TESTING FOR REAL INTEREST RATE CONVERGENCE OVER THE LONG-RUN: SOME FURTHER EVIDENCE

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Abstract

This study presents additional evidence on the convergence speeds of real interest rate differentials. Using median unbiased estimation and impulse response analysis, we estimate the speeds at which deviations from real interest parity (RIP) die out. Moreover, since reporting only point estimates provides an incomplete picture of the speed of convergence towards RIP, median unbiased confidence intervals are also computed. Our results show that the dynamics of deviations from parity, albeit rather persistent, exhibit mean-reversion. This implies that the domestic authorities still have some scope to alter real economic activity through the real interest rate channel.

Keywords: Real interest parity (RIP), persistence, half-life.

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I. Introduction

Real interest parity (RIP) is a simple theoretical proposition which states that in the absence of arbitrage costs for goods and financial assets, real interest rates for essentially identical securities should be equal across countries. Determining the extent to which real interest rates are equalised is important for a number of reasons. First, if real interest rates in an economy move one-for-one with those abroad, an important channel for monetary policy to influence the domestic economy is removed (Cumby and Mishkin, 1986). Second, unless real interest rates can differ across countries, policies aimed at increasing domestic savings cannot increase the rate of capital formation and, hence, productivity (Feldstein, 1982). Finally, this parity condition is a key working assumption in various models of exchange rate determination. Early monetary models, such as Frenkel (1976), assumed perfect price flexibility which implies that both RIP and purchasing power parity (PPP) hold instantaneously. However, given the large and volatile short-run deviations from PPP documented in the literature, other models, such as Dornbusch (1976), were developed based on the assumption of sticky prices. Price stickiness, in this context, is expected to cause real interest rates to differ in the short-term,

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whilst over the long-run, flexible prices imply equal real interest rates. For these reasons, an examination of RIP is warranted.

Despite ongoing financial market integration, characterized by the significant relaxation or complete abolition of capital controls, which has typified the recent floating exchange rate period, most empirical studies to date have found at best ambiguous evidence on the equalization of real interest rates, nonetheless. Starting with Meese and Rogoff (1988) and Edison and Pauls (1993), the empirical literature has focused on investigating the time-series properties of deviations from parity (the real interest rate differential, RIRD hereafter)[1]. This is achieved through the use of unit root tests to examine whether these differentials are mean-reverting. However, conventional unit root tests are only concerned with the narrow question of whether or not deviations from parity contain a unit root. But, rejection of the unit root hypothesis is not necessarily evidence in favor of RIP as it is possible that tests reject the nonstationarity hypothesis but deviations are still persistent. In this case, national monetary authorities can still exercise independent influence over their financial markets. Moreover, studies, which do not account for the persistence of parity deviations, may draw the wrong inferences about the extent of market integration across countries. Consequently, we believe that a powerful test of RIP requires a detailed examination of the persistence and the speeds of mean-reversion of RIRDs based on the estimation of confidence intervals for the half-life which has become the standard tool for measuring persistence.

In this paper, we bring a recent empirical innovation to data for six country pairs and a sample spanning the period from 1965 to 1998 to investigate the empirical validity of RIP. The innovation is the median unbiased estimation (MUE) method of Gospodinov (2004) which allows for the construction of confidence intervals for the half-life based on impulse response analysis. This is particularly important in the present context for the following two reasons. First, Murray and Papell (2002) illustrate the existence of a substantial amount of sampling variability in measuring the half-life and, as a result, the point estimate alone does not provide a complete description of the persistence of deviations from RIP. Therefore, it needs to be supplemented with confidence intervals in order to gauge the precision of the estimates. Second, the commonly used estimate of the half-life, $H.L=\ln(1/2)/\ln(\rho)$, which is based on an autoregressive (AR) model

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[1] Numerous other studies have tested RIP employing a variety of econometric techniques. For example, Marston (1995) concludes that RIP is soundly rejected since RIRDs are systematically related to variables in the current information set. This is despite the fact that on average real interest differentials are close to zero. Kugler and Neusser (1993), on the other hand, investigate the validity of real interest parity using ex-post real interest data for several countries in a stationary multivariate time-series approach and provide evidence in favor of RIP. Further, Wu and Fountas (2000) test RIP using cointegration methods that allow for endogenously determined structural breaks and find a lack of real interest rate convergence towards the US in some countries. More recent work allows for the possibility of nonlinear dynamics. Mancuso et al. (2003) consider two nonlinear approaches to testing RIP, namely threshold autoregression (TAR) models and flexible nonparametric regressions. Their results suggest that important nonlinearities may characterize real interest rate linkages. Taken as a whole, the evidence on real interest rate equalisation is mixed, and there appears to be room for further research. This is indeed the focus of this paper.

of order 1, assumes that shocks decay monotically, but for higher order AR processes this may not be the case. To remedy this, Cheung and Lai (2000) recommend using impulse response analysis.

The remainder of this paper is set as follows. The next section describes RIP and explains the econometrics of local-to-unity processes. In Section III we discuss the data and report the empirical results. The last section concludes.

II. Real interest rate equalization and empirical methodology

RIP is the condition where real rates of return on identical assets are equalized across countries and its confirmation or rejection provides an indication of whether countries are financially integrated[2]. In an integrated economy, real rates of return on physical assets will tend to converge, as also will real rates of return on financial assets. Real interest rate equalization is, therefore, the broadest and the most theoretically appealing of the various measures of financial integration (Goldberg et al., 2003). Previous studies have used Eq. (1) as the basis for testing RIP[3]:

$$r_{it}=r_{\mathrm{USA}t} \quad (1)$$

where r_{it} and $r_{\mathrm{USA}t}$ are the domestic and US real interest rates, respectively. The real interest rate is defined using the Fisher hypothesis as the difference between the nominal interest rate and the inflation rate. The correct econometric approach to testing for the one-to-one relationship between real interest rates implied by Eq. (1) depends upon the time-series properties of these variables. Specifically, if these variables are stationary or I (0), as suggested by the consumption based capital asset pricing model (CCAPM) (Rose, 1988) and the Fisher hypothesis, then standard regression theory is valid and both the ordinary least squares (OLS) estimates of the intercept and slope coefficients in Eq. (2) and their standard errors will be consistent:

$$r_{it}=\mu+\beta r_{\mathrm{USA}t}+\varepsilon_t \quad (2)$$

Nonetheless, given the nonstationary behaviour which typically characterises real interest rates (Goodwin and Grennes, 1994; Phylaktis, 1999; Rapach and Wohar, 2004), it has become customary to use cointegration methods to test for a long-run equilibrium relationship between r_t and $r_{\mathrm{USA}t}$. In this case, RIP holds if the error term, ε_t, is stationary and μ=0 and β=1, or, equivalently, if the real interest differential, r_{it}-$r_{\mathrm{USA}t}$, is stationary.

The stationarity of RIRDs can be verified by applying unit root tests to determine whether they contain a unit root. However, if unit root is rejected, but the true value of the largest root of the autoregressive representation of the differential is close to unity, shocks will be slow to dissipate, and this stationary process

[2] See, inter alia, Goldberg et al. (2003), Goodwin and Grennes (1994), Lothian (2002), Obstfeld and Taylor (2002), Sekioua (2005a b) and Taylor (2002) for various empirical approaches to the questions of financial integration and real interest rate equalization.

[3] We follow much of the extant literature by using ex-post inflation data. This allows us to sidestep the thorny issue of specifying explicitly how inflationary expectations are formed and given that we are interested in measuring persistence, a long-run notion, this should not be crucial. If expectations are rational, actual and expected inflation will only differ by a white-noise error term (Rapach and Wohar, 2004).

may not be significantly different from a true unit root process in the economic sense. As a result, the emphasis should not be on whether RIRDs 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 RIRDs. A measure of persistence typically applied in the literature is the half-life, which indicates how long it takes for the impact of a unit shock to dissipate by half.

Median unbiased estimation

The method employed in this paper is due to Gospodinov (2004) and is based on inverting the likelihood ratio (LR) statistic of the largest root under a sequence of null hypotheses of possible values for the impulse response and the half-life. Starting from the following augmented Dickey-Fuller (ADF) regression which includes lagged first differences to account for serial correlation[4]:

$$y_t = \alpha y_{t-1} + \sum_{i=1}^{k-1} \psi_i \Delta y_{t-i} + \varepsilon_t \qquad (3)$$

where α (the largest root) is a measure of the persistence of the series (Andrews and Chen, 1994) and is cast as local-to-unity ($\alpha=1+c/T$ and holding c fixed as $T\to\infty$), $\varphi = (\alpha,\psi')' \in \Xi \subset R^p$ and the maximum likelihood estimator of φ is $\hat{\varphi}$. Suppose that we are interested in the null hypothesis that the impulse response function at horizon l, denoted by θ_l, is 0.5 (the half-life), versus the alternative $\theta_l \neq 0.5$, then this null or restriction can be written as $h(\varphi)=0$, where $h \equiv \theta_l - 0.5$: $R^p \to R$ is a polynomial of degree l. Let $\tilde{\varphi}$ denote the restricted maximum likelihood estimator and LR_T the likelihood ratio statistic of the null. Gospodinov (2004) shows that the restricted estimator of α converges at a faster rate than the unrestricted estimator and this helps obtain a consistent estimate of the nuisance parameter c under the imposed restriction (null hypothesis). Moreover, the restricted estimation provides consistent estimates of the impulse response functions and, thus, the half-lives[5].

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[4] Because neither theory nor empirics support the idea of trends in RIRDs, the tests performed are based on demeaned data.

[5] The standard method for estimating Eq. (3) is OLS and the conventional asymptotic interval is based on the asymptotic $N(0,1)$ approximation to the t-statistic which is valid only if $|\alpha|<1$. This approximation is poor in practice especially when the persistence parameter $|\alpha|$ is close or equal to unity. Specifically, if the true persistence parameter is not unity, OLS estimates are biased downwards and confidence intervals based on asymptotic methods have poor coverage properties. When persistence is unity, the coverage problems of the asymptotic intervals stem from the fact that the asymptotic distribution of α is non-standard. Bootstrap methods are also poor. This is because the percentile-t bootstrap is based on the assumption that the bootstrap quantile functions are constant, which is false for the AR model. This nonconstancy persists in large samples if we cast α as local-to-unity as $\alpha=1+c/T$. In this case, the asymptotic distribution of the t-statistic depends on α through the nuisance parameter c that is not consistently estimable (Hansen, 1999). Thus, in the near unit setting, the interval does not properly control for Type I error (Basawa et al., 1991).

The restricted LR estimator of Eq. (3) under the null hypothesis $h(\varphi)$=0 is:

$$LR_T \Rightarrow \left[\int_0^1 J_c^\tau(s) dW(s)\right]^2 / \int_0^1 J_c^\tau(s)^2 ds \tag{4}$$

where $J_c^\tau(r) = J_c(r) - \int_0^1 J_c(s) ds$, $J_c(r) = \int_0^r \exp[(r-s)c] dW(s)$ is a homogenous Ornstein-Uhlenbeck process and $\Rightarrow$ denotes weak convergence. The limiting theory of LR is dominated by the near nonstationary component and is not affected by the presence of stationary components as measured by the second term in regression, $\sum_{i=1}^{k-1} \psi_i \Delta y_{t-i}$.

The method of Gospodinov (2004) has many interesting features. First, contrary to standard asymptotic and bootstrap methods, which have been shown to have poor coverage properties, this method parameterizes α as a function of T and is expected to yield better small-sample and coverage performance. Second, the LR statistic does not require variance estimation for studentization. It is criterion function-based and is tracking closely the profile of the objective function. Also, the inversion of the LR statistic appears to shift the confidence intervals away from the nonstationarity region much more often compared to methods based on inverting the OLS estimator of α such as the grid bootstrap of Hansen (1999). Further, using a series of Monte Carlo experiments, Gospodinov (2004) shows that the inversion of the LR statistic appears to be controlling the coverage over a wide range of parameter configurations and across different forecasting horizons. This method is also expected to yield tight confidence intervals, which makes them highly informative.

Another statistic which takes into account the restricted and the unrestricted estimates is also proposed:

$$LR_T^{\pm} = \mathrm{sgn}\left[h(\hat{\varphi}) - h(\tilde{\varphi})\right]\sqrt{LR_T} \tag{5}$$

where sgn(.) is the sign of $\left[h(\hat{\varphi}) - h(\tilde{\varphi})\right]$. This statistic can be used for constructing two-sided, equal-tailed confidence intervals and median unbiased estimates. Finally, the 100η% confidence interval for the half-life, which is based on impulse response analysis, is: $C_\eta(l) = \{l \in L : LR_T \leq q_\eta(c)\}$, where $q_\eta(c)$ is the η^{th} quantile of the asymptotic distribution, l is the lead time of the impulse response function and $\tilde{\varphi} = \arg\max l_T(\varphi)$ subject to θ_l-0.5=0. The confidence interval for the half-life can be constructed using either $LR_T^{\pm}$ or LR_T .

III. Data and empirical results

The data utilized in this paper is extracted from the International Financial Statistics (IFS) of the International Monetary Fund (IMF) database and includes monthly long-term government bond yields and consumer price index (CPI) series for the USA, UK, France, Germany, Switzerland, Canada and Japan spanning from 1965:01 to 1998:12. The long-term government bond yields are preferred to short-term rates because these rates are closely linked to the cost of long-lived capital. Also, there are unanswered questions about the impact of measurement errors in prices. Because some consumer prices are sampled infrequently, short-term changes imperfectly measure actual price changes. This may cause biases in short-term tests of RIP, which is why long-term tests are useful (Jorion, 1996). Further, the USA is chosen as the reference country because of the fact that it is the main trading partner of the countries involved.

First, we test for the stationarity of RIRDs using the efficient generalized least squares (GLS) version of the Dickey-Fuller (DF) test due to Elliott et al. (1996) whose results are reported in table 1. While most unit root tests are only concerned with testing the null that the largest root is unity against the alternative that it is less than one, the DF-GLS test tests the null against a specific alternative H_1:α<1 where α=1+c/T. Further, using a sequence of tests of the null of a unit root against a set of stationary persistent alternatives, Elliott et al. (1996) showed substantial power gain from the DF-GLS method over the conventional ADF test (which has low power against close alternatives so that the unit root null can seldom be rejected for highly persistent variables). The lag length is chosen using the modified AIC (MAIC) of Ng and Perron (2001) which produces the best combination of size and power. From table 1, one can see that the DF-GLS test rejects the unit root null for all series at the 1% level of significance with the notable exception of France for which the null is rejected at the 5% level. This is an important result since it provides support for the equality of real interest rates in the long-run.

Quantifying the speed of mean-reversion: the half-life

Although the evidence presented in the previous section supports the validity of long-run RIP, it offers little information about the speed at which deviations die out. To obtain such information, computation of persistence is needed and the half-life is used to quantify such persistence. Table 1 reports the median unbiased estimates and the 68%, 80%, 90% and 95% MUE confidence intervals for this measure of persistence. The intervals are constructed by inverting the acceptance region of the powerful DF-GLS test of Elliott et al. (1996). Whilst the methodology in Section 2.1 is based on an ADF regression, the extension of this method to the DF-GLS test is simple. Instead of working with the data in levels as in Eq. (3), we simply work with the GLS demeaned data in the DF-GLS regression. Moreover, the finite-sample distribution of the DF-GLS test is obtained using the grid bootstrap of Hansen (1999)[6].

However, prior to constructing confidence intervals for the half-life, it is important to determine an appropriate range for this measure of mean-reversion, one that is consistent with the theory of RIP. Early exchange rate determination models such as those of Frenkel (1976) and Bilson (1978) assume real interest rate equality so that RIP holds. Others, such as Dornbusch's (1976) overshooting model, predict that nominal rigidities, in the form of sticky prices and wages, would cause real interest rates to diverge across countries. In this case, if the failure of RIP in the short-term is attributed to nominal rigidities, then one would expect

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[6] We also estimated α and the nuisance parameter c. However, given that the interpretation of these two coefficients is invariably conjectural, unlike the half-life which is based on the theory of RIP, they are not reported to save space. They remain, however, available from the author request.

substantial convergence to RIP over 1 to 2 years, as prices and wages adjust to shocks. Therefore, our theoretical range would have 2 years as an upper bound.

The MUE point estimates and confidence intervals for the half-life based on the impulse response functions are shown in table 1. The point estimates of the half-life range from 0.9837 years for France to 2.2084 years for Germany. The point estimates are on the whole consistent with the half-life implied by models with nominal rigidities. But, although one cannot reject the idea that the half-life is in line with theory these results do not determine how low the lower bound or high the upper bound can be. To answer this question, we now look at the constructed confidence intervals which are robust to the presence of persistent data. The lower bounds are generally indicative of fast convergence of real interest rates. For example, adjustment may take place in as much as half a year with 95% probability. These bounds are evidently all within our theoretical benchmark of 1 to 2 years. This is also precisely what could, in principle, be expected to happen in a highly integrated world where economic forces act rapidly; hence, discrepancies in RIRDs do not grow systematically over time. The upper bounds of the 95% confidence intervals, on the other hand, are, with the exception of the UK, not consistent with these models. This is especially true for France which has an infinite upper bound corresponding to no convergence of real interest rates; although, as demonstrated later, the presence of a structural break around the beginning of the floating exchange rate experience is a possible explanation for the infinite half-life. Besides France, none of the MUE confidence intervals are found to contain an infinite upper bound though. This is in line with the results of the DF-GLS test which rejected the unit root null hypothesis. Nonetheless, while the upper bounds are high and not all in the vicinity of two years as predicted by theory, they still imply mean-reversion with half of the adjustment taking place within less than 5 years. It is slightly faster if one considers the 90% confidence intervals for the half-life. Finally, the MUE confidence intervals from the powerful DF-GLS test appear to be rather tight and this demonstrates the potential for sharper inference from this test (Elliott and Stock, 2001).

Further evidence from impulse response analysis

As an extra exercise, we construct median unbiased confidence intervals of the impulse response functions derived from the inversion of the *LR* statistic. The graphs of the first 120 responses are displayed in Fig. 1. Whilst, impulse response analysis can be performed for even longer horizons, we report results up to 10 years since this is quite a close approximation to the infinite horizon. According to the point estimates, RIRDs have zero persistence in the long-run, confirming the absence of a unit root. The speed of convergence differs from one country pair to another with the UK having the quickest total response. Nonetheless, the upper limits of the confidence intervals suggest a high degree of persistence with some total lives exceeding 10 years. These upper bounds give the impression that deviations from parity mean-revert very slowly due to their long tails; however, this is to overlook the fact that they adjust rapidly in the months following a shock and mean-revert slowly thereafter. Moreover, the shape of the impulse response function emphasises the importance of calculating the half-life from the impulse response function. Cheung and Lai (2000) argue that, if the function is hump-shaped, rising before falling, it is preferable to estimate the half-life from the impulse function rather than from the largest root.

Recently, Mancuso et al. (2003) have explored the possibility of nonlinearities in real interest differentials using threshold autoregressive (TAR) tests. Nonlinearity, in this context, materializes itself by a variable speed of adjustment towards equilibrium: the larger the deviation, the faster it will be driven back to its equilibrium value. This variable speed of adjustment might be due to the fact that small deviations are not considered important by the market, whereas for larger deviations, the market pressure becomes stronger (Allen and Taylor, 1990; Taylor and Allen, 1992). Furthermore, contractual arrangements, which require assets to be held for a given period of time, may make it costly to act quickly to eliminate any profit opportunities. Nonlinearities may also result from automatic trading rules, heterogeneous beliefs and the tendency of traders to wait for large arbitrage opportunities to open up before entering the market (Sarno et al., 2004). Another likely explanation is transaction costs (Dumas, 1992). Specifically, transaction costs create a band of inaction within which no adjustment in deviations from equilibrium takes place and deviations may follow a persistent or even a unit root process, while outside the band, as the benefit of arbitrage exceeds the cost, the process switches abruptly to become mean-reverting towards the transaction cost band. In this framework, the real interest differential follows a nonlinear process that is mean-reverting towards the threshold band. This sluggish band of inaction may indeed explain some of the long tails observed in the impulse responses and the combination of rapid and slow mean-reversion. However, future research to identify the source of the long-tails should be of interest for a better understanding of the RIRD behavior.

Is there a break after Bretton-Woods?

In this paper, we have used a data sample which starts during the fixed exchange rate period of Bretton-Woods and continues during the recent floating experience. However, one concern that has been raised in the literature is that the long samples required for generating a reasonable level of test power may be inappropriate because of differences in the behavior of real variables across different exchange rate regimes. To investigate whether regimes matter for the persistence of deviations from RIP and the adjustment to shocks, we run the following ADF regression:

$$y_t = \alpha y_{t-1} + \sum_{i=1}^{k-1} \psi_i \Delta y_{t-i} + D_1\left(\alpha' y_{t-1} + \sum_{i=1}^{k-1} \psi_i' \Delta y_{t-i}\right) \tag{6}$$

where y_t is the RIRD for the whole sample and D_1 is a dummy that equals 1 for the recent float from 1974 to 1998 and 0 otherwise. We then test the following null, $H_0 : \alpha' = \psi_1' = \ldots = \psi_{p-1}' = 0$.

Table 2 reports the p-values of the F-test of the null hypothesis which indicate that we cannot reject the null that the coefficients of the dummies for the recent float are insignificantly different from zero. Indeed, in no case is there a break after Bretton-Woods except for France. For this country, the presence of a structural break is a possible explanation for the substantial amount of uncertainty associated with the estimates of the half-life observed earlier. Indeed, unit root tests are of notoriously low power in small samples. In the presence of breaks, this is particularly true. Overall, despite differences in nominal exchange rate regimes and market integration across time, the deterministic aspects of RIRD persistence appear to have been fairly uniform. From a theoretical point of view, this provides support for the neutrality of regimes proposition since there is no broad-based difference in the pattern of adjustment to shocks across fixed and floating regimes.

Conclusion

In this paper, we have assessed the empirical validity of RIP by using a newly developed unit root test which allows for the construction of confidence intervals for the half-life and the impulse response functions in the presence of persistent data. The constructed confidence intervals for these measures of mean-reversion provide strong parity support for the UK only. Specifically, the effect of a shock lasts for about 1 year with an interval comprising a maximum of 1.70 years. For the other countries, however, deviations from parity are stationary, albeit rather persistent. This indicates that for these countries, capital is mobile but not sufficiently enough as to lessen domestic authorities' control over their real interest rates. Finally, an important message of this paper is that, in the context of testing for the equality of real interest rates in the long-run, unit root tests and point estimates of the half-life alone are simply not informative enough. They need to be supplemented with confidence intervals in order to measure the precision of the estimates

References

Allen, H. and M.P. Taylor, 1990, "Charts, Noise and Fundamentals in the Foreign Exchange Market," Economic Journal, 100, pp. 49-59.

Andrews, D.W.K. and Chen, H.Y. (1994). "Approximately Median-Unbiased Estimation of Autoregressive Models," Journal of Business and Economic Statistics, 12, pp. 187-204.

Basawa, I.V., Mallik, A.K., McCormick, W.P., Reeves, J.H. and Taylor, R.L. (1991) "Bootstrapping Unstable First-Order Autoregressive Processes," The Annals of Statistics 19, pp. 1098-1101.

Bilson, J.F.O. (1978). "Rational Expectations and the Exchange Rate," in Frenkel, J.A. and Johnson, H.G., eds., the Economics of Exchange Rates: Selected Studies, Reading, Mass.: Addison-Wesley.

Cumby, R. and Mishkin, F. (1986). "The International Linkage of Real Interest Rates: the European-US Connection," Journal of International Money and Finance, 5, pp. 5-23.

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 Spatially Separated World," 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.

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

Elliott, G. and Stock, J.H. (2001). "Confidence Intervals for Autoregressive Coefficients near One," Journal of Econometrics, 103, pp. 155-181.

Feldstein, M. (1982). "Domestic Savings and International Capital Movements in the Long-Run and the Short-Run," NBER working paper 947.

Frenkel, J. (1976). "A Monetary Approach to the Exchange Rate: Doctrinal Aspects of Empirical Evidence," Scandinavian Journal of Economics, 78, pp. 200-224.

Goldberg, L. Lothian, J.R. and Okunev, J. (2003). "Has International Financial Integration Increased," Open Economies Review, 14, pp. 299-317.

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.

Gospodinov, N. (2004). "Asymptotic Confidence Intervals for Impulse Responses of Near-Integrated Processes: An Application to Purchasing Power Parity," Econometrics Journal, 7, pp. 505-527.

Hansen, B. (1999). "The Grid Bootstrap and the Autoregressive Model," The Review of Economics and Statistics, 81, pp. 594-607.

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

Kugler, P. and Neusser, K. (1993). "International Real Interest Rate Equalisation," Journal of Applied Econometrics, 8, pp. 163-174.

Lothian, J.R. (2002). "The Internationalization of Money and Finance and the Globalization of Financial Markets," Journal of International Money and Finance, 21, pp. 699-724.

Mancuso, A.J., Goodwin, B.K. and Grennes, T.J. (2003), "Nonlinear Aspects of Capital Market Integration and Real Interest Rate Equalization," International Review of Economics and Finance, 12, pp. 283-303.

Marston, R.C. (1995). "International Financial Integration: A Study of Interest Differentials between the Major Industrial Countries," Cambridge University Press.

Meese, R.A. 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.

Murray, C.J. and Papell, D.H. (2002). "The Purchasing Power Parity Persistence Paradigm," Journal of International Economics, 56, pp. 1-19.

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 Taylor, A.M. (2002). "Globalization and Capital Markets," NBER working paper 8846.

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

Rapach, D.E. and Wohar, M.E. (2004). "The Persistence in International Real Interest Rates," International Journal of Finance and Economics, forthcoming.

Rose, A.K. (1988). "Is the Real Interest Rate Stable?" Journal of Finance, 43, pp. 1095-1112.

Sarno, L., Taylor, M.P. and Chowdhury, I. (2004). "Nonlinear Dynamics in the Deviations from the Law of One Price: A Broad based Empirical Study," Journal of International Money and Finance, 23, pp. 1-25.

Sekioua, S.H. (2005a). "Real Interest Parity (RIP) over the 20th Century: New Evidence based on Confidence Intervals for the Largest Root and the Half-life," Unpublished manuscript.

Sekioua, S.H. (2005b). "What are the Determinants of Reversion towards Real Interest Parity (RIP)?" Unpublished manuscript.

Taylor, A.M. (2002). "A Century of Purchasing Power Parity," The Review of Economics and Statistics, 84, pp. 139-150.

Taylor, M.P. and Allen, H. (1992), "The Use of Technical Analysis in the Foreign Exchange Market," Journal of International Money and Finance, 11, pp. 304-314.

Wu, J.L and Fountas, S. (2000). "Real Interest Rate Parity under Regime Shifts and Implications for Monetary Policy," The Manchester School, 68, pp. 685-700.

Table 1 Unit root test and median unbiased confidence intervals for the half-life

	DF-GLS	H.L	68_{lower}	68_{upper}	80_{lower}	80_{upper}	90_{lower}	90_{upper}	95_{lower}	95_{upper}
Germany	[0.0005]	2.2084	1.6288	2.6034	1.5761	2.9318	1.4683	3.6155	1.3843	4.9067
Japan	[0.0015]	1.5756	0.9981	1.8361	0.9952	2.2070	0.9907	2.6211	0.9857	3.7026
France	[0.0240]	0.9837	0.6465	2.3071	0.5994	5.4812	0.5659	$+\infty$	0.5425	$+\infty$
UK	[0.0000]	1.1027	0.9729	1.1910	0.9538	1.3466	0.9233	1.4526	0.7430	1.7068
Switzerland	[0.0005]	2.1358	1.4977	2.5558	1.4087	2.8390	1.2403	3.4294	1.1956	4.3313
Canada	[0.0025]	1.1900	0.9827	1.4186	0.9715	1.6912	0.9534	2.5570	0.9345	4.0404

Figures in square brackets are p-values. The half-lives estimated from the impulse response functions are measured in years.

Table 2 Testing for breaks with respect to the pattern of adjustment to shocks:

$$y_t = \alpha y_{t-1} + \sum_{i=1}^{k-1} \psi_i \Delta y_{t-i} + D_1 \left(\alpha' y_{t-1} + \sum_{i=1}^{k-1} \psi_i' \Delta y_{t-i} \right)$$

	F-test	p-value
Germany	1.0806	[0.3736]
Japan	0.6052	[0.8609]
France	2.1818	[0.0099]
UK	1.8152	[0.1311]
Switzerland	0.5589	[0.8745]
Canada	0.7188	[0.7555]

The dependent variable is the real interest differential and D_1 is a dummy variable corresponding to the recent floating experience.

Fig. 1 Median unbiased impulse response functions estimated from the DF-GLS regressions. The unbroken line indicates the point estimates of the impulse responses. The dashed and dotted lines give the corresponding confidence intervals