Why Are Real Interest Rates Not Equalized Internationally?

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Abstract

In this study we specify a set of necessary parity conditions for real interest rates to be equalized internationally, what we call real interest parity (RIP). These necessary conditions imply that a system of domestic and foreign nominal interest rates and inflation rates must be driven by one common stochastic trend if any single variable exhibits non-stationarity. Using multivariate unit root tests, which have significantly greater power than univariate alternatives, we demonstrate that these necessary conditions are not satisfied for the G-5 over the period of 1960 to 1996. We then examine each necessary parity condition individually to shed some light on the source of the rejection of RIP. Our results suggest that no single violation can explain the failure of RIP in all cases. It does appear, however, that the Fisher relation is the least likely to violate the RIP equilibrium while uncovered interest parity (UIP) appears to be the most commonly violated. This suggests the existence of non-stationary risk premia in international capital markets.
1 Introduction

International financial economists are intensely concerned with several parity conditions that relate goods prices and asset returns across countries. These include purchasing power parity (PPP), interest rate parity (IRP) and variations of these two, such as relative PPP, ex-ante PPP and real interest rate parity. Each of these parity conditions measures a degree of integration between the economies of the world. The greater the economic integration across countries, the greater the likelihood that the more restrictive of these parity conditions will have empirical support. For example, the strict PPP hypothesis implies that aggregate price levels will be equal in terms of a common currency. This condition is very stringent since it does not allow for even temporary deviations from the equilibrium condition. A less restrictive version is long-run PPP where price levels, measured in a common currency, across countries converge over some period of time, maybe a very long period of time. Recognizing all of the impediments that keep strict PPP from holding, economists have put their efforts in studying the long-run version of this equilibrium condition.

There is a relatively large literature examining the equality of real interest rates internationally. In all of these studies the real interest rate is defined by the Fisher equation. These studies then go on to estimate and test for the simultaneous existence of uncovered interest parity (UIP), which is also called the open-economy Fisher relation, and ex-ante PPP (EAPPP), the other conditions necessary for RIP. The problem in this strategy is that the Fisher equation may not hold. Many studies have been devoted to examining the validity of the Fisher relationship for both the U.S. and other economies. The results of this line of inquiry have been decidedly mixed. Even if the Fisher relation does hold, it is unlikely that the Fisher effect will be one-for-one as implied by defining the real interest rate from the Fisher equation. Both Mundell-Tobin and tax effects can drive a wedge between inflation expectations and the effect on nominal interest rates. It would seem more

1 Presumably goods arbitrage will be instantaneous in such a world preventing price levels from diverging.
4 For an examination of the importance of tax effects in the Fisher relation see Crowder
correct then to treat the Fisher relation as another parity condition to be tested than as an assumption.

In this study we examine all of the underlying parity conditions necessary for (RIP). We treat all parity conditions as hypotheses to be tested. We demonstrate that once it is recognized that RIP is predicated on four parity or equilibrium relationships among four variables, uncovered interest parity (UIP), ex-ante PPP (EAPPP) and the Fisher relation in each country, it is straightforward to derive time series implications for all four variables. These properties implied by RIP represent necessary conditions for RIP to hold. Specifically, for RIP to be true, the nominal interest rates and inflation rates in all countries must have an underlying common trend if any single variable is integrated. So a simple test of the necessary conditions for RIP is a test for the stationarity of these variables and then if any are integrated test for a single common stochastic trend.

Our results for the G-5 countries reject conclusively the RIP conditions. This implies that RIP does not hold between these economies. But we extend the analysis to determine which of the parity conditions may be failing resulting in the violation of RIP. Our results tell us that no single parity condition can explain the failure of RIP in all cases. But it does appear that the Fisher relation is the least likely to violate the RIP equilibrium. On the other hand UIP appears to be the most commonly violated of the four with EAPPP lying somewhere between these.

The rest of the paper is organized as follows: section 2 discusses the theoretical conditions necessary for RIP and the time series implications of these conditions. Section 3 describes the econometric methodology and discusses the various hypothesis tests. Section 4 presents the empirical results and section 5 concludes.

2 Real Interest Rate Parity

2.1 Four Parity Conditions

Real interest parity (RIP) is the condition where real rates of return on essentially identical assets are equalized across countries. There are many reasons why real interest rates will not always be equal across countries, i.e.,

and Wohar (1999).
country-specific risk, transactions costs, information asymmetries, differential tax treatment, etc. For this reason our focus is on long-run RIP.

The nominal interest rate in country $j$ can be related to the real interest rate by the Fisher relationship given in equation (1),

$$i_{j,t} = r_{j,t} + \pi_{j,t+k}^e$$  \hspace{1cm} (1)

where $i_{j,t}$ is the nominal interest rate of country $j$ in period $t$ and $\pi_{j,t+k}^e$ is the ex-ante inflation in country $j$ over the holding period $k$ of the asset.

Nominal interest rates across countries are related via the UIP relationship. UIP relates the nominal interest rate differential to the expected exchange rate depreciation as in equation (2)

$$i_{j,t} - i_{i,t} = \Delta s_{t+k}^e$$  \hspace{1cm} (2)

where $s_{t+k}^e$ is the log of the ex-ante spot exchange rate in period $t + k$ and $\Delta$ is the first difference operator.

Inflation rates are related across economies via the EAPPP relationship. EAPPP relates the inflation differential to the expected exchange rate depreciation as in equation (3).

$$\pi_{j,t+k}^e - \pi_{i,t+k}^e = \Delta s_{t+k}^e$$  \hspace{1cm} (3)

Combining equations (1), (2) and (3) yields the real interest parity relationship.

$$r_{j,t} = r_{i,t}$$  \hspace{1cm} (4)

Equations (1), (2) and (3) must be true for real interest rates to be equalized internationally. It should be clear at this point that RIP is itself based upon the validity of four equilibrium relationships, a Fisher relation in both the foreign and domestic countries relating nominal rates to real rates, the UIP equation relating nominal rates across countries and EAPPP relating expected inflation rates across countries.$^5$

### 2.2 Time Series Implications

We start the discussion of the time series implications of the previous section from the assumption that the observable variables $i_{j,t}$, $\pi_{j,t+k}$ and $s_{t+k}$ are

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$^5$It is possible that differential tax rates internationally will lead to equilibrium relationships that do not imply proportionality between the variables, i.e. that the relationships are not necessarily one-to-one. We allow for this possibility in the empirical section.
integrated of order one or I(1). This assumption implies that these variables have no tendency toward an equilibrium or constant value, they will wander aimlessly over time.

Previous studies have used equation (1) as the basis for testing the Fisher relation by noting that economic theory suggests that $r_{j,t}$ should be a stationary or I(0) variable. This result is derived from the implications of the consumption based CAPM model. In order for the Fisher relation to hold empirically when $i_{j,t}$ and $\pi_{j,t+k}$ are I(1) processes, they must be cointegrated. Cointegration implies that while both individual variables may exhibit stochastic trends through time, some linear combination of the two tends toward a constant value, i.e. they share a long-run equilibrium.

Similarly, if nominal interest rates and the spot exchange rate are first order integrated or I(1), then the log difference of the spot exchange rate is I(0) and equation (2) implies that the two nominal interest rates are also cointegrated.

Finally, if $\pi_{j,t+k}$ and $\pi_{i,t+k}$ are both I(1) variables, then they must also be cointegrated if the EAPPP relation is to be empirically valid. So for real interest rates to be equalized internationally there are four equilibrium relationships or parity conditions that must be satisfied in each two-country pairing, a Fisher relation in each country, UIP and EAPPP. There are four observable variables in these four parity relationships, $i_{j,t}$, $i_{i,t}$, $\pi_{j,t+k}$ and $\pi_{i,t+k}$. Since only three of the four equilibrium relationships are independent, i.e. if any three hold then the fourth must also be true, there are three stationary

\[ \text{We will present evidence supporting this assumption in the empirical portion of the study.} \]
\[ \text{This result has been exploited by Mishkin (1992), Crowder and Hoffman (1996) and Crowder (1998), among others.} \]
\[ \text{See Rose (1988) and Crowder and Hoffman (1996) for derivation of this result.} \]
\[ \text{While cointegration is a necessary condition for the Fisher relation to be true when nominal interest rates and inflation rates are I(1), it is not sufficient. A generalized version} \]
\[ \text{of equation (1) will include a risk premium term that reflects the conditional covariance} \]
\[ \text{between inflation and the market portfolio return or consumption growth and Siegel’s paradox.} \]
\[ \text{Again, this cointegration restriction represents a necessary but not sufficient condition} \]
\[ \text{for UIP to hold. It is possible that a time-varying but stationary risk premium exists in} \]
\[ \text{the foreign exchange market. Such a risk premium would violate the UIP relationship but} \]
\[ \text{still allow nominal interest rates to be cointegrated across countries.} \]
\[ \text{Since we are using actual inflation as a proxy for expected inflation, the parity condition} \]
\[ \text{that we are in fact testing is relative PPP. It is well understood that if relative PPP holds} \]
\[ \text{then ex-ante PPP will also be true.} \]
equilibrium relationships among the four variables. As demonstrated by Stock and Watson (1988), there must be one common stochastic trend among the four variables.

In order to see this result more clearly it is convenient to appeal to a vector autoregressive (VAR) model in the four observable variables as in equation (5),

\[
\Phi(L)X_t = \mu_t + u_t
\]

where \(X_t\) is the 4 \(\times\) 1 vector of the observable variables, \(i_{j,t}, i_{i,t}, \pi_{j,t+k}, \pi_{i,t+k}\), and \(\Phi(L)\) is a \(p^{th}\)-order matrix polynomial in the lag operator, \(\mu_t\) is a vector of deterministic components and \(u_t\) is a vector of white noise error terms. Equation (5) can be rewritten in error-correction model form as in equation (6),

\[
\Delta X_t = \mu_t + \Pi X_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + u_t
\]

where the long-run impact matrix \(\Pi\) is of particular interest. There are three possibilities with respect to the rank of the \(\Pi\) matrix; 1) \(\Pi\) is of zero rank implying that all of the variables in \(X_t\) are I(1) and not cointegrated, 2) \(\Pi\) is of reduced rank implying that at least some of the variables in \(X_t\) are integrated and that there exists a number of stationary relationships equal to the rank of \(\Pi\), and 3) \(\Pi\) is of full rank implying that all of the variables in \(X_t\) are stationary in levels. The long-run multiplier matrix \(\Pi\) can be decomposed into two matrices \(\alpha \beta'\) where \(\beta\) represent the co-integrating or stationary relationships among \(X_t\) and \(\alpha\) is the matrix of error-correction coefficients.

Equations (1), (2) and (3) imply the following representation for \(\beta' X_{t-1}\),

\[
\begin{bmatrix}
1 & 0 & 0 & -\eta \\
0 & 1 & 0 & -\gamma \\
0 & 0 & 1 & -\delta \\
\end{bmatrix}
\begin{bmatrix}
i_{j,t-1} \\
i_{i,t-1} \\
\pi_{j,t+k-1} \\
\pi_{i,t+k-1} \\
\end{bmatrix}
\]

where \(\eta, \gamma\) and \(\delta\) should be equal to one under a strict RIP interpretation but may differ from unity for reasons suggested earlier.\(^{12}\)

\(^{12}\)\(\beta'\) is only identified up to a matrix normalization. The normalization depicted in (7) is convenient as it implies a triangular representation of the system. Others are also used frequently in the literature. Note that the RIP conditions do not imply a direct relationship between \(i_t\) and \(\pi_{t+k}\) but the indirect relationship can be determined.
From (7) it can be seen that $\Pi$ is of reduced rank equal to three, then the necessary conditions for RIP hold, implying that all of the variables in $X_t$ are stationary in levels. In other words, the four equilibrium relationships necessary for RIP to hold imply that the underlying observable variables in the relationship must exhibit three cointegrating vectors or one common trend. Thus a simple test for the necessary condition for RIP is simply to test each of the individual series for a unit root. If any one is found to be non-stationary, RIP implies that they must share a common stochastic trend.

It is possible though that only some of the necessary equilibrium relationships are violated while others continue to be true. By analyzing each of these relationships individually it is possible to shed light on the sources of RIP violations. In the next section we present the empirical methodology used to test the necessary conditions for RIP and those methods used to analyze the sources of the rejection of RIP.

3 Empirical Methodology

3.1 Unit Root Tests

As a preliminary step we apply unit root tests to the individual variables. But it is well known that the power of such tests is notoriously weak when the autoregressive root is close to, but still less than, unity. Recently several researchers have proposed using a panel unit root testing approach that significantly increases the power of the test against the null.\textsuperscript{13} We apply two common panel unit root tests to our data.

The first is a test suggested by Levin and Lin (1992) based on the model in (8),

$$\Delta y_{it} = \rho_i y_{it-1} + z_{it}^\prime \gamma + u_{it}, \ i = 1, \ldots, N; t = 1, \ldots, T, \quad (8)$$

where $z_{it}$ is the deterministic component and $u_{it}$ is a stationary process. The Levin and Lin test assumes that $\rho_i = \rho$ for all $i$. Levin and Lin suggest a $t$-statistic calculated under the null as

$$t_{\rho} = \frac{(\hat{\rho} - 1) \sqrt{\sum_{i=1}^{N} \sum_{t=1}^{T} \bar{y}_{i,t-1}^2}}{s_u} \quad (9)$$

\textsuperscript{13}See Papell (1998) and the references therein for a discussion of the panel unit root tests as applied to the PPP hypothesis.
where

\[
    s^2_u = \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} \tilde{u}_{it}^2
\]

(10)

and \( \tilde{y}_{it} \) and \( \tilde{u}_{it} \) are simply \( y_{it} \) and \( u_{it} \) corrected for the deterministic components \( z_{it} \).

The second panel test is attributed to Im, Pesaran and Shin (1997) (hereafter IPS). It is based on averaging the individual unit root test statistics from standard augmented Dickey-Fuller regressions. Define the \( t \)-bar statistic as

\[
    \bar{t} = \frac{1}{N} \sum_{i=1}^{N} t_{\rho_i},
\]

(11)

where \( t_{\rho_i} \) is the individual \( t \)-statistic for testing the unit root null hypothesis. IPS suggest a statistic normalized by the expected value and variance of the individual \( t \)-statistics and call this statistic \( t_{IPS} \). They derive critical values for this statistic via Monte Carlo simulation.

O’Connell (1998) showed that the Levin and Lin test suffers from extreme size distortion (rejects a true null too often) when the contemporaneous error terms are correlated across groups (referred to as spatial correlation in the literature). O’Connell further demonstrated that once this spatial correlation was controlled for, the power of these tests dropped significantly. Furthermore, Taylor and Sarno (1998) have shown that rejection of the null hypothesis in the panel unit root tests cannot be interpreted as stationarity of all the series in the panel. These tests are uninformative about the number of series that are stationary versus the number that are non-stationary.

Since the univariate unit root tests have low power and the panel unit root tests have serious size distortions and an uninformative alternative hypothesis, we employ two other multivariate unit root tests that have been suggested in the literature. The first is the covariate ADF (CADF) test of Hansen (1995). Consider the simple ADF regression of (12),

\[
    \Delta y_t = \rho y_{t-1} + \sum_{i=1}^{k-1} \zeta_i \Delta y_{t-i} + v_t.
\]

(12)

Usually we observe related time series, say \( x_t \), assuming \( x_t \) is I(1) so that \( \Delta x_t \) is I(0), where the relationship between \( x_t \) and \( y_t \) can be captured as \( e_t = v_t - \Delta x_t \gamma \) so that the process for \( y_t \) can be alternatively written as,

\[
    \Delta y_t = \rho y_{t-1} + \Delta x_t \gamma + \sum_{i=1}^{k-1} \zeta_i \Delta y_{t-i} + e_t.
\]

(13)
The variance of the error in (13) is $\sigma_e^2 = \sigma_v^2 - \sigma_{xv}^2$ and will be smaller than the error variance in (12) unless $\sigma_{xv} = 0$. Thus hypothesis tests in (13) should have greater power than tests using (12). Hansen demonstrates that the power of the CADF test is inversely related to two factors. The first is the long-run (frequency zero) squared correlation between $v_t$ and $e_t$ denoted as $\eta^2$.

$$\eta^2 = \frac{\sigma_{v e}}{\sigma_v^2 \sigma_e^2}$$  \hspace{1cm} (14)

The second factor relating to the power of the CADF test is the variance ratio defined as,

$$R^2 = \frac{\sigma_e^2}{\sigma_v^2}. \hspace{1cm} (15)$$

Hansen (1995) demonstrates that with a sample size of 100 observations and an autoregressive root of 0.95 the power envelope for the standard ADF test is 33%. This increases to 51% when $\eta^2 = 0.7$ and to 90% when $\eta^2 = 0.3$ The distribution of the CADF test depends on the value of $\eta^2$ such that,

$$t_{CADF} = \eta(DF) + (1 - \eta^2)^{1/2} N(0, 1)$$  \hspace{1cm} (16)

where $DF$ represents the Dickey-Fuller distribution and $N(0, 1)$ is the standard normal.

The second multivariate unit root test we employ is Johansen’s (1988, 1991) test for the number of or cointegration vectors in a vector of I(1) variables. The Johansen procedure is based upon a $p^{th}$-order vector autoregressive (VAR) process transformed into error-correction model (VECM) form as in (6).

Maximum likelihood estimation of (6) can be carried out by applying reduced rank regression. Johansen (1988, 1991) suggests first concentrating out the short-run dynamics by regressing $\Delta X_t$ and $X_{t-1}$ on $\Delta X_{t-1}$, $\Delta X_{t-2}$, ..., $\Delta X_{t-k+1}$ and $\mu_t$ and saving the residuals as $R_{0t}$ and $R_{1t}$, respectively. Calculate the product moment matrices $S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it} R_{jt}'$ and solve the eigenvalue problem $|\lambda S_{11} - S_{10} S_{00}^{-1} S_{01}| = 0$. The likelihood ratio statistic testing the rank of $\Pi$ is given by $-T \sum_{i=r+1}^{p} \ln(1 - \hat{\lambda}_i)$ and is called the trace statistic by Johansen (1988, 1991).

The distribution of this statistic is non-standard and depends upon, among other things, the specification of the deterministic components $\mu_t$ in the VAR.
If the series exhibit deterministic trends in the levels and non-zero means in the differences, then it will be appropriate to include a constant in \( \mu_t \). If, however, the data are better characterized as having a non-zero mean in the levels and a zero mean in the first differences it will be appropriate to set \( \mu_t = 0 \) and include a constant term in the \( X_t \) vector. The third possibility is that there are no statistically significant deterministic terms in the cointegration vectors implying that \( \mu_t = 0 \) and the \( X_t \) vector will not contain a constant term.\(^{14}\)

This hypothesis can be formally tested using a LR test procedure discussed in Johansen (1994). We consider the hypothesis of \( H_1^*(r) \), no trends in the data but a constant term in the cointegrating relation, versus \( H_1(r) \), trends in the levels data. The form of the LR test is given in equation (17).

\[
-2 \ln \mathcal{L} \{H_1^*(r) \mid H_1(r)\} = T \sum_{i=r+1}^{p} \ln \left( \frac{1 - \hat{\lambda}_{1i}^*}{1 - \hat{\lambda}_{2i}} \right)
\]  

(17)

Note that the calculated test statistic depends on the value of \( r \), the number of cointegrating relations hypothesized. Johansen shows that this statistic has a \( \chi^2 \) distribution with \( p - r \) degrees of freedom. If we cannot reject this restriction, i.e., no trends in the levels data, we should test for the significance of the constant terms in various cointegrating relations since the trace test distribution will be different if this restriction is also true. This is a test of the hypothesis of \( H_2(r) \), no trends in the data and no constant term in the cointegrating relation, versus \( H_1^*(r) \), no trends in the levels data but a constant in the cointegration vectors. The form of this LR test is given in equation (18),

\[
-2 \ln \mathcal{L} \{H_2(r) \mid H_1^*(r)\} = -T \sum_{i=1}^{r} \ln \left( \frac{1 - \hat{\lambda}_{1i}^*}{1 - \hat{\lambda}_{2i}} \right)
\]  

(18)

where the distribution of the calculated statistic depends on the hypothesized number of cointegrating vectors \( r \) such that the statistic is distributed as a \( \chi^2(r) \) variate. Once the appropriate specification for the deterministic

\(^{14}\)There are other possible specifications for the deterministic components in the VAR. One entails including a trend in the VAR implying a quadratic trend in the data and the other includes a trend in the cointegrating vectors. Both of these specifications were rejected by the data using the proposed LR tests. The results, available upon request, are omitted for brevity.
components is established, we may proceed to test the rank of $\Pi$ using the appropriate set of critical values.\textsuperscript{15}

### 3.2 Tests of Parity Conditions

We use two methodologies to examine the individual parity conditions themselves for each bivariate pairing. The first technique employed in this phase of the analysis is the Johansen procedure. This methodology has the benefit of making no a priori restrictions or normalizations on the possible equilibrium relationships. If and when cointegration has been established, hypothesis tests on the cointegrating parameters can be conducted in order to determine if the particular parity condition in fact holds. For example, when testing for UIP, not only is it necessary that nominal interest rates are cointegrated across countries but the cointegrating vector $\beta$ must itself span the space $[1 \ -1]\textsuperscript{16}$ The estimate of $\beta$, $\hat{\beta}$, is given by the $r$-largest eigenvectors associated with the eigenvalues $\tilde{\lambda}$. Hypothesis tests on $\hat{\beta}$ can be conducted using likelihood ratio (LR) tests with standard $\chi^2$ inference. Let the form of the linear restrictions on $\beta$ be given by $\beta = H\varphi$ where $H$ is a $p \times s$ matrix of restrictions and $\varphi$ is a $s \times r$ matrix of unknown parameters. The LR test statistic is given by,

$$T \sum_{i=1}^{r} \ln \left[ \frac{(1 - \tilde{\lambda}_i)}{(1 - \lambda_i)} \right] \sim \chi^2_{r(p-s)}$$

where $\tilde{\lambda}_i$ are the eigenvalues from the restricted MLE.

The second technique used to examine the individual parity relationships is a cointegration test proposed by Horvath and Watson (1995) (henceforward HW) in which the cointegration parameters are pre-specified. This technique is based on estimating (6) while imposing a restricted $\beta$. It has the advantage of increased power against the null of no cointegration, but does require knowledge of the cointegrating vector(s), usually from theoretical restrictions derived from economic theory. Specifically, the HW test is a generalization of the Wald test for the significance of $\alpha$ in (6) when $\beta$ is prespecified.

\textsuperscript{15} We use critical values that have been tabulated by MacKinnon et al.\textsuperscript{(1996)} using response surface regressions. These have been shown to approximate the true asymptotic distributions more closely than standard Monte Carlo derived critical values.

\textsuperscript{16} Again it should be kept in mind that differential tax treatment, among other reasons, may cause the cointegration vectors to differ from the theoretical values without necessarily invalidating RIP.
Our strategy for the empirical analysis is straightforward. The first step is to test the univariate orders of integration for each of the variables in the RIP relationship. Once it has been established that we can reject the necessary conditions for RIP, we proceed to analyze the individual equilibrium conditions of UIP, EAPP, and the Fisher relation in each country, testing not only for an equilibrium between the germane variables but also testing the parameter restrictions implied by each of these parity conditions. The next section will discuss the data and the results from the empirical analysis.

4 Empirical Results

4.1 Data

The data used in our study consist of monthly observations on twelve-month eurocurrency deposit rates and annualized inflation rates for the United States (US), United Kingdom (UK), Canada (CN), Germany (GM), and Japan (JP) over the period February 1960 to April 1996.\footnote{The inflation rates are calculated as twelve times the log difference of the respective CPIs. All data are taken from the IFS database of the IMF.} This sample represents five of the largest and most open industrialized economies in the world, which we expect would be as likely as any to yield results favorable to the RIP hypothesis.

4.2 Testing the Necessary Conditions for RIP

The first step in our empirical analysis is to test the order of integration for each of the variables. The first test we present is the panel unit root tests from (9) and (11). Table 1 presents results from treating all variables as a panel and breaking the variables into two groups, an inflation and interest rate group. The results from all panel unit root tests tell the same story, namely that all the series appear to be stationary. But as discussed earlier the interpretation of these results is made difficult by the large size distortions in these tests when the error terms from the each of the groups in the panel are contemporaneously correlated. Furthermore, it is impossible to determine if the rejection of the null hypothesis means all variables are stationary or just some are.\footnote{For a discussion of this problem with panel unit root tests see Breuer et al. (2001).}
Table 2 displays the results of the CADF tests for unit roots in each individual series. Since the asymptotic distribution of the test statistic depends on the value of $\eta^2$, the 5% critical values are also presented.\textsuperscript{19} The results in table 2 were estimated from a model that included a constant and time trend in the deterministic portion of the model, seventeen lags of the dependent variable and six lags of the covariates.\textsuperscript{20} The results presented in table 2 are in stark contrast to those in table 1. There is no evidence from the CADF tests against the unit root null. To get an idea of the power increases associated with the CADF tests versus conventional unit root tests we can examine the results from table 5 of Hansen (1995). For example, the U.S. nominal interest rate CADF model yielded $\eta^2$ and $R^2$ estimates of 0.30 each. These values yielded power of 59% to 88% depending on the empirical model specifications (included lags, etc.). Even for the German inflation rate, which yielded the lowest power gains with estimates of $\eta^2$ and $R^2$ on the order of 0.7, the power gain of the CADF test versus the ADF test is 40%, (20% power for the ADF test versus 28% power for the CADF test).\textsuperscript{21}

The next unit root test we present is Johansen’s test of the rank of $\Pi$ from (6). Since the distribution of the test statistics, specifically the trace test, depends upon the specification of the deterministic components in the VECM, we must establish the proper specification for $\mu_t$ in order to make valid inference using this test.\textsuperscript{22} We consider the most relevant general model to be one that includes a constant in the VAR of (6) implying that the individual series have linear trends in their levels representation under the null hypothesis of a unit root. This hypothesis is denoted $H_1(r)$. The alternative to this model is one which restricts the deterministic components of the level series to include a non-zero mean. We denote this hypothesis as $H^*_1(r)$. Table 3 presents the LR test statistics described in (17) which tests $H_1(r)$ versus

\textsuperscript{19}The covariates used in the analysis were the log differences of industrial production for Canada, the U.S., the U.K, Japan, and France and West Texas Intermediate oil price and unemployment from Canada, the U.S. and the U.K. These variables were chosen on the basis of data availability over the entire sample period.

\textsuperscript{20}We estimated models with from 0 to 20 AR lags and up to 8 lags and leads of the covariates and with and without a time trend. We settled on the model with 17 AR lags by utilizing Schwert’s (1987) suggestion of setting the lag truncation at $t_{12} = \text{Int}[12(T/100)^{1/4}]$. Results from all specifications are available upon request.

\textsuperscript{21}Hansen’s table 4 shows that the CADF tests are correctly sized as well.

\textsuperscript{22}Crowder (2001) presents simulation evidence of serious size distortion in the trace tests when the empirical model allows for linear trends in the data but the true data generating process has only non-zero means.
There is no evidence against the restricted model for any choice of \( r \), implying that the data do not contain linear trend components.

Table 4 presents the calculated LR tests of the hypothesis that there are no significant deterministic components in the VAR, i.e. a test that each series has a zero mean. The evidence is somewhat mixed but it generally suggests that it is inappropriate to eliminate the constant from the estimated stationary relationships. Therefore, we will conduct our tests of the rank of \( \Pi \) based on a specification that does not include a constant in the VAR but does allow a constant term in the cointegration vector.

We use the Johansen trace test statistics to determine the rank of \( \Pi \) as outlined in the previous section. We calculate a recursive estimate of the trace statistic in order to determine the sensitivity of our conclusions to the particular sample period. These also are a convenient form to display the results of many tests in a compact manner. We present the results graphically in figure 3 where the test statistics have been normalized by the appropriate 10\% critical value such that values greater than one imply statistical significance.

Figure 3 presents the recursive trace statistics estimated under \( H_1^*(r) \), i.e. a constant term in the cointegrating vector but not in the VAR, which has been determined to be the appropriate specification. In no case is there evidence of more than two cointegrating relationships, implying the necessary conditions for RIP are violated.

The evidence presented from examining the order of integration of the individual RIP component variables strongly implies that the data are non-stationary in levels. Furthermore, the number of estimated cointegrating relationships is less than three for every system. From this we can conclude that the necessary conditions for RIP to hold are not met. In the next section we examine the various parity conditions that form the basis of RIP to see

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23 The lag length in the VAR was chosen such that it was the minimum necessary to eliminate all statistically significant residual serial correlation. This resulted in a lag length of 12 in all cases except for Germany-Canada which required 13 lags.

24 These recursive estimates of the trace statistics are calculated using the model in (6) with the short-run dynamics held constant and equal to their full sample estimates as suggested by Hansen and Johansen (1993).

25 Imposing three cointegrating vectors and examining the normalized estimates from each system generates even more evidence against RIP. Of the 10 four-variable systems analyzed only half yielded all three imposed cointegrating relationships of the appropriate sign. None of these five had all three cointegrating vectors statistically different from \([1 0]^\prime\).
if we can determine a source of the RIP violations.

4.3 Testing Individual Parity Conditions

Table 5 presents the results from our analysis of the Fisher relation in each country. We present result from applying the Johansen tests and tests suggested by Horvath and Watson (1995). The first column of table 5 references the country for which the Fisher relation is being tested. Columns two and three present the calculated trace test statistics for the null hypotheses of \( r = 0 \) and \( r \leq 1 \), respectively. The results from columns 2 and 3 imply strong evidence in favor of cointegration for all but the United Kingdom.\(^{26}\) The tests of the restriction that the cointegrating relationship is one-for-one are displayed in the fourth column. Of the countries with evidence of cointegration between inflation and nominal interest rates, only for Japan can the Fisher restriction be rejected and then only at the 10% level.\(^{27}\)

Columns 5 through 7 of table 5 present results associated with the HW procedure. Column 5 displays HW test statistics when the cointegrating vector between the nominal interest rate and inflation is assumed to be known a priori. In this case we assume that \( \beta = [1, -1]' \). The result from this test are in complete agreement with the Johansen test results. All but the United Kingdom reject the null of zero cointegrating vectors in favor of the alternative of one. Column 6 shows HW tests when the cointegrating relation is treated as unknown. Here the evidence is not as strong in favor of the Fisher relation, but the estimated Fisher effects in column 7 are consistent with the Fisher relationship in four out of the five countries examined.

Table 6 presents the results from testing the ten UIP relationships possible. While there appears to be some strong evidence of cointegration in many of the relationships, there are only three bilateral relationships consistent with the UIP restrictions. These are the United States and the United Kingdom, Canada and the United Kingdom, and the United States and Canada. In the other seven bilateral relations, either the evidence in favor

\(^{26}\)It turns out that the data for the Fisher relationships generally favor the \( H_2(r) \) specification, that is no deterministic components in the equilibrium relationship. When this specification is used the cointegration results are stronger and parameter restrictions are more consistent with the theoretical model.

\(^{27}\)Even in this case, however, the estimated cointegrating relation is consistent with a Fisher effect that includes the effects of taxes on interest income. The so-called Darby effect. See Crowder and Wohar (1999) for a more detailed discussion of this effect.
of cointegration is very weak or the UIP parameter restrictions are strongly rejected. The HW tests generally reinforce the Johansen results. The fact that Germany has a significant UIP cointegrating relationship with all four countries yet none of the UIP coefficient restrictions are supported can be explained by testing the null hypothesis that the German nominal interest rate is stationary by itself or that it forms a trivial, $\beta = [1, 0]^t$, cointegrating relation. This restriction cannot be rejected in any of the German systems. The likelihood ratio tests of this restriction are 0.946 for the United States system, 0.773 for the Canada system and 1.662 and 1.067 for the United Kingdom and Japanese systems, respectively. This statistic is distributed as a $\chi^2(1)$ variate. It would appear that UIP only holds in three out of ten bilateral relationships.

Table 7 presents the evidence for bilateral EAPPJ relationships. The Johansen test evidence here is quite mixed. Half of the bilateral relations exhibit evidence of cointegration at the 10% level but only three at the 5% level. Of the five only two cannot reject the EAPPJ parameter restrictions at the 10% level. The EAPPJ relation seems to be strongest for the Japan - United Kingdom and the United States and Canada systems. Germany seems the least likely to share an EAPPJ equilibrium relationship with any of the other countries. Again, as in tables 5 and 6, the HW tests tend to confirm the Johansen results.

5 Conclusions

In this study we set out to achieve a couple of objectives. First we wanted to demonstrate that real interest rate parity and the underlying parity conditions, uncovered interest parity, ex-ante PPP and the Fisher relation in each country, imply that the underlying variables must all share one common stochastic trend if RIP is to be a valid long-run characterization of the data. Using several procedures we reject this characterization of the data from the G-5 nations.

Our second objective was to identify which of the four underlying parity conditions may be responsible for the RIP rejections. We estimated and tested for twenty-five bilateral equilibrium relationships. From these we were

\[\text{28} \text{This result is at odds with the German Fisher relation results where the stationarity restriction on the German nominal interest rate can be rejected with a calculated LR statistic of 7.298.}\]
able to conclude that the Fisher relation has wide support empirically but that UIP and EAPPP have much more limited support.
References


Table 1: Panel Unit Root Tests

<table>
<thead>
<tr>
<th>Group</th>
<th>LL</th>
<th>t</th>
<th>t_{IPS}</th>
</tr>
</thead>
<tbody>
<tr>
<td>All</td>
<td>-6.99*</td>
<td>-2.62*</td>
<td>-4.84*</td>
</tr>
<tr>
<td>Inflation</td>
<td>-5.78*</td>
<td>-2.25*</td>
<td>-3.05*</td>
</tr>
<tr>
<td>Interest</td>
<td>-6.77*</td>
<td>-2.99*</td>
<td>-5.91*</td>
</tr>
<tr>
<td>5 % critical value</td>
<td>-5.43</td>
<td>-2.15</td>
<td>-1.96</td>
</tr>
</tbody>
</table>

* Denotes significance at the 5% level.

Note: Entries represent the Levin and Lin (1992) and Im et al. (1997) panel unit root test statistics calculated from (9) and (11).

Table 2: Covariate Unit Root Tests

<table>
<thead>
<tr>
<th>Variable</th>
<th>CADF</th>
<th>\hat{\eta}^2</th>
<th>\hat{R}^2</th>
<th>5 % CV</th>
</tr>
</thead>
<tbody>
<tr>
<td>\pi_{US}</td>
<td>-0.07</td>
<td>0.28</td>
<td>0.30</td>
<td>-2.69</td>
</tr>
<tr>
<td>i_{US}</td>
<td>0.02</td>
<td>0.07</td>
<td>0.17</td>
<td>-2.31</td>
</tr>
<tr>
<td>\pi_{UK}</td>
<td>-1.92</td>
<td>0.46</td>
<td>0.34</td>
<td>-2.94</td>
</tr>
<tr>
<td>i_{UK}</td>
<td>-1.33</td>
<td>0.57</td>
<td>0.56</td>
<td>-3.06</td>
</tr>
<tr>
<td>\pi_{CN}</td>
<td>-1.52</td>
<td>0.58</td>
<td>0.52</td>
<td>-3.07</td>
</tr>
<tr>
<td>i_{CN}</td>
<td>0.18</td>
<td>0.27</td>
<td>0.42</td>
<td>-2.68</td>
</tr>
<tr>
<td>\pi_{GM}</td>
<td>-2.34</td>
<td>0.78</td>
<td>0.60</td>
<td>-3.25</td>
</tr>
<tr>
<td>i_{GM}</td>
<td>-1.66</td>
<td>0.44</td>
<td>0.52</td>
<td>-2.92</td>
</tr>
<tr>
<td>\pi_{JP}</td>
<td>-2.38</td>
<td>0.46</td>
<td>0.45</td>
<td>-2.93</td>
</tr>
<tr>
<td>i_{JP}</td>
<td>-1.36</td>
<td>0.40</td>
<td>0.32</td>
<td>-2.87</td>
</tr>
</tbody>
</table>

Note: Results calculated from CADF model in (13) with a constant and trend included. The number of included lagged \( \Delta y_t \) determined by suggestion in Schwert (1987) and set at 17. Six lags and zero leads of the covariates, i.e. \( \Delta X_t \) were included. \( \hat{\eta}^2 \) and \( \hat{R}^2 \) estimated using the Quadratic Spectral kernel with 24 included autocorrelations. * Denotes significance at the 5% level.
Table 3: Test for Linear Trends in Data

<table>
<thead>
<tr>
<th>Country Pair</th>
<th>r = 0</th>
<th>r ≤ 1</th>
<th>r ≤ 2</th>
<th>r ≤ 3</th>
</tr>
</thead>
<tbody>
<tr>
<td>JP-UK</td>
<td>0.747</td>
<td>0.500</td>
<td>0.500</td>
<td>0.306</td>
</tr>
<tr>
<td>CN-UK</td>
<td>0.034</td>
<td>0.034</td>
<td>0.034</td>
<td>0.028</td>
</tr>
<tr>
<td>CN-JP</td>
<td>0.549</td>
<td>0.452</td>
<td>0.311</td>
<td>0.290</td>
</tr>
<tr>
<td>CN-US</td>
<td>0.190</td>
<td>0.188</td>
<td>0.153</td>
<td>0.034</td>
</tr>
<tr>
<td>JP-US</td>
<td>0.790</td>
<td>0.669</td>
<td>0.551</td>
<td>0.496</td>
</tr>
<tr>
<td>UK-US</td>
<td>0.128</td>
<td>0.116</td>
<td>0.105</td>
<td>0.102</td>
</tr>
<tr>
<td>GM-US</td>
<td>0.315</td>
<td>0.290</td>
<td>0.172</td>
<td>0.169</td>
</tr>
<tr>
<td>GM-CN</td>
<td>0.095</td>
<td>0.048</td>
<td>0.015</td>
<td>0.006</td>
</tr>
<tr>
<td>GM-JP</td>
<td>0.466</td>
<td>0.456</td>
<td>0.020</td>
<td>0.000</td>
</tr>
<tr>
<td>GM-UK</td>
<td>0.194</td>
<td>0.109</td>
<td>0.066</td>
<td>0.062</td>
</tr>
<tr>
<td>5 % critical value</td>
<td>9.49</td>
<td>7.81</td>
<td>5.99</td>
<td>3.84</td>
</tr>
</tbody>
</table>

**Note:** Entries represent the calculated LR test statistics from (17) where the null hypothesis is no linear trends in the data, only a non-zero mean in the stationary relationships. r represents the hypothesized number of cointegrating vectors under the null hypothesis. ∗— Denotes significance at the 5% level.

Table 4: Test for Non-Zero Means in Data

<table>
<thead>
<tr>
<th>Country Pair</th>
<th>r = 0</th>
<th>r ≤ 1</th>
<th>r ≤ 2</th>
<th>r ≤ 3</th>
</tr>
</thead>
<tbody>
<tr>
<td>JP-UK</td>
<td>0.880</td>
<td>4.904</td>
<td>8.112*</td>
<td>11.746*</td>
</tr>
<tr>
<td>CN-UK</td>
<td>4.308*</td>
<td>5.159</td>
<td>9.599*</td>
<td>12.575*</td>
</tr>
<tr>
<td>CN-JP</td>
<td>5.432*</td>
<td>13.379*</td>
<td>16.293*</td>
<td>20.186*</td>
</tr>
<tr>
<td>CN-US</td>
<td>0.038</td>
<td>4.488</td>
<td>6.218</td>
<td>10.847*</td>
</tr>
<tr>
<td>UK-US</td>
<td>5.770*</td>
<td>6.028*</td>
<td>8.712*</td>
<td>12.926*</td>
</tr>
<tr>
<td>GM-US</td>
<td>0.514</td>
<td>8.568*</td>
<td>11.017*</td>
<td>16.113*</td>
</tr>
<tr>
<td>GM-CN</td>
<td>0.180</td>
<td>12.235*</td>
<td>12.424*</td>
<td>17.476*</td>
</tr>
<tr>
<td>GM-JP</td>
<td>0.057</td>
<td>3.437</td>
<td>7.205</td>
<td>11.247*</td>
</tr>
<tr>
<td>GM-UK</td>
<td>0.884</td>
<td>2.731</td>
<td>5.635</td>
<td>10.887*</td>
</tr>
<tr>
<td>5 % critical value</td>
<td>3.84</td>
<td>5.99</td>
<td>7.81</td>
<td>9.49</td>
</tr>
</tbody>
</table>

**Note:** Entries represent the calculated LR test statistics from (18) where the null hypothesis is a zero mean in the stationary relationships. r represents the hypothesized number of cointegrating vectors under the null hypothesis. ∗— Denotes significance at the 5% level.
<table>
<thead>
<tr>
<th>Country</th>
<th>$r = 0$</th>
<th>$r \leq 1$</th>
<th>$H_0 : \beta = [1 \ -1]'$</th>
<th>$W_{0,1}(0, \alpha_k)$</th>
<th>$W_{0,1}(0,0)$</th>
<th>$\hat{\beta}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>US</td>
<td>17.11**</td>
<td>7.37**</td>
<td>0.01</td>
<td>21.25*</td>
<td>12.21</td>
<td>-0.95</td>
</tr>
<tr>
<td>UK</td>
<td>12.84</td>
<td>4.70</td>
<td>3.20**</td>
<td>12.70</td>
<td>8.06</td>
<td>2.24</td>
</tr>
<tr>
<td>CN</td>
<td>16.77**</td>
<td>5.08</td>
<td>0.71</td>
<td>16.56**</td>
<td>11.51</td>
<td>-1.32</td>
</tr>
<tr>
<td>GM</td>
<td>33.82*</td>
<td>7.61</td>
<td>1.08</td>
<td>23.47*</td>
<td>16.62*</td>
<td>-1.46</td>
</tr>
<tr>
<td>JP</td>
<td>24.69*</td>
<td>7.09</td>
<td>3.66**</td>
<td>17.62**</td>
<td>12.62**</td>
<td>-1.73</td>
</tr>
<tr>
<td>5 % critical value</td>
<td>17.98</td>
<td>7.56</td>
<td>3.84</td>
<td>18.17</td>
<td>14.18</td>
<td>NA</td>
</tr>
</tbody>
</table>

**Note:** Entries in columns two and three represent the calculated trace test statistics where $r$ is the hypothesized number of cointegrating vectors. These statistics were estimated from a VECM that included a constant and lags determined by minimizing the SBC criterion. Column four is the LR test statistic that the cointegrating vector is equal to $[1, -1]'$ of which the null hypothesis $H_0$ is distributed as a $\chi^2_1$ variate. Column five is the HW statistic calculated under the assumed null of one known cointegrating vector equal to $[1, -1]'$ and one unknown vector under the alternative. Column six is the HW statistic calculated under the null of zero known cointegrating vectors and one unknown vector under the alternative. Column seven is the estimated cointegrating parameter from the estimated model of column six. HW estimated model included a constant and lags determined by minimizing the SBC criterion. *— Denotes significance at the 5% level and **— denotes significance at the 10% level.
Table 6: Test of UIP Relation

<table>
<thead>
<tr>
<th>Country Pair</th>
<th>$r = 0$</th>
<th>$r \leq 1$</th>
<th>$H_0: \beta = [1 - 1]'$</th>
<th>$W_{0.1}(0, \alpha_k)$</th>
<th>$W_{0.1}(0, 0)$</th>
<th>$\hat{\beta}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>JP-UK</td>
<td>12.21</td>
<td>4.32</td>
<td>3.56**</td>
<td>14.40</td>
<td>8.22</td>
<td>1.80</td>
</tr>
<tr>
<td>CN-UK</td>
<td>22.81*</td>
<td>6.24</td>
<td>0.03</td>
<td>34.48*</td>
<td>29.76*</td>
<td>-1.00</td>
</tr>
<tr>
<td>CN-JP</td>
<td>13.58</td>
<td>5.85</td>
<td>1.82</td>
<td>14.01</td>
<td>8.69</td>
<td>-1.54</td>
</tr>
<tr>
<td>CN-US</td>
<td>17.97**</td>
<td>7.05</td>
<td>0.23</td>
<td>21.49*</td>
<td>14.89*</td>
<td>-1.00</td>
</tr>
<tr>
<td>JP-US</td>
<td>14.38</td>
<td>4.84</td>
<td>4.41*</td>
<td>14.01</td>
<td>8.51</td>
<td>-1.68</td>
</tr>
<tr>
<td>UK-US</td>
<td>18.76*</td>
<td>6.63</td>
<td>0.18</td>
<td>23.42*</td>
<td>18.59*</td>
<td>-1.34</td>
</tr>
<tr>
<td>GM-US</td>
<td>30.27*</td>
<td>4.23</td>
<td>18.91*</td>
<td>23.52*</td>
<td>17.90*</td>
<td>-25.54</td>
</tr>
<tr>
<td>GM-CN</td>
<td>30.78*</td>
<td>5.65</td>
<td>16.44*</td>
<td>21.22*</td>
<td>16.40*</td>
<td>-6.55</td>
</tr>
<tr>
<td>GM-JP</td>
<td>34.30*</td>
<td>7.41</td>
<td>16.56*</td>
<td>22.00*</td>
<td>15.88*</td>
<td>0.21</td>
</tr>
<tr>
<td>GM-UK</td>
<td>30.55*</td>
<td>6.19</td>
<td>13.71*</td>
<td>22.46*</td>
<td>15.89*</td>
<td>-8.35</td>
</tr>
<tr>
<td>5 % critical value</td>
<td>17.98</td>
<td>7.56</td>
<td>3.84</td>
<td>18.17</td>
<td>14.18</td>
<td>NA</td>
</tr>
</tbody>
</table>

Note: See notes to table 5

Table 7: Test of EAPP Relation

<table>
<thead>
<tr>
<th>Country Pair</th>
<th>$r = 0$</th>
<th>$r \leq 1$</th>
<th>$H_0: \beta = [1 - 1]'$</th>
<th>$W_{0.1}(0, \alpha_k)$</th>
<th>$W_{0.1}(0, 0)$</th>
<th>$\hat{\beta}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>JP-UK</td>
<td>19.62*</td>
<td>3.61</td>
<td>1.03</td>
<td>19.30*</td>
<td>15.72*</td>
<td>-1.34</td>
</tr>
<tr>
<td>CN-UK</td>
<td>17.87**</td>
<td>2.45</td>
<td>3.36**</td>
<td>17.58**</td>
<td>15.14*</td>
<td>-0.65</td>
</tr>
<tr>
<td>CN-JP</td>
<td>12.32</td>
<td>3.79</td>
<td>0.01</td>
<td>12.19</td>
<td>8.42</td>
<td>-1.07</td>
</tr>
<tr>
<td>CN-US</td>
<td>24.30*</td>
<td>4.38</td>
<td>0.87</td>
<td>23.78*</td>
<td>19.46*</td>
<td>-1.14</td>
</tr>
<tr>
<td>JP-US</td>
<td>12.22</td>
<td>5.68</td>
<td>0.12</td>
<td>12.07</td>
<td>6.43</td>
<td>-1.63</td>
</tr>
<tr>
<td>UK-US</td>
<td>20.33*</td>
<td>6.41</td>
<td>3.11**</td>
<td>20.03*</td>
<td>13.69**</td>
<td>-1.72</td>
</tr>
<tr>
<td>GM-US</td>
<td>16.12</td>
<td>5.67</td>
<td>2.92**</td>
<td>15.91**</td>
<td>10.30</td>
<td>-0.36</td>
</tr>
<tr>
<td>GM-CN</td>
<td>14.39</td>
<td>3.65</td>
<td>4.01*</td>
<td>14.22</td>
<td>10.59</td>
<td>-2.77</td>
</tr>
<tr>
<td>GM-JP</td>
<td>17.78**</td>
<td>4.55</td>
<td>5.32*</td>
<td>17.49**</td>
<td>12.97**</td>
<td>-2.92</td>
</tr>
<tr>
<td>GM-UK</td>
<td>14.69</td>
<td>4.45</td>
<td>4.68*</td>
<td>14.51</td>
<td>10.09</td>
<td>-4.09</td>
</tr>
<tr>
<td>5 % critical value</td>
<td>17.98</td>
<td>7.56</td>
<td>3.84</td>
<td>18.17</td>
<td>14.18</td>
<td>NA</td>
</tr>
</tbody>
</table>

Note: See notes to table 5
Nominal Interest Rates and Inflation Rates

Levels

US Nominal Rate

US Inflation

UK Nominal Rate

UK Inflation

Canada Nominal Rate

Canada Inflation

German Nominal Rate

German Inflation

Japan Nominal Rate

Japan Inflation
Nominal Interest Rates and Inflation Rates

First Differences

US Nominal Rate

US Inflation

UK Nominal Rate

UK Inflation

Canada Nominal Rate

Canada Inflation

German Nominal Rate

German Inflation

Japan Nominal Rate

Japan Inflation
Rolling Trace Test Statistics

Constant Term in Cointegrating Vector

US-UK System

0.0 0.2 0.4 0.6 0.8 1.0 1.2 1.4 1.6 1.8 2.0

US-CN System

0.0 0.25 0.50 0.75 1.00 1.25 1.50 1.75 2.00 2.25 2.50

US-JP System

0.0 0.25 0.50 0.75 1.00 1.25 1.50 1.75 2.00 2.25 2.50

US-GM System

0.0 0.2 0.4 0.6 0.8 1.0 1.2 1.4 1.6 1.8 2.0

JP-UK System

0.0 0.2 0.4 0.6 0.8 1.0 1.2 1.4 1.6 1.8 2.0

JP-CN System

0.0 0.25 0.50 0.75 1.00 1.25 1.50 1.75 2.00 2.25 2.50

JP-GM System

0.0 0.2 0.4 0.6 0.8 1.0 1.2 1.4 1.6 1.8 2.0

UK-CN System

0.0 0.25 0.50 0.75 1.00 1.25 1.50 1.75 2.00 2.25 2.50

UK-GM System

0.0 0.2 0.4 0.6 0.8 1.0 1.2 1.4 1.6 1.8 2.0

CN-GM System

0.0 0.2 0.4 0.6 0.8 1.0 1.2 1.4 1.6 1.8 2.0