The relationship between nominal interest rates and inflation: international evidence

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Abstract

This paper examines the long-run bivariate relationship between the short-term Eurocurrency interest rate and the inflation rate for nine European countries and the US. Application of cointegration methods reveals that in the majority of cases, there is a one-to-one relationship between Eurocurrency rates and rationally expected inflation. Moreover, the 1-month Eurocurrency rate contains information about the future path of the inflation rate. This finding supports the belief that market participants incorporate a predictable portion of the inflation rate into the nominal interest rate. © 2001 Elsevier Science B.V. All rights reserved.

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1. Introduction

In a recent contribution, Crowder and Hoffman (1996) examine the long-run dynamic relationship between the short-term nominal interest rate and inflation. Consistent with the implications of the Fisher hypothesis (FH), using quarterly data
they document that the 3-month US T-bill rate and the US inflation rate are cointegrated and, thus, share a common stochastic trend. They also find that the long-run Granger-cause ordering is from the inflation rate to the nominal rate, which implies that the inflation rate contains information about the future path of the nominal interest rate. Although their finding of cointegration is consistent with Fama’s (1975) theory that nominal interest rates are optimal predictors of future inflation, their Granger-cause finding is not. Relying on the observation that the inflation rate is somewhat predictable and on the notion that an informationally efficient market uses all relevant information when setting prices, Fama (1975) argues that participants in credit markets incorporate the predictable portion of future inflation into current rates. Thus, in one respect Crowder and Hoffman (1996) and Fama (1975) are at odds with each other.

The purpose of this paper is to resolve this controversy and, in doing so, provide an international perspective to the dynamic relationship between the short-term nominal interest rate and the inflation rate. This task is accomplished using the economic framework provided by the FH. The implications of this hypothesis are investigated by using the cointegration methods promulgated by Johansen (1991) and Gonzalo and Granger (1995). The latter offer an approach to decompose a cointegrated system into its common and transitory components; hence, the common stochastic trend shared by the nominal interest rate and the inflation rate can be identified. This decomposition permits two working research statements. First, if the nominal interest rate incorporates the rationally expected inflation rate and the inflation rate contains little or no information about the future nominal interest rate, the nominal interest rate drives the common stochastic trend in the long run. Second, if instead the reverse is the case so that the inflation rate contains information about the future nominal interest rate and that the nominal interest rate reflects little or any of the expected inflation rate, inflation is the driving force of the cointegrated system.

These statements are explored using short-term nominal Eurocurrency interest rates and inflation rates for 10 countries, which include nine European countries and the US. Consistent with the FH, except for France, country Eurocurrency rates are cointegrated with their respective inflation rates. To determine the long-run Granger-cause ordering between the variables, the common stochastic trend for nominal yields and inflation rates is identified, and it is examined which one of the two variables drives this long-run trend. The empirical findings support Fama’s (1975) position that the nominal interest rate drives the system.

New evidence is provided on the response of nominal rates to expected inflation. According to the FH, a one-to-one relationship exists between changes in expected inflation and changes in nominal interest rate in the long run. In prior work using US data, this FH restriction is consistently rejected. Evans and Lewis (1995) attribute these rejections to small sample bias created by the infrequent shifts in the inflation process in postwar US data. Crowder and Hoffman (1996) and Crowder and Wohar (1999), however, argue that the rejections may be a manifestation of the Darby (1975) effect, which states that when nominal interest rate is taxed, the FH
implies a response from nominal interest rates that is greater than change in expected inflation in order to maintain the constant ex-ante real interest rate.¹

The results show that for the majority of the countries in the sample, a one-point increase in the expected inflation is associated with a one-point increase in the nominal Eurocurrency rate, although this does not hold true for the US, its relationship being greater than one-to-one. Since Eurocurrency rates closely follow domestic rates and US taxation of interest income is not unique, this finding suggests that the rejection of this FH restriction for the US may result from factors other than taxes.

The organization of the paper is as follows. The next section briefly describes the method of analysis and integrates the theoretical model and its statistical tests. The third section describes the data. Empirical findings are presented in the fourth section. The last section offers a summary and concluding remarks.

2. Method of analysis

The FH, which maintains that the nominal interest rate is the sum of the constant real rate and the expected change in the purchasing power of money over the life of the nominal interest rate, acts as the theoretical underpinning of this paper. A decline (increase) in the purchasing power of money can be measured by an increase (decrease) in prices so that the FH can be stated as

\[ R_t = r_t + \pi_{t+1} \]  

where \( R_t \) is the nominal interest rate, \( r_t \) is the real interest rate, \( \pi_{t+1} \) is the expected inflation rate from period \( t \) to \( t + 1 \). Assuming the ex-ante real rate is constant, using realized rates for expectations, and recognizing that the nominal interest rate and the inflation rate can generally be characterized as unit root processes, the FH has two cointegration implications.² First, the nominal interest rate \( R_t \) and the expected inflation rate \( \pi_{t+1} \) are cointegrated and share a common stochastic trend. Second, a one-point increase in the nominal interest rate is associated with a one-point increase in the expected inflation in the long run and, hence, the cointegration vector between \( R_t \) and \( \pi_{t+1} \) is \([1, -1]\). In other words, the cointegration vector implied by the FH should obey a zero-sum restriction for its elements.

To test the FH cointegration restrictions, Johansen’s (1991) full information

¹ Other explanations for this restriction not holding are that economic agents suffer from some sort of fiscal illusion (Tanzi, 1980) and that price changes cause portfolio shifts that perversely impact the prices of nominal assets (Tobin, 1969). Similar to Evans and Lewis (1995) ‘peso problem’ explanation, these two tend to bias the Fisher relationship downward.

² Three economic interpretations of cointegration have been offered. First, Engle and Granger (1988) suggest that it measures the way in which an economic system reacts to reach equilibrium after a perturbation. Second, Granger (1988) demonstrates that it measures the relationship among control, target and dependent variables. Finally, Campbell and Shiller (1988) maintain that cointegration obtains if one time series anticipates another. It is this last view that underpins this analysis.
maximum likelihood method is used. This involves estimating the rank of the impact matrix $II$, which summarizes the long-run relationships among $n$ variables. Cointegration is equivalent to a reduced rank of $II$. A rank of $r$ implies that there are $n \times r$ matrices $\alpha$ and $\beta$ such that $II = \alpha \beta'$, where $\beta$ is the matrix of cointegration vectors and $\alpha$ is a corresponding matrix of factor loadings. Solving Johansen’s (1991) eigenvalue specification yields estimates of the eigenvalues $\lambda_1 > \ldots > \lambda_r > 0$ and the associated eigenvectors $\beta = (v_1, \ldots, v_r)$.

Two likelihood ratio tests may be constructed for the number of cointegration vectors $r$. The maximum eigenvalue test,

$$\hat{\lambda}_{\text{max}} = -T \ln(1 - \hat{\lambda}_{r+1})$$

(2)

tests the null hypothesis of $r$ cointegrating vectors against the alternative hypothesis of $r + 1$ cointegration vectors. The trace test,

$$\hat{\lambda}_{\text{trace}} = -T \sum_{j=r+1}^{n} \ln(1 - \hat{\lambda}_j)$$

(3)

tests the null hypothesis that ‘at most’ $r$ cointegration vectors, with ‘more than’ $r$ vectors being the alternative hypothesis. Although $\hat{\lambda}_{\text{max}}$ and $\hat{\lambda}_{\text{trace}}$ do not have standard asymptotic distributions, critical values are available from Monte Carlo simulations. For the first FH restriction to hold, these cointegration tests should indicate the presence of a cointegration vector between $R_t$ and $\pi_{t+1}$.

Johansen (1991) also develops a likelihood ratio test to examine whether cointegration vectors obey the zero sum restriction. In particular,

$$L_j = -T \sum_{i=1}^{n} \ln \left( \frac{(1 - \hat{\lambda}_i^*)}{(1 - \hat{\lambda}_j)} \right)$$

(4)

where $L_j$ is $\chi^2$ distributed and the asterisk denotes estimates from the restricted model. For the second FH restriction to hold, the zero-sum restriction must not be rejected.

As subsequently reported, the data generally confirm these restrictions. Thus, following Stock and Watson (1988), it is possible to identify the common factor driving the changes in the system by decomposing the cointegrating system into permanent and transitory components, and Park (1990) offers a simple decomposition based on the estimators from Johansen’s (1991) cointegration analysis. Gonzalo and Granger (1995) express this decomposition as a linear factor model and show that the common factor is a function of $\alpha_\perp$. The solution to their eigenvalue problem gives the eigenvalues $\lambda_1 > \ldots > \lambda_n$ and eigenvectors $M = (m_1, \ldots, m_n)$. The estimate of $\alpha_\perp$ is found from eigenvectors $m_{r+1}, \ldots, m_n$. Significance of $\alpha_\perp$ shows which variable plays the dominant role in driving the common trend. Statistical significance tests of $\alpha_\perp$ are conducted using Gonzalo and Granger’s (1995) likelihood ratio specification, i.e.

$^3$Gonzalo (1994) suggests that, when the number of data points exceeds 100, Johansen’s estimator performs better than a range of other cointegration vector estimators. He also shows that the finite sample properties of this method are consistent with the asymptotic results even when the errors are non-Gaussian and the dynamics are unknown.
\[ L_{GG} = -T \sum_{j=r+1}^{n} \ln \left( \frac{1 - \lambda_j^*}{1 - \lambda_j} \right) \]  

where the asterisk denotes estimates from the restricted model. \( L_{GG} \) is asymptotically distributed as \( \chi^2 \) with \((n - r) \times (r)\) degrees of freedom, where \( n - r \) is the number of common trends. If \( L_{GG} \) is significant, the null hypothesis of \( \tau \perp = 0 \) is rejected.

Within this context, the Granger-cause ordering between nominal interest rate and inflation rate may be examined. If interest rates reflect rationally expected inflation rates, the nominal rate drives the system in the long run. If this is not the case, the nominal rate eventually adjusts to the inflation rate. These two statements are examined using \( L_{GG} \) to determine which one of the two variables drives the common stochastic trend in the long run.

3. The data

The data consist of monthly 1-month nominal Eurocurrency interest rates and monthly inflation rates for 10 countries: Belgium, Denmark, France, Germany, Italy, the Netherlands, Norway, Sweden, the UK and the US. The sample period is from January 1978 to February 1997. The Eurocurrency rates are end-of-month bid rates quoted at approximately 10:00 h Swiss time and are provided by the Bank for International Settlements. Eurocurrency rates are used because they are less affected than domestic rates by capital controls and legal regulations.\(^4\) Inflation rates are proxied by monthly changes in the CPI indices obtained from IMF’s International Financial Statistics tapes. Following Mishkin (1992), the time ordering of variables is as follows. Month \( t \) interest rates are month \( t - 1 \) quotes and month \( t - 1 \) inflation rates are calculated from month \( t \) and month \( t - 1 \) CPI data. All rates are annualized percentages.

4. Empirical findings

4.1. Cointegration test of the FH

It is firmly established in the empirical literature that interest rates may be approximated as non-stationary variables (see, for instance, Engle and Granger, 1987; Campbell and Shiller, 1987; Hall et al., 1992, among others).\(^5\) For Eurocurrenc-
rency rates, Arshanapalli and Doukas (1996) find non-stationarity to be the norm. Augmented Dickey Fuller (ADF) test results, which are not reported but are available upon request, indicate that, consistent with these and other studies, the Eurocurrency rates examined herein are unit root processes. For the inflation rates, the existence of a unit root has been questioned, although there is substantial evidence to the contrary. Unit root tests conducted herein support, however, the contention that the inflation rate of each country contains a unit root, thereby

<table>
<thead>
<tr>
<th>Country</th>
<th>Rank</th>
<th>$k$</th>
<th>Johansen statistics</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium</td>
<td>$r \leq 1$</td>
<td>0</td>
<td>1.61, 1.61</td>
</tr>
<tr>
<td>Denmark</td>
<td>$r \leq 1$</td>
<td>4</td>
<td>41.17, 39.55</td>
</tr>
<tr>
<td>France</td>
<td>$r \leq 1$</td>
<td>5</td>
<td>35.43, 33.35</td>
</tr>
<tr>
<td>Germany</td>
<td>$r \leq 1$</td>
<td>0</td>
<td>2.77, 2.77</td>
</tr>
<tr>
<td>Italy</td>
<td>$r \leq 1$</td>
<td>3</td>
<td>53.14, 51.32</td>
</tr>
<tr>
<td>The Netherlands</td>
<td>$r \leq 1$</td>
<td>6</td>
<td>22.02, 20.18</td>
</tr>
<tr>
<td>Norway</td>
<td>$r \leq 1$</td>
<td>0</td>
<td>1.84, 1.84</td>
</tr>
<tr>
<td>Sweden</td>
<td>$r \leq 1$</td>
<td>0</td>
<td>21.02, 19.17</td>
</tr>
<tr>
<td>UK</td>
<td>$r \leq 1$</td>
<td>0</td>
<td>3.00, 4.00</td>
</tr>
<tr>
<td>US</td>
<td>$r \leq 1$</td>
<td>0</td>
<td>26.22, 23.21</td>
</tr>
</tbody>
</table>

This table shows the results of testing for cointegration between the Eurocurrency rate and the inflation rate for each country. $\lambda_{\text{trace}}$ and $\lambda_{\text{max}}$ are calculated for lag length, $k$, with the value of $k$ selected to render serially uncorrelated residuals, as suggested by Johansen (1988). The $P = 0.05$ critical values for the test statistics depend on $n-r$, i.e. the number of common trends, where $n$ is the number of variables in the data vector and $r$ is the number of cointegration vectors. The Hamilton (1994) critical values for $\lambda_{\text{trace}}$ are 15.19 for $n-r = 1$ and 3.96 for $n-r = 0$. The $\lambda_{\text{max}}$ critical values are 14.03 for $n-r = 1$ and 3.96 for $n-r = 0$. France is the only country for which the null hypothesis of no cointegration cannot be rejected.

For instance, Rose (1988) finds the inflation rates of all OECD countries to be stationary, but Haldrup (1994) and Mishkin (1992) argue that the inflation rates of Denmark and the US, respectively, are non-stationary. Engsted (1995) examines the inflation rates of 13 OECD countries and concludes that their inflation rates are better characterized as non-stationary variables. Similarly, Crowder and Hoffman (1996) find the US inflation rate to be non-stationary.
giving it the potential to be cointegrated with the short-term Eurocurrency rate in every instance.\footnote{Because Schwert (1989) documents that ADF tests have greater than actual sizes when the variables have a significant moving average component, which is the case for many inflation variables in this study, the existence of a unit root in each series is tested within the context of Johansen’s multivariate analysis by imposing the restriction that the vector $[0,1]$ spans the cointegration space. The null hypothesis of this test is stationary as opposed to non-stationarity in univariate ADF tests. A likelihood ratio statistic, that is $\chi^2$ distributed is used to test this restriction. In every instance, the results indicate that the null hypothesis of stationary inflation rates is soundly rejected ($P < 0.001$).}

Johansen’s $\hat{\lambda}_{\text{trace}}$ and $\hat{\lambda}_{\text{max}}$ statistic for each country is reported in Table 1. Before discussing the empirical findings, however, a background on the behavior of Eurocurrency rates in the latter part of the sample period is helpful. In 1992 and 1993, European Monetary System (EMS) exchange rates came under severe pressure, a situation often attributed to the German reunification. The initial result was the departure of the Italian lira and the British pound from the system in September 1992. The pressure waned but returned in the summer of 1993, resulting in a substantial widening of the exchange rate bands for the French franc and almost all the remaining members of the EMS in August 1993. To control for the effect of these market pressures and institutional changes on Eurocurrency rates, the European country cointegration structures are estimated using dummy variables for the September 1992–August 1993 period.\footnote{Although Norway is not a member of the EMS, a dummy variable is used because its interest rate was also affected by the EMS crisis.} This is not done for the US because its Eurocurrency rates did not experience any unusual behavior during this period.

The first restriction of the long-run FH states that the nominal rate and the expected inflation rate are cointegrated. With one exception, the twin Johansen statistics, which are presented in Table 1, support the conclusion that the Eurocurrency rate and the inflation rate share a common stochastic trend, i.e. they are cointegrated. This means that nominal rates and inflation rates move together in the long run even though there may not be an exact relationship in the short run. For France, however, the nominal rate and the inflation rate are not cointegrated, indicating that the expected inflation rate and the nominal rates do not move together. In other words, in the long run the FH does not hold in this case.\footnote{This conclusion holds at the $P = 0.05$ significance level. At the $P = 0.10$ level, however, the FH is not rejected. Nevertheless, another candidate for the French common factor is the German nominal interest rate. According to the German dominance hypothesis, Germany determines the monetary policy in the European Monetary Union and the other countries simply adapt to this policy. An implication of this hypothesis is that Germany money market rates should dominate the long-run behavior of the money market rates of other European Monetary System countries. The likelihood ratio test statistic of Gonzalo and Granger (1995) cannot reject the null hypothesis ($P = 0.219$) that the German Eurocurrency rate solely determines the long-run path of the French common factor, which is consistent with the German dominance hypothesis.}

The second restriction of the long-run FH states that there is a one-to-one relationship between changes in nominal interest rate and inflation rate, a relationship that implies a zero-sum restriction for the cointegration vector. This restriction, which involves eigenvalues $v_1$, is tested by the likelihood ratio test statistic $L_1$ and the associated $P$-values are reported in Table 2. For six currencies out of nine,
Table 2
Cointegration structure between eurocurrency interest rates and inflation rates

<table>
<thead>
<tr>
<th>Country</th>
<th>Rates</th>
<th>Eigenvectors ($c$)</th>
<th>Eigenvectors ($m$)</th>
<th>Avg. contrib.</th>
<th>Common trend</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Std. $c_1$</td>
<td>Zero-sum $P$-value</td>
<td>$m_2$</td>
<td></td>
<td>$P$-value</td>
</tr>
<tr>
<td>Belgium</td>
<td>Euro. 1.000</td>
<td>1.400</td>
<td>0.972</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-1.505$</td>
<td>0.104</td>
<td>0.028</td>
<td>0.027</td>
<td></td>
</tr>
<tr>
<td>Denmark</td>
<td>Euro. 1.000</td>
<td>0.406</td>
<td>0.879</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-1.128$</td>
<td>0.151</td>
<td>0.121</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>Germany</td>
<td>Euro. 1.000</td>
<td>1.732</td>
<td>0.961</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-1.667$</td>
<td>0.117</td>
<td>0.039</td>
<td>0.002</td>
<td></td>
</tr>
<tr>
<td>Italy</td>
<td>Euro. 1.000</td>
<td>0.259</td>
<td>0.710</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-0.921$</td>
<td>0.199</td>
<td>0.290</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>The Netherlands</td>
<td>Euro. 1.000</td>
<td>1.253</td>
<td>0.954</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Norway</td>
<td>Euro. 1.000</td>
<td>0.177</td>
<td>0.046</td>
<td>0.022</td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-1.273$</td>
<td>0.055</td>
<td>0.029</td>
<td>0.017</td>
<td></td>
</tr>
<tr>
<td>Sweden</td>
<td>Euro. 1.000</td>
<td>0.039</td>
<td>0.030</td>
<td>0.005</td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-0.949$</td>
<td>0.065</td>
<td>0.029</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>UK</td>
<td>Euro. 1.000</td>
<td>1.142</td>
<td>0.971</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-1.007$</td>
<td>0.065</td>
<td>0.029</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>US</td>
<td>Euro. 1.000</td>
<td>0.896</td>
<td>0.882</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Inf. $-1.610$</td>
<td>0.210</td>
<td>0.118</td>
<td>0.001</td>
<td></td>
</tr>
</tbody>
</table>

* This table provides selected cointegration structure information for the Eurocurrency rate (Euro.) and the inflation rate (Inf.) of nine countries. France is not included because her Eurocurrency rate and inflation rate are not cointegrated. The std. $c_1$ statistics are the coefficients of the cointegration vector, which have been standardized so that the coefficient for the Eurocurrency rate equals one. Zero-sum $P$-values represent the significance levels for testing the null hypothesis that the sum of the eigenvector ($c_1$) elements equals zero. The common stochastic trend for the Eurocurrency rate and the inflation rate is calculated using column $m_2$. The Avg. contrib. figures are the average proportional contributions of the Eurocurrency rate and the inflation rate to the common trend. For each country, the sum of the individual contributions equals one. The common trend $P$-values represent the significance levels for testing the null hypothesis that the Eurocurrency rate is the sole variable driving the system’s common stochastic trend.

this restriction is not rejected. It is, however, rejected for the Eurodollar, a finding that is consistent with prior studies. Hence, the results of testing the zero-sum restriction indicate stronger support for the long-run FH in international data.

Can the rejection of the FH zero-sum restriction for the US be attributed to the existence of the Darby (1975) effect? The reported findings do not permit an unambiguous ‘yes’ to this question. For example, consider the contrast provided by the UK and the US. Others have shown that the Eurocurrency rates of these two countries are related to their respective domestic money market interest rates.10

10 For instance, Lin and Swanson (1993) find that UK and US Eurocurrency and domestic interest rates are cointegrated and short- and long-run bidirectional causality is present. Tse and Booth (1995) report that the US Treasury bill futures contract is cointegrated with the Eurodollar futures contract, with causality running from Treasury bill to Eurodollar interest rates.
Thus, the Darby (1975) effect, if it is present, should be observable regardless of whether domestic or off-shore rates are used to measure it. Moreover, the US and the UK tax interest rate income in a similar way.\textsuperscript{11} As a result, if the Darby (1975) effect is indeed a plausible explanation for the greater than one-to-one relationship between inflation and interest rates, the following should obtain for each country. First, the zero-sum restriction (i.e. the FH is the null) should be rejected. In other words, the standardized $v_1$ for inflation should not be significantly different from $-1$. Second, the restriction that this $v_1$ equals $-1/(1 - \tau)$, where $\tau$ is the implied tax rate, (i.e. the Darby hypothesis is the null) should not be rejected.\textsuperscript{12} The average highest federal corporate tax rate in force during the study period is used to proxy $\tau$.\textsuperscript{13} The FH is rejected for the US ($P = 0.033$) but not for the UK ($P = 0.603$). The Darby hypothesis, however, is not rejected for the US ($P = 0.896$) but is rejected for the UK ($P = 0.026$). This inconsistency indicates that the explanation relating interest rates and inflation is possibly more complicated than simply a tax effect, a supposition supported, e.g. by Peek and Wilcox (1983) who show that the Fisher equation is a reduced-form specification with the coefficient of expected inflation being determined not only of the tax rate but also of the structural relationships among output, money, and prices.\textsuperscript{14}

4.2. Properties of the common trend

Recall that the main purpose of this study is to investigate (1) if interest rates reflect rationally expected inflation rates and, consequently, contain information about the long-run path of future interest rates or (2) the opposite is true, i.e. inflation rates contain information to predict future interest rates. Investigating the common stochastic trend between the nominal interest rate and the inflation rate enables a distinction between these two hypotheses to be made. The results of estimating the unrestricted cointegration structure between the Eurocurrency rate and the inflation rate, all countries save France are reported in Table 2. The common stochastic trend driving this system is calculated using the

\textsuperscript{11} Details pertaining to UK taxing of interest rates are provided on the PricewaterhouseCooper web page, http://www.pwcglobal.com/uk/eng/ms-sol/publ/pockettax/PTB98-IT_allsalliukeng.

\textsuperscript{12} The US implied tax rate is 37.9\%. A similar estimate (32.9\%) is provided by Crowder and Wohar (1999, Table II, MLE column) for the relationship between US Treasury bills and inflation during 1950–1995. The UK implied tax rate is 0.7\% and the implied tax rates for the other seven countries run from a low of $-8.6\%$ for Italy to a high of 40\% for Germany.

\textsuperscript{13} Obtaining an economically precise value of $\tau$ is not possible. To do so requires, among other things, knowing who participates in the Eurocurrency markets and their federal and municipal tax status. The $\tau$ proxy is biased upward for those participants who have taxable profits from this source that fall in a tax bracket below the maximum. It is biased downward for those participants who pay municipal income taxes. Nevertheless, the $\tau$ values used, 39.9\% for US and 40\% for the UK, are plausible.

\textsuperscript{14} Shen (1998) finds an inflation risk premium in the UK gilt market and suggests that this premium may also affect the relationship between interest rates and inflation, especially for long-term gilts.
$m_2$ eigenvector values for each country. For instance, the common trend between the Eurocurrency rate and the inflation rate of the US is calculated to be $0.896$ (Eurocurrency rate) $+ 0.210$ (inflation rate). In all cases, the $m_2$ values indicate that the Eurocurrency rate dominates the long-run behavior of the common trend. This dominance is even more pronounced when the $m_2$ values are combined with the associated Eurocurrency and inflation rates. The average contributions of these two variables to the common trend are shown in the next to the last column in Table 2. The contribution of the Eurocurrency rate to this trend ranges from $71.0\%$ for Italy to $97.1\%$ for Norway and the UK. The nine country mean (median) is $91.9\%$ (96.1\%). Whether the Eurocurrency rate is completely responsible for the long-run path of the cointegrated variables is tested using the $L_{GG}$ statistic. The common trend $P$-values, which are reported in the last column of Table 2, indicate that in every instance the null hypothesis (as defined by the restriction matrix), that the Eurocurrency rate is the sole contributor is rejected.

The interpretation of these findings is consistent with Fama’s (1975) argument that nominal interest rates should contain useful information to predict the future path of the inflation rate and with the contention that interest rates to a large extent reflect rationally expected future inflation. Examination of the data supports the notion that participants in credit markets display forward looking behavior (i.e. anticipation), responding to news (innovations) about future inflation and incorporating this information into prices when setting interest rates.

5. Summary and concluding remarks

This paper investigates the long-run implications of the FH to characterize the changes in the level of nominal short-term Eurocurrency interest rates. The analysis indicates that, with but one exception, each country’s Eurocurrency rate is cointegrated with its inflation rate. That this long-run stable relationship exists is consistent with the long-run FH. Furthermore, for the majority of the countries in the sample, it is not rejected that there is a long-run, one-to-one relationship between changes in the expected inflation rate and the Eurocurrency rate. The latter finding indicates that the long-run FH receives more support from the international data than is found using US data.

A special emphasis is given to the long-run causal relationship between the Eurocurrency rates and inflation rates. The common stochastic trend that these variables share is identified and the Eurocurrency rate is found to be mainly responsible for the behavior of this common trend. This implies that there is information contained in the nominal rate concerning the future path of the inflation rate. This implications is consistent with Fama’s (1975) position that

15 For each country, the monthly common trend is calculated and the portions attributable to the interest rate and inflation rate computed. These monthly values are averaged.
rational agents in a well-functioning market incorporate the predictable portion of the inflation rate into nominal rates when setting prices.

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References