSS also suggest the following framework to correct for plausible serially correlation errors:

$$\Delta y_t = \sum_{j=1}^{p} \rho_j \Delta y_{t-j} + \delta y_{t-1}^3 + error \qquad (4.3)$$

In both cases, the null hypothesis to be tested is H_0: δ=0 against the alternative H_1: δ>0. We refer to the test given by (4.2) and (4.3) as KSS(A) and KSS(B) hereafter. Table 4.3 presents the results of the KSS tests for the US pairs when short-term *ex post* real interest rates are used.

[27] The test is obtained using the first-difference approximation of the ESTAR model.

Table 4.3: Nonlinear Unit Root Test Result

Short-Term Ex Post Real Interest Rate	US-based KSS(A)	Lag	KSS(B)
Australia	-1.665	2	-1.804
Belgium	-3.381*	4	-3.742*
Canada	-5.118*	3	-2.453
France	-4.102*	4	-4.932*
Italy	-2.089	2	-4.483*
Japan	-3.014*	4	-3.545*
Luxembourg	-2.905	1	-2.393
Netherlands	-3.571*	1	-4.596*
New Zealand	-0.474	3	-0.442
Norway	-2.236	1	-3.094*
Spain	-2.549	2	-2.929
Switzerland	-3.958*	1	-3.585*
UK	-3.696*	2	-3.926*

Notes
:KSS(A) and KSS(B) denote KSS test as specified in Equation (4.2) and (4.3) respectively. The 5 percent asymptotic null critical value for both KSS tests is –2.93.Asterisks*denotes rejection of the unit roots at the 5% significance level.

As can be observed from Table 4.3, the null of a unit root was easily rejected against the nonlinear stationary alternative for most of the cases. It turns out that Australian, Luxembourgish, New Zealand and Spanish interest rates failed to reject the null at the 5% significance level by both the KSS(A) and KSS(B) tests. In these cases, different types of nonlinearity may render the adjustments to equilibrium. Thus, it appears that all the remaining countries are integrated with the major financial market, namely, the US. Similar results are obtained for short-term *ex post* real interest rates and, hence, we do not report them.

Hence, the above findings demonstrate the problem with using linear unit root tests as reported in earlier studies, that is, they tend to reject the stationary null in favour of the alternative hypothesis. Specifically, the classical linear unit root tests are not capable of rejecting the null hypothesis in the presence of nonlinearities in the adjustment process because they lack the power. Similar observations are made in Holmes and Maghrebi (2004) using nonlinear cointegration tests for the OECD countries. In their study, there is strong evidence in favor of the RIP as real interest rate differentials display non-linear mean reversion when using both the US as the base country.

Table 4.4 presents the results for KSS when long-term *ex post* real interest rates are used.

Table 4.4: Nonlinear Unit Root Test Result

Long-Term Ex Post Real Interest Rate	**US-based**		
	KSS(A)	**Lag**	**KSS(B)**
Australia	-1.878	2	-2.129
Belgium	-2.524	4	-2.577
Canada	-2.148	1	-2.284
France	-3.141*	5	-3.262*
Italy	-1.958	2	-2.491
Japan	-2.082	2	-2.320
Luxembourg	-2.294	2	-2.462
Netherlands	-3.262*	1	-3.652*
New Zealand	-1.226	2	-1.363
Norway	-1.527	4	-2.007
Spain	-1.901	2	-2.251
Switzerland	-1.908	4	-2.232
UK	-2.430	1	-2.516

Notes:
KSS(A) and KSS(B) denote KSS test as specified in Equation (4.2) and (4.3) respectively. The 5 percent asymptotic null critical value for both KSS tests is –2.93.Asterisks*denotes rejection of the unit roots at the 5% significance level.

Table 4.4 indicates that the null of a unit root was not rejected against the nonlinear stationary alternative for most of the cases. It turns out that only interest rates of France and the Netherlands reject the null at the 5% significance level by both the KSS(A) and KSS(B) tests. In other words, we found that the hypothesis of real interest rate convergence is rejected after allowing for nonlinearity in the real interest rate adjustments in all but two countries, France and the Netherlands. We obtain similar results for long-term *ex ante* real interest rates and, hence, we do not report them.

CHAPTER 5: CONCLUSION

Previous studies on real interest rate parity (RIP) provide inconclusive results as to whether or not real interest rates across countries are connected. One would expect that deviations between international real interest rates would lead to arbitrage trading and hence the parity condition would be restored quickly in the well-integrated financial markets. However, empirical findings do not necessarily agree with the theoretical prediction. The early RIP literature assumed that real rates were stationary, and thus used standard regression techniques to test whether the computed real interest rate in one country was closely linked with another country's real interest rate. Mostly these tests provided very limited evidence for real interest rate parity. More recent tests have allowed for the possibility of nonstationary time series process of the real interest rates. Therefore, they have examined a potential common long-run relationship between two random walk real interest rate series. Such studies generally found evidence of mean-reverting real interest rate differentials and thus supported the validity of long-run RIP hypothesis. Cumby and Mishkin (1986), Fraser and Taylor (1990), and Chung and Crowder (2004) are examples of studies that reject the validity of the RIP hypothesis, where as Goodwin and Grennes (1994), Moosa and Bhatti (1996), and Holmes and Maghrebi (2004) support the RIP relation. More sophisticated testing methodologies have been proposed to reconcile the empirical results with the RIP theory. However, the recent methodologies do not resolve the conflicting results.

In this study, we presented evidence on the RIP hypothesis for a sample of industrialized countries for the period that spans from the beginning of 1967 until the end of 2008. The sample period was slightly different for each country. We investigated the existence of *ex post* and *ex ante* real interest rate differentials in Australia, Belgium, Canada, France, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway using the US as the reference large economy.

According to the Fisher equation, the *ex ante* real interest rates can be defined as the difference of nominal rate of returns on a bond and the expected rate of inflation at the time the bond matures. To obtain the inflationary forecasts, which are not directly observed from the data, authors have to impose assumptions on how economic agents form their expectations. It is possible that different ways of measuring the expected rates of inflation may have significant impacts on the derived real interest rates and hence the RIP testing. We presented briefly the different methods of constructing real interest rates that have been used in previous literature and applied them to various types of RIP testing, such as linear regression testing, cointegration testing, and testing for nonlinear adjustments. The results were similar for both approaches.

Specifically, the unit root tests provided that each country's real interest rates were I(1). Therefore, since the real interest rates were found stationary in their first differences, we continued our analysis with the cointegration techniques (Engle-Granger and Johansen's cointegration tests). The results of the linear regression tests were robust to both approaches. However, the estimated coefficients were quite different, indicating some sensitivity to the underlying methods of construction of the real rates. The results of the Engle-Granger cointegration test indicated rejection of RIP for most of the countries in both approaches.

The results of the Johansen cointegration test indicated that the RIP condition does not hold between most of the countries examined and were consistent with those found in Cumby and Mishkin (1986) and Mishkin (1984). However, the results obtained for the short-term real interest rates of the Netherlands, Switzerland and the UK supported the weak version of the RIP hypothesis for the first two countries and the strong version of the RIP for the latter. These findings were the same for both approaches. Moreover, the results for France were mixed, since those obtained for the short-term *ex post* real rates supported the strong form of the RIP, while those obtained for the short-term *ex ante* real rates supported the weak form of it. The results obtained for the long-term real interest rates of Luxembourg supported the strong version of the RIP hypothesis for both approaches.

We also used cointegration tests that determine endogenously the regime shift. In other words, we explored the stability of the real interest rates with the Gregory-Hansen techniques. When short-term rates were used, we found evidence (at 5% level) for cointegration between US rates and rates in Canada, Italy, the Netherlands, New Zealand, Norway, Spain, Switzerland and the UK. For long-term real rates, we found that the null of a lack of cointegration was rejected (at 5%) for Australia, Belgium, Canada, France, Japan, Luxemburg, New Zealand, Norway, Spain and Switzerland.

Overall, we found that the results differ depending on the type of tests used: standard cointegration tests do not support the hypothesis of real interest rate convergence whereas tests that determine endogenously potential structural breaks imply that real interest rate convergence has taken place in most of the industrialized countries, particularly for long-term real rates. These results have important implications for the effectiveness of domestic stabilization policies. In particular, for those countries where real long-term interest rate convergence applies, domestic monetary policy would be expected to have lost some of its effectiveness as a long-run stabilization policy tool.

We proceeded to the Granger causality test between real interest rates of each country and the US and found unidirectional causality relationships for several industrialized countries. The finding that the causal relationship flows from real rates of the US to real rates of the other industrialized country, but not vice versa was expected to hold, due to the large size of the US economy and its outstanding role of activity.

Moreover, the estimates of the half-lives were supportive of reversion towards parity in all cases. We applied the generalized impulse to both short-term and long-term real interest rates. The time paths of the response to shocks confirmed that the real interest rate might not revert to its pre-shock equilibrium.

The last essay allowed for possible nonlinearities in international real interest rate dynamics. When an equilibrium between two real interest rates is disturbed, the adjustments toward the parity will occur only if the differences between the rates are large enough to compensate for transactions costs of trading. When such deviations are small, arbitrage trading does not occur and thus there is no tendency for real rates to revert back to their parity relation. Therefore, the adjustments of the difference of real interest rates are not linear. This nonlinear behavior was captured by an application of approximation of the smooth transition autoregressive (STAR) framework. The results seemed to indicate the existence of nonlinearities in most of the real interest rate differentials.

REFERENCES

- Adam, K., Japelli, T., Menichini, A., Padula, M. and Pagano M. (2002), “Analyze, Compare, and Apply Alternative Indicators and Monitoring Methodologies to Measure the Evolution of Capital Market Integration in the European Union”, *Report to the European Commission*
- Al-Awad, M.A. and Goodwin, B. (1998), “ Dynamic Linkages Among Real Interest Rates in International Capital Markets”, *Journal of International Money and Finance,* 17, pp. 881-907
- Al-Awad, M.A. and Grennes, T.J. (2002), “Real Interest Parity and Transactions Costs for the Group of 10 Countries”, *International Review of Economics and Finance*, 11, pp.363-372
- 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*, 3(7), pp.277-287
- Allen, S.D., (1990), “The Effect of Federal Deficits and Debt on The Tax-Adjusted, Short-Term, Real Interest Rate”, *Economics Letters*, 34 (2), pp.169-173
- Baharumshah, A. Z., Hawa, C. T., and Fountas, S. (2005), “A panel Study on Real Interest Rate Parity in East Asian Countries: Pre- and Post-Liberalization Era”, *Global Finance Journal*, 16, pp. 69-85
- Balke, N. and Wohar, M. (1998), “Nonlinear Dynamics and Covered Interest Rate Parity”, *Empirical Economics,*23, pp. 535-559
- Barassi M. R., Caporale, G. M. and Hall, S.G. (2000), “Interest Rates Linkages: Identifying Structural Relations”, *Discussion Paper*, no. 2000, *Center for International Macroeconomics*
- Benninga, S. and Protopapadakis, A. (1983), “Real and Nominal Interest Rates under Uncertainty: The Fisher Theorem and the Term Structure”, *Journal of Political Economy*, 91, pp. 856-867
- Cavaglia, S. (1992), “The Persistence of Real Interest Differentials”, *Journal of Monetary Economics*, 3(29), pp. 429-443

- Chen, L.H., (2001), "A Model for Ex Ante Real Interest Rates", *Applied Economics Letters*, 8, pp. 713-718
- Cheung, Y-W., Chinn, M.D. and Fujii, E. (2003), "China, Hong Kong and Taiwan: A Quantitative Assessment of Real and Financial Integration", *China Economic Review*, 14, pp. 281-303
- Chung, S.W. and Crowder, W. (2003), "Why Are Real Interest Rates Not Equalized Internationally?", *University of Texas at Arlington, working paper, July*
- Chung, S.Y. and Crowder, W.J. (2004), "Why Are Real Interest Rates Not Equalized Internationally?" *Southern Economic Journal*, 71, pp. 441-58
- Coakley, J. and Fuertes, A.M. (2002), "Asymmetric Dynamics in UK Real Interest Rates", *Applied Financial Economics,* 12, pp. 397-387
- Cover, J. (1992), "Asymmetric Effects of Positive and Negative Money Supply Shocks", *Quarterly Journal of Economics,* 107(4), pp. 1261-1282
- Crowder, W. and Hoffman, D. (1996), "The Long-Run Relationship between Nominal Interest Rates and Inflation: the Fisher Equation Revisited", *Journal of Money, Credit and Banking*, 28(1), pp. 102-118
- Cumby, R.E. and Mishkin, F.S. (1984), "The International Linkage of Real Interest Rates: The European-U.S. Connection", *NBER working paper,* no. 1423, pp.1-35
- Cumby, R.E. and Mishkin, F.S. (1986), " The International Linkage of Real Interest Rates: the European-US Connection", *Journal of International Money and Finance,* 5, pp.5-23
- 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
- De Bondt, G. (2005), "Interest Rate Pass-Through: Empirical Results for the Euro Area", *German Economic* Review, 6, pp. 37-78
- Doornik, J. and Hendry, D. (1994), "Pc-Give 8.0: An Interactive Econometric Modelling System", *London: International Thompson Publishing*
- Dornbusch, R. (1976), "Expectations and Exchange Rate Dynamics", *Journal of Political Economy*, 84, pp. 1161-1176

- Dumas, B. (1992), "Dynamic Equilibrium and the Real Exchange Rate in a Spatially Separated World", *The Review of Financial Studies*, 5, pp. 153–180
- 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, pp. 165-187
- Elmendorf, D.W. and Mankiw, G.N. (1998), "Government Debt", *NBER working paper*, no. 6470, pp.1-94
- Elliot, G., Rothenberg, T.J. and Stock, J.H. (1996), "Efficient Tests for an Autoregressive Unit Root", *Econometrica*, 64, pp. 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, pp.304-311
- Enders, W. and Siklos, P.L. (2001), "Cointegration and Threshold Adjustment", *Journal of Business and Economic Statistics*, 19, pp. 166-176
- Evans, M. and Lewis, K. (1995), "Do Expected Shifts in Inflation Affect Estimates of the Long-Run Fisher Relation?", *The Journal of Finance*, 50, pp. 225-253
- Fama, E. (1977), "Interest Rates and Inflation: the Message in the Entrails", *American Economic Review*, 67(3), pp. 487-496
- Feldstein, M. and Horioka, C. (1980), "Domestic Saving and International Capital Flows", *The Economic Journal*, 90, pp. 314-329
- Ferreira, A.L. (2004), *University of Kent, working paper, August* "Does the Real Interest Parity Hypothesis Hold? Evidence for Developed and Emerging Markets",
- Ferreira, A.L. and León-Ledesma, M.A. (2007), "Does the Real Interest Parity Hypothesis Hold? Evidence for Developed and Emerging Markets", *Journal of International Money and Finance*, 26, pp. 364-382
- Fountas, S. and Wu, J.-L. (1999), "Testing for Real Interest Rate Convergence in European Countries", *Scottish Journal of Political Economy*, 46(2), pp. 158-174
- 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, pp.1-23

- Frankel, J.A. and Rose, A.K. (1996), “A Panel Project on Purchasing Power Parity: Mean Reversion within and between Countries”, *Journal of International Economics,* 40, pp. 209-224
- Fraser, P. and Taylor, M.P. (1990), “Some Efficient Tests of International Real Interest Rate Parity”, *Applied Economics*, 22, pp. 1083-1092
- Frenkel, J.A. (1976), “A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence”, *Scandinavian Journal of Economics*, 78, pp. 200-224
- Fujii, E. and Chinn, M.D. (2000), “Fin de Siecle Real Interest Parity”, *NBER working paper,* 7880, pp. 1-38
- Gagnon, J.E. and Unferth, M.D. (1995), “ Is There a World Real Interest Rate?”, *Journal of International Money and Finance*, 14, pp.845-855
- Goldberg, L.G., Lothian, J.R. and Okunev, J. (2003), “Has International Financial Integration Increased?”, *Open Economies Review*, 14, pp. 299-317
- Goodhart, C.A.E. (1999), “Central Bankers and Uncertainty”, *Bank of England Quarterly Bulletin*, 39, pp. 102–114
- Goodwin, B.K. and Grennes, T.J. (1994), “Real Interest Rate Equalization and the Integration of International Financial Markets”, *Journal of International Money and Finance*, 13, pp.107-124
- Granger, C.W.J. (1980), “Testing for Causality: a Personal Viewpoint”, *Journal of Economic Dynamics and Control*, 2, pp. 329-352
- Granger, C.W.J. and Newbold, P. (1974), “Spurious Regression in Econometrics”, *Journal of Econometrics*, 2, pp. 111-120
- Granger, C.W.J and Teräsvirta, T. (1999), “A Simple Nonlinear Time Series Model with Misleading Linear Properties”, *Economics Letters*, 62,pp. 161-165
- Gregory, A. and Hansen, B. (1996), “Residual-Based Tests for Cointegration in Models with Regime Shifts”, *Journal of Econometrics*, 70, pp. 99-126
- Haldane, A.G. and Pradhan, M. (1992), “Testing Real Interest Parity in the European Monetary System”, *Bank of England*
- Haug, A., MacKinnon, J. and Michelis, L. (2000), “European Monetary Union: a Cointegration Analysis”, *Journal of International Money and Finance*, 19, pp. 419-432

- Holmes, M. and Maghrebi, N. (2004), “ Asian Real Interest Rates, Nonlinear Dynamics and International Parity”, *International Review of Economics and Finance,* 13, pp. 387-405
- Holmes, M. and Wang, P. (2008), “Real Convergence and the EU Accession Countries: A New Perspective on Real Interest Parity”, *Journal of Emerging Market Finance,* 7(3), pp. 215-236
- Huizinga, J. and Mishkin, F. (1984), “Inflation and Real Interest Rates on Assets with Different Risk Characteristics”, *Journal of Finance,* 39(3), pp.699-712
- Im, K., Pesaran, H. and Shin, Y. (1997) "Testing for Unit Roots in Heterogeneous Panels”, *Discussion Paper, University of Cambridge, June.*
- Jenkins, M. and Madzharova, P. (2007), “Real Interest Rate Convergence Under the Euro”, *Applied Economics Letters,* 15, pp.473-476
- Johansen, S. (1988), “Statistical Analysis of Cointegrating Vectors”, *Journal of Economic Dynamics and Control,* 12, pp. 231-254
- Johansen, S. (1991), “Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models”, *Econometrica,* 59, pp. 205-230
- Jorion, P. (1996), “Does Real Interest Parity Hold at Longer Maturities?”, *Journal of International Economics,* 40, pp. 105-126
- Kaminsky, G.L. and Leiderman, L. (1998), “High Real Interest Rates in the Aftermath of Disinflation: Is it a Lack of Credibility?”, *Journal of Development Economics,* 55, pp. 189-212
- Kapetanois, G., Shin, Y. and Snell, A. (2003), “Testing for a Unit Root in the Nonlinear STAR Framework”, *Journal of Econometrics,* 112, pp. 359-379
- Katsimbris, G.M. and Miller, S.M. (1993), “Interest Rate Linkages within the European Monetary System: Further Analysis”, *Journal of Money, Credit and Banking,* 25, pp.771-779
- Kim, D.H., Osborn, D. and Sensier M. (2005), “Nonlinearity in the Fed’s Monetary Policy Rule”, *Journal of Applied Econometrics, in press*
- Kugler, P. and Neusser, K. (1993), “International Real Interest Rate Equalization: A Multivariate Time-Series Approach”, *Journal of Applied Econometrics,* 8, pp. 163-174

- Kwiatkowski, D., Phillips, P.C.B., Schmidt, P. and Shin, Y. (1992), “Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure Are We that Economic Time Series Have a Unit Root?”, *Journal of Econometrics,* 54, pp. 159-178
- Liew, V. K. S., Baharumshah, A. Z. and Lau, E. (2004), “Nonlinear Adjustment towards Purchasing Power Parity: Empirical Evidence from ASEAN Exchange Rates”, *ICFAI Journal of Applied Economics,* 3, pp. 7–18
- Lothian, J.R. (2000), “Capital Market Integration and Exchange-Rate Regimes in Historical Perspective”, *Research in Banking and Finance,* 1, pp. 141-176
- Maddala, G.S. and Kim, I.M. (1998), “Unit Roots, Cointegration and Structural Change”, *Cambridge University Press, Cambridge*
- Mancuso, A., Goodwin, B. and Grennes, T. (2003), “ Nonlinear Aspects of Capital Market Integration and Red al Interest Rate Equalization”, *International Review of Economics and Finance,* 12, pp. 283-303
- Mark, N. (1985), “A Note on International Real Interest Rate Differentials”, *Review of Economics and Statistics,* 67(4), pp. 681-684
- Meese, R. and Rogoff, K. (1988), “Was it Real? The Exchange Rate-Interest Differential Relation over the Modern Floating Rate Period”, *Journal of Finance,* 43, pp. 933-948
- Michael, P., Nobay, R. and Peel, D. (1997), “Transactions Costs and Nonlinear Adjustment in Real Exchange Rates: An Empirical Investigation”, *Journal of Political Economy,* 105, pp. 862-879
- Mishkin, F.S. (1984), “Are Real Interest Rates Equal Across Countries? An Empirical Investigation of International Parity Conditions”, *The Journal of Finance*, 39 (5), pp. 1345-1357
- Mishkin, F.S. (1992), “Is the Fisher Effect for Real? A Re-Examination of the Relationship between Inflation and Interest Rates”, *Journal of Monetary Economics*, 30, pp. 195-215
- Moosa, I. and Bhatti, R. (1996), “Some Evidence on Mean Reversion in Ex Ante Real Interest Rates”, *Scottish Journal of Political Economy,* 43(2), pp. 177-191

- Mussa, M. (1986), “Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Applications”, *Carnegie Rochester Series on Public Policy,* 25, pp.117-213
- Nakayawa, H. (2002), “Real Exchange Rates and Real Interest Rate Differentials: Implications of Nonlinear Adjustment in Real Exchange Rates”, *Journal of Monetary Economics,* 49, pp.629-649
- Neftci, S. N. (1984), “Are Economic Time Series Asymmetric Over the Business Cycles?”, *Journal of Political Economy,* 85, pp. 281-291
- Ng, S. and Perron, P. (1995), “Unit Root Tests in ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag”, *Journal of the American Statistical Association,* 90, pp.268-281
- Ng, S. and Perron, P. (2001), “Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power”, *Econometrica,* 69, pp. 1519-1554
- Obstfeld, M. and Rogoff, K. (2000), “The Six Major Puzzles in International Finance: Is there a Common Cause?” *NBER Macroeconomics Annual,* 15, pp. 339-390
- Obstfeld, M. and Taylor, A.M. (1997), “Nonlinear Aspects of Goods-Market Arbitrage and Adjustment: Heckscher’s Commodity Points Revisited”, *National Bureau of Economic Research, working paper, June*
- Obstfeld, M. and Taylor, A.M. (2002), “Globalization and Capital Markets”, *NBER working paper,* no. 8846, pp. 1-69
- Perron, P. (1989), “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis”, *Econometrica,* 57(6), pp. 1361-1401
- Perron, P. and Ng, S. (1996), “Useful Modifications to Some Unit Root Tests with Dependent Errors and their Local Asymptotic Properties”, *Reviews of Economic Studies,* 63, pp. 435-463
- Pesaran, M. and Shin, Y. (1998), “Generalized Impulse Response Analysis in Linear Multivariate Models”, *Economics Letters,* 58, pp. 17–29
- Phylaktis, K. (1999), “Capital Market Integration in the Pacific Region: An Impulse Response Analysis”, *Journal of International Money and Finance,* 18, pp.267-287

- Pipatchaipoom, O. and Norrbin, S. (2008), "Is the Real Interest Rate Parity Condition Affected by the Method of Calculating Real Interest Rates?", *Applied Economics,* pp.1466-4283
- Ramsey, J.B. and Rothman, P. (1996), "Time Irreversibility and Business Cycle Asymmetry", *Journal of Money, Credit and Banking*, 28, pp. 1-21
- Roll, R. (1979), "Violation of Purchasing Power and Their Implications for Efficient International Commodity Markets" in Sanat, M. and Szego, G.P. (eds), *International Finance and Trade*, vol.1, Cambridge, Mass: Ballinger, pp.133-176
- Rose, A. (1988), "Is the Real Interest Rate Stable?", *Journal of Finance,* 43, pp. 1095-1112
- Scholnick, B. (1996), "Asymmetric Adjustment of Commercial Bank Interest Rates: Evidence from Malaysia and Singapore", *Journal of International Money and Finance,*15, pp. 485-496
- Seddighi H., Lawler, K. and Katos, V. (2000), "Econometrics: A Practical Approach", *Routledge, London*
- Sekioua, S.H. (2004a), "Real Interest Parity (RIP) over the 20th century: New Evidence Based on Confidence Intervals for the Dominant Root and Half-Lives of Shocks", *Money Macro and Finance (MMF) Research Group Conference 2004*, 91, Money Macro and Finance Research Group, pp.1-49
- Sekioua, S.H. (2004b), "What are the Determinants of Reversion Towards Real Interest Parity (RIP)?", mimeo, Warwick Business School, pp.1-31
- Shiller, R. and Perron, P. (1985), "Testing the Random Walk Hypothesis Power versus Frequency of Observation", *Economics Letters,* 18(4), pp. 381-386
- Siklos, P. and Wohar, M. (1997), "Convergence in Interest Rates and Inflation Rates across Countries and over Time", *Review of International Economics,* 5(1), pp. 129-141
- Sims, C. (1980), "Macroeconomics and reality", *Econometrica,* 48, pp. 1–48
- Swanson, P., (1987), "Capital Market Integration over the Past Decade: the Case of the US Dollar", *Journal of International Money and Finance*, 6, pp.215-225

- Taylor, A. (2001), “Potential Pitfalls for the Purchasing Power Parity Puzzle? Sampling and Specification Biases in Mean-Reversion Tests of the Law of One Price”, *Econometrica,* 69(2), pp. 473-498
- Taylor, M.P., Peel, D.A. and Sarno, L. (2001), “Nonlinear Mean-Reversion in Real Exchange Rates: Towards a Solution to the Purchasing Power Parity Puzzles”, *International Economic Review,* 42, pp. 1015–1042
- Teräsvirta, T. (1994), “Specification, Estimation, and Evaluation of Smooth Transition Autoregressive Models”, *Journal of the American Statistical Association,* 89, pp. 208-218
- Venetis, I.A., Duarte, A. and Paya, I. (2004), “The Long Memory Story of Ex Post Real Interest Rates. Can it be Supported?”, *Econometrics 0404004, Economics working paper archive at WUSTL*
- Wu, J.-L. and Chen, S.-L. (1998), “A Re- Examination of Real Interest Rate Parity”, *Canadian Journal of Economics,* 31(4), pp. 837-851
- Wu, J.-L. and Chen, S.-L. (2001), “Mean Reversion of Interest Rates in the Eurocurrency Market”, *Oxford Bulletin of Economics and Statistics*, 63(4), pp. 459-473
- Wu, J.-L. and Fountas, S. (2000), “ Real Interest Rate Parity Under Regime Shifts and Implications for Monetary Policy”, *The Manchester School,* 68 (6), pp. 685-700
- Zhou, S. (2003), “Interest Rate Linkages within the European Monetary System: New Evidence Incorporating Long-Run Trends”, *Journal of International Money and Finance*, 22, pp.571-590
- Zivot, E. and Andrews, D. (1992), “Further Evidence on the Great Crash, the Oil-Price Stock, and the Unit Root Hypothesis”, *Journal of Business and Economic Statistics*, 10, pp. 251-270

APPENDIX

Table A1: KPSS Unit Root Test For Short-Term Ex Post Real Interest Rates

Country	**KPSS**	**Critical value 5%**
Level		
Australia	0.248763* (T)	0.146000
Belgium	0.298891* (T)	0.146000
Canada	0.120202 (T)	0.146000
France	0.278987* (T)	0.146000
Italy	0.151125* (T)	0.146000
Japan	0.140588 (T)	0.146000
Luxemburg	0.089693 (T)	0.146000
Netherlands	0.110314 (T)	0.146000
New Zealand	0.116436 (T)	0.146000
Norway	0.176397* (T)	0.146000
Spain	0.075194 (T)	0.146000
Switzerland	0.093141 (T)	0.146000
UK	0.260160	0.463000
US	0.206541* (T)	0.146000
1st difference		
Australia	0.089272 (T)	0.146000
Belgium	0.025958 (T)	0.146000
France	0.104934 (T)	0.146000
Italy	0.094224 (T)	0.146000
Norway	0.066232 (T)	0.146000
US	0.037938 (T)	0.146000

Notes:
* indicates rejection of the null at 5%. (T) indicates the trend is included as indicated by the significant of the trend terms in the estimation.

Table A2: DF-GLS Unit Root Test For Short-Term Ex Post Real Interest Rates

Country	**DF-GLS**	**Critical value 5%**
Level		
Australia	-1.597694 (T)	-2.991000
Belgium	-2.579078 (T)	-2.991000
Canada	-1.874463 (T)	-3.000000
France	-0.879600 (T)	-2.967000
Italy	-1.691939 (T)	-3.074800
Japan	-2.667353 (T)	-2.968000
Luxemburg	-1.583986 (T)	-3.106800
Netherlands	-1.125645 (T)	-3.023000
New Zealand	-0.806627 (T)	-3.042800
Norway	-1.684983 (T)	-3.025000
Spain	-3.006610 (T)	-3.016000
Switzerland	-2.846489 (T)	-3.024000
UK	-1.172868	-1.943385
US	-2.886511 (T)	-2.974000
1st difference		
Australia	-7.659993* (T)	-2.991000
Belgium	-8.091525* (T)	-2.989000
Canada	-15.17642* (T)	-3.000000
France	-11.04011* (T)	-2.968000
Italy	-6.333794* (T)	-3.074800
Japan	-8.779100* (T)	-2.968000
Luxemburg	-8.144036* (T)	-3.110000
Netherlands	-8.900547* (T)	-3.024000
New Zealand	-7.133186* (T)	-3.042800
Norway	-4.245700* (T)	-3.025000
Spain	-8.491661* (T)	-3.016000
Switzerland	-3.784661* (T)	-3.025000
UK	-8.438556*	-1.943385
US	-3.249997* (T)	-2.972000

Notes:
* Implies significance at 5%. (T) indicates the trend is included as indicated by the significant of the trend terms in the estimation.

Table A3: KPSS Unit Root Test For Long-Term Ex Post Real Interest Rates

Country	KPSS	Critical value 5%
Level		
Australia	0.340553* (T)	0.146000
Belgium	0.317547* (T)	0.146000
Canada	0.322994* (T)	0.146000
France	0.299382* (T)	0.146000
Italy	0.327214* (T)	0.146000
Japan	0.200098* (T)	0.146000
Luxemburg	0.286132* (T)	0.146000
Netherlands	0.251945* (T)	0.146000
New Zealand	0.335579	0.463000
Norway	0.348412* (T)	0.146000
Spain	0.139698 (T)	0.146000
Switzerland	0.141196 (T)	0.146000
UK	0.282568* (T)	0.146000
US	0.266399* (T)	0.146000
1st difference		
Australia	0.069485 (T)	0.146000
Belgium	0.063302 (T)	0.146000
Canada	0.043789 (T)	0.146000
France	0.068428 (T)	0.146000
Italy	0.066130 (T)	0.146000
Japan	0.051577 (T)	0.146000
Luxembourg	0.067018 (T)	0.146000
Netherlands	0.041781 (T)	0.146000
Norway	0.137536 (T)	0.146000
UK	0.074227 (T)	0.146000
US	0.042754 (T)	0.146000

Notes:
* Indicates rejection of the null at 5%. (T) indicates the trend is included as indicated by the significant of the trend terms in the estimation.

Table A4: DF-GLS Unit Root Test For Long-Term Ex Post Real Interest Rates

Country	DF-GLS	Critical value 5%
Level		
Australia	-1.123300 (T)	-2.967000
Belgium	-1.402411 (T)	-2.968000
Canada	-1.381306 (T)	-2.967000
France	-1.526833 (T)	-2.968000
Italy	-1.370225 (T)	-2.968000
Japan	-1.911060 (T)	-2.968000
Luxemburg	-1.742002 (T)	-2.980000
Netherlands	-2.081424 (T)	-2.970000
New Zealand	-1.284406	-1.942805
Norway	-1.103917 (T)	-2.968000
Spain	-2.326258 (T)	-3.013000
Switzerland	-2.011546 (T)	-2.960000
UK	-1.109007 (T)	-2.967000
US	-1.880574 (T)	-2.970000
1st difference		
Australia	-11.01743* (T)	-2.968000
Belgium	-7.780794* (T)	-2.968000
Canada	-11.37008* (T)	-2.968000
France	-7.671089* (T)	-2.968000
Italy	-7.208107* (T)	-2.968000
Japan	-10.58362* (T)	-2.968000
Luxemburg	-10.10320* (T)	-2.980000
Netherlands	-5.651366* (T)	-2.970000
New Zealand	-10.60714*	-1.942805
Norway	-8.874388* (T)	-2.968000
Spain	-6.562006* (T)	-3.013000
Switzerland	-8.345616* (T)	-2.968000
UK	-10.74279* (T)	-2.968000
US	-5.874725* (T)	-2.970000

Notes:

* Implies significance at 5%. (T) indicates the trend is included as indicated by the significant of the trend terms in the estimation.

CUMULATIVE BIBLIOGRAPHY

ARTICLE	WRITER	COUNTRIES	DATA	TECHNIQUES/ METHODOLOGY	RESULTS
✧ Are Real Interest Rates Equal Across Countries? An Empirical Investigation of International Parity Conditions (1984)	Frederic S. Mishkin	OECD countries: the US, Canada, the UK, France, West Germany, the Netherlands, Switzerland	Three-month euro deposit rates, CPI, WPI, 1967Q2-1979Q2	Linear regressions, tests of equality	• Strongly rejects the hypothesis of the equality or real euro rates
✧ The international linkage of real interest rates: the European-US connection (1986)	Robert Cumby and Frederic Mishkin	Canada, Italy, the Netherlands, France, West Germany, the UK, the US	Euro deposit and domestic money markets, CPI, from 1973M6-1983M12	Linear regressions of *ex post* and *ex ante* real interest rates	• Statistical association between real rates in nearly all pairs of countries
✧ . International Real Interest Rate Equalization: A Multivariate Time-Series Approach (1993)	Peter Kugler and Klaus Neusser	OECD countries	Monthly ex post real interest rates, CPI, 1980-1991	Technique of codependence developed by Gouriéroux and Peaucelle	• RIP holds in the long-run but not in the short-run
✧ Real interest rate equalization and the integration of international financial markets (1994)	Barry K. Goodwin and Thomas J. Grennes	US, Canada, UK, Belgium, Italy, Germany, the Netherlands, Switzerland, Japan, France	Monthly 90-day Eurocurrency rates, 90-91 day t-Bill rates, Jan. 1975-Feb. 1987	Bivariate and multivariate cointegration tests (Engle and Granger, Johansen)	• Support for RIP
✧ Does real interest parity hold at longer maturities? (1996)	Philippe Jorion	US, Britain and Germany	Monthly observations, Treasury Bill rates for 3-month maturities, CPI, WPI, 1973-1991	Assuming rational expectations, differences in unobservable ex ante real rates are expressed as a linear projection: $[i_t^{m}-\pi_{t+m}^{m}]-[i_t^{*m}-\pi_{t+m}^{*}]=\alpha_m+\beta_m(i_t^{m}-i_t^{m*})+\varepsilon_t$	• RIP was rejected • Across these countries and sample periods, variations in nominal interest rate differentials seem to mirror variations in real interest rate differentials at all horizons.
✧ Some Evidence on Mean Reversion in Ex Ante Real Interest Rates (1996)	Imad A. Moosa and Razzaque H. Bhatti	12 major industrial countries	Three-month treasury Bill rates, CPI, 1972Q1-1993Q3	Alternative unit root tests	• Goods, capital and foreign exchange markets have become highly integrated
✧ Dynamic linkages among real interest rates in international capital markets (1998)	Mouawiya Al Awad, Barry K. Goodwin	G-10 countries: US, Canada, UK, Belgium, France, Germany, Italy, the Netherlands, Switzerland, Japan.	Weekly Eurocurrency rates on 3-month and 12-month bonds, monthly CPI, 1976-1994	Cointegration tests (Johansen), out of sample Granger causality tests.	• Strong evidence that real interest rates among the G-10 countries are linked both in the short-run and the long-run. • Real interest equalization is rejected
✧ A re-examination of real interest rate	Jyh-Lin Wu and Show-Lin Chen	9 OECD countries	Monthly observations on Euro-market rates,	Three-panel-based unit root tests	• Support for RIP

parity (1998)			1979-1996		
✧ Testing for real interest rate convergence in European countries (1999)	Stilianos Fountas and Jyh-lin Wu	8 European countries: Belgium, Denmark, France, Germany (base country), Ireland, Italy, the Netherlands, UK.	Short-term (3-month Eurocurrency rate) and long-term (the government bond yield) interest rates, CPI, Quarterly data, 1979-1993	Cointegration tests that allow for structural shifts in the cointegrating vector (Engle and Granger methodology)	• Real interest rate convergence has taken place in several European countries, particularly for long-term real rates.
✧ Real interest rate parity under regime shifts and implications for monetary policy (2000)	Jyh-lin Wu and Stilianos Fountas	G7: Canada, France, Germany, Italy, Japan, UK, USA	Short-term (overnight money market rate and the TB rate) and long-term (government bond yield) interest rates, CPI, Quarterly data, 1973Q2-1995Q1	Use of developed cointegration techniques (Engle and Granger methodology) that allow for structural shifts in the cointegrating vector. $r_t=\alpha+bD_t+cr_t^*+d(Trend)+u_t$ $t=1,\ldots,n$	• The results differ depending on the type of tests used. • Standard cointegration tests support the hypothesis of real interest rate convergence only for Germany (short-term rates) • Cointegration tests that determine potential structural breaks endogenously imply the existence of bilateral real interest rate convergence
✧ Fin de Siècle real interest parity (2000)	Eiji Fujii, Menzie Chinn	G-7 countries	Short-term interest rates (3, 6 and 12 month maturity eurocurrency yields), long-horizon interest rate data, CPI,WPI, Quarterly frequency, 1976Q1-2000Q1	Tested the RIP hypothesis first by assuming that expectations are rational and then by using time series forecasts of future inflation rates	• RIP holds better at long horizons than at short • RIP cannot be rejected for the selected G-7 countries at a 5 and/or a 10 year horizon
✧ Real exchange rates and real interest differentials: implications of nonlinear adjustment in real exchange rates (2002)	Hironobu Nakagawa	US, Germany, Japan, Canada, UK	Long-term interest rates(5-10 year government bonds), short-term rates (3-month interbank rates), CPI, Monthly and quarterly observations for the period 1974.1-1997.12 and 1974Q1-1997Q4, respsctively	Incorporates nonlinear real exchange rate adjustment into the Mundell-Fleming-Dornbusch model	• Results support the evidence of the desired link between real exchange rates and real interest differentials.
✧ Nonlinear aspects of capital market integration and real interest rate equalization (2003)	Anthony J. Mancuso, Barry K. Goodwin, Thomas J. Grennes	Major G-5 countries (US, UK, Canada, Germany, Switzerland, Japan)	12- and 3- month nominal Eurocurrency rates, monthly CPI, 1979-2001	Basic TAR model, fully flexible nonparametric version of the TAR model	• Strong indications of nonlinearities in patterns of adjustment • The results provide mixed evidence for real interest rate

					equality • The results imply that extreme deviations among rates do tend to provoke equilibrating responses and that these adjustments tend to occur faster as the shocks are more extreme.
✧ Are international real interest rate linkages characterized by asymmetric adjustments? (2004)	Mark J. Holmes, Nabil Maghrebi	9 OECD countries: Belgium, Canada, France, Germany, Italy, Japan, Netherlands, UK, US	3-month Treasury bill rates, CPI, Monthly data, June 1973-Jan. 2004	Unit root test (DFGLS), method of nonlinear cointegration (TAR and MTAR models)	• There is stronger evidence of long-run cointegrating relationships when an explicit distinction is made between decreasing and increasing deviations from equilibrium.
✧ Real interest parity (RIP)over the 20th century: New evidence based on confidence intervals for the dominant root and half-lives of shocks. (2004)	Sofiane H. Sekioua	UK, Japan, France relative to the US.	Monthly long-term government bond yields and CPI, 1923-2000 (sub-periods: the interwar 1923-1938, the Bretton-Woods fixed exchange rate period 1950-1973, the recent floating rate experience 1974-2000.	Unit root tests (DF-GLS) Confidence intervals for the dominant root: -grid bootstrap method, -median unbiased estimation (MUE) method.	• The results are, on the whole, supportive of reversion towards parity • Apart from the volatile wartime period, RIRPs appear to be uniform across nominal exchange rate regimes, especially Bretton Woods and the recent float→support for nominal exchange rate neutrality
✧ Does the Real Interest Parity Hypothesis Hold? Evidence for Developed and Emerging Markets (2004)	Alex Luiz Ferreira	Emerging markets: Argentina, Brazil, Chile, Mexico and Turkey Developed countries: France, Italy, Spain, the UK and Germany	Treasury Bill rates, deposit rates, CPI, 1995M3-2002M5	Unit root tests, TAR representation	• Mean reversion especially for the emerging markets • High degree of market integration • Existence of asymmetries
✧ Why Are Real Interest Rates Not Equalized Internationally? (2004)	S. Young Chung, William J. Crowder	US, UK ,Canada, Germany, Japan	Monthly observations on 12-month Eurocurrency deposit rates, Feb. 1960-April 1996	Cointegration technique (Johansen)	• RIP does not hold.
✧ The Long Memory Story of Ex Post Real Interest Rates. Can it be Supported (2004)	Ioannis A. Venetis, Agustin Duarte and Ivan Paya	14 European countries (Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain, Sweden, the UK) and the US	3-month money rates, CPI, 1973-1999	Non-tapered and tapered versions of the GPH estimator are applied on ex-post real interest rate series	• Stationarity of real interest rates differentials is not independent of the time lag structure. • Most European countries show higher speed of real rates equalization with Germany rather than the US.

✧ A panel study on real interest rate parity in East Asian countries:Pre-and post-liberalization era (2005)	A.Z.Baharumshah, Tze Haw, S.Fountas	10 Asian countries(Indonesia, Philippines, Sri Lanka, Taiwan, Thailand, South Korea, Hong Kong, Singapore, Malaysia) Base country:Japan	Quarterly,1977-2001(3 sub-periods)	Panel unit root tests (LL, HT), half life measurement	• RIP holds strongly between Japan and Asian emerging countries. • Deviations from RIP have a half-life of 6-7 months.
✧ Does the real interest parity hypothesis hold? Evidence for developed and emerging markets (2007)	Alex Luiz Ferreira, Miguel A. León-Ledesma	Emerging markets: Argentina, Brazil, Chile, Mexico and Turkey Developed countries: France, Germany, Italy, Spain, UK, US.	Monthly data, Treasury Bill Rate, CPI, 1975.5-2003.8 (δεν συμπ. high inflation years)	Unit root tests, TAR representation	• High mean reversion especially for the emerging markets (short-run) • Reject the unit root hypothesis for all countries using different methods evidence supporting the existence of a positive long-run mean in the rids of, especially, emerging markets
✧ Real interest rate convergence under the euro (2007)	Michael A. Jenkins and Petya Madzharova	15 EU member states	Monthly observations of the HICP and interest on government bonds with 10 years maturity, Jan. 1999-Dec.2004	ADF, cointegration technique (Johansen), Pedroni panel cointegration tests	• RIP does not hold in the post euro period.
✧ Is the real interest rate parity condition affected by the method of calculating real interest rates? (2008)	Onsurang Pipatchaipoom and Stefan C. Norrbin	Four OECD countries: Japan, Switzerland, the UK, the US.	3-month Eurocurrency deposit rates, monthly CPI, 1978.9-2004.7	Linear method, stationarity tests, cointegration techniques (Johansen), STAR methodology	• Using linear methods, the RIRP is soundly rejected. • Different results between methods of computing real interest rates. • Using nonlinear method→ RIRP hypothesis holds.
✧ Real Convergence and the EU Accession Countries: A New Perspective on Real Interest Parity (2008)	Mark J. Holmes and Ping Wang	10 European countries that joined the EU on 1 May 2004: Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia and Slovenia. The real interest rate differentials are defined with respect to US, UK and Germany.	Monthly observations on three deposit rates, CPI, July 1993-Dec.2005	Unit root tests (NP, DF-GLS, KPSS) , panel unit root tests (Im , Pesaran and Shin, 2003), SURADF	• Univariate unit root testing indicates the general absence of real convergence • The panel tests indicate that RIP holds in the majority of the cases.