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by Andreas Beyer[2], Alfred A. Haug[3] and William G. Dewald[4]

In 2009 all ECB publications feature a motif taken from the €200 banknote.

1 The views expressed in this paper do not necessarily reflect those of the European Central Bank or the Eurosystem. The authors thank, without implicating, seminar participants at the universities of Auckland and Otago for helpful comments.

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Abstract

There is scant empirical support in the literature for the Fisher effect in the long run, though it is often assumed in theoretical models. We argue that a break in the cointegrating relation introduces a spurious unit root that leads to a rejection of cointegration. We applied new break tests and tested for nonlinearity in the cointegrating relation with post-war data for 15 countries. Our empirical results support cointegration, after accounting for breaks, and a linear Fisher relation in the long run. This is in contrast to several recent studies that found no support for linear cointegration.

Keywords: Fisher effect; linear and nonlinear cointegration; structural change.

JEL Classification: E43; C32.

Non-technical summary

The long-run Fisherian theory of interest states that a permanent shock to inflation will cause an equal change in the nominal interest rate so that the real interest rate is not affected by monetary shocks in the long run. Fisher's model determines the real interest rate as the difference between the nominal interest rate and the expected inflation rate. If the nominal interest rate and the inflation rate are each integrated of order one, denoted I(1), then the two variables should cointegrate with a slope coefficient of unity so that the real interest rate is covariance stationary. If the Fisher effect holds, a permanent change in inflation will lead to a one-for-one change in the nominal interest rate in the long run. Inflation then exhibits long-run neutrality with respect to real interest rates. A relatively large number of theoretical models assume that the Fisher hypothesis holds. However, empirical support has been difficult to establish despite numerous attempts. In this paper, we provide an explanation for the apparent failure of the Fisher hypothesis in the earlier literature. We argue that the finding in the literature of no cointegration for the Fisher hypothesis may be due to structural changes in the cointegrating vector. We use a recently developed test by Carrion-i-Sylvestre and Sans (2006) for the null hypothesis of cointegration against the alternative hypothesis of no cointegration in the presence of a structural break under both hypotheses. Their methodology allows us to determine whether the finding of no cointegration for the Fisher effect is due to structural change or not. We study the long-run Fisher effect for 15 countries over the post-war period. Many previous studies found no empirical support for the long-run Fisher effect. These findings would imply that money was not super-neutral and that there was money illusion, if one assumes that money growth drives inflation, because real interest rates would be affected by inflation. Our results are different. We found for most of the countries evidence of a break in the cointegrating relationship. Once we account for these breaks, the pre-break and post-break samples reveal clear evidence in favor of cointegration. In addition we found scant evidence against our linear specification and conclude that the linear model of cointegration for the Fisher effect is well supported by the empirical evidence presented.

1. Introduction

The long-run Fisherian theory of interest states that a permanent shock to inflation will cause an equal change in the nominal interest rate so that the real interest rate is not affected by monetary shocks in the long run.[1] Fisher's model determines the real interest rate as the difference between the nominal interest rate and the expected inflation rate. If the nominal interest rate and the inflation rate are each integrated of order one, denoted I(1), then the two variables should cointegrate with a slope coefficient of unity so that the real interest rate is covariance stationary. Or, alternatively, if they do not cointegrate, tests for long-run neutrality developed by Fisher and Seater (1993) can be applied as long as both variables are I(1). If the Fisher effect holds, a permanent change in inflation will lead to a one-for-one change in the nominal interest rate in the long run. Inflation then exhibits long-run neutrality with respect to real interest rates. A relatively large number of theoretical models assume that the Fisher hypothesis holds. However, empirical support has been difficult to establish despite numerous attempts.

Weber (1994), King and Watson (1997), Koustas and Serletis (1999), and Rapach (2003) rejected long-run neutrality of inflation with respect to real interest rates, using the framework of Fisher and Seater (1993) and a large number of OECD countries.[2] Engsted (1995), Koustas and Serletis (1999), Atkins and Serletis (2003), and Rapach (2003), among others, formally tested for cointegration and found no support for cointegration between inflation and nominal interest rates. On the other hand, Mishkin (1992), Evans and Lewis (1995), and Crowder and Hoffman (1996), among others, found evidence in favor of cointegration with post-war United States data. Also, Beyer and Farmer (2007) found cointegration between the Federal Funds rate and inflation, but the cointegrating vector has a break in 1979. Before the break, the strict Fisher hypothesis cannot be rejected, but after the break the coefficient on inflation is significantly different from and larger than unity. Beyer and Farmer interpreted this finding as a shift in US monetary policy in 1979 when Paul Volcker became chairman of the Federal Reserve. Recently, researchers have applied to the Fisher hypothesis various new alternative econometric models and tests.

[1] See Fisher (1930).

[2] All of these studies used a bi-variate model with inflation and nominal interest rates, except for Rapach who used a tri-variate model that includes in addition real GDP.

Westerlund (2008) tested the Fisher effect in a cointegrated panel of 20 OECD countries with quarterly data from 1980 to 2004 and could not reject the Fisher hypothesis. Panels generally add power to cointegration tests due to the added cross-sectional dimension, however, they also impose restrictions, particularly on the cross-sectional dependencies, that may not hold in the data. On the other hand, Jensen (2008) found that inflation follows a mean-reverting, fractionally integrated, long-memory process and not an I(1) process as supported by previous studies. However, the estimates and associated likelihood ratio tests rely on asymptotic results that may be unreliable for inference and may lack power in small samples. Lastly, Christopoulos and León-Ledesma (2007) argued instead that the failure of finding support for the Fisher effect may be due to nonlinearities in the long-run relationship between inflation and nominal interest rates.

In this paper, we provide an alternative explanation for the apparent failure of the Fisher hypothesis in the earlier literature. In a seminal paper, Perron (1989) showed that a break in the deterministic time trend dramatically reduces the power of standard unit root tests because the possibility of a break changes the (asymptotic) distribution of the test. Similarly, Gregory, Nason and Watt (1996) demonstrated that the rejection frequency of the ADF test for cointegration falls substantially in the presence of a structural break in the cointegrating relation. This means that the null hypothesis of no cointegration is not rejected when the cointegrating relation is unstable. Therefore, the finding in the literature of no cointegration for the Fisher hypothesis may be due to structural changes in the cointegrating vector.

Carrion-i-Sylvestre and Sansó (2006) developed a test for the null hypothesis of cointegration against the alternative hypothesis of no cointegration in the presence of a structural break under both hypotheses. They derive the test for a known and an unknown break with strictly exogenous and also with endogenous regressors. For unknown breaks, an efficient procedure for estimating the break date is proposed. This methodology allows us to determine whether the finding of no cointegration for the Fisher effect is due to structural change or not. We briefly describe the break tests of Carrion-i-Sylvestre and Sansó (2006) in Section 2 and present the data and empirical results in Section 3. Section 4 summarizes the results and suggests directions for future research.

2. Testing for Structural Breaks in Cointegrating Relations

In this section we briefly outline the features of the test for structural change proposed by Carrion-i-Sylvestre and Sansó (2006). The test is a supplementary test to the standard tests of the null hypothesis of no cointegration against the alternative hypothesis of cointegration. It is particularly useful in cases where standard cointegration tests (that do not allow for breaks) lead to the finding of no cointegration. The problems with standard cointegration tests in the presence of breaks were carefully illustrated by Gregory et al. (1996). A break introduces spurious unit root behaviour in the cointegrating relationship so that the hypothesis of no cointegration is difficult to reject.[3]

The advantage of the testing approach developed by Carrion-i-Sylvestre and Sansó is that it is not embedded in either a model that assumes covariance stationary variables or a cointegrated system.[4] Furthermore, it avoids the problem of disentangling a regime shift from a stable cointegrating relationship as is the case in the set-up of Gregory and Hansen (1996a, 1996b). Gregory and Hansen's alternative hypothesis allowed for a breaking cointegrating relation but also included the stable cointegrating relation in the alternative hypothesis.[5]

Carrion-i-Sylvestre and Sansó's test allows for a possible structural change in the parameters of the cointegrating vector, including but not limited to the deterministic parts of the vector. The test is based on a multivariate extension of the KPSS test (Kwiatkowski et al., 1992). The null hypothesis of a cointegrating relation in the possible presence of one break is tested against the alternative hypothesis of no cointegration (with possibly one break). Therefore, the null hypothesis allows for the possibility that the cointegrating relation shifts from one long-run equilibrium to another one, i.e., for a change in the parameters of the cointegrating vector.

[3]The rejection frequency of standard tests of the null hypothesis of no cointegration falls dramatically in the presence of a break in a cointegrating relationship. A researcher would therefore correctly conclude that there is no evidence for standard (without a break) cointegration.

[4]For example, the popular tests of Bai, Lumsdaine and Stock (1998) require choosing either a stationary system or a cointegrated system within which testing for breaks takes place. Similarly, the tests of Hansen (1992) and Seo (1998) require such a choice a priori.

[5]A very general testing framework, not pursued here, with unknown multiple breaks in cointegrated systems with I(1) and I(0) regressors has been suggested by Kejriwal and Perron (2008a).

For regressors that are endogenous with respect to the parameters in the cointegrating vector, dynamic ordinary least squares (DOLS) is applied to efficiently estimate the model under the null hypothesis.[6] The break test proposed by Carrion-i-Sylvestre and Sansó is a Lagrange-Multiplier-type (LM) test that is calculated from the DOLS residuals and the associated long-run variance-covariance matrix. This matrix is estimated non-parametrically with a Bartlett kernel and a data-dependent procedure to select the optimal spectral bandwidth, following Kurozumi (2002).

The LM-test requires imposing a break date. If the break date is unknown, Carrion-i-Sylvestre and Sansó suggested to estimate the break date consistently with the dynamic algorithm of Bai and Perron (1998, 2003). It minimizes the sum of squared residuals from DOLS regressions over a closed subset of break fractions.

We choose a model without deterministic time trends, which is consistent with the Fisher hypothesis. We allow for structural change in the constant term as well as in the cointegrating slope parameters, at an unknown point in time. Critical values for the LM test depend on the break fraction and the number of regressors. We pick appropriate critical values from Table 2 in Carrion-i-Sylvestre and Sansó. In addition, these authors carried out a Monte Carlo study that showed that their testing methodology leads to break tests with good power and size properties.

3. Data and Empirical Results

3.1 Data

We retrieved quarterly data, on various days in March 2008, for the consumer price index (CPI) and short-term interest rates from the International Monetary Fund's IFS online data base for 15 OECD countries: Australia, Belgium, Canada, Denmark, France, Germany, Italy, Japan, New Zealand, the Netherlands, Norway, Sweden, Switzerland, the United Kingdom and the United States.[7] The periods for which data were available are listed in Table 1. Inflation rates were calculated

[6] See Saikkonen (1991) and Stock and Watson (1993).

[7] The CPI for the Netherlands shows a large increase from 63.1 in 1980:4 to 78.7 in 1981:1, followed by a large drop from 90.0 in 1984:1 to 73.9 in 1984:2, causing big spikes in the inflation series. The IFS file has markers for these dates indicating that "multiple time series versions are linked by ratio splicing using first annual overlap". We cross-checked the IFS series with the OECD's online "OECD.Stat Extracts" (webnet.oecd.org/wbos/index.aspx) that did not have these jumps. We therefore used the OECD's CPI time series for the Netherlands, retrieved on 24 October 2008.

from the first differences of the natural logarithm of the CPI, multiplied by 400 to get annualized rates in percent. We picked for the short-term interest rate the 3-month Treasury bill rate where available for sufficiently long spans (Belgium, Canada, France, Sweden, Switzerland, the UK and the US). Otherwise, the money market rate was used (Australia, Denmark, Germany, Italy, Japan and New Zealand), or the deposit rate if the other two rates were unavailable (the Netherlands and Norway).

Figure 1 depicts the nominal interest rates, the CPI-based inflation rates and the ex-post real interest rates for all 15 countries over the available sample periods. We also show the average over all countries for each quarter (thicker line). The graphs generally show an upward movement for the nominal interest rate and the inflation rate for most countries from the 1950s to the 1970s and 1980s, followed by a downward movement afterwards. On the other hand, the real interest rates show an average level that holds pretty steady in the 1950s and 1960s, followed by a more turbulent period in the mid-1970s that last till the mid-1980s. Afterwards, the average real rate moves to a higher level for some time, compared to the 1950s and 1960s. This may indicate possible structural changes in the Fisher relation in at least some countries, however, a formal statistical analysis is required before drawing conclusions.

Figure 2 graphs the average for the 15 countries for the nominal interest rate, the inflation rate and the ex-post real interest rate against the five-year moving average for each series. Cointegration is a concept that captures the co-movement of variables towards a long-run equilibrium. We depict the five-year moving average to provide some intuition for long-run movements. One can see that nominal interest rates and inflation rates move broadly together upwards till the mid-1970s, followed by a downward movement from roughly the mid 1980s onwards. One can also see that this broad movement is out of synchronization in the period from the mid 1970s to the mid 1980s, with nominal interest rates lagging behind inflation. This discord is reflected in the movement of the real ex-post interest rate.

3.2 Test Results for Unit Roots, Cointegration and Breaks

In this subsection we present our empirical findings.[8]We first test for unit roots in the inflation and interest rate series for each country. We apply the DF-GLS test of Elliott, Rothenberg and Stock (1996) to locally demean or demean and detrend by generalizes least squares (GLS) before applying the Dickey-Fuller (DF) unit root test. We choose the lag augmentations based on the modified Akaike criterion (MAIC) advocated by Ng and Perron (2001). However, we calculated MAIC based on OLS detrending, instead of GLS, as suggested by Perron and Qu (2007) in order to correct for a power reversal problem with the Ng and Perron (2001) procedure. We base our inference on the 5% level of significance.[9] We cannot reject the null hypothesis of a unit root for all cases when only a constant is considered in the test regression, except for the interest rate for Germany and the Netherlands and the inflation rate for Germany. The interest rate for Denmark is a borderline case. Once a deterministic time trend is added, a unit root is no longer rejected for inflation and interest rates for all countries, except for the interest rate for the Netherlands. We therefore exclude the Netherlands from the analysis, which leaves us with 14 countries.

The Fisher hypothesis states that the nominal interest rate and the expected inflation rate move one-for-one so that the real interest rate is determined by real factors only:

$$i_t^s = r_t + E_t\pi_t^s,$$

where i_t^s is the nominal interest rate paid on a bond that matures in period $t+s$; r_t is the real interest rate; E_t is the expectations operator conditional on information available at time t. Therefore, $E_t\pi_t^s$ is the inflation rate expected over the life of the bond. As usual, we assume rational expectations so that expected inflation equals actual inflation plus a Gaussian mean-zero forecast error term e_t, i.e., $e_t = E_t\pi_t^s - \pi_t^s$

[8]We used code written in GAUSS 8.0 for all applications, except for the Johansen cointegration tests that we carried out in EViews 6. We used for the break tests in large parts the GAUSS code made available by J.L. Carrion-i-Sylvestre on his web site (riscd2.eco.ub.es/∼carrion/). The GAUSS code for the linearity tests was downloaded from I. Choi's web site (ihome.ust.hk/∼inchoi/).

[9]The critical values for the DF-GLS test with a constant are the same as the Dickey-Fuller critical values without a constant and are calculated with the program from MacKinnon (1996). The DF-GLS critical values with a deterministic time trend are reported in Table 1 in Elliott et al. (1996).

with $e_t \sim$ iid N(0, σ^2). Hence, we can write the Fisher equation as

$$i_t^s = r_t + \pi_t^s + e_t.$$

When nominal interest rates and inflation behave each as an I(1) process in our (finite) samples, then they should be cointegrated with a slope coefficient β equal to unity in the case of linear cointegration, if the Fisher effect holds:

$$i_t^s = \alpha + r_t + \beta\pi_t^s + e_t. \tag{1}$$

Cointegration of nominal interest rates and inflation, regardless of the value of the slope coefficient β, implies that the real (ex-post) interest rate should be covariance stationary. If the Fisher effect holds, then proportionality holds so that $\beta = 1$, and $r_t = i_t^s - \pi_t^s$ should follow a mean-reverting process and not exhibit unit root behaviour.[10] However, we do impose $\beta = 1$ from the start and instead estimate the slope coefficient in order to assess in what ways it differs from unity, if it does.

We proceed next to tests for linear cointegration. We apply the maximum likelihood-based trace and maximum-eigenvalue (λ-max) tests of Johansen (1995) and calculate the p-values according to MacKinnon, Haug and Michelis (1999). Results are reported in Table 1. Consistent with the findings of Koustas and Serletis (1999), among others, we find mostly evidence against cointegration.[11] We find support for cointegration between inflation and interest rates only for 5 countries out of 14, i.e., we reject the null hypothesis of no cointegration for Denmark, Germany, Japan, New Zealand and Sweden. For the other 9 countries, we cannot reject the null hypothesis of no cointegration with the trace and maximum-eigenvalue tests.

In the next step, we test for breaks by applying the test of Carrion-i-Sylvestre and Sansó (2006) to the 9 countries for which we cannot reject the null hypothesis of no cointegration. We use DOLS for calculating the break tests.[12] The last two

[10]We do not pursue here this line of investigation. See Rose (1988) for a pioneering study on ex-post real interest rates. Neely and Rapach (2008) surveyed the empirical literature on real interest rate persistence.

[11]See, for example, Haug (1996) on the relative performance of cointegration tests in Monte Carlo simulations.

[12]Throughout the paper, unless otherwise indicated, we allowed for 4 leads and lags in the DOLS regressions. We also replaced 4 leads and lags with 6 in the larger samples and got essentially the same qualitative results. Similarly, switching the right hand side and left hand side variables in the DOLS regressions does not materially affect the results. Also, long-run variances are estimated with

columns of Table 1 give results for the test statistic values and the estimated break dates. All break dates are estimated to occur in the late 1970s to the mid-1980s and differ across countries. Figure 3 depicts the ex-post real interest rate along with the five-year moving average for each country separately. The vertical dashed lines in the individual country graphs indicate the break dates for those countries for which we uncovered structural changes in the Fisher relation. For the break tests based on the estimated break dates, we cannot reject the null hypothesis that there is a cointegrating relationship with a break at the estimated dates for all 9 countries, except for Canada, using a 5% significance level. Though, Canada is a borderline case and we will consider it further.[13]

If the strict version of the Fisher effect holds in the long run, then monetary shocks should not cause structural change in the cointegrating relation because inflation will not affect real interest rates when inflation and nominal interest rates move one-for-one in the log run. However, real shocks can affect the real interest rate and lead to structural changes in the cointegrating relationship. Supply shocks, as for example the oil price hikes in 1973 and 1979, may cause a level shift in the cointegrating relation. The same holds true for technology and preference shocks. On the other hand, fiscal shocks that affect marginal income tax rates can lead to a structural change in the slope coefficient of the cointegrating vector because we use nominal interest rates that are not tax-adjusted.[14] Identification of the sources of shocks would require a structural analysis, as for example an analysis based on a structural vector-autoregression of fiscal and monetary policy transmission. This is beyond the scope of this paper.

Following the results for the break tests, we now test for cointegration in the pre- and post break samples for all 9 countries. Table 2 reports results. We ignore the pre-break sample for Switzerland because it contains only 18 observations and is therefore too small to draw any reliable inference. We find now clear evidence in favor of cointegration for all pre-break and post-break samples. Out of the 17 sub-samples,

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a quadratic spectral kernel and automatic bandwidth selection after autoregressive pre-whitening, as suggested by Andrews (1991) and Andrews and Monahan (1992).

[13]If one takes a 10% level, the null hypothesis is in addition also rejected for the UK, however, we rely on the 5% level for our inference.

[14]Padovano and Galli (2001) studied decadal average marginal income tax rates for OECD countries. They found breaks for the decades of the 1970s and/or 1980s for the countries that we consider here, but no breaks for the 1950s and 1960s.

only the pre-break result for France is a borderline case with the trace test (p-value = 0.076), however, not with the maximum eigenvalue test (p-value = 0.023).

Our results provide a good empirical example for the assertion that breaks can lead to the spurious finding of no cointegration by introducing unit root behavior into the cointegrating relationship. After accounting for the breaks, we find support for cointegration. We therefore confirm the simulation results of Gregory et al. (1996) as being relevant in practice.

The Fisher model imposes a coefficient restriction on the cointegrating vector and we proceed to testing it. The Fisher hypothesis in its strict form states that interest rates move one-for-one with inflation rates.[15] This implies that the slope coefficient in the cointegrating relationship, β in equation (1), should be equal to 1. We test this hypothesis by restricting the coefficient to 1 (or -1 for the normalized cointegrating vector) and report results in Table 3 for likelihood ratio (LR) tests.[16] For the 5 countries for which we found cointegration over the full sample, the strict version of the Fisher hypothesis cannot be rejected. Once we move to the pre-break and post-break sub-samples for the remaining 9 countries, results for the strict Fisher effect are not very favorable. In the vast majority of cases, the restriction that inflation rates and interest rates move one-for-one is rejected. There are only 5 cases out of the 17 test statistics for the sub-samples where the restriction is not rejected, including 2 borderline cases.

The results for testing restrictions on the cointegrating vectors may suffer from small sample problems. Haug (2002) compared the performance of Wald and likelihood-ratio tests of restrictions in cointegrating vectors with Monte Carlo methods and recommended using Bartlett corrections suggested by Johansen (2000). However, when we applied the Bartlett corrections to testing the strict version of the Fisher hypothesis, we got the same qualitative results as with the unadjusted statistics.

Stock and Watson (1993) have presented Monte Carlo evidence that the Johansen method can give unreliable cointegration coefficient estimates in small samples. We therefore followed their advise and used DOLS for the pre-break and post-break samples for the countries with breaks.[17] Table 3 reports coefficient estimates

[15]Some researchers refer to this case as a "full Fisher effect", following Owen (1993).

[16]See Johansen (1995) for details on testing restrictions on cointegrating vectors.

[17]DOLS results are somewhat sensitive to the exact number of leads and lags included in samples with relatively small numbers of observations. We therefore chose the leads and lags for the sub-

along with appropriately calculated standard errors. We exclude the pre-break samples for Norway and Switzerland due to the sample size being below 20 after allowing for leads and lags. That leaves us with 16 sub-samples for the DOLS regressions. Of the 16, we find 6 cases for which the hypothesis of a unit slope coefficient is not rejected. For the remaining 10 cases, the slope is significantly larger than unity for 7 cases. There are only 3 cases for which the slope coefficient is significantly smaller than unity.

The previous literature has provided explanations why the strict version of the Fisher hypothesis may not hold. Crowder and Hoffman (1996) pointed out that pre-tax nominal interest rates will not move one-for-one with inflation in the long run if real interest rates are supposed to be unaffected by permanent shocks to inflation.[18] It is therefore necessary to calculate post-tax real returns for each country considered, which should show a one-for-one relationship with inflation rates if the Fisher effect is to hold. Crowder and Hoffman used time-varying average marginal tax rates for the US based on the methodology of Barro and Sahasakul (1986). Using pre-tax nominal interest rates leads to a slope coefficient larger that 1 in the Fisher relation for positive tax rates. On the other hand, a Mundell-Tobin effect implies a slope coefficient less than 1, because inflation leads to real balance effects that lower the marginal product of capital and hence the real interest rate. Our coefficient estimates are generally larger than 1, providing strong support for Crowder and Hoffman's argument and empirical evidence against the Mundell-Tobin effect. It is beyond the scope of this paper to calculate appropriate variable marginal tax rates for the 15 countries considered.[19]

The period over which we found breaks, 1976:1 to 1985:2, more or less corresponds with the period when many industrial countries substantially lowered tax

samples based on the Schwarz information criterion. See Kejriwal and Perron (2008b) on the good performance of the Schwarz criterion for leads and lags selection in DOLS regressions.

[18]The slope coefficient of the cointegrating vector in the Fisher relation is not unity but instead equal to $1/(1-\tau_t)$, where τ_t is the relevant average, time-varying, marginal income tax rate. A summary of the simple analytics of the effect of taxation on the relationship between inflation and interest rates is presented in Dewald (1986). In general an increase in inflation will affect real interest rates. It will reduce the after tax interest rate for savers and reduce saving at every interest rate whereas fully taxed investment will not change inasmuch as an increase in inflation will increase both investment returns and taxes on them proportionately.

[19]Padovano and Galli (2001) calculated marginal income tax rates for 23 OECD countries, including the ones considered here, however, only by decade, with four tax rates for each country (1950s, 1960s, 1970s and 1980s). A proper analysis would require calculations as in Barro and Sahasakul.

rates. It is consistent with the findings of Padovano and Galli (2001) that decadal average marginal income tax rates show significant structural changes in the 1970s and 1980s. This is a likely cause for the changes in the cointegrating relationships that we uncovered in Table 1. On the other hand, it is also possible that the oil shocks in the 1970s are responsible for changes in the Fisher relation due to permanent effects on real interest rates.[20]

3.3 Nonlinearity Tests

Nonlinear models may be able to represent certain economic relationships better than linear models. There is an extensive recent literature dealing with alternative nonlinear econometric models. In particular, the purchasing power parity puzzle literature, surveyed by Taylor and Taylor (2004), provides numerous examples, however, there are also applications to the term structure of interest rates (e.g., Haug and Siklos, 2006), among several other areas. With respect to the Fisher hypothesis, Christopoulos and León-Ledesma (2007) argued for the US over the 1960-2004 period that the long-run relationship between inflation and nominal interest rates is nonlinear with an exponential or logistic smooth transition function, i.e., with ESTR or LSTR functions. However, it may well be the case that their finding in favor of nonlinearity is due to structural breaks. In other words, a linear relationship with breaks is mistakenly approximated by a nonlinear relationship. This concern has been raised by Koop and Potter (2001). Indeed, we find basically no evidence of nonlinear cointegration once the breaks are accounted for. We test for nonlinearities in all samples for which we found linear cointegration, i.e., we test for nonlinearities after accounting for breaks. If there were nonlinearities in the full samples, these should show up as well in the sub-samples, but they do not. As a general rule, researchers should always test first for breaks (at an unknown point in time) when linear cointegration is not supported by the data, before applying nonlinearity tests.

We apply the LM-type tests of Choi and Saikkonen (2004) to the cointegrated models over the full sample for Denmark, Germany, Japan, New Zealand and Sweden, and over the sub-samples for the remaining 9 countries. The null hypothesis is the standard linear model of cointegration. The alternative hypothesis is a nonlinear

[20]As we pointed out earlier, in order to pin down the sources of the shocks one would need a structural model with fiscal and monetary transmission channels.

model that includes ESTR and LSTR models, among other nonlinear specifications. The tests can be regarded as general tests for nonlinearity. Based on the simulation results in Choi and Saikkonen, we apply two tests, the T_1 test and T_2 test. The tests are based on auxiliary regressions and are asymptotically distributed as χ^2 with degrees of freedom determined by the number of restrictions imposed under the null hypothesis. The first test uses a first-order Taylor series approximation of the underlying nonlinear functional form, whereas the second test uses a third-order approximation instead. We allow for a constant only (i.e., there is no deterministic time trend) in the auxiliary regressions. We allow for one transition variable only and choose, following Christopoulos and León-Ledesma (2007), the inflation rate for this purpose. In other words, the nominal interest rate is modelled as adjusting to inflation in a nonlinear way. In order to correct for the endogeneity of the regressor, DOLS is applied again.

Table 3 report results for the linearity tests. We exclude the pre-break samples for Norway and Switzerland because the sample size is below 20 after accounting for leads and lags. All T_1 and T_2 tests cannot reject the null hypothesis of linearity for all countries, expect for Australia. We find two cases for this country for which we reject the null hypothesis at the 5% significance level: for the pre-beak sample for T_1 and T_2. We therefore argue that accounting for breaks avoids spurious nonlinearities. The overwhelming evidence is in favor of a linear Fisher relationship.

4. Conclusion

In this paper, we studied the long-run Fisher effect for 15 countries over the post-war period. Many previous studies found no empirical support for the long-run Fisher effect. These findings would imply that money was not super-neutral and that there was money illusion, if one assumes that money growth drives inflation, because real interest rates would be affected by inflation. Our results are different. We argue that it is essential to establish the time series properties of the variables involved, which are inflation and interest rates. In particular, it is crucial for inference whether there is cointegration between the variables or not.

We applied improved unit root tests (Perron and Qu, 2007) and did not reject a unit root for inflation and nominal interest rates for all 15 countries, except for the nominal interest rate for the Netherlands. We next tested for cointegration between inflation and nominal interest rates. Consistent with previous studies by Koustas and Serletis (1999), among others, we found evidence mostly against cointegration (for 9 of all countries considered). However, this finding was due to breaks in the cointegrating relationship that might introduce spurious unit roots, as argued in a theoretical context by Gregory et al. (1996). Furthermore, such spurious unit roots may also lead to finding spurious nonlinear cointegration, as argued by Koop and Potter (2001). Therefore, we applied recently developed tests for this scenario by Carrion-i-Sylvestre and Sansó (2006). These tests consider the null hypothesis of cointegration with a break at an unknown time against the alternative hypothesis of no cointegration with a break. We found for all 9 countries evidence of a break in the cointegrating relationship. Once we account for these breaks, the pre-break and post-break samples reveal clear evidence in favor of cointegration.

In addition to testing for breaks, we checked for nonlinearities in the cointegrating relationships that we found between interest rates and inflation. The tests include nonlinear models of the exponential and logistic smooth transition type. The tests (Choi and Saikkonen, 2004) also have power against general mis-specification of a model. We found scant evidence against our linear specification and conclude that the linear model of cointegration for the Fisher effect is well supported by the empirical evidence presented.

There is one shortcoming in the empirical performance of the Fisher hypothesis. Though we find support for cointegration between inflation and nominal interest rates, the two variables do not move one-for-one in the long run for all cases. Out of 21 cases considered, including pre-break and post-break sub-samples, the slope coefficient is not significantly different from unity for 11 cases, as required by the Fisher hypothesis, but it is significantly larger than unity for 7 cases, which is evidence against a Mundell-Tobin effect. For the US, Beyer and Farmer (2007) related a shift of the slope coefficient from unity to larger than unity in the post-1979 period to a tighter monetary policy. An alternative explanation was put forward by Crowder and Hoffman (1996) based on tax effects in the US. Their explanation suggests calculating variable marginal tax rates for the various countries and test the Fisher effect with tax-adjusted interest rates. We leave this for future research.

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Table 1. Cointegration and break test results for the full sample period

Country	Sample	Johansen cointegration tests			Break test[c]	Break date
		trace test[a]	λ-max test[a]	lags[b]	H_0: cointegration with break at unknown time	
Australia	1969:3-2007:4	0 (0.13)	0 (0.23)	2	0.077	1983:4
Belgium	1957:1-2007:4	0 (0.18)	0 (0.11)	2	0.071	1976:1
Canada	1957:1-2007:4	0 (0.08)	0 (0.11)	3	0.153**	1982:3
Denmark	1972:1-2007:4	1 (0.03)	1 (0.03)	2	--	--
France	1970:1-2004:3	0 (0.24)	0 (0.27)	1	0.104	1981:4
Germany	1957:1-2007:4	1 (0.001)	1 (0.003)	3	--	--
Italy	1971:1-2007:4	0 (0.10)	0 (0.15)	2	0.075	1981:1
Japan	1957:1-2007:4	1 (0.02)	1 (0.01)	4	--	--
Netherlands	1981:1-2007:4	--	--	--	--	--
New Zealand	1985:1-2007:4	1 (0.001)	1 (0.002)	3	--	--
Norway	1979:1-2007:3	0 (0.86)	0 (0.96)	3	0.075	1985:2
Sweden	1963:1-2006:3	1 (0.01)	1 (0.03)	2	--	--
Switzerland	1980:1-2007:4	0 (0.11)	0 (0.12)	3	0.086	1985:1
UK	1957:1-2007:4	0 (0.35)	0 (0.63)	3	0.123*	1980:1
US	1957:1-2007:4	1 (0.03)	0 (0.07)	2	0.076	1977:2

Note: [a] The column lists the number of cointegrating vectors found at the 5% significance level. The figure in parentheses is the p-value calculated according to MacKinnon, Haug and Michelis (1999).
[b] The number of lags in the VAR is chosen with Schwarz's information criterion and reported for first-differences of the VECM specification.
[c] Critical values are from Carrion-i-Sylvestre and Sansó's (2006) Table 2. Significance at the 10% level is indicated by *, and at the 5% level by **.

Table 2. Cointegration tests for the pre- and post-break samples

Country	Sample[a]	Johansen cointegration tests trace test[b]	λ-max test[b]	Lags[c]
Australia	1969:3-1983:3	1 (0.003)	1 (0.001)	0
	1984:3-2007:4	1 (0.000)	1 (0.000)	0
Belgium	1957:1-1975:4	1 (0.005)	1 (0.002)	0
	1977:1-2007:4	1 (0.001)	1 (0.000)	1
Canada	1957:1-1982:2	1 (0.001)	1 (0.000)	0
	1982:3-2007:4	1 (0.000)	1 (0.000)	0
France	1970:1-1981:3	0 (0.076)	1 (0.023)	0
	1982:3-2004:3	1 (0.000)	1 (0.000)	0
Italy	1971:1-1980:4	1 (0.000)	1 (0.000)	0
	1981:4-2007:4	1 (0.000)	1 (0.001)	0
Norway	1979:1-1985:1	1 (0.046)	1 (0.019)	0
	1986:2-2007:3	1 (0.000)	1 (0.000)	1
Switzerland	1980:1-1984:4	--	--	
	1986:4-2007:4	1 (0.009)	1 (0.005)	4
UK	1957:1-1979:4	1 (0.000)	1 (0.000)	0
	1981:1-2007:4	1 (0.000)	1 (0.000)	1
US	1957:1-1977:1	1 (0.003)	1 (0.002)	1
	1978:3-2007:4	1 (0.016)	1 (0.024)	2

Note: [a] The start date after the break excludes the break date itself and accounts for lags in the construction of the inflation rate and the VECM.
[b] The column lists the number of cointegrating vectors found at the 5% significance level. The figure in parentheses is the p-value calculated according to MacKinnon, Haug and Michelis (1999).
[c] The number of lags in the VAR is chosen with Schwarz's information criterion and reported for first-differences of the VECM specification.

Table 3. LR tests for restrictions on the cointegrating vector that inflation and interest rates move one-for-one and LM tests for linearity

Country	Sample (for VECM)[a]	LR test of restriction[b] (p-value)	LM linearity tests[c, d]		DOLS /VECM slope estimates[d] (standard errors)
			T_1	T_2	
Australia	1969:3-1983:3	0.043	8.6**	8.6**	0.69 (0.09)***
	1984:3-2007:4	0.000	0.72	4.64*	1.64 (0.14)***
Belgium	1957:1-1975:4	0.001	0.08	0.47	0.93 (0.30)
	1977:1-2007:4	0.000	0.15	0.16	1.63 (0.08)***
Canada	1957:1-1982:2	0.052	0.03	0.18	0.97 (0.10)
	1982:3-2007:4	0.000	0.0005	0.10	1.61 (0.16)***
Denmark	1972:1-2007:4	0.141	0.14	0.15	0.71 (VECM)
France	1970:1-1981:3	0.490	0.03	2.0	0.93 (0.04)*
	1982:3-2004:3	0.000	0.34	0.36	2.15 (0.17)***
Germany	1957:1-2007:4	0.686	0.23	0.46	1.10 (VECM)
Italy	1971:1-1980:4	0.731	0.17	0.93	0.79 (.04)***
	1981:4-2007:4	0.001	2.0	2.6	1.91 (.11)***
Japan	1957:1-2007:4	0.885	0.62	1.52	0.97 (VECM)
New Zealand	1985:2-2007:4	0.055	0.19	0.22	0.81 (VECM)
Norway	1979:1-1985:1	0.005	--	--	--
	1986:2-2007:3	0.000	0.02	0.03	1.39 (0.29)
Sweden	1963:1-2006:3	0.121	0.01	0.08	0.85 (VECM)
Switzerland	1980:1-1984:4	--	--	--	--
	1986:4-2007:4	0.000	1.79	4.26	1.0 (0.17)
UK	1957:1-1979:4	0.000	1.48	1.73	0.63 (.09)***
	1981:1-2007:4	0.000	0.000	0.25	1.35 (0.08)***
US	1957:1-1977:1	0.873	2.15	2.31	0.94 (0.12)
	1978:3-2007:4	0.050	0.04	0.06	1.27 (0.13)**

Note: [a] The start date after the break excludes the break date itself and accounts for lags in the construction of the inflation rate and the VECM.
[b] An entry of 0.000 indicates a value smaller than 0.0005.
[c] The T_1 test follows an asymptotic χ^2-distribution with 1 degrees of freedom and the T_2 test one with 2 degrees of freedom.
[d] Significance at the 10% level is indicated by *, at the 5% level by **, and at the 1% level by ***.

Figure 1.
Three-month interest rate, CPI inflation, and real ex-post three-month interest rate

Fifteen OECD countries and cross-country average, quarterly, 1957-2007

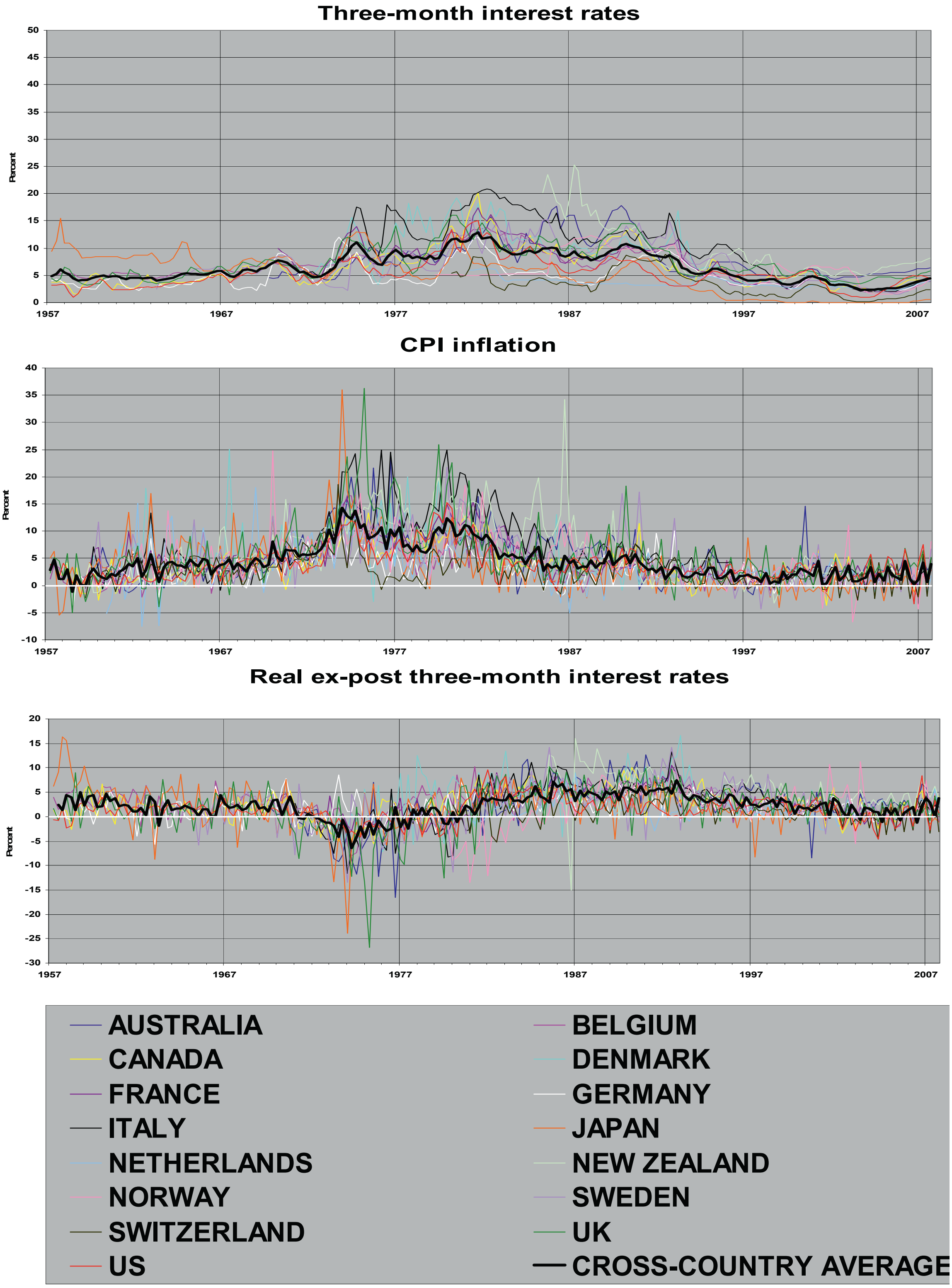
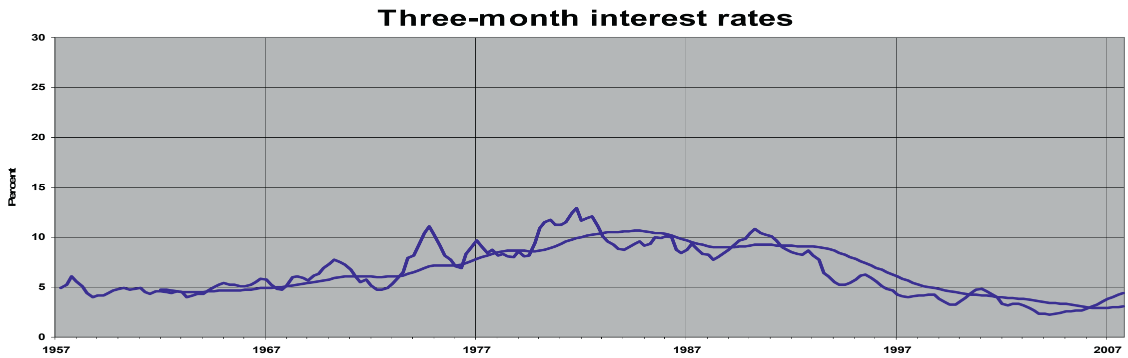
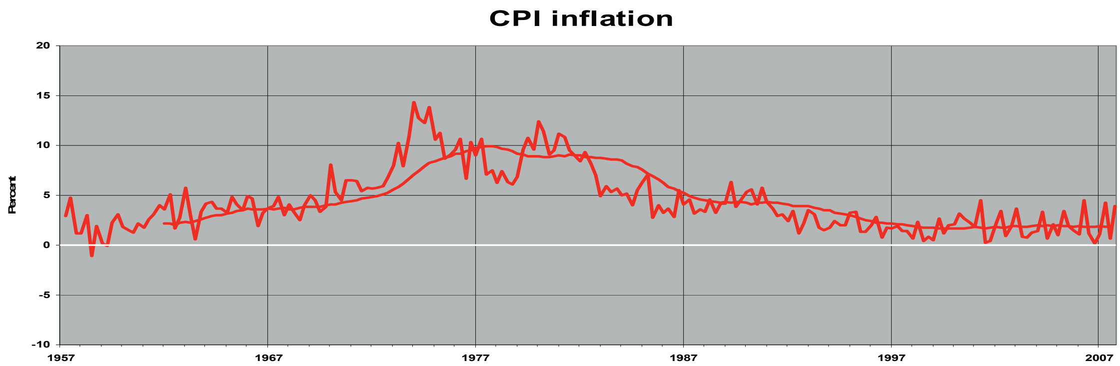


Figure 2.
Three-month interest rate, CPI inflation, and real ex-post three-month interest rate

Fifteen OECD countries, quarterly and five-year cross-country average, 1957-2007





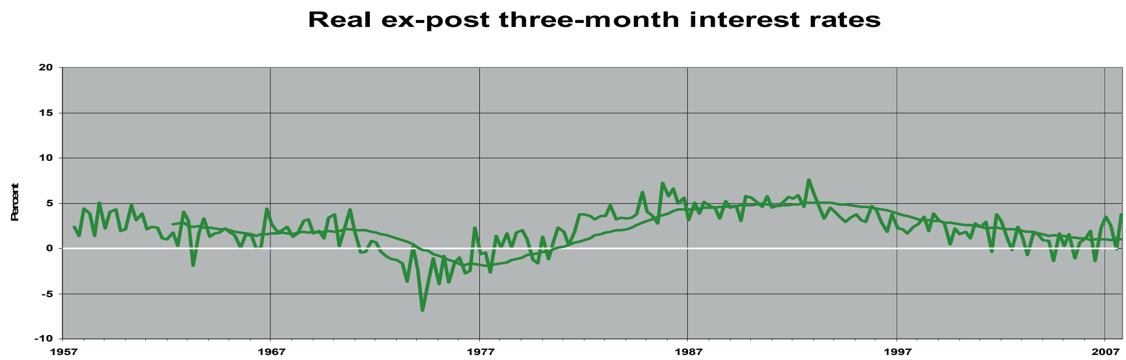
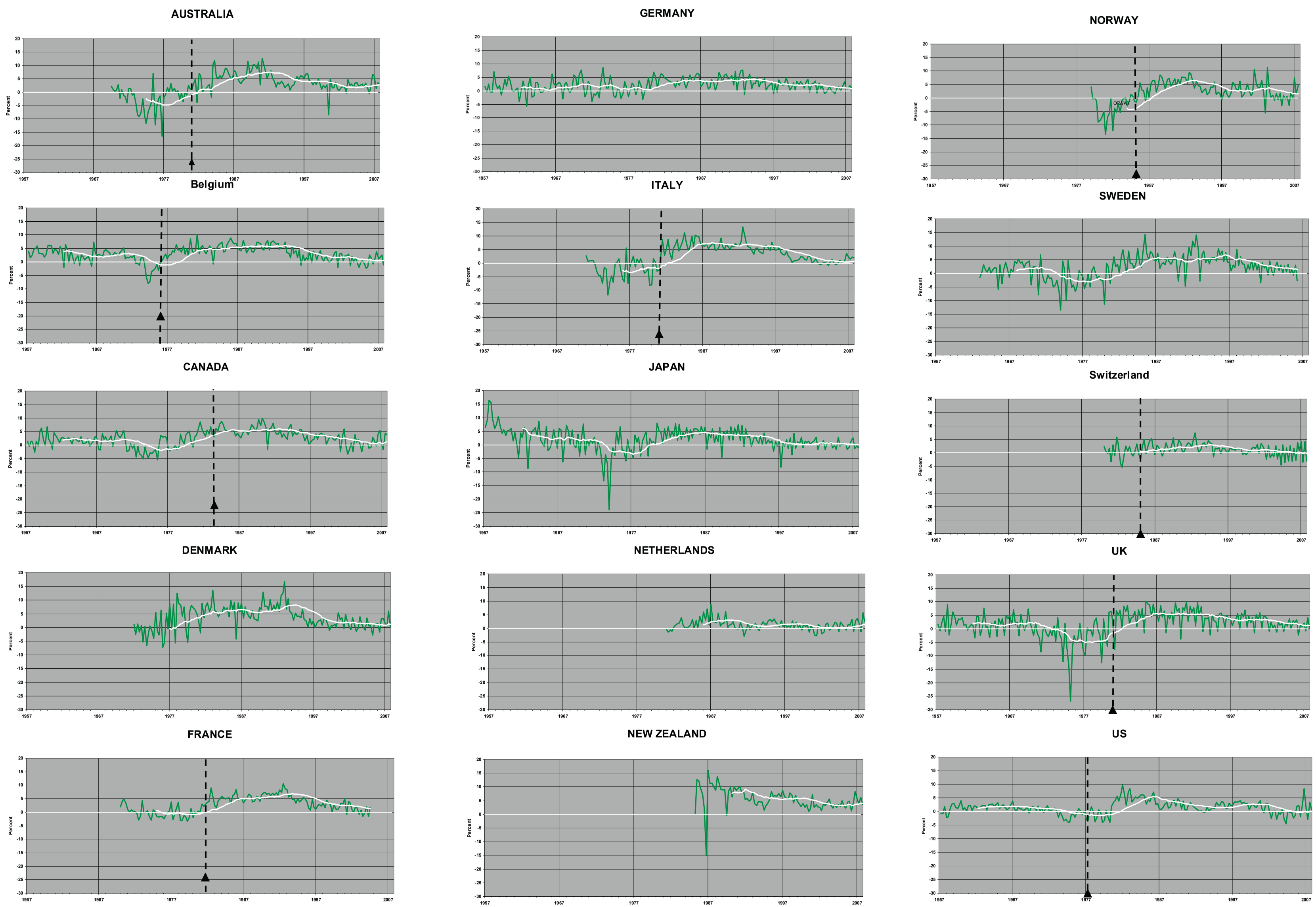


Figure 3. Real short-term interest rates

Fifteen OECD countries, quarterly and five-year averages, 1957-2007*



* Vertical dashed lines indicate break dates.

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