DIW BERLIN

1667

Discussion Papers

Deutsches Institut für Wirtschaftsforschung 2017

Testing the Fisher Hypothesis in the G-7 Countries Using I(d) Techniques

Guglielmo Maria Caporale and Luis Gil-Alana

IMPRESSUM

DIW Berlin
German Institute for Economic Research
Mohrenstr. 58
10117 Berlin

Tel. +49 (30) 897 89-0
Fax +49 (30) 897 89-200
http://www.diw.de

ISSN electronic edition 1619-4535

Papers can be downloaded free of charge from the DIW Berlin website:
http://www.diw.de/discussionpapers

Discussion Papers of DIW Berlin are indexed in RePEc and SSRN:
http://ideas.repec.org/s/diw/diwwpp.html
http://www.ssrn.com/link/DIW-Berlin-German-Inst-Econ-Res.html

Testing the Fisher Hypothesis in the G-7 Countries

Using I(d) Techniques

Guglielmo Maria Caporale

Brunel University London, CESifo and DIW Berlin

Luis Gil-Alana

University of Navarra

May 2017

Abstract

This paper revisits the Fisher hypothesis by estimating fractional integration and cointegration models that are more general than the standard ones based on the classical I(0)/I(1) dichotomy. Two sets of results are obtained under the alternative assumptions of white noise and Bloomfield (1973) autocorrelated errors respectively. The univariate analysis suggests than the differencing parameter is higher than 1 for most series in the former case, whilst the unit root null cannot be rejected for the majority of them in the latter case. The multivariate results imply that there exists a positive relationship, linking nominal interest rates to inflation; however, there is no evidence of the full adjustment of the former to the latter required by the Fisher hypothesis.

Keywords: Fisher effect, fractional integration, long memory, G7 countries.

JEL classification: C22, C32, E43

Corresponding author: Professor Guglielmo Maria Caporale, Research Professor at DIW Berlin. Department of Economics and Finance, Brunel University London, Kingston Lane, Uxbridge, UB8 3PH, United Kingdom. Tel.: +44 (0)1895 266713. Fax: +44 (0)1895 269770.

E-mail: Guglielmo-Maria Caporale@brunel.ac.uk

The second-named gratefully acknowledges financial support from the Ministerio de Economía y Competitividad (ECO2014-55236).

1. Introduction

The long-run behaviour of interest rates is normally analysed using the so-called Fisher relationship (see Fisher, 1930) linking nominal rates and expected inflation and requiring a one-to-one adjustment of the former to the latter, in the absence of which permanent shocks to either inflation or nominal rates have permanent effects on real rates themselves, which is inconsistent with standard models of intertemporal asset pricing as well as superneutrality (to the extent that inflation is a monetary phenomenon), both requiring stationarity of real interest rates. Neoclassical models of dynamic growth also have the property that consumption growth and real rates should both be constant in the steady state.

However, the Fisher relationship is often not supported by the empirical evidence. Many studies, going back to Mishkin (1992), find that the slope coefficient in a regression of inflation against nominal rates is significantly different from one, and the real rates exhibit a unit root. This "paradox" (see Carmichael and Stebbing, 1983) might be due to various reasons, such as not measuring accurately inflationary expectations (see, e.g. Woodward, 1992), overlooking taxation (see Darby, 1975), using short rather than long rates (see Gilbert and Yeoward, 1994), not distinguishing between short- and long-run Fisher effects (see Mishkin, 1992), and finally differences in estimation procedures; in particular, Crowder and Hoffman (1996) and Caporale and Pittis (2004) both showed that using the estimators with the best small sample properties and appropriately computed empirical critical values (rather than the asymptotical ones) produces evidence more supportive of the Fisher effect. All such studies, though, analyse the Fisher relationship on the basis of the classic dichotomy between I(0) and I(1) variables, which imposes rather restrictive assumptions on the stochastic behaviour of the variables of interest.

By contrast, in the present paper we examine the Fisher effect in the G-7 by adopting a fractional integration/cointegration framework that does not restrict the order of integration d to be 0 or 1 and allows it instead to take any real value; this is clearly a much more general specification allowing for a richer dynamic structure. In particular, consider the following model:

$$R_t \;=\; \alpha \;+\; \beta\,\pi_{t+1} \;+\; \varepsilon_t, \qquad (1)$$

where R_t is the nominal interest rate and π_t is the inflation rate (under the implicit assumption of rational expectations, i.e., $\pi_{t+1} = \pi_t^e$).

If both R_t and π_t are I(0) variables, standard regression methods can be applied. On the other hand, if they are I(1) a cointegration approach becomes necessary. Unlike in the classical case, we allow these variables to be I(d) with 0 < d < 1 or d > 1. First we estimate d_R and d_π, i.e., their respective orders of integration, then we make the series stationary I(0) by differencing them to obtain $\tilde{R}_t = (1-B)^{\tilde{d}_R}$ and $\tilde{\pi}_t = (1-B)^{\tilde{d}_\pi}$, where B stands for the backshift operator, i.e. $Bz_t = z_{t-1}$, and $\tilde{d}_R$ and $\tilde{d}_\pi$ are the estimated orders of integration of the two variables. In other words, we test for the Fisher effect in the following model:

$$\tilde{R}_t \;=\; \alpha \;+\; \beta\tilde{\pi}_t \;+\; \varepsilon_t, \qquad t = 1, 2, \qquad (2)$$

where the null hypothesis in (2) is:

$$H_o: \;\beta \;=\; 1\;.$$

and both variables are I(0) as a result of taking first differences. We also examine the cases where $\tilde{d}_R$ and $\tilde{d}_\pi$ are restricted to be either 1 or 0 as in the classical cointegration framework, but also test the order of integration of the errors in (2) to allow for a greater degree of generality as explained in Section 3.

The layout of the paper is the following. Section 2 briefly reviews the literature on the Fisher effect. Section 3 outlines the empirical methodology. Section 4 presents the main empirical results. Section 5 offers some concluding remarks.

2. The Fisher Effect: A Brief Literature Review

Early studies analysed the Fisher effect without considering stationarity issues. These include Fama (1975), Nelson and Schwert (1977) and Garbade and Wachtel (1978). Gilbert and Yeoward (1994) argued that papers using short rates are not informative about the Fisher effect unless short and long rates are strongly correlated; examples are the papers by Summers (1983) and Barsky (1987).

The following generation of studies took a cointegration approach (see, e.g., Mishkin, 1992; Evans and Lewis, 1995; Wallace and Warner, 1993). Engsted (1995) looked at the spread between the long-term (multi-period) interest rate and the one-period inflation estimating a VAR model and found considerable cross-country differences.

Crowder and Wohar (1997) analysed both taxable US Treasury and tax exempt municipal bond interest rates and found a "Darby effect", i.e. evidence that rational agents require nominal rates to adjust in response to movements in "tax-adjusted" expected inflation. Crowder and Hoffman (1996) showed that the choice of estimator is crucial, an issue further investigated by Caporale and Pittis (2004), who provided extensive evidence that if the estimators with the best small sample properties are used and statistical tests are carried out with appropriate empirical critical values the data are supportive of the Fisher effect.

More recently, fractional integration and fractional cointegration techniques have been used to analyse the long memory properties of inflation and interest rates. For example, Shea (1991) investigated the consequences of long memory in interest rates for tests of the expectations hypothesis of the term structure. Phillips (1998), found stationarity but with a high degree of dependence in US interest rates. Tsay (2000) modelled interest rates as AutoRegressive Fractionally Integrated Moving Average (ARFIMA) processes and concluded that the ex-post real interest rate can be well described as a fractionally integrated or I(d) process. Further evidence of long-memory behaviour in interest rates is provided by Barkoulas and Baum (1997), Meade and Maier (2003), Gil-Alana (2004a,b), Couchman, Gounder and Su (2006), Gil-Alana and Moreno, 2012, Haug, 2014, Apergis et al., 2015, Abbritti et al. (2016), etc. As for inflation rates, evidence of long memory has been reported in many papers including Hassler (1993), Delgado and Robinson (1994), Hassler and Wolters (1995), Baillie et al. (1996), Baum et al. (1999), Hyung et al. (2006), Kumar and Okimoto (2007), etc. Lardic and Mignon (2003) found some evidence for the Fisher hypothesis in the G7 countries using semi-parametric I(d) techniques based on log-periodogram regressions. The opposite conclusion was reached by Ghazalia and Ramlee (2003) by estimating fully parameterised AutoRegressive Fractionally Integrated Moving Average (ARFIMA) models for the same set of countries. Kasmanmet et al. (2006) examined the Fisher relationship with fractional cointegration techniques in 33 developed and developing countries. They found no evidence of cointegration when using classical methods (i.e., Johansen, 1996); however, they found fractional cointegration by using the Geweke and Porter-Hudak (1982) approach on the estimated errors from the cointegrating relationship. Similar conclusions were reached in the case of Turkey by Burcu (2013) and for Nigeria by Etuk et al. (2014).

3. The Empirical Methodology

As mentioned before in the present study we apply fractional integration methods allowing both nominal interest rates and inflation to be I(d), where d can be a fractional value. The starting point is the estimation of the differencing parameter d in:

$$(1 - B)^d x_t = u_t, \quad t = 1, 2, \ldots, \qquad (3)$$

where B stands for the backshift operator, u_t is an I(0) process, and in order to allow for deterministic terms x_t are assumed to be the errors in a regression model of the form:

$$y_t = \alpha + \beta t + x_t, \qquad t = 1, 2, \ldots, \qquad (4)$$

where y_t is the original series (in our case, interest rates or inflation). We use a testing approach suggested by Robinson (1994) that, unlike other methods, does not require stationarity and allows d to take any real value. More specifically, it tests the null hypothesis,

$$H_o: \; d = d_o,$$

in (3) and (4) for any real value d_o, including the stationary ($d_o < 0.5$) and nonstationary ($d_o \geq 0.5$) cases.[1]

To analyse the relationship between the two variables we run a regression of the following form:

$$(1-L)^{d_R} R_t \;=\; \alpha \;+\; \beta (1-L)^{d_\pi} \pi_t \;+\; u_t, \qquad t = 1, 2, \tag{5}$$

$$(1-L)^{d} u_t \;=\; \varepsilon_t, \qquad t = 1, 2, \tag{6}$$

estimating simultaneously α, β and d, for different cases. In particular, we examine three cases: i) imposing a priori $d_R = d_\pi = 1$ in (5); ii) imposing $d_R = d_\pi = 0$ in (5); and iii) using for d_R and d_π the estimated values obtained using the univariate methods, first without any restrictions and then setting d in (6) equal to 0. We assume the error terms to be in turn uncorrelated (white noise) and autocorrelated, in the latter case using the exponential spectral model of Bloomfield (1973), which performs very well in the context of fractional integration.

4. Empirical Results

We start with the univariate analysis. Tables 1 and 2 display the estimates of d (along with the 95% bands corresponding to their non-rejection values using Robinson's (1994) method), for inflation and interest rates respectively. Following common practice in the unit root literature we consider the three cases of i) no deterministic terms, ii) an intercept and iii) an intercept with a linear time trend. The estimated model is:

[1] This method is based on the Whittle function in the frequency domain (Dahlhaus, 1989).

$$y_t = \alpha + \beta t + x_t; \qquad (1 - L)^{d_o} x_t = u_t; \qquad t = 1, 2, ..., \qquad (7)$$

with uncorrelated and Bloomfield (1973) errors in turn. In the case of inflation rates, a time trend is required only for the US with autocorrelated errors, and no deterministic terms are required for Japan.[2] The estimated values of d under the assumption of white noise errors are much higher than 1, and evidence of a unit root (i.e., d = 1) is only obtained in the cases of Canada and Germany. With autocorrelated errors, evidence of a unit root is found in all cases except for Italy where the estimated value of d is much higher than 1.

[Insert Tables 1 and 2 about here]

Moving on to interest rates, a time trend seems to be appropriate for Japan with uncorrelated errors, and in all cases except Italy with autocorrelated ones. With white noise errors, the estimated values of d are much higher than 1 in all cases except for Japan and the US where the unit root null cannot be rejected. Under the assumption of autocorrelation, we obtain evidence of unit roots in all cases except for Japan, where the series seems to be stationary (i.e. with $d < 0.5$).

[Insert Table 3 about here]

Table 3 summarizes the estimates of d with autocorrelated errors, a more appropriate assumption given the results of the diagnostic tests carried out on the residuals (not reported). The unit root null is almost never rejected. The only exceptions are the inflation rate in Italy, with an order of integration much higher than 1, and interest rates in Japan, with a value significantly smaller than 1. However, in the case of Italy the confidence bands for the two variables overlap suggesting that the estimated values are not significantly different.

Tables 4, 5 and 6 show the estimated coefficients from equations (5) and (6) under three different assumptions, specifically, in Table 4 d_R and d_π are both set equal to 1; in Table

[2] This is based on the t-values of the corresponding estimated coefficients, noting that equation (7) can be written as $(1-L)^{d_o} y_t = \alpha(1-L)^{d_o} 1_t + \beta(1-L)^{d_o} t + u_t$, and u_t is I(0) by construction.

5, despite the evidence from Table 3, both variables are assumed to be I(0), and therefore $d_R = d_\pi = 0$, while in Table 6 the estimated values from Table 3 are used for d_R and d_π.

In Table 4, we impose $d_R = d_\pi = 1$ - this is a common assumption in the empirical literature that usually treats interest rates and inflation as being non-stationary I(1). Therefore, both variables in the regression model, $\tilde{R}_t$ and $\tilde{\pi}_t$ are expected to be I(0). In the case of Japan, given the stationarity of interest rates implied by Table 3, we also display the results with $d_R = 0$ and $d_\pi = 1$. It can be seen that with uncorrelated errors (Table 4i) the estimated value of d is significantly positive for all the countries except Japan and USA, while β is only significant for France, Great Britain and the USA, whilst with autocorrelated errors the estimated value of d is smaller than in the previous case, and the I(0) hypothesis for the error cannot be rejected for any country except Japan and the US, and the estimated value of β is significantly positive in all cases except Japan.

[Insert Tables 4 – 6 about here]

In Table 5 we assume that $d_R = d_\pi = 0$, i.e. we run the regression model using the original data without taking differences. In the autocorrelation case, we cannot reject the null that the estimated errors are I(1) with the only exception of Japan, where d is found to be smaller than 1.[3] Once more, all the estimated values of β are significantly positive except in the case of Japan, though they are rather small. Finally, in Table 6 we take differences using the estimated values of d reported in Table 3. With uncorrelated errors (Table 6i), the estimated value of d is found to be positive in all cases, ranging from 0.22 (Italy) to 0.61 (Japan), whilst with autocorrelated errors the I(0) hypothesis cannot be rejected for Italy, d being positive in the remaining cases and ranging from 0.23 (Germany) to 0.51 (Japan).The estimates of β are very similar in the two cases of white noise and autocorrelated errors, and the values are substantially higher than previously and positive, except for Japan.

[3] Note that in this case, given the I(1) evidence provided by the univariate analysis, an estimated value of d smaller than 1 implies fractional cointegration, whilst non-fractional cointegration holds if d = 0.

[Insert Table 7 about here]

Finally, in Table 7, we compare the estimates of β with those obtained when imposing d = 0 in (6). They are generally higher than in the previous cases, especially when $d_R = d_\pi = 0$. In fact, in the I(0) case, even the hypothesis β = 1 cannot be rejected for some countries (France, Germany and Italy). However, it should be noted that this specification is incorrect since the null hypothesis of d = 0 is rejected in favour of d > 0 in all the cases shown in Table 5.

Overall, the evidence based on our preferred model (Table 6ii) suggests that there exists a positive relationship between nominal interest rates and inflation, since the β coefficient is positive; however, it is statistically different from 1 (more precisely, it is smaller). Therefore, we do not find evidence of the full adjustment of nominal rates to inflation required by the Fisher hypothesis.

5. Conclusions

This paper revisits the Fisher hypothesis in the G7 countries using fractional integration and cointegration models that are more general than the standard ones based on the classical I(0) /I(1) dichotomy. Two sets of results are produced under the alternative assumptions of uncorrelated and autocorrelated errors. The univariate analysis suggests than the differencing parameter is higher than 1 for most series in the former case, whilst the unit root null cannot be rejected for the majority of them in the latter case. The multivariate results imply that there exists a positive relationship, linking nominal interest rates to inflation; however, the Fisher hypothesis is rejected since there is no full adjustment of the former to the latter. The implication of our findings is that the evidence in favour of the Fisher effect found in various studies is invalidated by their failure to take into account the fractional nature of the series of interest as well as of the regression errors.

References

Abbritti, M., L.A. Gil-Alana, Y. Lovcha and A. Moreno (2016). Term structure persistence, Journal of Financial Econometrics 14, 2, 331-352.

Apergis, N., Christou, C., Payne, J.E., and J.W. Saunoris (2015). The change in real interest rate persistence in OECD countries: evidence from modified panel ratio tests. Journal of Applied Statistics 42, 202-213.

Baillie, R.T., C.F. Chung, and M.A. Tieslau, (1996). Analysing inflation by the fractionally integrated ARFIMA-GARCH model, Journal of Applied Econometrics 11, 23-40.

Barkoulas, J.T. and C.F. Baum (1997). Fractional differencing modeling and forecasting of eurocurrency deposit rates. The Journal of Financial Research 20, 355-372.

Barsky, R.B. (1987). The Fisher hypothesis and the forecastability and persistence of inflation, Journal of Monetary Economics 19, 3-24.

Baum, C.F., Barkoulas, J. and Caglayan, M. (1999). Persistence in the international inflation rates, Southern Economic Journal 65, 900-913.

Bloomfield, P. (1973). An exponential model in the spectrum of a scalar time series, Biometrika 60, 217-226.

Burcu, K. (2013). A fractional cointegration analysis of Fisher hypothesis. Evidence from Turkey, Quality and Quantity 47, 2, 1077-1084.

Caporale, G.M. and N. Pittis (2004). Estimator choice and Fisher’s paradox: a Monte Carlo study. Econometric Reviews 23, 1, 25-52.

Carmichael, J. and P.W. Stebbing (1983). Fisher's paradox and the theory of interest, American Economic Review, 73, 4, 619-630.

Couchman, J., R. Gounder and J.J. Su (2006). Long memory properties of real interest rates for 16 countries. Applied Financial Economics Letters 2, 25-30.

Crowder, W.J. and D.L. Hoffman (1996). The long-run relationship between nominal interest rates and inflation: the Fisher equation revisited, Journal of Money, Credit and Banking 28, 1, 102-118.

Crowder, W.J. and M.E. Wohar (1999). Are tax effects important in the long-run Fisher relation? Evidence from the municipal bond market. Journal of Finance 54, 1, 307-317.

Dahlhaus, R. (1989). Efficient parameter estimation for self-similar process. Annals of Statistics 17, 1749-1766.

Darby, M.R. (1975). The financial and tax effects of monetary policy on interest rates, Economic Inquiry 13, 266-269.

Delgado, M. and P.M. Robinson (1994). New methods for the analysis of long memory time series. Application to Spanish inflation, Journal of Forecasting 13, 97-107.

Engsted, T. (1995). Does the long-term interest rate predict future inflation? A multi-country analysis, Review of Economics and Statistics 77, 1, 42-54.

Ettuk, I.E., T.D. James and B.K. Asare (2014). Fractional cointegration analysis of Fisher hypothesis in Nigeria. Asian Journal of Applied Sciences 2, 1,

Evans, M. and K. Lewis (1995). Do expected shifts in inflation affect estimates of the long-run Fisher relation?, Journal of Finance 50, 225-253.

Fama, E.F. (1975). Short-term interest rates as predictors of inflation, American Economic Review 65, 269-282.

Fisher, I. (1930). The Theory of Interest, New York, MacMillan.

Garbade, K. and P. Wachtel (1978). Time variation in the relationship between inflation and interest rates, Journal of Monetary Economics 4, 755-765.

Geweke, J. and S. Porter-Hudak (1983). The estimation and application of long memory time series models, Journal of Time Series Analysis 4, 221-238.

Ghazali, N. A. and S. Ramlee (2003). A long memory test of the long-run Fisher effect in the G7 countries.Applied Financial Economics,13, 763–769.

Gil-Alana, L.A. (2004a). Long memory in the interest rates in some Asian countries. International Advances in Economic Research 9, 257-267.

Gil-Alana, L.A. (2004b). Long memory in the US interest rate.International Review of Financial Analysis 13, 265-276.

Gil-Alana, L.A. and A. Moreno (2012). Uncovering the U.S. term premium.An alternative route.Journal of Banking and Finance 36, 1184-1193.

Gilbert, C.L. and S.M.J. Yeoward (1994). Is the Fisher effect for real? Evidence from the UK real and nominal bonds, D.P. no. 310, Department of Economics, Queen Mary and Westfield College, London.

Hassler, U. (1993). Regression of spectral estimators with fractionally integrated time series, Journal of Time Series Analysis 14, 369-380.

Hassler, U. and J. Wolters (1995). Long memory inflation rates. International evidence, Journal of Business and Economic Statistics 13, 37-45.

Haug, A. (2014). On Real Interest Rate Persistence: The Role of Breaks, Applied Economics46, 1058-1066.

Hyung, N., P.H. Franses and J. Penm (2006). Structural breaks and long memory in US inflation rates. Do they matter for forecasting?, Research in International Business and Finance 20, 1, 95-110.

Johansen, S. (1996). Likelihood based inference in cointegrated vector autoregressive models, Oxford University Press.

Kasman, S., A. Kasman and E. Turgutly (2006). Fisher hypothesis revisited: A fractional cointegration analysis, Emerging Markets Finance and Trade 42, 59-76.

Kumar, M.S. and T. Okimoto (2007). Dynamics of persistence in international inflation rates, Journal of Money, Credit and Banking 39, 6, 1457-1479.

Lardi, S. and V. Mignon (2003). Fractional cointegration between nominal interest rates and inflation: A re-examination of the Fisher relationship in the G7 countries. Economics Bulletin 3, 1–10.

Meade, N. and M.R. Maier (2003). Evidence of long memory is short term interest rates. Journal of Forecasting 22, 553-568.

Mishkin, F.S. (1992). Is the Fisher effect for real?, Journal of Monetary Economics 30, 195-212.

Nelson, C.R..and G.W. Schwert (1977). Short-term interest rates as predictors of inflation: on testing the hypothesis that the real rate of interest is constant, American Economic Review 67, 3, 478-486.

Phillips, P.C.B. (1998). Econometric analysis of Fisher's equation, Cowles Foundation Discussion Paper #1180, Yale University.

Shea, G. (1991). Uncertainty and implied variance bounds in long memory models of the interest rate term structure. Empirical Economics 16, 287-312.

Summers, L.H. (1983). The nonadjustment of nominal interest rates: a study of the Fisher effect, in J. Tobin (ed.), Macroeconomics, Price and Quantities, Oxford, Blackwells, 201-241.

Tsay, W.J. (2000). Long memory story of the real interest rate, Economics Letters 67, 3, 325-330.

Wallace, M.S. and J.T. Warner (1993). The Fisher effect and the term structure of interest rates: tests of cointegration, Review of Economics and Statistics 75, 2, 320-324.

Woodward, G.T. (1992). Evidence of the Fisher effect from UK indexed bonds, Review of Economics and Statistics, 74, 2, 315-320.

Table 1: Estimates of d and 95% confidence bands for the inflation series

i) Uncorrelated (white noise) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	0.95 (0.82, 1.12)	**0.97 (0.83, 1.15)**	0.97 (0.83, 1.15)
FRANCE	1.15 (1.04, 1.29)	**1.25 (1.13, 1.41)**	1.25 (1.13, 1.41)
GREAT BRITAIN	1.19 (1.07, 1.35)	**1.19 (1.06, 1.35)**	1.19 (1.06, 1.35)
GERMANY	1.04 (0.93, 1.22)	**0.96 (0.86, 1.09)**	0.96 (0.86, 1.09)
ITALY	1.20 (1.09, 1.34)	**1.25 (1.15, 1.38)**	1.25 (1.15, 1.38)
JAPAN	**1.18 (1.05, 1.36)**	1.18 (1.06, 1.36)	1.18 (1.06, 1.36)
U.S.A.	1.23 (1.07, 1.44)	**1.40 (1.20, 1.67)**	1.40 (1.20, 1.67)
i) Autocorrelated (Bloomfield) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	0.77 (0.51, 1.09)	**0.73 (0.43, 1.11)**	0.74 (0.43, 1.11)
FRANCE	1.11 (0.87, 1.46)	**1.18 (0.88, 1.52)**	1.19 (0.88, 1.54)
GREAT BRITAIN	1.12 (0.89, 1.45)	**1.12 (0.86, 1.54)**	1.12 (0.88, 1.54)
GERMANY	1.12 (0.90, 1.46)	**1.18 (0.93, 1.56)**	1.18 (0.93, 1.56)
ITALY	1.24 (1.00, 1.54)	**1.40 (1.08, 1.77)**	1.40 (1.08, 1.79)
JAPAN	**0.99 (0.76, 1.30)**	0.99 (0.76, 1.31)	0.99 (0.76, 1.31)
U.S.A.	0.83 (0.57, 1.15)	0.69 (0.46, 1.06)	**0.70 (0.42, 1.06)**

In bold the most adequate specification according to the t-values for the deterministic terms. In parenthesis the 95% confidence bands of the non-rejection values using Robinson's (1994) tests.

Table 2: Estimates of d and 95% confidence bands for the interest rate series

i) Uncorrelated (white noise) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	1.11 (0.98, 1.26)	**1.20 (1.03, 1.42)**	1.20 (1.02, 1.42)
FRANCE	1.12 (1.01, 1.27)	**1.23 (1.06, 1.46)**	1.23 (1.06, 1.46)
GREAT BRITAIN	1.09 (0.98, 1.24)	**1.29 (1.12, 1.51)**	1.29 (1.12, 1.51)
GERMANY	1.10 (0.99, 1.26)	**1.25 (1.09, 1.48)**	1.25 (1.09, 1.47)
ITALY	1.10 (0.99, 1.25)	**1.14 (1.02, 1.35)**	1.15 (1.02, 1.34)
JAPAN	1.03 (0.92, 1.18)	0.84 (0.74, 1.03)	**0.78 (0.56, 1.03)**
U.S.A.	1.12 (1.00, 1.27)	**1.17 (0.99, 1.38)**	1.17 (0.99, 1.38)
i) Autocorrelated (Bloomfield) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	1.02 (0.83, 1.32)	0.78 (0.63, 1.13)	**0.72 (0.41, 1.13)**
FRANCE	1.09 (0.90, 1.38)	0.81 (0.70, 1.06)	**0.73 (0.49, 1.06)**
GREAT BRITAIN	1.07 (0.88, 1.42)	0.82 (0.63, 1.13)	**0.77 (0.50, 1.13)**
GERMANY	1.03 (0.85, 1.34)	0.85 (0.72, 1.18)	**0.80 (0.53, 1.17)**
ITALY	1.11 (0.92, 1.37)	**0.93 (0.79, 1.16)**	0.92 (0.78, 1.14)
JAPAN	1.00 (0.81, 1.31)	0.66 (0.58, 0.78)	**0.20 (0.03, 0.48)**
U.S.A.	1.05 (0.84, 1.38)	0.73 (0.58, 1.12)	**0.71 (0.49, 1.12)**

In bold the most adequate specification according to the t-values for the deterministic terms. In parenthesis the 95% confidence bands of the non-rejection values using Robinson's (1994) tests.

Table 3: Estimates of the orders of integration for each series

	Inflation rates	Interest rates
CANADA	**0.73 (0.43, 1.11)**	**0.72 (0.41, 1.13)**
FRANCE	**1.18 (0.88, 1.52)**	**0.73 (0.49, 1.06)**
GREAT BRITAIN	**1.12 (0.86, 1.54)**	**0.77 (0.50, 1.13)**
GERMANY	**1.18 (0.93, 1.56)**	**0.80 (0.53, 1.17)**
ITALY	1.40 (1.08, 1.77)	**0.93 (0.79, 1.16)**
JAPAN	**0.99 (0.76, 1.30)**	0.20 (0.03, 0.48)
U.S.A.	**0.70 (0.42, 1.06)**	**0.71 (0.49, 1.12)**

In bold, evidence of unit roots at the 5% level.

Table 4: Estimates of d, α and β in the long run equilibrium relationship using d = 1

i) Uncorrelated (white noise) errors			
	d (and 95% conf. band)	α (t-value)	β (t-value)
CANADA	0.19 (0.03, 0.40)	-0.00867 (-0.28)	0.04628 (1.48)
FRANCE	0.20 (0.05, 0.42)	-0.00582 (-0.16)	**0.12112 (1.99)**
GREAT BRITAIN	0.26 (0.10, 0.48)	-0.00644 (-0.14)	**0.14159 (2.95)**
GERMANY	0.24 (0.06, 0.46)	-0.01053 (-0.26)	0.03486 (0.77)
ITALY	0.13 (0.02, 0.35)	-0.00604 (-0.16)	0.06315 (0.65)
JAPAN	-0.18 (-0.33, 0.05)	-0.01226 (-3.39)	-0.00465 (-0.22)
JAPAN (*)	-0.19 (-0.34, 0.03)	-0.01151 (-3.16)	-0.00287 (-0.64)
U.S.A.	0.14 (-0.01, 0.34)	-0.00696 (-0.20)	**0.09376 (2.54)**
ii) Autocorrelated (Bloomfield) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	-0.24 (-0.49, 0.16)	-0.02014 (-4.54)	**0.06398 (2.13)**
FRANCE	-0.19 (-0.35, 0.12)	-0.02063 (-3.48)	**0.14878 (2.94)**
GREAT BRITAIN	-0.17 (-0.37, 0.12)	-0.02118 (-3.18)	**0.16265 (3.85)**
GERMANY	-0.16 (-0.35, 0.19)	-0.02635 (-3.90)	**0.09898 (2.13)**
ITALY	-0.11 (-0.28, 0.12)	-0.01154 (-0.89)	**0.19044 (2.22)**
JAPAN	-0.54 (-0.67, -0.29)	-0.01163 (-16.09)	0.00745 (0.52)
JAPAN (*)	-0.55 (-0.69, -0.30)	-0.01082 (-13.80)	-0.00354 (-1.98)
U.S.A.	-0.19 (-0.44, 0.20)	-0.01682 (-2.20)	**0.09810 (3.11)**

In bold, statistically significant positive coefficients at the 5% level.

Table 5: Estimates of d, α and β in the long run equilibrium relationship using d = 0

i) Uncorrelated (white noise) errors			
	d (and 95% conf. band)	α (t-value)	β (t-value)
CANADA	1.19 (1.02, 0.41)	3.86977 (22.33)	0.04656 (1.49)
FRANCE	1.20 (1.04, 1.43)	3.04058 (14.93)	**0.12176 (2.01)**
GREAT BRITAIN	1.26 (1.10, 1.48)	3.77301 (20.42)	**0.14142 (2.96)**
GERMANY	1.17 (1.02, 1.37)	3.16939 (17.61)	0.05260 (1.12)
ITALY	1.13 (0.97, 1.35)	3.36384 (10.44)	0.06364 (0.66)
JAPAN	0.84 (0.73, 1.04)	1.56111 (16.82)	-0.00652 (-0.30)
U.S.A.	1.13 (0.98, 1.34)	4.01174 (15.88)	**0.09432 (2.57)**
ii) Autocorrelated (Bloomfield) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	0.82 (0.66, 1.18)	3.85681 (23.44)	**0.06705 (2.19)**
FRANCE	0.83 (0.71, 1.10)	3.07621 (16.69)	**0.16417 (3.19)**
GREAT BRITAIN	0.85 (0.68, 1.12)	3.78324 (21.14)	**0.16681 (3.91)**
GERMANY	0.88 (0.74, 1.19)	3.16978 (17.97)	**0.09953 (2.13)**
ITALY	0.88 (0.72, 1.14)	3.13425 (10.92)	**0.20962 (2.44)**
JAPAN	0.65 (0.57, 0.78)	1.50744 (19.51	-0.00320 (-0.17)
U.S.A.	0.79 (0.60, 1.16)	4.05142 (17.92)	**0.10027 (3.22)**

In bold, statistically significant positive coefficients at the 5% level.

Table 6: Estimates of d, α and β in the long run relationship using estimated d

i) Uncorrelated (white noise) errors			
	d (and 95% conf. band)	α (t-value)	β (t-value)
CANADA	0.37 (0.26, 0.93)	0.77014(4.40)	**0.35377 (6.36)**
FRANCE	0.42 (0.32, 0.99)	0.52693 (4.03)	**0.51634 (7.32)**
GREAT BRITAIN	0.30 (0.21, 0.45)	0.26939 (2.55)	**0.54364 (6.74)**
GERMANY	0.28 (0.18, 0.42)	0.18328 (2.07)	**0.32952 (4.71)**
ITALY	0.22 (0.11, 0.38)	0.10150 (1.28)	**0.55540 (6.02)**
JAPAN	0.61 (0.54, 0.83)	1.15207 (12.84)	-0.01522 (-0.58)
U.S.A.	0.46 (0.32, 0.95)	0.74281 (3.89)	**0.39219 (7.93)**
ii) Autocorrelated (Bloomfield) errors			
	No deterministic terms	An intercept	A linear time trend
CANADA	0.08 (0.25, 0.61)	0.23491 (2.52)	**0.38721 (6.72)**
FRANCE	0.32 (0.19, 0.55)	0.28027 (3.08)	**0.53874 (7.42)**
GREAT BRITAIN	0.27 (0.10, 0.52)	0.21799 (2.35)	**0.54961 (6.84)**
GERMANY	0.23 (0.09, 0.47)	0.12122 (2.34)	**0.34014 (4.79)**
ITALY	0.15 (-0.01, 0.39)	0.05864 (1.00)	**0.56673 (6.06)**
JAPAN	0.51 (0.45, 0.60)	0.92435 (13.46)	-0.01240 (-0.49)
U.S.A.	0.31 (0.11, 0.50)	0.29458 (2.62)	**0.39703 (8.25)**

In bold, statistically significant positive coefficients at the 5% level.

Table 7: Estimates of α and β in equations (5) and (6) with d imposed to be equal to 0

	$d_R = d_\pi = 1$		$d_R = d_\pi = 0$		d_R and d_π estimated	
	α	β	α	β	α	β
CANADA	-0.0170 (-1.30)	**0.0544 (1.75)**	**2.2081 (33.49)**	**0.3354 (9.57)**	**0.0483 (1.66)**	**0.4248 (7.30)**
FRANCE	-0.0182 (-1.31)	**0.1383 (2.44)**	**1.7197 (34.14)**	**0.8459*** (25.92)**	**0.0479 (1.89)**	**0.5828 (6,90)**
GR. BRITAIN	-0.0198 (-1.38)	**0.1578 (3.47)**	**2.1473 (29.40)**	**0.4250 (15.62)**	**0.0493 (1.70)**	**0.6097 (6.90)**
GERMANY	-0.0251 (-1.84)	0.0697 (1.49)	**1.1298 (13.62)**	**0.8756*** (17.29)**	0.0169 (0.60)	**0.3989 (4.92)**
ITALY	-0.0109 (-0.52)	0.1276 (1.49)	**2.5005 (49.38)**	**0.9152*** (35.24)**	0.0232 (0.75)	**0.5912 (5.94)**
JAPAN	-0.0114 (-1.40)	-0.0121 (-0.54)	**1.0062 (52.26)**	**-0.0929 (-5.95)**	**0.3708 (20.29)**	0.0263 (0.52)
U.S.A.	-0.0133 (-0.74)	**0.0977 (2.81)**	**2.4165 (51.07)**	**0.3072 (5.67)**	**0.0662 (2.13)**	**0.3783 (8.64)**

In bold, significant coefficients at the 5% level.