Matti Virén*
Bank of Finland Research Department
8.6.1989 20/89

THE LONG-RUN RELATIONSHIP BETWEEN INTEREST RATES AND INFLATION: SOME CROSS-COUNTRY EVIDENCE

Forthcoming in Journal of Banking and Finance.

* I am indebted to Maritta Lavikko for research assistance and to an anonymous referee and the participants of the first meeting of the Euro Working Group on Financial Modelling, in Rotterdam, and the 14th meeting of the European Finance Association, in Madrid, for helpful comments on an earlier draft. Financial support from the Nordic Economic Research Council and the Yrjö Jahnsson Foundation is gratefully acknowledged.

Suomen Pankin monistuskeskus
Helsinki 1989
ISBN 951-686-207-1
ISSN 0785-3572

ABSTRACT

This paper analyzes the effect of anticipated inflation on nominal interest rates, paying special attention to the stability of this relationship over various regimes. These regimes are analyzed to find out whether the inflation forecastability proposition recently advanced by Barsky explains the observed poor performance of the Fisher equation with historical data. The analyses make use of the so-called threshold and digression models and data from 11 western countries for 1875 - 1984. The main result of this paper is that the forecastability proposition does not explain the weakness of the anticipated inflation effect on interest rates. Rather, various institutional factors seem to be of more importance.

CONTENTS

1 INTRODUCTION

The belief that movements in nominal interest rates are mainly determined by expected inflation seems to be more or less at variance with empirical evidence. Practically the only piece of evidence supporting this proposition is the study of Fama (1975). Fama's results were, however, challenged by Summers (1983), who showed that the U.S. data for the period 1860 - 1979 (and for different subperiods) were clearly not in accordance with the underlying Fisher hypothesis. What was particularly striking in Summers's results was the fact that a hypothesis on the existence of a long-run Fisher relationship was also decisively rejected.[1]

Obviously, the result that the Fisher hypothesis cannot be used even as a long-run approximation is a little bit hard to swallow. One may therefore ask whether Summers's result is in fact due to the particular estimation method used (see fn. 1), the way inflation expectations are measured, or the data sample (Summers used U.S. data from the period 1860 - 1979). These are the questions we try to answer in this paper. To do so, we make use of relatively long time series on nominal interest rates and inflation for 11 countries, so that the longest sample period is 1875 to 1984. Given these data, we pay special attention to the properties of the time series process of the inflation and nominal interest rate variables. In particular, we study the stability of the time series process of inflation and the forecastability of inflation. We use so-called threshold and digression models to find out how stable or unstable the Fisher equation is and what is the explanation for any instabilities.

This article is organised as follows: Section 2 presents a short summary of the basic definitions of the Fisher hypothesis and discusses the explanations put forward for the observed poor performance of this hypothesis. In particular, the forecastability proposition is analyzed and some alternative test procedures for this proposition are derived. Results of estimating the Fisher equation with the data from 11 countries covering the period 1875 - 1984 are presented in section 3. Also presented in this section are the test results for the overall stability of the Fisher relationship and the inflation forecastability proposition, in particular. Some concluding remarks are presented in section 4.

2 MODEL FORMULATION AND TEST PROCEDURES

The starting point for the subsequent empirical analysis is the simple Fisher equation which - assuming that the real interest rate is constant - reads:

$$(1) \qquad R_t = a + E_t(p_{t+1}) + e_t,$$

where R_t is the nominal interest rate, $E_t(p_{t+1})$ the expected inflation rate and e_t an uncorrelated disturbance term. A standard way of testing the Fisher hypothesis boils down to estimating the following simple equation:

$$(2) \qquad R_t = b_0 + b_1 p_t^e + u_t,$$

where p_t^e is some proxy for the expected inflation rate, and where the coefficient restriction $b_1 = 1$ is tested. Obviously, the stochastic properties of u_t - which, in fact, may correspond to the stochastic properties of the real rate - should also be scrutinized.

Now, as was mentioned in the introduction, empirical evidence obtained using (2) has more or less been at variance with the Fisher hypothesis. A number of explanations for this finding have been advanced. Without going into the details, we may summarize these explanations or arguments as follows:

1) The Fisher hypothesis cannot account for short-run movements in nominal interest rates. This is due to various legal or contractual restrictions (see e.g. Ito and Ueda (1981), Stiglitz and Weiss (1981) and Kugler (1987)). The substitutability of money and financial assets (together with the fact that the nominal rate of return on money is

constant) may also contribute to this result (see Carmichael and Stebbing (1983) for details).

2) The poor performance of the Fisher hypothesis, particularly in the early sample periods, follows from the fact that the expected rate of inflation is not forecastable. In such a situation the data would reject the coefficient restriction $b_1 = 1$ even if (1) held (see Barsky (1987) for details).

3) Equation (2) may suffer from omitted variables affecting the real rate, from measurement errors both in terms of the after-tax nominal interest rate and the price index, simultaneity problems and from all such kinds of "standard" problems.

In the present context, we completely ignore the standard explanations mentioned in 3) above. This is because these problems have already been studied in detail elsewhere, and it is probably fair to say that these studies have not produced very affirmative results.[2] Another reason is that the data base used here does not allow for an explicit treatment of potential additional variables. This is particularly true for taxes. Instead, we concentrate on examining the behavior of interest rates and inflation only. In particular, we examine whether there are different regimes in terms of the Fisher effects, and, if there are, whether these regimes are determined by the time series properties of inflation, or by other (institutional) factors. We first investigate the "inflation explanation", i.e. the explanation which suggests that inflation has not been forecastable. This hypothesis has been put forward by Klein (1975) and Barsky (1987). Both of them argue that the data for the 1880 - 1915 gold standard period represent a clear example of inflation unpredictability. This is something which is not necessarily completely at variance with data. The price level for this period is almost stationary and one may suspect that the first log differences are white noise.

Briefly, this explanation is based on the following argument. If the (actual) inflation rate is uncorrelated over time, equation (2) fails to support the Fisher hypothesis even if it is true. This is clearly revealed by examining a simple example where a regression of the nominal rate R_t is run on p_t, the actual rate of inflation. The corresponding regression coefficient is $\mathrm{cov}(R_t,p_t)/\mathrm{var}(p_t)$, which given (1) equals $\mathrm{cov}(p_{t+1},p_t)/\mathrm{var}(p_t)$, which by definition is zero (this has also been pointed out by Barsky (1987)).

If, in turn, the regression is run in terms of R_t and p_t^e, the situation is somewhat different. If again the inflation rate is serially uncorrelated and p_t^e is generated by some AR(p) model, the resulting time series is merely a constant. Thus, coefficient b_1 in equation (2) is zero, even if (1) is true. The same result applies in the case p_t^e is generated as a least squares projection from the variables which are included in the agents' information sets if p_{t+1} is a martingale difference with respect to all these variables. Thus, we should carefully examine the properties of the inflation rate series so as to ensure that the expected rate of inflation is indeed forecastable on the basis of ex post inflation. Moreover, we should know to what extent the results obtained by regressing equation (2), where p_t^e is the one-period-ahead forecast of inflation generated by an AR(1) or by an AR(p) model, are conditional on assumption that the information sets include only lagged values of p. Unfortunately, this is something we cannot know or test here; we do, however, know that the importance of the information set assumption becomes quite crucial in the case where inflation is just "white noise".

In the empirical analysis we will examine the role of inflation rate forecastability and the stability of the basic relationship by running the following k-period moving regressions:

$$
(3)\qquad
\begin{aligned}
p_t &= b_{10} + b_{11}p_{t-1} + u_{1t} \\
R_t &= b_{20} + b_{21}p_t + u_{2t},
\end{aligned}
$$

where the estimation periods are (1+i, 1+k+i; i = 0,1,...,T-k). Clearly, an (unconstrained) estimation of (3) consitutes a rather weak test for the forecastability proposition. Thus, we try to arrange a more powerful test by scrutinizing the performance of the Fisher equation in different inflationary regimes. The basic idea is that the forecastability or persistence of inflation is essentially different in low and high inflation regimes. With this aim in mind we estimate equation (2) using so-called threshold models. They are based on the assumption that the parameters of the explanatory variables change according to some threshold variable, which in this case is assumed to be the absolute value of the expected inflation rate, denoted by z. Thus, we fit the following type of nonlinear equation to the data:

$$R_t = c_0^{(1)} + c_1^{(1)} p_t^e + u_t^{(1)} \quad \text{for all t with } z < z^*$$

(4)

$$R_t = c_0^{(2)} + c_1^{(2)} p_t^e + u_t^{(2)} \quad \text{for all t with } z > z^*$$

where z* is the fixed threshold value of the threshold variable (see Tong and Lim (1980) for details of the methodology). This threshold model experiment is based on the following notion: If the inflation (or deflation) rate is very low - i.e. the price level is more or less constant - agents may interpret price increases as purely temporary, i.e. as some kind of measurement errors. Thus they do not lead to any revisions in terms of the inflation rate, so that in such a situation the coefficient of p^e, $c_1^{(1)}$, goes to zero (see e.g. Andersen (1985) and Turnovsky (1969) for formal derivation of the inflation expectations schemes). This argument is, in fact, basically the same as the "multi-geared adaptive expectations hypothesis" presented by Flemming (1976). This hypothesis states that if the price level variable has exhibited no trend in recent years, then this is the variable to which economic agents will apply the adaptive expectations hypothesis. However, if the price level variable has revealed a trend, while its rate of change variable, the inflation rate, has not, then the adaptive scheme will be used to predict the inflation rate variable. All this suggests that $c_1^{(1)}$ should be less than $c_1^{(2)}$. This result should also emerge because in the case of deflation periods nominal interest rates cannot go below zero (assuming no interest is paid on money balances).

Obviously, it may well be that our data sample(s) does, in fact, contain two different regimes and thus two different regression relationships but that regimes are not determined by the inflation (or alternatively inflation uncertainty) variables. In order to account for this possibility we also estimate a digression model specification which is similar to (4) except that now no threshold variables are used. In other words, we estimate the following system of equations:

$$(5) \quad \begin{aligned} R_t &= d_0^{(1)} + d_1^{(1)} p_t^e + u_t^{(1)} \\ R_t &= d_0^{(2)} + d_1^{(2)} p_t + u_t^{(2)} \end{aligned}$$

so that the data are assumed to be generated by two regression relationships (1) and (2), and each observation is attributed to the nearest regression curve. Either the selective least squares or the maximum likelihood method is used. In the former case, the following nonlinear optimization problem is solved: $\sum_t \min \{(R_t - d_0^{(1)} - d_1^{(1)} p_t^e)^2, (R_t - d_0^{(2)} - d_1^{(2)} p^e)^2\}$. See Mustonen (1978, 1982) for further details of the methodology. Because no conditionalization is made with respect to the type of any single observation, estimating (5) does not tell anything about the economic rationale of these two regimes. Something can be said about it, however, by scrutinizing the corresponding parameter estimates and the regime indicators.

Before we are able to estimate equations (2), (4) and (5), we should, of course, determine how to construct a proxy variable for the unobservable expected rate of inflation (see e.g. Barsky (1987) and Friedman and Schwartz (1982, Ch. X) for a discussion of various alternatives and problems of measuring inflation expectations in this kind of setting). Here it is done by using the predicted values of an AR(2) model estimated recursively, starting from the first observation and always adding one observation to the data sample. Alternatively, a standard constant-parameter AR(2) model, and also an adaptive-expectations model were applied. In the latter case, the parameter λ in $p_t^e = (1 - \lambda)(p_t - p_{t-1}^e)$ was also allowed to vary over time in an adaptive way. Because the recursive AR(2) model clearly outperformed the alternative specifications in the empirical analyses we report here only the corresponding estimation results.

3 EMPIRICAL RESULTS

3.1 The Data

First, we present some comments on the data. As was noted in the introduction, the data are derived for 11 countries. The annual observations (after data transformations) cover the period 1873(1913) - 1984 (the countries and the exact sample periods are reported in Table 1). The nominal interest rate series correspond to long-term government bond yields, whereas the consumer price index is used for p. In addition, two dummy variables were generated for the World Wars.

The main data sources are Maddison (1982), Mitchell (1980, 1983), the Statistical Year-Book of the League of Nations and the IMF International Financial Statistics. In addition, some national sources are also used. In the case of the Nordic countries, both the growth studies and the monetary histories and/or the histories of the central banks provide the main part of the data. As far as other countries are concerned, the following special studies could be referred to: Sheppard (1971), De Mattia (1978) and Butlin, Hall and White (1971).

If one scrutinizes the data for R_t and p_t, it is obvious that the data contains different subperiods: the international gold standard period 1880 - 1914, World War I 1914 - 1918, the 1920s and 1930s, World War II 1939 - 1945, the Bretton Woods period until 1973, and, finally, the "current" period until 1984. As mentioned earlier, there are several empirical studies which focus on the behavior of interest rates and inflation during these different periods (cf. e.g. Bloomfield (1959), Klein (1975), Friedman and Schwartz (1982) and Barsky (1987)). These analyses suggest that the behavior of interest rates and inflation has not been invariant over these subperiods. Given this observation and the institutional differences between subperiods one may argue that the Fisher relationship will ultimately fail in a sample like 1875 - 1984. Whether this true or not, is one question we will examine in the subsequent empirical analysis.[3]

Table 1. Estimation Results of Equation (2)

	$\hat{b}_0$	$\hat{b}_1$	SEE	D-W	ARCH4	ρ	t_0
Australia	4.32	.643	2.04	0.63	44.20	-	1875
	(18.74/18.44)	(8.34/4.66)					
	7.57	.040	0.79	1.88	4.56	.988	1875
	(1.91/2.67)	(1.19/1.31)					
Canada	4.27	.800	2.34	0.73	26.99	-	1903
	(12.19/13.40)	(5.50/2.90)					
	7.08	.014	0.67	1.56	24.11	.991	1903
	(1.70/1.98)	(0.44/0.89)					
Denmark	5.63	.437	2.94	0.39	56.92	-	1875
	(17.23/25.48)	(6.10/3.85)					
	6.63	.049	1.04	1.98	12.83	.960	1875
	(3.19/3.67)	(1.64/2.02)					
Finland	6.77	.028	1.74	0.24	75.37	-	1875
	(37.47/37.08)	(1.28/1.35)					
	7.11	-.021	0.69	2.04	7.10	.951	1875
	(6.16/5.82)	(2.97/2.63)					
France	5.80	.197	2.46	0.27	50.25	-	1913
	(13.26/18.46)	(2.85/2.48)					
	7.36	.018	0.82	1.25	13.12	.975	1913
	(2.73/3.18)	(0.89/1.19)					
Italy	5.38	.119	3.52	0.20	87.34	-	1875
	(14.27/16.83)	(3.61/2.41)					
	8.54	.003	0.87	1.14	31.91	.989	1875
	(1.75/2.45)	(0.37/0.45)					
Netherlands	4.77	.684	1.61	1.02	10.54	-	1913
	(20.22/26.54)	(6.51/5.83)					
	5.89	.010	0.62	1.69	20.47	.971	1913
	(3.23/4.43)	(0.28/0.49)					
Norway	5.46	.260	1.75	0.48	63.07	-	1875
	(29.90/33.92)	(4.75/3.59)					
	7.27	.003	0.65	1.93	17.07	.984	1875
	(2.77/3.69)	(0.19/0.19)					
Sweden	4.96	.347	2.02	0.52	60.21	-	1875
	(22.83/26.94)	(5.88/4.15)					
	7.36	.004	0.59	0.59	25.97	.991	1875
	(2.00/2.63)	(0.28/0.27)					
United Kingdom	4.64	.978	1.58	1.11	9.96	-	1875
	(28.46/29.84)	(18.06/16.13)					
	6.14	.068	0.61	1.40	6.57	.987	1875
	(2.07/2.53)	(2.02/2.10)					
United States	4.62	.463	1.91	0.43	82.68	-	1875
	(23.64/21.22)	(6.16/3.89)					
	8.28	.035	0.47	1.41	42.29	.997	1875
	(1.53/2.66)	(1.85/1.94)					

Numbers in parentheses are t-ratios; the first number is the standard t-ratio and the second number is White's heterocedasticity adjusted t-ratio. SEE is the standard error of the estimate, D-W the Durbin-Watson autocorrelation statistic, ARCH4 Engle's autocorrelation conditional heterocedasticity statistic (the lag length being 4), ρ the autocorrelation adjustment parameter used in the Cochrane-Orcutt procedure, and t_0 the first year of the sample period (the last year is in all cases 1984). All equations also include two dummy variables for the World Wars (i.e. D1 = 1: t = 1914 - 1918 and D2 = 1: t = 1939 - 1946). For reasons of space the respective parameter estimates are not, however, reported here.

3.2 Empirical Results

Turning now to the empirical results, presented in Table 1, we consider first the OLS estimates of equation (2). At least the following conclusions can be drawn from these results: First, as a rule the coefficient estimate of p_t^e deviates considerably from unity, even though there are some notable exceptions, for instance the United Kingdom. Because of strong autocorrelation of residuals, it is not, however, meaningful to test whether the coefficient restriction $\alpha_1 = 1$ actually holds for these countries. Second, the residuals of both equations are strongly autocorrelated and heteroscedastic. Thus, the D-W statistics suggest in most cases that the residuals have almost a unit root representation, and, thus, R_t and p_t^e do not seem to be co-integrated.

As far as residual autocorrelation is concerned, the standard way of eliminating it, i.e. the use of the Cochrane-Orcutt procedure, does not produce satisfactory results. The autocorrelation coefficient ρ, which is used in filtering the data, turns out to be close to unity but even then the residuals remain autocorrelated in the case of several countries. There is not much improvement if the data are filtered with the value of ρ obtained from the OLS residuals. In that case, autocorrelation is reduced only slightly - in some cases the values of the D-W statistic even decrease! All this should be borne in mind when examining the parameter estimates of p_t^e obtained by the Cochrane-Orcutt procedure. The estimates are, in fact, very low - it is simply not possible to reject the hypothesis that the respective coefficients equal zero. It does not really help very much that the coefficients (with the exception of Finland) remain positive because the numerical values are of the magnitude 0 to 0.1![4]

3.3 Examining the Role of Inflation Forecastability

As pointed out earlier, if the inflation rate is not forecastable the Fisher hypothesis would seem to fail to explain the data even if the hypothesis is true. To find out how relevant this point is we estimate (3) using the moving regression approach, so that the estimation period is 20 in each point of the data sample. The corresponding coefficient estimates are presented in Figure 1. It can be clearly seen that the inflation rate series is almost uncorrelated for the pre-1918 period, while correlation in the 1970's and the 1980's is almost unity. For the rest of the sample period correlation is somewhere between 0 and 1. It cannot therefore really be claimed that inflation has not been forecastable for the entire sample period in different countries. The "Fisher equation coefficient", $\hat{b}_{21}$, in turn, behaves in the same way as the autoregression coefficient, $\hat{b}_{11}$, for 1875 - 1918 and 1970 - 1984 <u>but not for the rest of the sample period</u>. Thus, $\hat{b}_{11}$ deviates rather clearly from zero both for the 1920's and 1930's and the 1950's and 1960's but the same cannot be said of $\hat{b}_{21}$. Accordingly, the hypothesis advanced by Barsky (1987) cannot be fully accepted.

The moving regression results suggest that there are some changes in the time-series process of the inflation rate series over time but that these changes do not seem to coincide with the changes in the structure of correlation between nominal interest rates and inflation. The results with the threshold model specification (4) are generally consistent with these findings. These results - which are presented in Table 2 - indicate that equation (2) is outperformed by a non-linear threshold model specification, and particularly by a threshold model specification which allows for different variances in the two regimes.[5] What is more interesting is the behavior of the coefficient of p_t^e. Only in 3 cases out of 11 does the coefficient increase when we move from a low inflation (or deflation) rate regime to a high inflation (or deflation) rate regime. So, <u>we cannot simply say that the higher is the inflation rate the higher is the Fisher equation coefficient!</u>

So, we are left with the result that the Fisher equation is very unstable but that the instability cannot be explained by the rate of inflation. The instability becomes even more evident when we estimate the digression model specification (5), which gives us two separate regression relationships for different regimes of the data sample. The respective estimation results, which are presented in Table 3, indicate that (except for Finland) these two regimes are completely different in the sense that the first regime (the order of the regime

Table 2. Estimation Results with the Threshold Model (6)

Country	Regime	$\hat{c}_0$	$\hat{c}_1$	SEE	SBIC	SBICT	$\tau(2)$	z*
Australia	1	3.68	.348	1.30	44.1	114.9	13.35	2.20
		(24.55)	(2.37)			150.6		(2.30)
	2	5.92	.470	3.02	70.8	155.0		
		(7.34)	(2.99)					
Canada	1	3.73	.873	1.31	33.4	94.9	7.29	1.51
		(13.10)	(2.79)			133.9		(2.50)
	2	6.08	.540	3.74	61.5	132.7		
		(4.83)	(1.71)					
Denmark	1	4.50	.155	0.61	-35.4	98.2	18.1	2.27
		(53.19)	(2.20)			208.0		(1.98)
	2	7.32	.357	3.98	133.8	217.1		
		(2.78)	(9.30)					
Finland	1	6.16	.314	1.31	45.7	93.3	28.30	4.85
		(37.22)	(4.42)			104.7		(3.02)
	2	8.01	-.015	2.02	47.6	123.9		
		(19.85)	(0.46)					
France	1	4.74	.334	1.07	12.1	75.8	14.96	4.27
		(17.17)	(3.38)			116.3		(7.46)
	2	9.16	-.095	3.49	63.8	123.1		
		(6.47)	(0.61)					
Italy	1	3.88	.426	1.42	63.4	168.7	17.0	5.01
		(20.54)	(6.13)			229.9		(2.95)
	2	8.06	.110	5.17	105.3	237.8		
		(5.38)	(1.07)					
Netherlands	1	4.31	.453	1.03	9.8	43.6	10.10	1.82
		(25.23)	(2.73)			57.8		(3.63)
	2	5.52	.759	1.93	33.8	59.7		
		(8.67)	(3.69)					
Norway	1	4.78	.322	0.71	29.1	51.6	17.34	1.67
		(49.35)	(2.96)			115.1		(2.44)
	2	.630	.241	2.34	80.7	123.3		
		(16.65)	(2.87)					
Sweden	1	4.46	.308	1.01	10.1	83.3	15.26	2.91
		(36.22)	(3.91)			144.7		(2.08)
	2	6.46	.250	3.15	73.2	150.8		
		(8.87)	(2.02)					
U.K.	1	4.10	.966	0.98	5.8	54.0	30.66	1.56
		(33.73)	(6.33)			81.7		(2.22)
	2	6.10	.860	2.04	48.2	103.2		
		(14.45)	(10.84)					
U.S.A.	1	3.81	.720	0.88	-2.6	103.2	11.38	1.25
		(26.73)	(3.12)			140.7		(1.12)
	2	5.13	.497	2.32	105.8	140.7		
		(16.77)	(5.15)					

SBIC indicates the value of the Schwartz information criterion for one equation of the threshold model, SBICT (first row) the corresponding value for the whole model, SBICT (second row) the corresponding value for the whole model estimated assuming that the error variances for both model structures are equal, and SBICT (third row) the corresponding value for a linear model (i.e. (2)). z* is the value of the (absolute) inflation threshold and $\tau(2)$ a chi square test statistic (with 2 degrees of freedom) for the hypothesis that the coefficients above and below the threshold are equal (this statistic is computed assuming equal error variances). The sample median values of p are reported in parentheses below z*. The estimation periods are the same as in Table 1 except for the World War years, which have been omitted here.

Table 3. Estimation Results with the Digression Model

Country	Regime	$\hat{d}_0$	$\hat{d}_1$	SEE	n
Australia	1	5.17	-.061	.40	43
		(80.73)	(2.56)		
	2	3.08	1.216	1.71	54
		(11.13)	(13.93)		
Canada	1	3.30	.054	0.58	34
		(22.99)	(0.87)		
	2	4.01	1.844	1.08	35
		(14.17)	(15.28)		
Denmark	1	4.53	.095	0.64	68
		(57.57)	(4.70)		
	2	4.76	1.399	2.40	29
		(5.46)	(7.71)		
Finland	1	5.21	.021	0.52	46
		(65.62)	(1.84)		
	2	8.13	.061	1.32	51
		(42.53)	(2.16)		
France	1	5.02	.058	0.62	36
		(37.91)	(2.89)		
	2	1.54	1.547	1.55	23
		(1.91)	(9.74)		
Italy	1	4.60	.027	0.58	45
		(50.39)	(2.99)		
	2	2.88	.920	1.81	52
		(8.90)	(19.00)		
Netherlands	1	3.25	-.019	0.13	12
		(84.09)	(0.77)		
	2	4.80	.924	1.50	47
		(17.75)	(7.27)		
Norway	1	4.96	.004	0.50	67
		(81.66)	(0.20)		
	2	5.05	.978	2.02	30
		(10.34)	(6.42)		
Sweden	1	4.27	.041	0.58	67
		(60.52)	(1.87)		
	2	5.09	.900	2.19	30
		(8.76)	(6.17)		
U.K.	1	3.05	.037	0.23	48
		(73.21)	(1.38)		
	2	5.36	.946	1.71	49
		(18.52)	(12.75)		
U.S.A.	1	3.21	-.116	0.34	55
		(60.81)	(5.12)		
	2	5.28	.698	2.37	42
		(13.28)	(4.73)		

The notation and data are the same as in Table 2. The estimates are maximum likelihood estimates obtained using the digression model specification (5).

is, of course, completely arbitrary) is characterized by no relationship between R_t and p_t^e while the second regime is more or less consistent with the Fisher hypothesis (notice that because of taxes the parameter c_1 should probably exceed one). Altogether, about one half of the observations belong to either of these two regimes. This does not mean that the regimes follow a completely arbitrary pattern. Almost all observations for the period 1956 - 1984 and about one half of the observations for the period 1919 - 1933 belong to the second ("Fisher equation") regime. The third period which could be mentioned in this context is the late 1890s.[6] This pattern suggests that various legal and/or contractual restrictions have crucially affected the behavior of interest rates. Obvious examples of such restrictions are the regulation Q, other restrictions on the operations of banks, as well as regulation of prices and credit rationing elements in the banking sector (recall the discussion in Section 2 above). Thus, the fact that interest rates have been so persistent is probably not so much due to inflation forecastability problems as to various market inefficiencies. This conclusion contains, however, one obvious caveat. It may well be that changes in the relationship between (our measure of) expected inflation and nominal interest rates follow with a long lag changes in the degree of autocorrelation of inflation and, hence, the periods of high forecastability of inflation do not overlap with periods during which the Fisher equation seems to hold.

4 CONCLUDING REMARKS

On the basis of the empirical evidence it seems justified to argue that nominal interest rates and inflation have been only weakly related during the past hundred years. This does not imply that the Fisher hypothesis is not true. The measures which we can develop for expected inflation are obviously deficient and may contain systematic measurement errors. Moreover, the historical data which we have used seem to include different regimes and the most current data seem mainly to belong to the "Fisher equation regime". As far as these regimes are concerned, it turns out that the inflation forecastability proposition cannot explain the observed poor performance of the Fisher hypothesis but only for a part of the data sample. Rather, various institutional factors seem to provide more insight into the relationship between inflation and interest rates over the past decade.

FOOTNOTES

1. The analytical method used by Summers - i.e. the band spectrum approach - has subsequently been criticized by McCallum (1984, 1986). McCallum's basic argument is that the low-frequency measures applied by Summers cannot take into account the distinction between anticipated and unanticipated fluctuations in various variables. This distinction, in turn, is crucial for relationships such as the Fisher equation. In commenting on these criticisms, Summers (see Summers (1986)) argues that the points raised by McCallum are not empirically important.

2. See, for instance, Levi and Makin (1979), Makin (1983) and Peek and Wilcox (1983).

3. One can, however, argue that even though there have been different regimes, people have not immediately realized that. Thus, following Klein (1975), one can argue that, for instance, altough the United States went off the gold standard de jure in 1933, gold standard expectations persisted de facto in the 1960s.

4. One may suspect here that the results can at least partially be explained by the omission of the tax variable, particularly for the post-WWII sample. We could not, however, directly test this hypothesis.

5. The estimation and testing procedures of the threshold models are explained in detail in Luukkonen (1983).

6. In addition to the maximum likelihood method the selective least squares method was also applied in computing the digression model. The results were practically identical with this method. The same was true with alternative initial values of the regime indicator. (cf. Mustonen (1978) for details.)

Figure 1 MOVING REGRESSION COEFFICIENTS

——— Inflation rate autoregression coefficient

- - - - Fisher equation coefficient

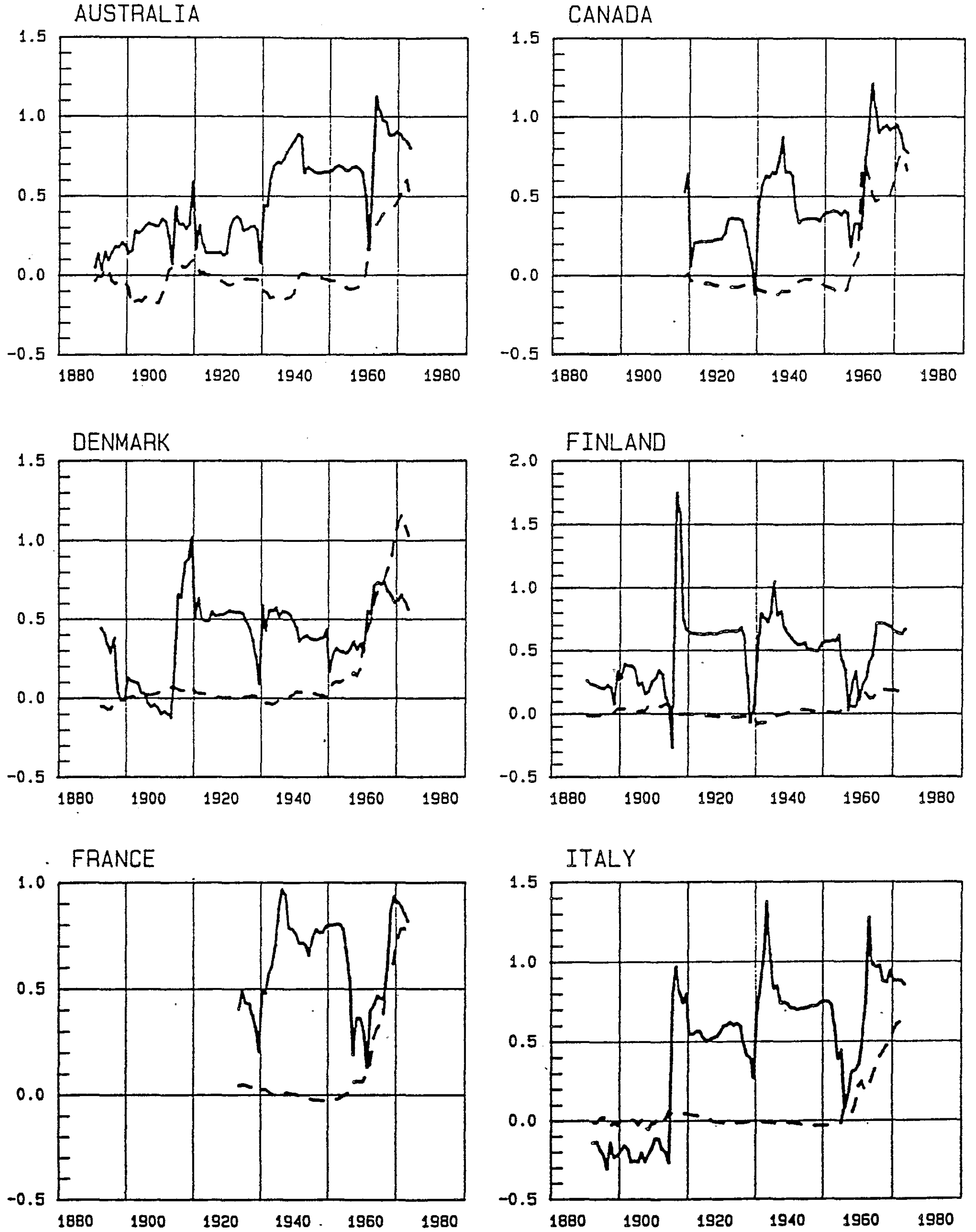
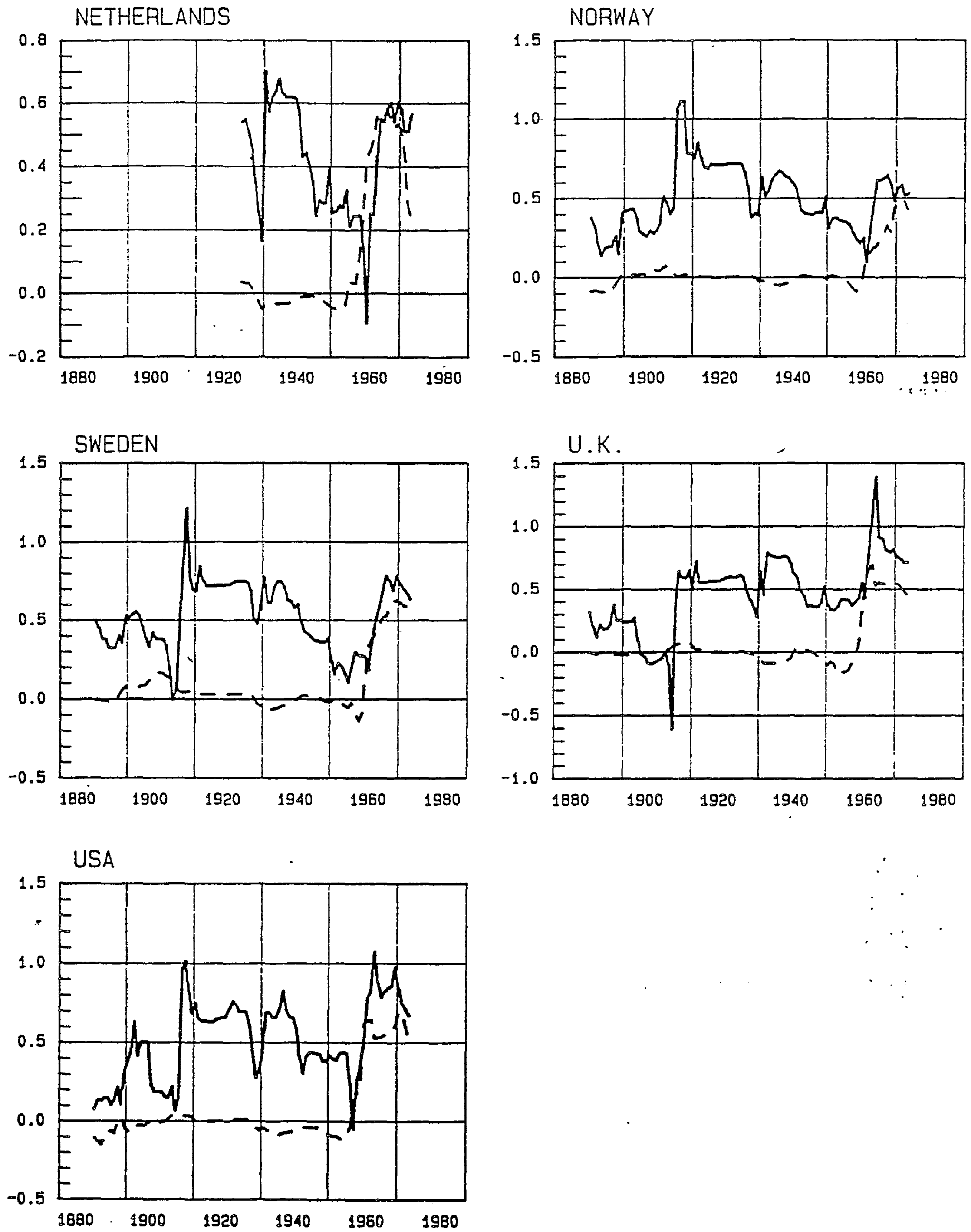


Figure 1 Continued

——— Inflation rate autoregression coefficient

- - - - Fisher equation coefficient



REFERENCES

ANDERSEN, T.M., 1985, Price Dynamics under Imperfect Information, Journal of Economic Dynamics and Control 9, 339 - 361.

BARSKY, R.B., 1987, The Fisher Hypothesis and the Forecastability and Persistence of Inflation, Journal of Monetary Economics 19, 3 - 24.

BLOOMFIELD, A.I., 1959, Monetary Policy under the International Gold Standard (Federal Reserve Bank of New York, New York).

BUTLIN, S.J., HALL, A.R. and R.C. WHITE, 1971, Australian Banking and Monetary Statistics 1817 - 1945, Occasional Paper No. 4A (Reserve Bank of Australia, Sidney).

CARMICHAEL, J. and P.W. STEPPING, 1983, Fisher's paradox and the Theory of Interest, American Economic Review 73, 619 - 630.

DE MATTIA, R., 1978, Storia del Capitale della Banca d'Italia e Degli Instituti Predecessori (Banca d'Italia, Roma).

FAMA, E.G., 1975, Short-Term Interest Rates as Predictors of Inflation, American Economic Review 65, 269 - 282.

FLEMMING, J., 1976, Inflation (Oxford University Press, Oxford).

FRIEDMAN, M. and A.J. SCHWARTZ, 1982, Monetary Trends in the United States and the United Kingdom (University of Chicago Press, Chicago and London).

ITO, T. and K. UEDA, 1981, Tests of the Equilibrium Hypothesis in Disequilibrium Econometrics: An International Comparison of Credit Rationing, International Economic Review 22, 691 - 708.

KLEIN, B., Our New Monetary Standard: The Measurement Effects of Price Uncertainty, 1880 - 1973, Economic Inquiry 13, 461 - 484.

KUGLER, P., 1987, Credit Rationing and the Adjustment of the Loan Rate: An Empirical Investigation, Journal of Macroeconomics 9, 505 - 525.

LEVI, M.D. and J.H. MAKIN, 1979, Fisher, Phillips, Friedman and the Measured Impact of Inflation on Interest, Journal of Finance 34, 35 - 52.

LUUKKONEN, R., 1983, SURVO: Programs for Time Series Analysis, Department of Statistics, University of Helsinki, Research Report, No. 40.

MADDISON, A., 1982, Phases of Capitalist Development (Oxford University Press, Oxford).

MAKIN, J.H., 1983, Real Interest, Money Surprises, Anticipated Inflation and Fiscal Deficits, Review of Economics and Statistics 65, 374 - 384.

McCALLUM, B.T., 1984, On Low-Frequency Estimates of Long-Run Relationships in Macroeconomics, Journal of Monetary Economics 14, 1 - 14.

McCALLUM, B.T., 1986, Estimating the Long-Run Relationship between Interest Rates and Inflation: A Reply, Journal of Monetary Economics 18, 87 - 90.

MITCHELL, B.R., 1980, European Historical Statistics 1750 - 1975 (Macmillan, London).

MITCHELL, B.R., 1983, International Historical Statistics: The Americas and Australasia (Macmillan, London).

MUSTONEN, S., 1978, Digression Analysis: Fitting Alternative Regressions to Heterogenous Data, in Proceedings in Computational Statistics, Compstat 1978, (Vienna) 95 - 101.

MUSTONEN, S., 1982, Digression Analysis, Encyclopedia of Statistical Sciences 2 (John Wiley & Sons, New York), 373 - 374.

PEEK, J. and J.A. WILCOX, 1983, The Postwar Stability of the Fisher Effect, Journal of Finance 38, 1111 - 1124.

SHEPPARD, D.K., 1971, The Growth and Role of UK Financial Institutions 1880 - 1962 (Methuen, London).

STIGLITZ, J. and A. WEISS, 1981, Credit Rationing in Markets with Imperfect Information, American Economic Review 71, 393 - 410.

SUMMERS, L.H., 1983, The Nonadjustment of Nominal Interest Rates: A Study of the Fisher Effect, in J. Tobin (Editor): Macroeconomics, Prices and Quantities: Essays in Memory of Arthur M. Okun (Basil Blackwell, Oxford) 201 - 241.

SUMMERS, L.H., 1986, Estimating the Long-Run Relationship between Interest Rates and Inflation: A Response to McCallum, Journal of Monetary Economics 18, 77 - 86.

TONG, H. and K.S. LIM, Threshold Autoregression: Limit Cycles and Cyclical Data: Journal of the Royal Statistical Society, Series B, 245 - 292.

TURNOVSKY, S.T., 1969, Bayesian Approach to the Theory of Expectations, Journal of Economic Theory 1, 220 - 227.

BANK OF FINLAND DISCUSSION PAPERS

ISSN 0785-3572

1/89 PAULA LÄHDEMÄKI Economic indicators of the U.S.S.R. analysed on the basis of the National Accounts. 1989. 57 p. In Finnish. (ISBN 951-686-182-2)

2/89 MATTI VIRÉN A note on interest rate policy during the great depression. 1989. 20 p. (ISBN 951-686-183-0)

3/89 ERKKI KOSKELA - MATTI VIRÉN International differences in saving rates and the life cycle hypothesis: a comment. 1989. 20 p. (ISBN 951-686-184-9)

4/89 SAMPO ALHONSUO The efficiency of financing and banking in Finland. 1989. 81 p. In Finnish. (ISBN 951-686-185-7)

5/89 AMY SKOLNIK The U.S. - Canada free trade agreement: a model for Finland? 1989. 26 p. (ISBN 951-686-186-5)

6/89 JUHA TARKKA - ALPO WILLMAN - CHRIS-MARIE RASI Labour supply, wages and prices in the BOF4 quarterly model of the Finnish economy. 1989. 50 p. (ISBN 951-686-187-3)

7/89 JARMO KONTULAINEN Determination of exchange rates in a general equilibrium model. 1989. 80 p. In Finnish. (ISBN 951-686-188-1)

8/89 ESKO SYDÄNMÄKI New protectionism. 1989. 41 s. In Finnish. (ISBN 951-686-189-X)

9/89 JUHA TARKKA - ALPO WILLMAN - HANNA-LEENA MÄNNISTÖ Consumption and investment in the BOF4 quarterly model of the Finnish economy. 1989. 59 p. (ISBN 951-686-190-3)

10/89 SAMPO ALHONSUO - KJELL PETER SÖDERLUND - JUHA TARKKA Return on bonds in Finland in 1948 - 1986. 1989. 34 p. In Finnish. (ISBN 951-686-193-8)

11/89 PENTTI PIKKARAINEN - MATTI VIRÉN Granger causality between money, output, prices and interest rates: some cross-country evidence from the period 1875 - 1984. 1989. 19 p. (ISBN 951-686-195-4)

12/89 HELVI KINNUNEN Estimation of preliminary current account figures by means of short-term forecasting methods. 1989. 20 p. In Finnish. (ISBN 951-686-196-2)

13/89 PERTTI HAAPARANTA - JARMO KONTULAINEN Real exchange rate as an unobservable variable. 1989. 17 p. (ISBN 951-686-197-0)

14/89 MATTI VIRÉN Saving, investment and the current account: a review of recent evidence. 1989. 17 p. (ISBN 951-686-198-9)

15/89 HARRI LAHDENPERÄ The effect of information on the functioning of financial markets and on the efficiency of central bank policy - a survey of the literature. 1989. 55 p. In Finnish. (ISBN 951-686-199-7)

16/89 PAAVO PEISA Aggregate versus industry-specific sources of economic growth and business cycle fluctuations. 1989. 35 p. (ISBN 951-686-202-0)

17/89 TIMO TYRVÄINEN Unions, wages and employment in Finland. 1989. 56 p. (ISBN 951-686-203-9)

18/89 ANTTI ILMANEN Use of duration analysis in the estimation and management of interest rate risk of bonds. 1989. 116 p. In Finnish. (ISBN 951-686-204-7)

19/89 MATTI VIRÉN How does domestic and foreign money growth affect the U.S. economy? 1989. 24 p. (ISBN 951-686-205-5)

20/89 MATTI VIRÉN The long-run relationship between interest rates and inflation: some cross-country evidence. 1989. 24 p. (ISBN 951-686-207-1)