Real interest parity (RIP) over the 20th century: New evidence based on confidence intervals for the dominant root and half-lives of shocks

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Abstract

In this paper, we have employed local-to-unity asymptotics and data dating back to the first quarter of the 20th century to examine the empirical validity of real interest parity (RIP) for the UK, Japan and France relative to the US. The results are as follows. First, the evidence for the full-sample is favourable using two powerful unit root tests. Nevertheless, the half-life estimates provide parity support for the UK only; with the upper bound of the confidence interval for this measure of persistence being less than two years. This finding underscores the importance of supplementing the unit root test results with direct measures of persistence. Second, the persistence of deviations from parity appears to have been exaggerated by the inclusion of data for the non-market 1940s. Indeed, the omission of this period leads to a decrease in the estimates of the half-life and RIP support for both the UK and France. For Japan, the upper bound is too high to be consistent with price stickiness and a world that has experienced a significant reduction in capital controls. Finally, there is little variation in the persistence of deviations from parity across fixed and floating exchange rate regimes. This finding is in line with the nominal exchange rate neutrality proposition.

JEL Classification: C22, F31, F36, E43.

Keywords: Real interest parity (RIP), Local-to-unity asymptotics, Persistence, Half-life.

1. Introduction

Real interest parity (RIP) states that if agents form their expectations rationally and if there are no impediments to trade, then real interest rates should be equalized across countries. 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)[1]. 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).

The early literature tested RIP using classic regression techniques and provided little or no evidence on its validity. For example, Cumby and Obstfeld (1984), Mishkin (1984) and Mark (1985) test the RIP hypothesis using ordinary least squares (OLS) regression analysis[2] and find evidence against this hypothesis. Nevertheless, the evidence based on regression methods must be interpreted with caution regardless of whether or not it is positive. There is some evidence from unit root tests suggesting that real interest rates are integrated of order one variables, and hence standard statistical inference may be invalid in these regression models. Even if the nonstationary real interest rates move together in the long-run, *i.e.* they are cointegrated; classical statistical inference is invalid since the estimated standard errors are inconsistent (Stock, 1987). Moreover, these studies failed to test for the stationarity of the residuals. If the residuals are nonstationary, then shocks to the real interest rate differential (RIRD(s) hereafter), which incidentally represents deviations from RIP, are permanent and the validity of the RIP hypothesis is rejected.

[1] Whether or not real interest rates are equal across countries is an issue closely related to the operation of activist stabilisation policy in the open economy since one channel through which monetary policy is thought to influence real economic variables is through the real interest rate. This channel would not be available if real rates are equal across countries since the ability of the authorities to influence their own real rate would be limited to the extent to which they could influence the world rate (Mark, 1985).

[2] This analysis is based on regressing the domestic real interest rate on its foreign counterpart and testing the hypothesis that the intercept and slope coefficients are equal to 0 and 1, respectively.

More recent research has focused on investigating the time-series properties of RIRDs[3]. 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 use data for Japan, Germany, the UK and Canada against the US dollar and are 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. To circumvent this problem, 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, Rothenberg and Stock (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 favorable 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

[3] For the most recent literature, we are only focusing on studies using unit root tests. 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, Goodwin and Grennes (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 our study.

papers, which were based on shorter samples and weaker statistical tests than those we have used."

An important criticism of the recent literature is that it is only concerned with the question of whether or not RIRDs contain a unit root. However, rejection of the unit root hypothesis is not necessarily evidence in favor of RIP since it is possible that unit root tests reject the nonstationarity hypothesis but the process is still persistent in the sense that deviations are slow to die out[4]. If deviations from RIP are persistent, this would indicate that national monetary authorities can still exercise independent influence over their domestic financial markets. Studies, which do not account for the persistence of real interest differentials, may lead to the incorrect acceptance of the hypothesis of market integration. Therefore, instead of focusing on the unit root question only, we believe that a powerful test of RIP requires a more thorough examination of the persistence and the speed of mean-reversion of RIRDs based on interval estimation of the dominant roots of autoregressive (AR) models and the half-lives of deviations from RIP.

In this paper, we bring two recent empirical innovations to a long span of historical data, both to investigate the validity of RIP and to study its historical evolution. The two empirical innovations are Hansen's (1999) grid bootstrap method and Gospodinov's (2004) median unbiased estimation (MUE) technique. These techniques are based on the construction of confidence intervals for the dominant root of local-to-unity AR processes and are particularly useful for estimating the half-lives of deviations which have become the standard tool for measuring persistence. To the best of our knowledge, this is the first paper to construct confidence intervals for the dominant root and the half-lives of deviations from RIP.

The main objective of this paper is to measure the persistence of RIRDs for a number of countries using long span data going back to the beginning of the 20th century. However, we realise also that studies using long span data may be subject to aggregation

[4] This problem is highlighted in the purchasing power parity (PPP) literature, in particular. A number of studies have tested the validity of this parity condition using long span data and have been able to reject the unit root hypothesis. This would normally be regarded as a validation of the PPP condition. Nevertheless, on closer inspection, the half-lives of deviations from PPP are found to be in the range of 3 to 5 years (Rogoff, 1996). This range is problematic since it is inconsistent with models of nominal rigidities which predict deviations from PPP to last between 1 and 2 years. Therefore, the lesson learnt from the PPP literature is that unit roots tests alone are simply not informative enough.

bias. This is due to the possible adverse effects of data heterogeneity, of combining data for varied nominal exchange rate regimes, and of the applicability to the current float of results obtained with such data (Lothian and McCarthy, 2002). Consequently, we divide the full-sample into a number of sub-samples associated with different nominal exchange rate arrangements. Then, we compare the behavior and, most importantly, the persistence of RIRDs in each of these samples. Finally, a formal test of whether or not the regime matters will be provided. There are two conflicting theories on this issue. The nominal exchange rate neutrality condition views the regime as being neutral, only affecting nominal economic variables in the countries involved, and not the behavior of real variables. On the other hand, some modern stochastic macroeconomic theory asserts that the channels of transmission of economic disturbances and the way in which economies respond to shocks may depend on the exchange rate regime (Caporale *et al*., 1994).

We employ monthly interest rate and price data dating back to the first quarter of the 20th century for the UK, Japan and France relative to the US. The efficient DF-GLS unit root test results are supportive of reversion towards parity for each country. However, point and interval estimates of the dominant root and the half-life of deviations from RIP, obtained using Hansen's (1999) grid bootstrap and the MUE method of Gospodinov (2004), indicate that RIRDs are strongly persistent. Further, the confidence intervals for the half-life of shocks provide parity support for the UK only. Nonetheless, the persistence of RIRDs, reflected in the estimates of the half-life, appears to have been exaggerated by the inclusion of the non-market wartime period from 1939 to 1949. Indeed, the omission of this period leads to a significant reduction in the estimates of the half-life and RIP support for both the UK and France. Specifically, shocks will last for 13 months on average and the upper bound of the confidence interval is in the vicinity of 24 months for the UK and France. For Japan, the upper bound is too high to be consistent with nominal price stickiness, which is viewed as the major cause of deviations from real interest parity, and a world that has experienced a dramatic reduction in capital controls. Finally, there is little variation in the persistence of RIRDs across fixed and floating exchange rate regimes. This finding is in line with the nominal exchange rate neutrality proposition.

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 3 we discuss the data and report the empirical results which include unit root tests and confidence intervals for the dominant root and half-life of deviations from parity. In the following section, we provide a test of regime neutrality. The final section concludes.

2. Real interest parity and empirical methodology

Ex-ante real interest parity rests on two important conditions, uncovered interest parity (UIP) and relative PPP:

$$i_t - i_t^* = \Delta s_t^e \qquad (1) \qquad \text{(UIP)}$$

$$\Delta p_t^e - \Delta p_t^{*e} = \Delta s_t^e \qquad (2) \qquad \text{(PPP)}$$

where s_t is the spot exchange rate defined as the domestic price of foreign currency, i_t is the nominal interest rate and p_t is the price level. Δ is the difference operator and the superscripts *e* and * denote expected values and foreign variables respectively.

Substituting (2) into (1) yields:

$$i_t - i_t^* = \Delta p_t^e - \Delta p_t^{*e} \qquad (3)$$

This can be rewritten as:

$$i_t - \Delta p_t^e = i_t^* - \Delta p_t^{*e} \qquad (4)$$

Utilizing the Fisher relationship, we derive an expression for real interest parity:

$$r_t = r_t^* \qquad (5) \qquad \text{(RIP)}$$

where r_t and r_t^* are the home and foreign real interest rates. The real interest rate differential (RIRD) is simply the deviation from RIP expressed as:

$$r_t - r_t^* = RIRD_t \qquad (6)$$

If expectations are rational, then:

$$\Delta p_t + \varepsilon_t = \Delta p_t^e \qquad (7)$$

$$\Delta p_t^* + \varepsilon_t^* = \Delta p_t^{*e} \qquad (8)$$

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).

For long-run RIP to hold the real interest differential should be a zero mean stationary process. The stationarity of RIRDs 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. 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 RIRDs' 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 months 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. Suppose the deviation of the RIRD y_t from its long-run level y_0, which is constant under RIP, follows an AR (1) process:

$$y_t - y_0 = \alpha(y_{t-1} - y_0) + e_t \tag{9}$$

where e_t is a white noise error term and the slope coefficient is estimated by OLS. At horizon *h*, the half-life deviation from RIP is the smallest value of *h* such that the percentage deviation from equilibrium α^h is reduced by one half. That is:

$$\alpha^h = 0.5 \Rightarrow h_{0.5} = \ln(0.5)/\ln(\alpha) \tag{10}$$

Note that this point estimate alone does not provide a complete description of the persistence of RIRDs. It needs to be supplemented with confidence intervals in order to measure the precision of the estimates (Lopez *et al*., 2003). The construction of confidence intervals for the slope coefficient and the half-lives using OLS and asymptotic distribution poses a number of problems, however. These intervals are not valid under the unit root null hypothesis and, even if long-run RIP holds, are biased downwards in small samples. Moreover, the estimate of the half-life in (10) is based on an autoregressive model of order 1 and assumes that shocks to real interest differentials decay monotically, but for higher order AR (*p*) processes this may not be the case. Therefore, estimating the half-life using (10) may lead the researcher to draw the wrong inference about the speed of adjustment. To remedy this problem, Cheung and Lai (2000) recommend measuring the half-life using impulse response analysis so that $h_{0.5}$ is:

$$h_{0.5} = \sup_{l \in L} \left| \partial y_{t+l} / \partial e_t \right| \geq 0.5 \tag{11}$$

Now, to address the problems associated with the estimation of the confidence intervals for the dominant root, we use two methods that have recently been developed. The first method is Hansen's (1999) grid bootstrap and the second one is Gospodinov's (2004) median unbiased estimation (MUE) technique. These methods are outlined next.

2.1. The grid bootstrap method

The starting point of this analysis is an autoregressive (AR) process of order p:

$$y_t^d = \gamma_1 y_{t-1}^d + \gamma_2 y_{t-2}^d + \ldots + \gamma_p y_{t-p}^d + \varepsilon_t \tag{12}$$

where y_t^d denotes the demeaned process[5] and ε_t is a serially uncorrelated error term. This AR (p) can be rearranged to obtain the following ADF regression:

$$y_t^d = \alpha y_{t-1}^d + \psi_1 \Delta y_{t-1}^d + \ldots + \psi_{p-1} \Delta y_{t-p+1}^d + \varepsilon_t = \alpha y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i \Delta y_{t-i}^d + \varepsilon_t \tag{13}$$

where $\alpha = \gamma_1 + \gamma_2 + \ldots + \gamma_p$ and $\psi_j = -\sum_{i=j+1}^{p} \gamma_i$ for j=1,…, p-1. The standard method to estimate regression (13) is by OLS, where $\hat{\alpha}$ and $s.e.(\hat{\alpha})$ are the OLS estimates of α and its standard error, respectively. The conventional asymptotic interval is based on the asymptotic $N(0, 1)$ approximation to the t-statistic:

$$t(\alpha) = (\hat{\alpha} - \alpha) / s.e.(\hat{\alpha}) \tag{14}$$

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,

[5] Because neither theory nor empirics support the idea of a trend in RIRDs, the tests performed in this paper are based on demeaned data.

the coverage problems of the asymptotic confidence intervals stem from the fact that the asymptotic distribution of the persistence estimate is non-standard (Clark, 2003). Bootstrap methods, on the other hand, 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 the persistence parameter as local-to-unity as $\alpha = 1 + c/T$ and holding c fixed as $T \to \infty$. 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). Since c is not consistently estimable, it follows that the percentile-*t* interval has incorrect first-order asymptotic coverage. Thus, in the near unit setting, the interval does not properly control for Type I error (Basawa *et al*., 1991; Hansen, 1999).

To overcome these problems, Hansen (1999) proposed a grid bootstrap method that has been shown, using Monte Carlo simulations, to yield accurate confidence intervals and unbiased estimates[6]. This method is implemented as follows. First, since the aim is to construct confidence intervals for α, we estimate this parameter by OLS, where $\hat{\alpha}$ and $s.e.(\hat{\alpha})$ are the OLS estimates of α and its standard error respectively. Next, we select a grid for α ($\alpha \in A_G = [\alpha_1, \alpha_2, ..., \alpha_G]$) in the relevant range of the OLS estimate, $\hat{\alpha}$. For any value of α in the grid, we regress $y_t^d - \alpha y_{t-1}^d$ on $\left(\Delta y_t^d, \Delta y_{t-1}^d, ..., \Delta y_{t-p+1}^d\right)$ and obtain the coefficients of this regression. Using the estimated coefficients, we then generate B (=1999) random time-series for the variable of interest y_t^d from the bootstrap distribution of the sample. For each sample, we calculate the test statistic $S_n(\alpha) = t(\alpha) = (\hat{\alpha} - \alpha)/s.e.(\hat{\alpha})$. Then, we sort the B test statistics $S_n(\alpha)$. The $100\theta\%$ order statistic $\hat{q}_n(\theta|\alpha)$ is the simulation estimate of the bootstrap quantile function. This procedure must be repeated for all the values of α in the grid and $\hat{q}_n(\theta|\alpha)$ is calculated at each $\alpha \in A_G$.

[6] Indeed, this method is shown to have first-order correct asymptotic coverage for both stationary and near unit root models. Thus, it asymptotically controls Type I error globally in the parameter space (Hansen, 1999).

Finally, the β-level grid bootstrap confidence interval for α is defined as the set of points α for which $S_n(\alpha)$ lies between $\hat{q}_n(\theta_1|\alpha)$ and $\hat{q}_n(\theta_2|\alpha)$ and is given by:

$$C_g = \{\alpha \in R : \hat{q}_n(\theta_1|\alpha) \leq S_n(\alpha) \leq \hat{q}_n(\theta_2|\alpha)\} \tag{15}$$

where $\theta_1 = 1-(1-\beta)/2$ and $\theta_2 = (1-\beta)/2$; so $\beta = \theta_2 - \theta_1$.

To calculate the half-life of shocks using Hansen's (1999) grid bootstrap method, we must first obtain the confidence intervals for the dominant root of the autoregressive process. Then, we simply apply (10) to the lower and upper bounds of the confidence intervals for the dominant root. However, this is only valid if the autoregressive process is of order 1. For higher order AR (p) models, the estimate of the half-life in (10) is problematic since it may be a non-monotonic function of the AR parameters. In this situation, one can either rely on impulse response analysis to estimate the half-life, as in Cheung and Lai (2000), or use an analytical measure due to Rossi (2003). This measure takes into account short-run dynamics and is useful for estimating the half-life for local-to-unity AR (p) processes. The bias-corrected half-life is:

$$h_{0.5}^{*} = \ln((0.5)\times b)/\ln(\alpha) \tag{16}$$

where b is a correction factor which can be consistently estimated from the ADF regression (13) and is equal to:

$$\hat{b} = \left(1-\sum_{i=1}^{p-1}\hat{\psi}_i\right) \tag{17}$$

2.2. The median unbiased estimation (MUE) method

The second method employed in this paper is the one proposed by Gospodinov (2004). Unlike Hansen's (1999) method, which is based on inverting the OLS estimator of the dominant root, this method is based on inverting the likelihood ratio (LR) statistic of the dominant root under a sequence of null hypotheses of possible values for the impulse response and the half-life. Using the same autoregressive AR (p) process in (13) we have:

$$y_t^d = \alpha y_{t-1}^d + \psi_1 \Delta y_{t-1}^d + \ldots + \psi_{p-1} \Delta y_{t-p+1}^d + \varepsilon_t = \alpha y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i \Delta y_{t-i}^d + \varepsilon_t \qquad (18)$$

where $\alpha = \gamma_1 + \gamma_2 + \ldots + \gamma_p$, $\psi_j = -\sum_{i=j+1}^{p} \gamma_i$ for j=1,…, p-1 and the parameter of interest α is cast as local-to-unity as $\alpha = 1 + c/T$ and holding c fixed as $T \to \infty$. Let $\rho = (\alpha, \psi_1, \ldots, \psi_{p-1})' = (\alpha, \psi')' \in \Xi \subset R^p$. The quasi likelihood estimator of the AR (p) process is shown in (19) as:

$$l_T(\rho) = -(T/2)\ln(\sigma^2) - 1/2 \sum_{t=1}^{T} e_t^2 / \sigma^2 \qquad (19)$$

and the unrestricted maximum likelihood estimator of ρ is $\hat{\rho} = \arg\max_{\rho \in \Xi} l_T(\rho)$. If we are interested in testing the null hypothesis that $h(\rho) = 0$, where $h : R^p \to R$ is a polynomial of degree l, then the restricted maximum likelihood estimator is given by $\tilde{\rho} = \arg\max_{h(\rho)=0} l_T(\rho)$ and $LR_T = T \ln(SSR_0 / SSR)$ is the likelihood ratio statistic of the null hypothesis, where SSR_0 and SSR are the sum of squares of the restricted and estimated residuals, respectively. Under a number of assumptions[7], Gospodinov (2004) shows that the restricted estimator of the persistence parameter converges at a faster rate than the unrestricted estimator and this helps obtain a consistent estimate of the nuisance parameter c. Moreover, the restricted estimation provides a consistent estimate of the impulse response function which is employed to measure the half-life of deviations.

The restricted LR estimator of (18) under the null hypothesis $h(\rho) = 0$ is:

[7] For a more detailed description of this method, see section 2.1 in Gospodinov (2004).

$$LR_T \Rightarrow \left[\int_0^1 J_c^{\mu}(s)dW(s)\right]^2 \Big/ \int_0^1 J_c^{\mu}(s)^2 ds \tag{20}$$

where $J_c^{\mu}(r) = J_c(r) - \int_0^1 J_c(r)dr$ and $J_c(r) = \int_0^r \exp[(r-s)c]dW(s)$ is a homogenous Ornstein-Uhlenbeck process generated by the stochastic differential equation $dJ_c(r) = cJ_c(r) + dW(r)$ with $J_c(0) = 0$, $W(r)$ is the standard Brownian motion defined on $[0,1]$ and $\Rightarrow$ denotes weak convergence.

The method proposed in Gospodinov's (2004) paper 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 like Hansen's (1999) OLS estimator. It is criterion function-based and is tracking closely the profile of the objective function.

Another statistic which takes into account the restricted and the unrestricted estimates of (16) is also proposed. This statistic is:

$$LR_T^{\pm} = \operatorname{sgn}\left[h(\hat{\rho}) - h(\tilde{\rho})\right]\sqrt{LR_T} \tag{21}$$

where sgn(.) is the sign of $\left[h(\hat{\rho}) - h(\tilde{\rho})\right]$ and $\hat{\rho}$ and $\tilde{\rho}$ are the unrestricted and restricted estimates. This statistic can be used for constructing two-sided, equal-tailed confidence interval and median unbiased estimate.

Finally, the $100\theta\%$ confidence interval for the half-life in (11), which is based on impulse response analysis, is:

$$C_{h_{0.5}} = \{l \in L : LR_T \leq q_\theta(c)\} \tag{22}$$

where $q_\theta(c)$ is the θ^{th} quantile of the limiting distribution, l is the lead time of the impulse response function and $\tilde{\rho} = \arg\max l_T(\rho)$ subject to $\alpha^l + l\alpha^{l-1}\psi + ... + \psi^l - 0.5 = 0$. The confidence interval for the half-life can be constructed using either $LR_T^{\pm}$ or LR_T.

3. Empirical results

3.1. Data and preliminary analysis

The data utilized in this paper is extracted from the www.globalfindata.com database and includes monthly[8] long-term government bond yields and consumer price indices (CPIs) for the US, UK, France and Japan (see table 1). The long-term government bond yields data, which applies to bonds of maturities of seven years or longer, is preferred to data for short-term rates for two reasons. First, firms do not usually make their investment decisions on the basis of short-term rates. Indeed, to the extent that firms borrow in bond markets, long-term yields will be more informative (Fujii and Chinn, 2000). Second, 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).

The inflation rate is defined as the rate of growth of the CPI which was seasonally adjusted by taking the average value for the previous 12 months. Further, some data points were missing for Japan's long-term government bond yield during the war period (1940s). This was corrected using linear interpolation[9]. 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. Finally, the empirical analysis is carried out using RIRDs computed as in equation (6).

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 RIRDs in level and first difference form for the UK, France and Japan relative to the US are plotted in figure 1. These graphs indicate that RIRDs were relatively volatile

[8] Since the aim of this paper is to measure the half-life of deviations from parity, by using high frequency monthly data, we are able to avoid the temporal aggregation bias analysed by Taylor (2001). Indeed, Taylor (2001) showed that the half-life, at least in the case of PPP, can be seriously over-estimated if adjustment takes place during a time frame that is shorter than the sampling frequency of the data.

[9] Although such an interpolation may be ad hoc as argued by Taylor (2002), it was considered necessary to give the empirical analysis a fair chance on this historical data. Without interpolation, any mean-reversion of the real interest differential for Japan would be missed and a bias against stationarity would result.

during the interwar (1923-1938), a period when exchange rates began to float or stay fixed for only a few years; and even more so during the volatile non-market wartime period between 1939 and 1949, especially for Japan and France. It is apparent also that departures from a zero mean, the value predicted by RIP, were associated with these volatile periods. During the same period, the differentials for Japan and France showed a positive trend until the middle of the 1930s and a negative mean afterwards. For the UK it is difficult to see a clear pattern. However, beyond the volatile period of the 1940s, the volatility of RIRDs declined significantly and they became fairly stable over the long-run with no apparent trend. Overall, the evidence is indicative of reversion towards a zero mean with occasional departures, confined mainly to the interwar period (1923-1938) and the 1940s, which appear to last for a considerable amount of time.

Having analyzed the behaviour of RIRDs graphically, it is now useful to ask: how large are these differentials on average?[10] To answer this question, we report in table 2 the average RIRDs estimated for all three countries with respect to the United States and their standard errors and *p*-values[11]. The null hypothesis being tested is that of zero differentials. It is clear from this table that the null hypothesis cannot be rejected for the UK, whereas for Japan and France the differentials appear to be significantly different from zero. The positive differential for the UK indicates that it commands a higher real interest rate relative to the US. Japan and France, on the other hand, command significantly lower real interest rates. Note, however, that the large and significant differentials for these two countries may be due to the inclusion of data for the unstable interwar period (1923-1938) and the 1940s. These two periods were characterized by large[12] and highly volatile differentials which may have driven up the differentials for the full-sample. In fact, when one looks at the second longest period only, from 1950 to 2000, which excludes the two volatile periods, one notices that the differentials are not significantly different from zero, consistent with RIP, and their standard deviations are

[10] In the absence of persistent risk premiums or any exogenous barriers to capital movement, RIRDs should be tightly clustered around a mean of zero.
[11] These estimates must be interpreted cautiously given the large variances.
[12] The significant differentials for these two periods may also be due to a risk premium which drives a wedge between the real interest rates for Japan and France relative to the US (see Marston, 1992).

fairly low and identical across both fixed and floating exchange rate regimes[13]. As a whole, apart from the unstable interwar and wartime periods, regime changes seem to have had very little, or no, effect on the behaviour of real differentials. Indeed, there is no clear cut difference between fixed and floating regimes since the differentials are of the same magnitude and are essentially close to zero. This provides tentative support for both RIP and regime neutrality.

Nevertheless, analysing just the average differentials is not very informative because it can hide substantial fluctuations. More insight into the behaviour of these differentials might be gained by applying unit root tests.

3.2. Unit root tests

The results of the efficient Dickey-Fuller generalized least squares (DF-GLS) unit root test recommended by Elliott *et al*. (1996) are reported in table 3[14]. While most unit root tests are only concerned with testing the null hypothesis that the dominant root of an AR (*p*) process is unity ($H_0 : \alpha = 1$) against the alternative that it is less than one ($H_1 : \alpha < 1$), the DF-GLS method tests the null against a specific alternative $H_1 : \alpha < 1$ where $\alpha = 1 + c/T$ and holding c fixed as $T \to \infty$. Using a sequence of tests of the null hypothesis of a unit root against a set of stationary local alternatives, Elliott *et al*. (1996) showed substantial power gain over the conventional ADF test, that has low power against close alternatives so that the null hypothesis of a unit root can seldom be rejected for highly persistent variables, could be obtained from using the DF-GLS test. This test is based on the following regression:

$$y_t^d = \alpha y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i \Delta y_{t-i}^d + \varepsilon_t \qquad (23)$$

[13] The volatility of RIRDs declined during the Bretton-Woods period during the 1960s. Once the floating rate began in the early 1970s, there was very little change in the volatility of real differentials.

[14] In this table, we also report the results of the Ng and Perron (2001) unit root test. The results which are reported but not analysed are quantitatively similar to those of the DF-GLS test.

where y_t^d is the GLS demeaned RIRD, *i.e.* $y_t^d = y_t - z_t\omega$, where $z_t = 1$. ω is the vector of OLS coefficients of $\widetilde{y}_t = \left[y_1, (1-\overline{\alpha}L)y_2, ..., (1-\overline{\alpha}L)y_T\right]'$ on $\widetilde{z}_t = \left[z_1, (1-\overline{\alpha}L)z_2, ..., (1-\overline{\alpha}L)z_T\right]'$ and L is the lag operator.

The lag length for the DF-GLS test is chosen using the modified Akaike Information Criterion (MAIC) of Ng and Perron (2001). Ng and Perron (2001) showed that this criterion is particularly useful because it produces the best combination of size and power for the DF-GLS test. Given the monthly frequency of the data, we have allowed for a maximum lag length of 18. It must be stressed, however, that the long lags selected by the modified AIC, and shown in the third column of table 3, are not surprising. This criterion is designed to select relatively long lag lengths in the presence of roots near unity and shorter lags in the absence of such roots.

With the lag selected by MAIC, the DF-GLS test rejects the null hypothesis for the full-sample at the 1% significance level for all three series. This is an important result since it offers support for the RIP hypothesis over the long-run. Next, we examine the behaviour of RIRDs during three different sub-samples which are associated with different nominal exchange rate arrangements. These periods are: the interwar (1923-1938), a period characterised by unstable exchange rate arrangements; the Bretton-Woods fixed exchange rate period (1950-1973); and the recent floating rate experience (1974-2000). We also analyse RIRDs during the non-market period from 1939 to 1949. The idea here is to determine whether RIRDs behave differently under these historical periods and nominal exchange rate regimes. It is not uncommon that studies using long historical samples are criticised for combining data for different nominal exchange rate regimes. If the variable under study behaves differently under these regimes, then the pooling of data will lead to invalid inferences. Thus, this idea deserves to be investigated.

Unsurprisingly, the evidence for the four sub-samples is weaker than for the full-sample. For example, during the recent float the null hypothesis cannot be rejected for the UK and Japan and it is only rejected at the 10% significance level for France. This result, which is in fact consistent with much of the previous literature, implies that even the powerful DF-GLS test cannot reject the unit root hypothesis. The same result is found for the interwar, the 1940s and Bretton-Woods. In fact, the only favourable evidence comes from the data for France where we have able to reject the null hypothesis in more than one period. In all, RIRDs appear to behave similarly in most regimes, at least in fixed and floating.

However, the negative results for the sub-samples are most likely to be a reflection of the low power of the tests, especially in short-samples, than evidence against RIP. Indeed, this was confirmed when we studied the full-sample. Nevertheless, a formal test of regime neutrality will be provided in section 4.

3.3. Confidence intervals for the dominant root

3.3.1. Grid bootstrap

While the focus of this paper is on measuring the persistence of RIRDs rather than the rejection of unit root, we have addressed the latter subject first. Now we address the former. We start our analysis by constructing confidence intervals for the dominant root using Hansen’s (1999) grid bootstrap method. The lag length is chosen using the modified Akaike information criterion (MAIC) as in section 3.2. Table 4 reports the lag length, the OLS estimate of α and its standard error and the 90% and 95% bootstrap confidence intervals for this persistence parameter. These bootstrap intervals were constructed using 1999 replications at each of 200 grid points.

The results for the full-sample are consistent with those of the DF-GLS test which strongly rejected unit root. The confidence intervals for the persistence parameter do not include unity in any case suggesting that the differentials are mean-reverting. Nevertheless, the point estimates imply that the differentials are highly persistent. In all cases, the OLS estimates of α , though biased downwards as one might expect, are 0.9 or higher. The RIRDs for Japan and France are the most persistent with point estimates equal to 0.9718 and 0.9747, respectively. This is indicative of very slow mean-reversion. It is interesting to point out that the lower bounds of the confidence intervals are close to the point estimate and are never below 0.9.

The results also show that there are no material differences in the persistence of RIRDs for the sub-samples[15]. In fact, the differentials are strongly persistent in all four sub-samples; the point estimates of α are all essentially between 0.9203 and 0.9630. However, the only difference between the results for the sub-samples and the full-sample is the presence of unity in the confidence intervals for the UK and Japan. The evidence for these two countries is not surprising since we were unable to reject the hypothesis of unit root across the different regimes for these two countries. However, this is more likely to be the result of using shorter samples, *i.e.* low power, than evidence against RIP. For France, the upper bounds of the confidence intervals are consistent with the results of the DF-GLS test.

3.3.2. Median unbiased estimation (MUE)

Median unbiased estimates[16]and confidence intervals for the persistence parameter are reported in table 5. The confidence intervals are constructed by inverting the acceptance region of the powerful DF-GLS test of Elliott *et al*. (1996). Although, the methodology outlined in 2.2. is based on an ADF regression, the extension of this method to the DF-GLS test is straightforward. Instead of working with the data in levels as in (16), we simply work with the GLS demeaned data in the DF-GLS regression (23). Moreover, the finite-sample distribution of the DF-GLS test is obtained using Hansen's (1999) grid bootstrap method.

For the full-sample, the median unbiased estimates of the persistence parameter are slightly above the OLS estimates, and are indicative of very strong persistence. None of the MUE confidence intervals are found to contain unity as an upper bound. This is consistent with the results of the DF-GLS unit root test. Essentially, this finding is supportive of the idea that RIRDs are stationary, but highly persistent.

[15] Given the insufficient number of observations for the UK real interest rate differential during the interwar, we decided not to run the tests for this period.
[16] Median unbiased estimation was proposed by Andrews (1993) and Andrews and Chen (1994) to correct for the bias in the OLS estimate of the persistence parameter. In Gospodinov's (2004) paper, MUE is extended to the LR estimate of α in the DF-GLS regression.

Looking across the sub-samples, we find that the real interest differentials exhibit substantially greater persistence than OLS estimates suggest. And unlike the full-sample, the OLS bias now is quite large. For instance, during the recent float the OLS estimate of α for Japan is 0.9216, whereas the bias corrected estimate is almost unity. These findings run in accordance with the expectation that the OLS point estimate is a misleading indicator of the true value of α . Furthermore, for the full-sample there appears to be little difference between the confidence intervals based on the grid bootstrap method and those obtained using median unbiased estimation; and the bias in the OLS estimate, normally due to using short-samples, disappears almost completely. On the other hand, when we look at the sub-samples we find that Hansen's (1999) confidence intervals are very wide and this may make it difficult to make definitive statements one way or another regarding the unit root/stationarity question. In general, the tight MUE confidence intervals from the powerful DF-GLS test demonstrate the potential for sharper inference on the persistence of RIRDs.

3.4. Point estimates and confidence intervals for the half-life

Before constructing confidence intervals for the half-life, 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[17], 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) and Bilson (1978) 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, as prices adjust to shocks. In

[17] There is a large empirical literature on measuring the persistence of deviations from PPP. Recent examples include: Cheung and Lai (2000), Murray and Papell (2002), Gospodinov (2004), Lopez *et al*. (2003), Rossi (2003) and Sarno and Valente (2003) to name just a few. Generally, results in these papers are compared with a benchmark based on models with nominal rigidities that predict a range for PPP deviations between 1 and 2 years.

fact, this theoretical range for the half-life estimates of price convergence is supported by Cheung *et al*. (2002) 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 is also consistent with RIP since it implies rapid adjustment of RIRDs. Therefore, our range would have an upper bound of 24 months, but any value less than this is obviously acceptable.

We begin our analysis by reporting in table 6 the half-life estimates which are based on the OLS estimate of the dominant root and are computed using equation (10). Since reporting merely point estimates does not convey the inevitable imprecision with which the adjustment speed is measured, asymptotic confidence intervals are also reported. These asymptotic confidence intervals are based on normal sampling distributions and are obtained using a delta approximation method. Table 6 indicates that for the full-sample the point estimates of the half-life are 12.8 months for the UK, 24.3 months for Japan and 27 months for France. These estimates are supportive of reversion towards parity since they are all within or slightly above our benchmark which has an upper bound of 24 months. The lower bounds include a range of short half-lives, which can be much less than 12 months and, in the case of the UK, less than 8 months. The upper bounds, on the other hand, imply speeds of adjustment which appear relatively slow, approximately 36 months for Japan and almost 42 months for France. In all, the point estimates and confidence intervals for the full-sample are supportive of RIP and are remarkably consistent with the results of the unit root tests, especially for the UK.

The estimates for the sub-samples are puzzling, nevertheless. In fact, these estimates seem completely at odds with the results of the DF-GLS unit root test and the grid bootstrap and MUE confidence intervals for the dominant root which provided clear evidence in favor of the presence of unit roots in real interest differentials. If a unit root is present, then evidently we would expect deviations never to die out and the half-life, or at least the upper bound of the confidence interval for this measure, to be infinity for the sub-samples considered. This is not the case since table 6 shows that the highest upper bound is less than 43 months. This finding is not altogether surprising. The half-lives and their confidence intervals are based on the OLS estimation of the dominant root. However, the OLS estimate is significantly downward biased and this bias becomes more severe as the root α gets larger, *i.e.* gets closer to the nonstationarity region (Murray and Papell, 2003). As for the asymptotic confidence intervals, their coverage properties have been shown to be poor. Indeed, Rossi (2003) argued that these intervals might be unreliable when variables are highly persistent and the sample size is small. This is particularly relevant for the persistent real differentials during the different sub-samples.

Another drawback of the results in table 6 is that the method employed to calculate the half-lives of deviations is only valid if the process representing the RIRD is an autoregressive process of order 1. For higher order AR (p) models it is preferable to calculate the half-life from the impulse response function instead of (10). Alternatively, we can utilize the bias-correction-factor of Rossi (2003) which is given in (17). The associated half-life is shown in (16) and is appropriate for highly persistent AR (p) processes.

The bias-corrected half-lives are reported in table 7. Correcting for the bias yields significantly higher point estimates, not only for the full-sample but also for the other historical periods under examination. The higher point estimates are more reasonable than what is implied in table 6 because they support the expectation that real interest differentials are highly persistent, in particular for the sub-samples. For full-sample, the estimates are quite problematic, nonetheless. The estimates for the three time-series are outside our benchmark which requires the half-lives to be less than 24 months. Specifically, the point estimates are 30 months for the UK, 46 months for Japan and 37 months for France. These torpid rates of reversion are puzzling because they appear too slow to be explained by price stickiness. The asymptotic confidence intervals for the corrected half-life estimates are also reported in table 7. However, these asymptotic intervals are unreliable as previously stated[18].

The main purpose for reporting asymptotic confidence intervals for the point estimates of the half-life was to provide a complete picture of the speed of convergence towards RIP and a better indication of the uncertainty in obtaining these estimates. Yet, the results in table 6 and 7 have underlined the potential problems in relying on asymptotic confidence intervals that consistently understate the persistence of RIRDs.

[18] The highest upper bound of 74 months is still inconsistent with the unit root test results.

Therefore, we next report confidence intervals that are more robust to the presence of highly persistent variables.

The confidence intervals based on Hansen's (1999) method are shown in tables 8 and 9. These intervals are computed by initially inverting the OLS estimator of the dominant root, and then constructing two-sided confidence intervals using the grid bootstrap method. The confidence intervals for the half-life are calculated by applying equations (10) and (16) to the confidence intervals for α. Clearly, the point estimates of the half-life remain the same as before, hence we only focus on the lower and upper bounds of the confidence intervals. The results in table 8 show that for the full-sample, the lower bounds are indicative of reasonable speeds of adjustment towards RIP. Indeed, these bounds range from a low of 9.5 months to a high of 19.6 months and are consistent with models with sticky prices. With the notable exception of the UK, the upper bounds are less supportive of RIP, however. For instance, France has an upper bound of almost 96 months. If RIP reversion is tied to the price convergence speed, then real interest differentials should be found to converge faster than this.

Looking across the sub-samples, we find that the lower bounds are between 6 and 11 months. It is interesting to note the very little variation in the lower bounds across the different exchange rate regimes. For the UK, in particular, the lower bounds are remarkably close. So it appears that moving from one regime to another has limited effect on the persistence of RIRDs. Nonetheless, the upper bounds for the confidence intervals are infinite in the majority of cases. These particular upper bounds correspond to periods for which we were unable to reject the unit root null hypothesis. This finding emphasizes the unreliability of asymptotic confidence intervals which failed to reveal the substantial amount of sampling uncertainty associated with estimates of the half-life. Overall, the results for the sub-samples are not very informative with confidence intervals being too wide and an infinite upper bound.

The half-lives in table 8 were calculated from equation (10) which is based on an AR (1) model. This equation assumes that adjustment is monotonic, but this may not be true for higher order AR (*p*) processes. To deal with this problem, we use the bias-correction-factor of Rossi (2003). Predictably, the results in table 9 move us even further from our theoretical benchmark. Apart from the lower bound of the confidence intervals for the

UK, all other bounds are greater than 24 months. As for the sub-samples, the intervals paint a similar picture to that in table 8.

So far we have constructed confidence intervals for the half-lives by applying either (10) or (16) to the confidence intervals for the dominant root. However, this has turned out to be disappointingly misleading. Our last option is to draw on impulse response analysis. Moreover, by computing the half-life estimates from the inversion of the DF-GLS test statistic, the hope is that this powerful test will yield tighter confidence intervals than what we have estimated thus far.

The MUE point estimates and confidence intervals for the half-life based on impulse response analysis and the DF-GLS test statistic are shown in table 10. For the full-sample, the MUE point estimates for the UK and France are less than 24 months, whereas the estimate for Japan is slightly above the theoretical upper bound of 24 months. The lower bounds of the confidence intervals range from about 12 months to 14 months with most values being in the vicinity of 12 months. The point estimates and the lower bounds are consistent with the half-life implied by models with sticky prices. The upper bounds, on the other hand, are too high to be explained by these models, except for the UK. Overall, the MUE confidence intervals are slightly tighter than those obtained by the grid bootstrap method of Hansen (1999), and this demonstrates the gain from employing more powerful tests. Nevertheless, they are wide enough to make it difficult to support RIP for France, whereas for Japan the evidence is rather weak.

Consequently, on the basis of impulse response analysis, it appears that we can support reversion towards parity only for the UK. However, our analysis would be incomplete without noting that in table 2, we showed that RIRDs were highly volatile during the non-market 1940s. This volatile period was also associated with departures from a zero mean which appeared to last for a sizeable amount of time. In fact, this is confirmed in the estimates of the half-life, especially for Japan and France. More importantly, it seems that the inclusion of this volatile period might have exaggerated the extent of RIRD persistence for the full-sample. Indeed, when we look at the second longest sample, 1950-2000, for which we could easily reject the unit root null hypothesis for all three countries, we can clearly see that the point estimate and the upper bound for the confidence interval are only slightly lower than for the full-sample for the UK. In actual fact, the omission of the wartime period does not have much impact for the UK. However, the reduction for France is drastic; the upper bound falls from a high of 76 months to a reasonable value of 27 months. A timeframe that is consistent with our benchmark. For Japan, the upper bound is infinite. Moreover, the exclusion of the volatile period indicates that the half-life estimates for 1950-2000 are close to those for the floating and fixed exchange rate periods (in other words, regimes are neutral). On the whole, we now have favourable evidence not just for the UK but also for France.

Looking now at the sub-samples, we see that the overall picture obtained using MUE is one of noticeably greater persistence than with OLS estimation. Further, the results are similar to those in table 8 and reveal substantial uncertainty associated with the estimation of the half-life. With the inversion of the DF-GLS test being more powerful, we expected a tightening of the confidence intervals for the half-life estimates. However, similar to Hansen's (1999) grid bootstrap, the upper bound remained at infinity for most cases. These upper bounds are puzzling not only because they are inconsistent with sticky price models, but also because they are incompatible with a world which has experienced a significant relaxation of capital controls especially during the recent floating exchange rate period[19]. Japan's RIRD during the recent float is the exception, however. For this country, we could not reject the unit root hypothesis using the DF-GLS test. Further, the MUE confidence interval for the dominant root was found to have a lower bound equal to 0.9801 and upper bound of unity. The point estimate, on the other hand, is a staggering 0.9923. So it seems quite inconceivable that the half-life would have a maximum value of 21 months only; it should be infinity according to the unit root test results and MUE confidence intervals for the dominant root. The opposite is also true for the period from 1950 to 2000. Thus, although we might be tempted to conclude that RIP has held well for Japan during the recent float on the basis of the half-life estimates and there is no need to analyze the full-sample, we believe that the evidence for this country during the sub-samples should be interpreted cautiously.

[19] The presence of infinite upper bounds is likely to be due to the low power of the tests when analyzing short-samples. In general, it is only when we use long span data that the tests yield reasonable speeds of adjustment. This underscores the importance and the gain in test power from using long historical data.

Lastly, we previously established that for the full-sample the difference between the point estimates and confidence intervals for the dominant root constructed by OLS and the grid bootstrap and those computed by median unbiased estimation was negligible. This finding is substantiated in this section. Indeed, the difference between the point estimates and confidence intervals for the half-life estimated by either (10) or impulse response analysis is very small. Therefore, for long samples it does not matter which method we utilize, and the use of Rossi's (2003) bias-correction-factor is not required.

4. Testing nominal exchange rate neutrality

In this paper, it has been argued that the failure of the previous literature to find evidence of a long-run relationship between real interest rates is due to the low power of conventional univariate tests. The low power refers to the inability to reject a false unit root hypothesis with a sample corresponding to the length of the recent float. To remedy this, we have employed long-span data which offer a means of overcoming low test power problems. However, one concern that has been raised in the literature, especially with studies using long historical data, is that the long-samples required to generate a reasonable level of test power may be inappropriate because of differences in the behavior of real variables across different historical periods and across different nominal exchange rate regimes (Sarno *et al*., 2003; Sarno and Valente, 2003).

In the analysis thus far, we have established that there is little variation in the estimates of the persistence parameter across regimes. Specifically, the median unbiased estimates are found to be between 0.9378 and 0.9923[20], and this is indicative of strong persistence of RIRDs. Moreover, with the exception of the volatile wartime period, reversion is only slightly faster under the fixed exchange rate regime, 12 months for the UK and Japan and 17 for France, versus floating, 13 months for the UK, 18 months for Japan and 18 months for France. This is an interesting finding since it implies that despite fundamental differences in nominal exchange rate arrangements, institutional structure and market integration across time, the persistence of shocks to RIRDs has been fairly uniform. From a theoretical point of view, this provides some prima facie support for the

[20] See table 10.

nominal exchange neutrality proposition which states that the behaviour of real variables over time is not affected by the nominal exchange rate system in place.

Nevertheless, to investigate formally whether or not regime matters, we need to estimate the following regression[21] with dummy variables corresponding to different historical periods and exchange rate regimes:

$$
\begin{aligned}
y_t^d = \alpha y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i \Delta y_{t-i}^d + d_1\left(\alpha' y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i' \Delta y_{t-i}^d \right) + \\
d_2\left(\alpha'' y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i'' \Delta y_{t-i}^d \right) + \\
d_3\left(\alpha''' y_{t-1}^d + \sum_{i=1}^{p-1} \psi_i''' \Delta y_{t-i}^d \right) + \varepsilon_t
\end{aligned}
\tag{24}
$$

where the dependent variable is the RIRD for the full-sample and d_1 is a dummy variable that equals 1 for the recent float (1974-2000) and 0 otherwise, d_2 equals 1 for the Bretton-Woods phase of fixed exchange rate (1950-1973) and 0 otherwise and d_3 equals 1 for the non-market wartime period between 1939 and 1949 and 0 otherwise[22]. In order to determine if regime differences matter, we test the following null hypotheses:

$$
\begin{aligned}
H_{01} &: \alpha' = \psi_1' = \ldots = \psi_{p-1}' = 0 \\
H_{02} &: \alpha'' = \psi_1'' = \ldots = \psi_{p-1}'' = 0 \\
H_{03} &: \alpha''' = \psi_1''' = \ldots = \psi_{p-1}''' = 0
\end{aligned}
\tag{25}
$$

Table 11 reports the F and χ^2 statistics for the null hypotheses of interest with their associated p-values. The χ^2 statistic is equal to the F statistic times the number of restrictions under test. In this example, there is more than one restriction and so the two test statistics are different with the p-values of both statistics indicating that we cannot reject the null hypothesis that the coefficients of the dummies for the float, Bretton-Woods and the non-market period are insignificantly different from zero. This finding

[21] Note that this regression allows us to determine whether its parameters are different across regimes and not whether or not the half-life estimates are different.

[22] Since we have analysed the behaviour of RIRDs over three different nominal exchange rate arrangements and the non-market wartime period (the recent float, Bretton-Woods, 1940s and the interwar), we can only have 3 dummies. In addition, the lag length is the same as the one chosen for the full-sample using the modified AIC. Experimentation with other lag lengths, revealed no significant difference.

runs in accordance with the expectation that the exchange rate regime should not affect the behavior RIRDs. However, H_{03} is rejected for France. This implies that the volatile wartime period has undoubtedly affected the behaviour of the real differential for this country. This was evident in the estimate of the half-life which fell dramatically when we omitted this period from the full-sample. For the UK, the exclusion of the volatile period, led to only a small reduction in the point estimate and upper bound of the confidence interval for the half-life. This might explain why we could not reject H_{03}.

To corroborate our results, we now test the pooling restriction that all the coefficients are equally zero. This restriction is specified as:

$$H_{00} : \alpha' = \psi_1' = ... = \psi_{p-1}' = \alpha'' = \psi_1'' = ... = \psi_{p-1}'' = \alpha''' = \psi_1''' = ... = \psi_{p-1}''' = 0 \quad (26)$$

The results which are reported in table 11, suggest that we can reject the null hypothesis for the UK and France, but not for Japan. The result for France is not unexpected since we found that H_{03} is rejected and this might have affected the results for the full-sample. For the UK, it seems that the full-sample is not as homogeneous as we previously thought. Despite this finding, we believe that we have sufficient evidence to support regime neutrality. Throughout this paper, we have remarked that RIRDs are fairly stable across fixed and floating regimes. The only exception is the non-market wartime period (1939-1949) which was shown to have been the only major determinant of RIRD behavior; though for the UK, the effect of this period is less severe than in the case Japan and France. This may not be surprising given the effect of World War II and its aftermath right after. The evidence for Japan implies that we can treat the full-sample as homogeneous.

As a final point, we note that the observed empirical regularities in this paper provide evidence in support of the nominal exchange regime neutrality condition which states that the behaviour of real variables should not be substantially and systematically affected by the nature of the nominal exchange rate regimes. As a result, the idea that different monetary regimes may generate regime breaks in the structural dynamics of RIRDs, especially when using long-samples, remains unproven.

5. Conclusion

This paper builds and extends on previous literature of RIP in several new directions. We start our analysis by testing whether RIP has held over an almost century long sample using two powerful unit root tests. The results are, on the whole, supportive of reversion towards parity for the countries investigated which are the UK, Japan and France. However, unit root tests are only concerned with the specialised case in which the dominant root of the autoregressive process is unity. Here, we supplement the unit root test results with confidence intervals for the dominant root and half-life of shocks to RIRDs. These confidence intervals can be used to provide more information than that given by unit root tests. The constructed intervals are indicative of strong persistence of RIRDs, with estimates of the dominant root found to be in the vicinity of unity for the countries involved. Further, the estimated half-lives provide support for RIP for the UK only. During the full-sample the effect of a shock on the RIP equilibrium relationship for this country lasts for about 17 months with a confidence interval comprising a maximum of 21 months. This timeframe is consistent with both sticky price models and a world with significant capital market integration. In contrast, the upper bound for the half-life for France indicates that shocks may last for more than 75 months. This upper bound implies that monetary policy is still effective in influencing the economy through the real interest rate channel, since a necessary condition for such a policy to work is the failure of real interest parity.

Results of studies employing long-spans of data are often called into question. This centres on the idea that adjustment of real variables to shocks will be different under fixed and floating regimes, and combining of data for the two types of regimes will lead to invalid inferences (Lothian and McCarthy, 2002). To remedy this problem, we study in detail the behaviour of RIRDs across the unstable interwar regime of fixed and floating rates, the fixed exchange rate period of Bretton-Woods and the recent float. We also analyse RIRDs during the non-market wartime period during the 1940s. The results are interesting for a number of reasons. First, apart from the volatile wartime period, RIRDs appear to be uniform across nominal exchange rate regimes, especially Bretton-Woods and the recent float. This provides support for nominal exchange rate neutrality. Second, during the different sub-samples associated with fixed and floating regimes, we could not find decisive evidence in favour of reversion towards parity. In fact, our results, particularly for the recent float, replicate most of the evidence found in the previous literature. However, this is likely to be due to the limited power of the tests that arises when using short-samples rather than evidence against RIP, and this is what we have highlighted in this paper using long-samples. Third, the omission of the non-market wartime period leads to a significant reduction in the estimates of the half-life for France. Indeed, the upper bound of the confidence interval obtained using impulse response analysis falls from 76 months to a reasonable figure of 26 months. Fourth, the evidence for Japan, for at least two sub-samples, is rather inconsistent since the confidence intervals for the half-life are incompatible with the unit root test results and the confidence intervals for the dominant root which reveal strong persistence. Consequently, the evidence for Japan during the sub-samples must be interpreted carefully. Though, since the full-sample was found to be homogenous we consider the evidence for Japan as simply weak.

Overall, whilst the long-sample used in this paper mixes fixed and flexible nominal exchange rate regimes, the fact that regimes have been found not to matter indicates that the findings in this paper are robust with respect to structural stability. Put differently, although this paper cannot answer the question of whether RIP would hold with a century long flexible exchange rate arrangement, regime neutrality implies that it safely answers the question of whether RIP has held over our long sample.

Finally, given that RIP is based on two parity conditions which are UIP and PPP, rejection of real interest parity during the recent float can be attributed to the failure of financial and/or goods market integration. However, the fact we have been able to uncover positive evidence on RIP for the UK and France, and somehow weak evidence for Japan, using long spans of data is not surprising. Both UIP and PPP have been found to hold better over the very long-run and this essentially adds to the growing consensus that at long horizons, parity conditions exert greater force on international goods and assets markets so that fundamentals matter (Flood and Taylor, 1996; Fujii and Chinn, 2000).

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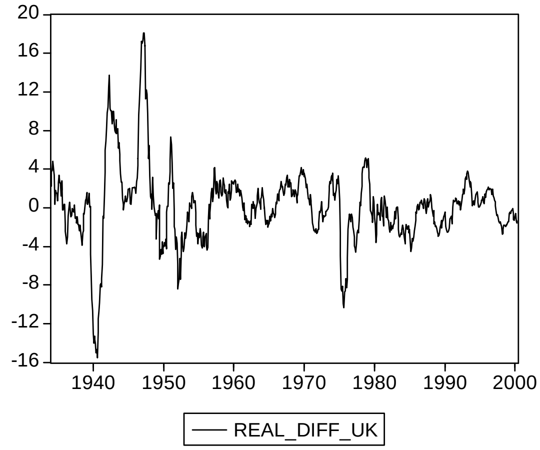
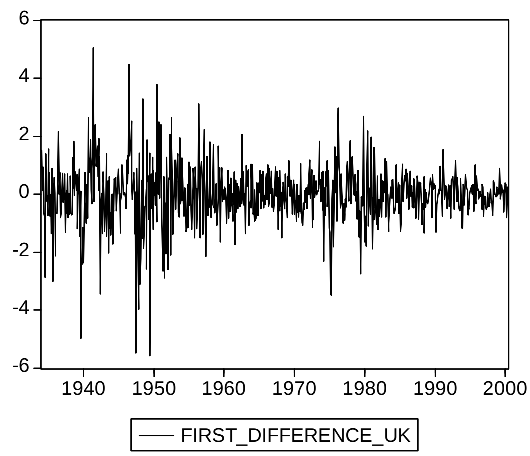
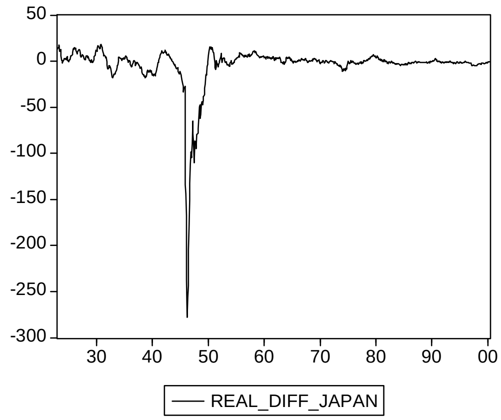
7. Empirical results

Figure 1 Real Interest differentials in level and first difference form.







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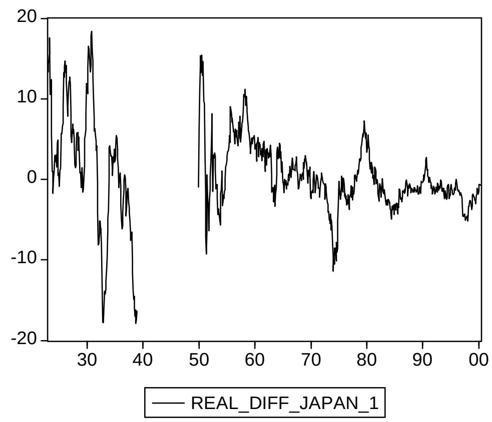

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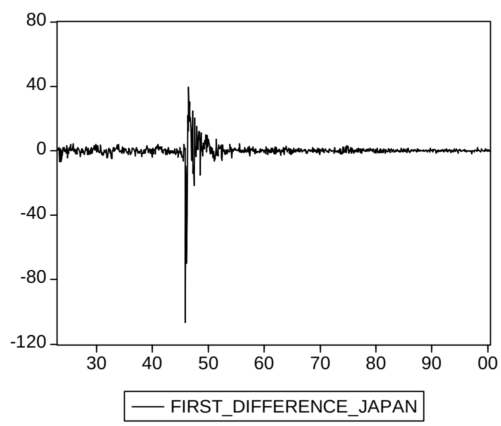

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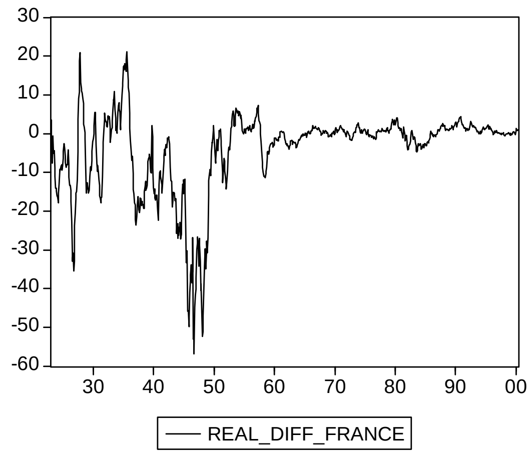
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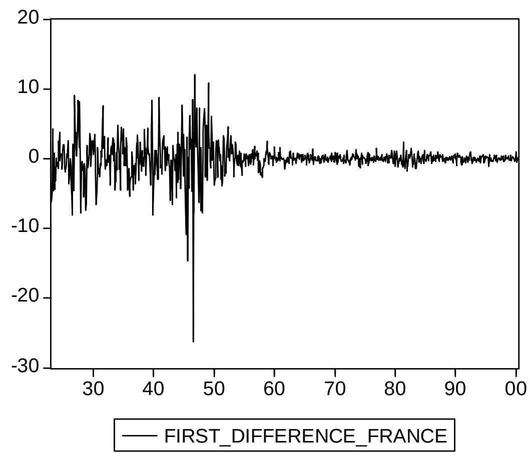
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Table 1 Data description.

	Consumer price index	Long term interest rate
	Interval	Interval
USA	1875- 2003	1800- 2000
UK	1906- 2003	1933- 2003
Japan	1922- 2003	1871- 2003
France	1915- 2003	1746- 2003

Table 2 Range of estimates of real interest rate differentials in %.

Country	Time period	*n*	average	std. error	*t*-statistic	*p*-value	std. dev (RIRD)
UK	1974:01- 2000:06	318	-0.5807	0.1408	-4.1238	0.0000***	2.5115
	1950:01- 1973:12	288	0.1197	0.1438	0.8324	0.4058	2.4410
	1950:01- 2000:06	606	-0.2478	0.1015	-2.4398	0.0150**	2.5008
	1939:01- 1949:12	132	2.1590	0.6839	3.1570	0.0020***	7.8572
	1934:01- 1938:12	60	-0.0019	0.2716	-0.0068	0.9946	2.1038
	1934:01- 2000:06	798	0.1687	0.1416	1.1914	0.2338	4.0016
Japan	1974:01- 2000:06	318	-1.3736	0.1463	-9.3868	0.0000***	2.6095
	1950:01- 1973:12	288	1.9601	0.2478	7.9098	0.0000***	4.2055
	1950:01- 2000:06	606	0.2107	0.1559	1.3514	0.1770	3.8385
	1939:01- 1949:12	132	-37.2090	5.0927	-7.3063	0.0000***	58.5107
	1923:01- 1938:12	192	1.5823	0.6201	2.5517	0.0115**	8.5925
	1923:01- 2000:06	930	-4.8111	0.8561	-5.6198	0.0000***	26.1075
France	1974:01- 2000:06	318	0.4138	0.0948	4.3649	0.0000***	1.6906
	1950:01- 1973:12	288	-0.7556	0.2157	-3.5023	0.0005***	3.6617
	1950:01- 2000:06	606	-0.1419	0.1163	-1.2207	0.2227	2.8635
	1939:01- 1949:12	132	-20.7525	1.2425	-16.7026	0.0000***	14.2749
	1923:01- 1938:12	192	-5.2227	0.8501	-6.1432	0.0000***	11.7803
	1923:01- 2000:06	930	-4.1162	0.3477	-11.8357	0.0000***	10.6060

Note: *, ** and *** indicate significance at 10%, 5% and 1%.

Table 3 Unit root tests: DF-GLS and Ng and Perron (2001).

Country	Time period	MAIC: *p*	DF-GLS	MZ(α)	MZ(*t*)
UK	1974:01- 2000:06	15	-1.4057	-2.2570	-0.9622
	1950:01- 1973:12	13	-1.5004	-1.1909	-0.5426
	1950:01- 2000:06	15	-1.9257*	-6.1257*	-1.7300**
	1939:01- 1949:12	12	-2.3391**	-5.9795*	-1.6911*
	1934:01- 1938:12	0	N/A	N/A	N/A
	1934:01- 2000:06	15	-3.9873***	-43.0360***	-4.6222***
Japan	1974:01- 2000:06	12	-0.5658	-0.6387	-0.4183
	1950:01- 1973:12	14	-1.2469	-1.7842	-0.7500
	1950:01- 2000:06	14	-2.3479**	-8.2454**	-2.0304**
	1939:01- 1949:12	12	-1.2469	-4.1509	-1.4345
	1923:01- 1938:12	17	-0.8236	-3.7382	-1.0622
	1923:01- 2000:06	17	-2.9067***	-19.4640***	-3.1127***
France	1974:01- 2000:06	12	-1.6842*	-6.6254*	-1.8161*
	1950:01- 1973:12	16	-3.2094***	-11.6545**	-2.4111**
	1950:01- 2000:06	14	-3.2750***	-20.2779***	-3.1815***
	1939:01- 1949:12	12	-1.0657	-3.0204	-1.1724
	1923:01- 1938:12	12	-1.4994	-5.2497	-1.5349
	1923:01- 2000:06	15	-2.7721***	-15.0111***	-2.7384***

Note: MAIC is the modified Akaike Information Criterion of Ng and Perron (2001). *, ** and *** indicate significance at 10%, 5% and 1%.

Critical values for the unit root tests.

	DF-GLS	MZ(α)	MZ(*t*)
1%	-2.5723	-13.8	-2.58
5%	-1.9418	-8.1	-1.98
10%	-1.616	-5.7	-1.62

Table 4 OLS estimation and confidence intervals based on Hansen's (1999) grid bootstrap method.

Country	Time period	MAIC: *p*	OLS	std. error	90%	90%	95%	95%
UK	1974:01- 2000:06	15	0.9350	0.0208	0.9078	0.9862	0.9009	1.0018
	1950:01- 1973:12	13	0.9342	0.0237	0.9043	1.0054	0.8961	1.0097
	1950:01- 2000:06	15	0.9249	0.0164	0.9021	0.9580	0.8966	0.9637
	1939:01- 1949:12	12	0.9203	0.0201	0.8910	0.9657	0.8835	0.9737
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	0.9470	0.0099	0.9327	0.9663	0.9292	0.9698
Japan	1974:01- 2000:06	12	0.9204	0.0191	0.8938	0.9622	0.8881	0.9678
	1950:01- 1973:12	14	0.9630	0.0271	0.9360	1.0266	0.9256	1.0320
	1950:01- 2000:06	14	0.9629	0.0136	0.9467	1.0035	0.9420	1.0058
	1939:01- 1949:12	12	0.9558	0.0233	0.9334	1.0192	0.9245	1.0239
	1923:01- 1938:12	17	0.9549	0.0178	0.9298	1.0070	0.9243	1.0091
	1923:01- 2000:06	17	0.9718	0.0069	0.9633	0.9843	0.9607	0.9866
France	1974:01- 2000:06	12	0.9541	0.0190	0.9318	1.0093	0.9256	1.0120
	1950:01- 1973:12	16	0.9383	0.0162	0.9151	0.9740	0.9100	0.9799
	1950:01- 2000:06	14	0.9490	0.0120	0.9324	0.9744	0.9283	0.9788
	1939:01- 1949:12	12	0.9328	0.0348	0.8924	1.0255	0.8783	1.0322
	1923:01- 1938:12	12	0.9301	0.0205	0.9018	0.9788	0.8943	0.9888
	1923:01- 2000:06	15	0.9747	0.0071	0.9652	0.9903	0.9631	0.9930

Note: Confidence intervals for OLS estimates are constructed using Hansen's (1999) grid bootstrap method. MAIC is the modified Akaike Information Criterion of Ng and Perron (2001).

Table 5 Confidence intervals for the dominant root based on Median Unbiased Estimation (MUE).

Country	Time period	MAIC: p	OLS	std. error	MUE	90%	90%	95%	95%
UK	1974:01- 2000:06	15	0.9350	0.0213	0.9789	0.9613	0.9987	0.9578	1.0000
	1950:01- 1973:12	13	0.9342	0.0243	0.9783	0.9580	1.0000	0.9539	1.0000
	1950:01- 2000:06	15	0.9249	0.0166	0.9737	0.9597	0.9882	0.9570	0.9910
	1939:01- 1949:12	12	0.9218	0.0210	0.9271	0.8940	0.9617	0.8872	0.9699
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	0.9470	0.0100	0.9539	0.9386	0.9687	0.9359	0.9720
Japan	1974:01- 2000:06	12	0.9216	0.0193	0.9923	0.9822	1.0000	0.9801	1.0000
	1950:01- 1973:12	14	0.9635	0.0278	0.9746	0.9406	1.0000	0.9334	1.0000
	1950:01- 2000:06	14	0.9631	0.0138	0.9661	0.9435	0.9885	0.9397	0.9929
	1939:01- 1949:12	12	0.9559	0.0244	0.9675	0.9353	1.0000	0.9285	1.0000
	1923:01- 1938:12	17	0.9544	0.0186	0.9792	0.9549	1.0000	0.9503	1.0000
	1923:01- 2000:06	17	0.9718	0.0069	0.9777	0.9692	0.9858	0.9675	0.9875
France	1974:01- 2000:06	12	0.9541	0.0194	0.9643	0.9360	0.9944	0.9301	1.0000
	1950:01- 1973:12	16	0.9383	0.0166	0.9533	0.9316	0.9756	0.9271	0.9795
	1950:01- 2000:06	14	0.9490	0.0122	0.9629	0.9469	0.9801	0.9442	0.9835
	1939:01- 1949:12	12	0.9327	0.0363	0.9599	0.9161	1.0000	0.9079	1.0000
	1923:01- 1938:12	12	0.9303	0.0212	0.9378	0.9059	0.9715	0.8992	0.9788
	1923:01- 2000:06	15	0.9747	0.0071	0.9795	0.9702	0.9895	0.9684	0.9914

Note: Median Unbiased Estimate (MUE) and confidence intervals are constructed using Hansen's (1999) grid bootstrap method using the efficiently demeaned DF-GLS statistic.

Table 6 Half-lives in months: Confidence intervals based on OLS and normal sampling distributions.

Country	Time period	MAIC: *p*	*h*	90_*l*	90_*u*	95_ *l*	95_ *u*
UK	1974:01- 2000:06	15	10.3133	4.6978	15.9289	3.6225	17.0042
	1950:01- 1973:12	13	10.1836	3.9398	16.4275	2.7441	17.6232
	1950:01- 2000:06	15	8.8786	5.5613	12.1958	4.9261	12.8310
	1939:01- 1949:12	12	8.3456	4.7355	11.9557	4.0442	12.6470
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	12.7285	8.7089	16.7481	7.9392	17.5178
Japan	1974:01- 2000:06	12	8.3565	4.9174	11.7956	4.2588	12.4542
	1950:01- 1973:12	14	18.3850	0	40.9590	0	45.2816
	1950:01- 2000:06	14	18.3345	7.0668	29.6021	4.9091	31.7598
	1939:01- 1949:12	12	15.3328	1.7316	28.9340	0	31.5385
	1923:01- 1938:12	17	15.0199	5.0398	25.0000	3.1287	26.9111
	1923:01- 2000:06	17	24.2315	14.3375	34.1255	12.4429	36.0201
France	1974:01- 2000:06	12	14.7520	4.4671	25.0368	2.4976	27.0063
	1950:01- 1973:12	16	10.8839	6.0301	15.7377	5.1006	16.6672
	1950:01- 2000:06	14	13.2415	7.9798	18.5033	6.9722	19.5109
	1939:01- 1949:12	12	9.9640	1.1737	18.7544	0	20.4377
	1923:01- 1938:12	12	9.5655	4.7794	14.3516	3.8629	15.2681
	1923:01- 2000:06	15	27.0491	14.4008	39.6974	11.9787	42.1194

Note: The half-lives are shown in months. Since the lower bound of the confidence interval cannot be negative, it is replaced by zero.

Table 7 Half-lives in months: Corrected confidence intervals based on OLS and normal sampling distributions.

Country	Time period	MAIC: *p*	*b*	*h**	90_*l*	90_*u*	95_ *l*	95_ *u*
UK	1974:01- 2000:06	15	0.6129	17.5965	11.9809	23.2120	10.9056	24.2873
	1950:01- 1973:12	13	0.8494	12.5822	6.3383	18.8261	5.1427	20.0218
	1950:01- 2000:06	15	0.6539	14.3190	11.0018	17.6363	10.3666	18.2715
	1939:01- 1949:12	12	0.0435	46.0865	42.4764	49.6966	41.7851	50.3879
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	0.3925	29.9036	25.8840	33.9232	25.1143	34.6929
Japan	1974:01- 2000:06	12	0.4839	17.1073	13.6682	20.5465	13.0097	21.2050
	1950:01- 1973:12	14	1.5073	7.5008	0	30.0748	0	34.3975
	1950:01- 2000:06	14	1.3503	10.3907	0	21.6584	0	23.8161
	1939:01- 1949:12	12	0.5085	30.2912	16.6901	43.8924	14.0856	46.4969
	1923:01- 1938:12	17	0.1158	61.7301	51.7500	71.7102	49.8389	73.6213
	1923:01- 2000:06	17	0.5494	45.1676	35.2736	55.0616	33.3790	56.9562
France	1974:01- 2000:06	12	0.9710	15.3788	5.0939	25.6637	3.1245	27.6332
	1950:01- 1973:12	16	0.2234	34.4167	29.5629	39.2705	28.6334	40.2000
	1950:01- 2000:06	14	0.5396	25.0282	19.7665	30.2900	18.7589	31.2976
	1939:01- 1949:12	12	0.5500	18.5594	9.7690	27.3498	8.0858	29.0330
	1923:01- 1938:12	12	0.1780	33.3855	28.5994	38.1716	27.6829	39.0881
	1923:01- 2000:06	15	0.7718	37.1559	24.5076	49.8042	22.0856	52.2262

Note: The half-lives are shown in months. Since the lower bound of the confidence interval cannot be negative, it is replaced by zero.

Table 8 Half-lives in months: Confidence intervals based on OLS and Hansen's (1999) grid bootstrap method.

Country	Time period	MAIC: p	h	90_l	90_u	95_ l	95_ u
UK	1974:01- 2000:06	15	10.3133	7.1657	49.8807	6.6418	∞
	1950:01- 1973:12	13	10.1836	6.8905	∞	6.3184	∞
	1950:01- 2000:06	15	8.8786	6.7276	16.1545	6.3507	18.7463
	1939:01- 1949:12	12	8.3456	6.0059	19.8598	5.5960	26.0073
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	12.7285	9.9488	20.2196	9.4394	22.6035
Japan	1974:01- 2000:06	12	8.3565	6.1738	17.9884	5.8409	21.1778
	1950:01- 1973:12	14	18.3850	10.4800	∞	8.9655	∞
	1950:01- 2000:06	14	18.3345	12.6549	∞	11.6008	∞
	1939:01- 1949:12	12	15.3329	10.0571	∞	8.8296	∞
	1923:01- 1938:12	17	15.0199	9.5231	∞	8.8054	∞
	1923:01- 2000:06	17	24.2315	18.5381	43.8020	17.2884	51.3800
France	1974:01- 2000:06	12	14.7520	9.8128	∞	8.9655	∞
	1950:01- 1973:12	16	10.8839	7.8126	26.3114	7.3496	34.1372
	1950:01- 2000:06	14	13.2415	9.9030	26.7280	9.3165	32.3478
	1939:01- 1949:12	12	9.9641	6.0887	∞	5.3415	∞
	1923:01- 1938:12	12	9.5655	6.7060	32.3478	6.2047	61.5409
	1923:01- 2000:06	15	27.0491	19.5694	71.1113	18.4357	98.6740

Note: The half-lives are shown in months.

Table 9 Half-lives in months: Corrected confidence intervals based on OLS and Hansen's (1999) grid bootstrap method.

Country	Time period	MAIC: *p*	*b*	*h**	90_*l*	90_*u*	95_ *l*	95_ *u*
UK	1974:01- 2000:06	15	0.6128	17.5998	12.2284	85.1220	11.3343	∞
	1950:01- 1973:12	13	0.8493	12.5835	8.5143	∞	7.8073	∞
	1950:01- 2000:06	15	0.6544	14.3101	10.8433	26.0370	10.2357	30.2144
	1939:01- 1949:12	12	0.0435	46.0865	33.1662	109.6709	30.9028	143.6190
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	0.3929	29.8836	23.3574	47.4709	22.1615	53.0678
Japan	1974:01- 2000:06	12	0.4935	16.8708	12.4641	36.3164	11.7921	42.7555
	1950:01- 1973:12	14	1.5279	7.1416	4.0710	∞	3.4826	∞
	1950:01- 2000:06	14	1.3528	10.3416	7.1380	∞	6.5434	∞
	1939:01- 1949:12	12	0.5085	30.2912	19.8685	∞	17.4436	∞
	1923:01- 1938:12	17	0.1299	59.2463	37.5642	∞	34.7331	∞
	1923:01- 2000:06	17	0.5493	45.1755	34.5612	81.6616	32.2314	95.7896
France	1974:01- 2000:06	12	0.9710	15.3783	10.2294	∞	9.3461	∞
	1950:01- 1973:12	16	0.2250	34.3061	24.6253	82.9337	23.1660	107.6005
	1950:01- 2000:06	14	0.5395	25.0305	18.7198	50.5241	17.6109	61.1472
	1939:01- 1949:12	12	0.5500	18.5594	11.3410	∞	9.9492	∞
	1923:01- 1938:12	12	0.1762	33.5244	23.5026	113.3698	21.7456	∞
	1923:01- 2000:06	15	0.7723	37.1321	26.8642	97.6193	25.3080	135.4564

Note: The half-lives are shown in months.

Table 10 Half-lives in months: Confidence intervals based on Median Unbiased Estimation (MUE).

Country	Time period	MAIC: *p*	MUE	90_*l*	90_*u*	95_ *l*	95_ *u*
UK	1974:01- 2000:06	15	13.1220	9.7975	33.2564	8.9535	40.8005
	1950:01- 1973:12	13	11.9236	11.5037	39.3252	6.9889	∞
	1950:01- 2000:06	15	11.7232	11.4373	13.2530	10.8947	15.1982
	1939:01- 1949:12	12	19.9678	13.4582	26.4618	12.7816	29.9158
	1934:01- 1938:12	0	N/A	N/A	N/A	N/A	N/A
	1934:01- 2000:06	15	16.8794	13.9022	20.2634	13.1497	20.8276
Japan	1974:01- 2000:06	12	18.1324	2.8830	26.5255	2.7880	29.7242
	1950:01- 1973:12	14	11.9537	9.0552	∞	8.1802	∞
	1950:01- 2000:06	14	11.8811	11.7693	42.1507	11.6383	∞
	1939:01- 1949:12	12	49.8082	7.9559	∞	7.6114	∞
	1923:01- 1938:12	17	28.9976	14.2387	∞	13.6260	∞
	1923:01- 2000:06	17	28.8446	14.5171	46.8784	11.9386	56.1716
France	1974:01- 2000:06	12	18.8322	6.9688	∞	6.8109	∞
	1950:01- 1973:12	16	17.0218	11.8520	26.3549	11.7646	34.0498
	1950:01- 2000:06	14	13.6822	11.8473	19.2421	11.7719	26.7814
	1939:01- 1949:12	12	26.9809	6.1450	∞	5.7427	∞
	1923:01- 1938:12	12	15.9487	11.3312	24.3972	10.7987	35.7289
	1923:01- 2000:06	15	22.9540	12.7594	56.7205	12.5366	75.5023

Note: The half-lives are shown in months.

Table 11 Testing regime neutrality.

	MAIC: p	Test	H_{00}	H_{01}	H_{02}	H_{03}
UK	15	1	0.0000***	0.5469	0.5894	0.7434
		2	0.0000***	0.5464	0.5893	0.7444
Japan	17	1	0.8185	1.0000	0.9999	0.6574
		2	0.8237	1.0000	0.9999	0.6579
France	15	1	0.0021***	0.9082	0.9523	0.0000***
		2	0.0015***	0.9092	0.9531	0.0000***

Note: Tests 1 and 2 correspond to the F-and χ^2 statistics respectively. The lag length is the same as the one for the full sample. *, ** and *** indicate significance at 10%, 5% and 1%.