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Future Inflation and the Information in International Term Structures

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Abstract: This paper extends previous work on the information content of the term structure of interest rates using a newly constructed dataset for the United States, Japan, Germany, Switzerland, France, Belgium and the Netherlands (1982–1991). Results significantly differ from Jorion and Mishkin (1991). Apparently, the relation between the term structure of interest rates and future inflation is highly period- and country-dependent. We provide new evidence that these results may be due to the inability of financial markets to accurately predict a term structure of inflation in combination with the conduct of monetary policy. This probably accounts for large variation in ex post real interest rate levels and the term structure of real interest rates. Consequently, it is unlikely that the term structure of nominal interest rates will serve as a good indicator of future inflationary developments.

JEL Classification System-Numbers: E43, E52

1 Introduction

In the past decade, much research has focused on the information in the term structure of nominal interest rates. Initially, most research has been directed at the power of the term structure to forecast future short- and long-term interest rates, see for example, Fama (1984), Shiller, Campbell and Schoenholtz (1983), Mankiw (1986) and Fama and Bliss (1987). Results have been mixed. On the one hand, Mankiw (1986) and Shiller, Campbell and Schoenholtz (1983) conclude that the term structure has little power to predict future interest rates. They, therefore, reject the simple expectations theory of the term structure and, moreover, fail to provide positive evidence on the existence of time-varying risk premia, which might explain the lack of success. Fama (1984) and Fama and Bliss (1987), on the other hand, suggest that implicit forward rates contain information that helps predicting future excess returns, especially for the longer maturities.

[1] Gerard Opsteeg importantly contributed in gathering and organizing the data. Francois Nissen provided skillful assistance in the computational analysis. Helpful comments from Manfred Neumann, participants of the Konstanz Seminar on Monetary Theory and Monetary Policy 1993, and from two anonymous referees are acknowledged. Any remaining errors are ours alone.

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Recently, research has taken another angle, focusing more on the power of the nominal term structure to predict future inflation and/or real interest rates, see for example Mishkin (1990a,b, 1991), Fama (1990), and Jorion and Mishkin (1991). Most empirical work in this direction has concentrated on the United States, though Mishkin (1991) and Jorion and Mishkin (1991) extend this research across countries. Empirical results are mixed, on average, and heavily depend on the existence of a constant term structure of real interest rates over time. Nevertheless, especially the longer maturity results seem to point at the presence of significant information in the term structure of nominal interest rates about the term structure of (future) inflation. See Estrella and Hardouvelis (1991) for a more extensive overview of this literature.

The above overview of the empirical literature shows that although the term structure of nominal interest rates may contain valuable information about future inflation, it is notoriously difficult to extract this information. Alternatively, a skeptic might on the basis of the same evidence conclude that there is not much to gain from studying the term structure of interest rates and that observed linkages are likely to be both period- and country-dependent.

In this paper, our purpose is to extend the available empirical evidence in a number of directions and to critically evaluate the claim that the term structure of interest rates indeed carries information about future inflation, which may be used for policy purposes.

First, we construct a new data set of yield to durations for 1 to 5 year fixed duration bonds for seven countries – the US, Japan, Germany, Switzerland, France, Belgium and the Netherlands – for the period 1982–1991. Especially for the last three European countries, it is to our knowledge the first representative data set of this kind, allowing a broader test of the theory than possible so far. We first apply the methodology as described by Jorion and Mishkin (1991) – henceforth JM – for each country. Subsequently, we explicitly analyze the failure of the term structure of interest rates to predict future inflation by distinguishing between variation in the level and term structure of real interest rates and investigate to what extent this variability may be explained by the impossibility to forecast longer-term inflation and by the conduct of monetary policy.

The paper is structured as follows. In section 2, the theoretical assumptions underlying the regressions are formulated, while section 3 contains information about the construction of the data and the statistical characteristics of the data. In section 4 the JM-regressions are replicated and additional experiments are presented. Section 5 contains a summary and conclusions.

2 A Theoretical Model of Interest Rates and Inflation

To investigate the information content of the term structure of interest rates in the United States, Belgium, France, Germany, Japan, the Netherlands and Switzerland, we follow Mishkin (1990a,b, 1991), and JM in estimating the following inflation change equation:

$$\pi_t^m - \pi_t^1 = \alpha^m + + \beta^m(i_t^m - i_t^1) + \varepsilon_t^m \tag{1}$$

In equation (1) the forward-looking annualized ex post m-year inflation rate minus the one-year ahead ex post inflation rate is regressed on the slope of the term structure, defined as the spread between the annualized m-year interest rate and the one-year interest rate. Both inflation rates and interest rates are measured at time t. Throughout the analysis, all inflation rates and interest rates are continuously compounded.

Underlying equation (1) is the Fisher equation, according to which the m-period nominal interest rate equals the expected inflation over m periods plus the m-period real interest rate:

$$i_t^m = E_t\pi_t^m + r_t^m \tag{2}$$

where E_t is the expectations operator, π_t^m is the inflation rate from time t to $t + m$, i_t^m is the nominal interest rate at time t with maturity m and where r_t^m is the ex-ante real interest rate at time t with maturity m. The realized inflation over the period from t to $t + m$ can be written as the expected inflation defined above plus a forecast error:

$$\pi_t^m = E_t\pi_t^m + \upsilon_t^m \tag{3}$$

where υ_t^m is the forecast error. From (2) and (3) the following equation for the difference between the m-year and one-year ahead realized inflation may be derived:

$$\pi_t^m - \pi_t^1 = (i_t^m - i_t^1) - (r_t^m - r_t^1) + \upsilon_t^m - \upsilon_t^1 \tag{4}$$

Under the assumptions that expectations are formed rationally and that the slope of the real term structure, $(r_t^m - r_t^1)$, is constant over time, equation (4) imposes restrictions on equation (1): β^m equals unity and the error term is uncorrelated. More generally, a positive coefficient β^m in equation (1) that is significantly different from zero implies that a steep slope of the term structure of interest rates predicts rising future inflation, see in Mishkin (1990a,b, 1991) and JM, so that the slopes of the real and nominal term structures do not move together perfectly. A value of β^m in equation (1) significantly different from one reveals that the slope of term structure of real interest rates is not constant over time and so, that the nominal term structure contains information on the term structure of real interest rates.

3 Data

Monthly data on individual bond prices, duration and yield to duration have been obtained from Datastream for Belgium, France, Germany, Japan, the

Netherlands, and Switzerland.[2] Similar data for the US were taken from the CRSP[3]-tape. The starting date differs across countries because of differences in availability of data. In the empirical analysis, January 1982 is chosen as the general starting date, as from that time onward data for all seven countries are available. In a number of cases, the same experiments are also performed for the US and Germany only, starting in April 1976. The sample ends in December 1992 for all countries, except for the US. The available CRSP tape ends in December 1991. For the period January–November 1992, monthly issues of the Federal Reserve Bulletin were used to update the US series. For our purpose, we needed constant-maturity yield-to-duration time series. The procedure is outlined in the appendix. Consumer price series have been taken from the IFS tape of the IMF and have been used to calculate 1-year to 5-year inflation rates as defined in section 2.

Table 1 and 2 contain information about the statistical characteristics of the one year interest rate and inflation rate levels and 2-1, 3-1, 4-1, and 5-1 interest rate and inflation spreads for the sample period starting in January 1982. Table 1 shows that inflation in general has been quite low across countries in the 1980s, ranging from a low of 1.8 percent for Japan to a high a 4.4 percent in France. Second, the term structure of inflation as measured by the calculated spreads has been approximately flat for Japan, Germany and the Netherlands, significantly declining for Belgium and France and significantly rising for the US and Switzerland. For Belgium and France it reflects the slow disinflation in the EMS in the mid-eighties, while for the US it is a sign of the low inflation level at the start of the sample, after the fast deflation in 1980–81. Third, inflation levels and spreads for each country show significant positive correlation with the corresponding US and German variables, indicating common trends in short and long term inflation behavior.

For the interest rates in table 2 a slightly different picture emerges. First, interest rate levels differ much more across countries than inflation rates, ranging from 4.8 percent in Switzerland to 10.2 percent in France. Although interest rate levels are significantly positively correlated, the correlations of interest spreads for Belgium, France and Switzerland with those in Germany and the US are low and often (insignificantly) negative. Also, the standard deviation of interest rate spreads is in general smaller than for inflation spreads.

By way of illustration, figures 1 and 2 display the 5-1 inflation and interest rate spread for the US and Germany, respectively, for the longest available sample period 1976–1992. These graphs show that the interest and inflation spreads for the US and Germany experience a similar downward swing in the

[2] Missing from this list is the United Kingdom. Initially, we sampled UK data from Datastream as well. We discovered large errors in many observations, however, due to a Datastream programming error. So far, Datastream has been unable to correct this problem. We, therefore, have excluded the UK from our analysis. Denmark, Italy and Spain were excluded because of lack of reliable data over most of the sample.

[3] Center for Research in Security Prices, University of Chicago.

Table 1. Summary statistics for inflation rates, 1982:1–1991:9

Period	Series	Mean	Std. dev.	ρ (US)	ρ (GE)
U.S.					
1982:1–1991:9	1 year	3.760	1.064	1.000	0.396
1982:1–1990:9	2-1 year	−0.012	0.734	1.000	0.338
1982:1–1989:9	3-1 year	0.101	0.855	1.000	0.575
1982:1–1988:9	4-1 year	0.226	0.923	1.000	0.726
1982:1–1987:9	5-1 year	0.417	1.011	1.000	0.691
Germany					
1982:1–1991:9	1 year	2.150	1.356	0.396	1.000
1982:1–1990:9	2-1 year	0.042	0.634	0.338	1.000
1982:1–1989:9	3-1 year	0.064	0.960	0.575	1.000
1982:1–1988:9	4-1 year	0.062	1.284	0.726	1.000
1982:1–1987:9	5-1 year	0.128	1.531	0.691	1.000
Netherlands					
1982:1–1991:9	1 year	1.899	1.565	0.206	0.871
1982:1–1990:9	2-1 year	0.075	0.680	0.183	0.648
1982:1–1989:9	3-1 year	0.086	1.124	0.403	0.821
1982:1–1988:9	4-1 year	−0.001	1.481	0.476	0.851
1982:1–1987:9	5-1 year	−0.155	1.706	0.571	0.916
Belgium					
1982:1–1991:9	1 year	3.462	2.086	0.186	0.556
1982:1–1990:9	2-1 year	−0.280	0.780	0.598	0.801
1982:1–1989:9	3-1 year	−0.495	1.266	0.601	0.909
1982:1–1988:9	4-1 year	−0.737	1.778	0.658	0.966
1982:1–1987:9	5-1 year	−1.017	2.090	0.676	0.979
France					
1982:1–1991:9	1 year	4.364	2.216	0.014	0.338
1982:1–1990:9	2-1 year	−0.351	0.667	0.512	0.748
1982:1–1989:9	3-1 year	−0.692	1.070	0.591	0.866
1982:1–1988:9	4-1 year	−1.046	1.473	0.675	0.930
1982:1–1987:9	5-1 year	−1.381	1.737	0.685	0.941
Switzerland					
1982:1–1991:9	1 year	3.118	1.644	0.568	0.789
1982:1–1990:9	2-1 year	0.085	0.803	0.580	0.555
1982:1–1989:9	3-1 year	0.317	1.113	0.567	0.718
1982:1–1988:9	4-1 year	0.476	1.409	0.509	0.881
1982:1–1987:9	5-1 year	0.444	1.563	0.667	0.927
Japan					
1982:1–1991:9	1 year	1.794	1.107	0.581	0.757
1982:1–1990:9	2-1 year	0.017	0.614	0.575	0.512
1982:1–1989:9	3-1 year	0.127	0.871	0.555	0.731
1982:1–1988:9	4-1 year	0.193	1.129	0.592	0.840
1982:1–1987:9	5-1 year	0.206	1.247	0.617	0.895

Table 2. Summary statistics for interest rates, 1982:1–1991:9

Period	Series	Mean	Std. dev.	ρ (US)	ρ (GE)
U.S.					
1982:1–1991:9	1 year	8.402	1.745	1.000	0.377
1982:1–1990:9	2-1 year	0.331	0.324	1.000	0.583
1982:1–1989:9	3-1 year	0.645	0.545	1.000	0.510
1982:1–1988:9	4-1 year	0.990	0.541	1.000	0.177
1982:1–1987:9	5-1 year	1.166	0.565	1.000	0.205
Germany					
1982:1–1991:9	1 year	6.332	1.759	0.377	1.000
1982:1–1990:9	2-1 year	0.400	0.258	0.583	1.000
1982:1–1989:9	3-1 year	0.802	0.424	0.510	1.000
1982:1–1988:9	4-1 year	1.210	0.445	0.177	1.000
1982:1–1987:9	5-1 year	1.402	0.520	0.205	1.000
Netherlands					
1982:1–1991:9	1 year	6.862	1.512	0.307	0.937
1982:1–1990:9	2-1 year	0.438	0.363	0.676	0.511
1982:1–1989:9	3-1 year	0.843	0.576	0.612	0.380
1982:1–1988:9	4-1 year	1.160	0.663	0.487	0.146
1982:1–1987:9	5-1 year	1.237	0.687	0.453	0.071
Belgium					
1982:1–1991:9	1 year	9.351	2.043	0.705	0.585
1982:1–1990:9	2-1 year	0.400	0.361	0.040	0.187
1982:1–1989:9	3-1 year	0.737	0.602	−0.028	0.088
1982:1–1988:9	4-1 year	0.958	0.774	−0.109	−0.014
1982:1–1987:9	5-1 year	0.946	0.882	0.026	−0.021
France					
1982:1–1991:9	1 year	10.206	2.206	0.680	0.439
1982:1–1990:9	2-1 year	0.334	0.442	0.472	0.306
1982:1–1989:9	3-1 year	0.658	0.723	0.362	0.153
1982:1–1988:9	4-1 year	0.919	0.880	0.060	−0.139
1982:1–1987:9	5-1 year	0.977	1.041	−0.082	−0.231
Switzerland					
1982:1–1991:9	1 year	4.818	1.322	−0.097	0.751
1982:1–1990:9	2-1 year	0.003	0.167	0.223	0.285
1982:1–1989:9	3-1 year	0.073	0.249	0.019	0.025
1982:1–1988:9	4-1 year	0.146	0.333	−0.272	−0.250
1982:1–1987:9	5-1 year	0.138	0.375	−0.255	−0.306
Japan					
1982:1–1991:9	1 year	5.915	1.316	0.468	0.772
1982:1–1990:9	2-1 year	0.138	0.156	0.679	0.706
1982:1–1989:9	3-1 year	0.318	0.233	0.616	0.597
1982:1–1988:9	4-1 year	0.507	0.269	0.369	0.333
1982:1–1987:9	5-1 year	0.628	0.307	0.227	0.265

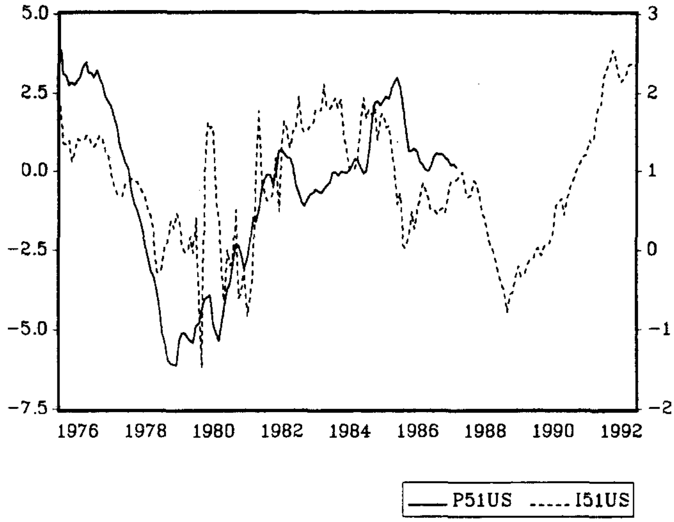

Fig. 1. (5-1)-year interest (dotted lines) and inflation (drawn lines) spreads: United States

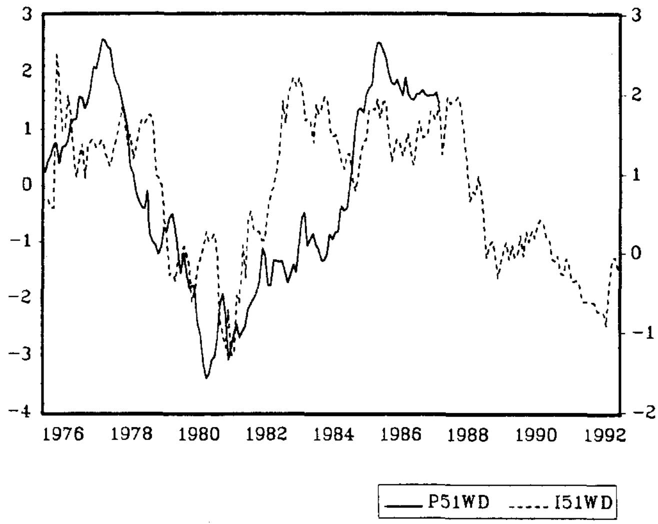

Fig. 2. (5-1)-year interest (dotted lines) and inflation (drawn lines) spreads: Germany

mid seventies followed by an upward movement in the late seventies and early eighties. Afterwards, co-movements appear less strong. From the graphs it is also clear that both interest rate and inflation spreads are highly persistent.[4]

It should be noted that – even though the amount of observations is rather limited due to the short period over which data are available –, the period 1982–1991 has been a period which a priori appears to be more in accordance with the assumptions underlying the JM-framework than the longer period starting in the seventies. On average, real interest rates have been high and positive in the eighties. For samples covering both the seventies and eighties, the jump from low (negative) real rates in the seventies to high (positive) real rates in the eighties often dominates empirical results.

4 Empirical Results

In this section, we present our empirical results. In paragraph 4.1, we follow JM (1991) in regressing inflation spreads for different maturities on corresponding interest rate spreads. The evidence, in general, does not support JM's conclusions. Results – in terms of sign and magnitude of the estimated β-coefficients – appear to depend on the sample period and country studied. In paragraph 4.2, we provide an explanation for this phenomenon. The dependency of the β-coefficient on the variability of the real term structure of interest rates is laid out. A relatively high variability of the term structure of real interest rates and a negative covariability between the term structures of inflation and real interest rates respectively lead to a downward bias of β-estimates. β even may become negative.

In paragraph 4.3 we attempt to determine whether the variability of the term structure of real interest rates comes from the short or long end of the maturity specter. For this purpose, we empirically investigate the link between inflation and interest rate levels (as opposed to spread) for different countries and maturities. The hypothesis that high nominal interest rates predict high future inflation cannot be rejected for a one year horizon in the case of the European countries and, marginally, Japan. For long horizons, high interest rates often predict low inflation. These results suggest that the variability of the term structure of real interest rates is mainly due to ex post variability of long-term real interest rates, which in turn is caused by unpredictable long-term developments in inflation. For the U.S., recursive estimation shows that the link between the short-term levels of inflation and interest rates varies as well.

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[4] Unreported unit root results generally are unable to reject nonstationarity of interest rate and inflation spreads (and levels). Caution is warranted though, both due to the short sample period and the complication of overlapping observations.

Paragraph 4.4 contains evidence that changes in current inflation have a similar impact on short- and long-term interest rates. This is supportive of our arguments in paragraph 4.3 that long-term inflation developments are hard to predict, so that ex post long-term real interest rates are almost perfectly negatively correlated with swings in inflation, leading to biased β-coefficients in JM-type regressions.

In paragraph 4.5, we introduce the conduct of monetary policy as a potential source of our findings. If monetary authorities react to either high inflation, high interest rates or a steep yield curve with contractionary monetary policy, a negative relation between the yield curve and the future term structure of inflation can arise. Necessary – and quite realistic – conditions in this respect are a considerable amount of discretion on the side of the monetary authorities, so that the public is unable to make out the precise timing and determinedness of policy changes, and the existence of long and variable lags between the implementation of monetary policy and its effect on future inflation.

4.1 A First Look: Jorion and Mishkin Revisited

To address the question whether the term structure of interest rates does indeed contain significant information about future inflation across a range of countries that differ widely in size and monetary policy – especially exchange rate regime –, single equation regressions[5] of equation (1) have been performed for the sample period starting in January 1982. The results are in part A of table 3. As the overlapping data generate a moving average error term of order $(12m - 1)$ in the error term, where m is maturity in years, the standard errors reported in the empirical work have been generated through the so-called Generalized Methods of Moments (GMM) estimator with Newey-West adjustment, as described in Campbell and Clarida (1987) that ensures the variance-covariance matrix to be positive by imposing linearly declining weights on autocovariance matrices.

Overall, the results do not support the hypothesis that there is a one-to-one (or even a less than proportional but significantly positive) relation between the term structure of nominal interest rates and the term structure of ex post inflation rates. Moreover, for countries that are both in our sample and in JM (1991), strongly deviating results are found. Especially for the US, we find

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[5] Additionally, SUR-regressions for equal maturities across countries have been performed. Qualitative conclusions are similar to single equation results. Moreover, no GMM standard errors could be computed, so that results must be interpreted with great caution. For both reasons, we do not report these results.

Table 3. Information about the term structure of inflation in interest rate spreads, $\pi_t^m - \pi_t^1 = \alpha^m + \beta^m(i_t^m - i_t^1) + \varepsilon_t^m$

PART A Period	Series	α^m (s.e.)	β^m (s.e.)	$\bar{R}^2$	SE
U.S.					
1982:1–1990:9	2-1 year	−0.013 (0.284)	0.003 (0.514)	−0.010	0.737
1982:1–1989:9	3-1 year	0.197 (0.389)	−0.149 (0.428)	−0.002	0.855
1982:1–1988:9	4-1 year	0.723 (0.316)	−0.503 (0.308)	0.075	0.888
1982:1–1987:9	5-1 year	0.869 (0.110)	−0.388 (0.146)	0.033	0.994
Germany					
1982:1–1990:9	2-1 year	0.187 (0.183)	−0.361 (0.440)	0.012	0.631
1982:1–1989:9	3-1 year	0.428 (0.311)	−0.453 (0.278)	0.029	0.945
1982:1–1988:9	4-1 year	−0.283 (0.677)	0.285 (0.354)	−0.003	1.286
1982:1–1987:9	5-1 year	−0.947 (0.555)	0.767 (0.277)	0.054	1.489
Netherlands					
1982:1–1990:9	2-1 year	0.398 (0.216)	−0.737 (0.254)	0.146	0.629
1982:1–1989:9	3-1 year	0.930 (0.464)	−1.001 (0.241)	0.255	0.970
1982:1–1988:9	4-1 year	1.449 (0.642)	−1.249 (0.230)	0.304	1.235
1982:1–1987:9	5-1 year	1.754 (0.739)	−1.543 (0.304)	0.378	1.346
Belgium					
1982:1–1990:9	2-1 year	−0.417 (0.175)	0.341 (0.367)	0.015	0.774
1982:1–1989:9	3-1 year	−0.938 (0.546)	0.602 (0.230)	0.072	1.219
1982:1–1988:9	4-1 year	−1.503 (0.763)	0.800 (0.181)	0.110	1.677
1982:1–1987:9	5-1 year	−1.656 (0.782)	0.676 (0.131)	0.068	2.018
France					
1982:1–1990:9	2-1 year	−0.127 (0.235)	−0.671 (0.324)	0.190	0.600
1982:1–1989:9	3-1 year	−0.418 (0.588)	−0.415 (0.239)	0.069	1.032
1982:1–1988:9	4-1 year	−0.742 (0.871)	−0.331 (0.277)	0.027	1.453
1982:1–1987:9	5-1 year	−1.118 (0.741)	−0.269 (0.132)	0.011	1.727
Switzerland					
1982:1–1990:9	2-1 year	0.080 (0.186)	1.736 (0.504)	0.121	0.753
1982:1–1989:9	3-1 year	0.195 (0.390)	1.673 (0.361)	0.131	1.037
1982:1–1988:9	4-1 year	0.293 (0.529)	1.251 (0.390)	0.076	1.354
1982:1–1987:9	5-1 year	0.345 (0.491)	0.715 (0.414)	0.015	1.551
Japan					
1982:1–1990:9	2-1 year	−0.031 (0.164)	0.348 (0.755)	−0.002	0.614
1982:1–1989:9	3-1 year	0.441 (0.146)	−0.986 (0.776)	0.059	0.844
1982:1–1988:9	4-1 year	1.190 (0.186)	−1.968 (0.434)	0.194	1.014
1982:1–1987:9	5-1 year	1.462 (0.217)	−2.001 (0.244)	0.232	1.093

Table 3 Continued

PART B Period	Series (m)	α^m (s.e.)	β^m (s.e.)	$\bar{R}^2$	SE
U.S.					
1976:4–1990:9	2-1 year	−0.251 (0.380)	0.428 (0.476)	0.006	1.114
1976:4–1989:9	3-1 year	−0.669 (0.740)	0.829 (0.694)	0.044	1.772
1976:4–1988:9	4-1 year	−1.606 (1.059)	1.546 (0.949)	0.150	2.113
1976:4–1987:9	5-1 year	−2.343 (0.936)	1.950 (0.719)	0.290	2.182
Germany					
1976:4–1990:9	2-1 year	−0.055 (0.259)	0.216 (0.505)	0.002	0.708
1976:4–1989:9	3-1 year	−0.312 (0.418)	0.515 (0.392)	0.056	1.044
1976:4–1988:9	4-1 year	−0.887 (0.334)	0.921 (0.205)	0.200	1.267
1976:4–1987:9	5-1 year	−1.359 (0.243)	1.170 (0.168)	0.341	2.113

β^m to be negative and (marginally) significant for the longer maturity spreads, whereas JM (1991) find significant positive values. For Germany and Switzerland, we do find significant positive estimates of β^m, but both the size of these estimates and the explanatory power of the regressions is considerably lower than in JM (1991).

Apart from this, the explanatory power of the regressions is highest for Japan and the Netherlands with large negative β-coefficients. For France, β^m-coefficients are negative and close to significance too. Only for Belgium, the hypothesized positive coefficients are found, though with little explanatory power.

To facilitate a better comparison with JM (1991), part B of table 3 contains results for the US and Germany for the longer sample starting in April 1976. Unfortunately, data problems prevent us from doing the same for the other countries. The length of our sample now is close to that in JM (1991), though they start and end approximately three years earlier. For the 5-1 spread regression, for instance, they use 73:8–84:6, while we use 76:4–87:9.

Especially for the US, the results for the longer sample are much closer to JM (1991) than to our results in part A of table 3. Now, β^m-estimates and explanatory power rise with maturity and are in excess of unity for the longest (5-1) maturity spread. The German results in part B are in between those in part A and the results in JM (1991).

Summarizing, the combined evidence from table 3 and JM (1991) suggests that the choice of sample period may dramatically change the results. The tentative conclusion by JM (1991) that the longer the maturity, the more informative the term structure is about future inflation (perhaps suggesting long-run mean reversion in real interest rates), cannot be sustained. We conclude that as yet the overall evidence fails to support the hypothesis that a steep slope of the term structure of interest rates reliably predicts rising future inflation. To the extent that the term structure does contain significant information about future inflation, it is in the opposite direction: a steep slope of the yield curve predicts lower inflation in the future.

4.2 A Proximate Explanation: Real Interest Rate Dynamics

In the above analysis we have to reject the null hypothesis $\beta^m = 1$. In earlier work along these lines, the most frequently observed explanation for such rejection is the inappropriateness of the assumption that the real term structure of interest rates is a constant. It can be shown, see for instance Mishkin (1990a) and JM (1991) that the estimate of β^m is downward biased when the assumption of a constant real term structure is violated. β^m can be derived to be:

$$\beta^m = (\sigma^2 + \rho\sigma)/(1 + \sigma^2 + 2\rho\sigma) \tag{5}$$

where σ is the ratio of the standard deviation of the expected term structure of inflation $E_t(\pi_t^m - \pi_t^1)$, to the standard deviation of the slope of the real term structure $(r_t^m - r_t^1)$, and ρ is the correlation between $E_t(\pi_t^m - \pi_t^1)$ and $(r_t^m - r_t^1)$. Even when ρ equals zero, β^m will be below unity unless the standard deviation of the slope of the real term structure is relatively small. Large negative values of ρ in combination of small values of σ will result in negative β estimates. Figure 3 shows theoretical values of β as a function of σ when ρ equals -0.1, -0.5, and -0.9 respectively. Mishkin (1990a) and JM (1991) document empirical values of ρ between -0.5 and -0.9, in general, thus justifying the finding of negative βs.

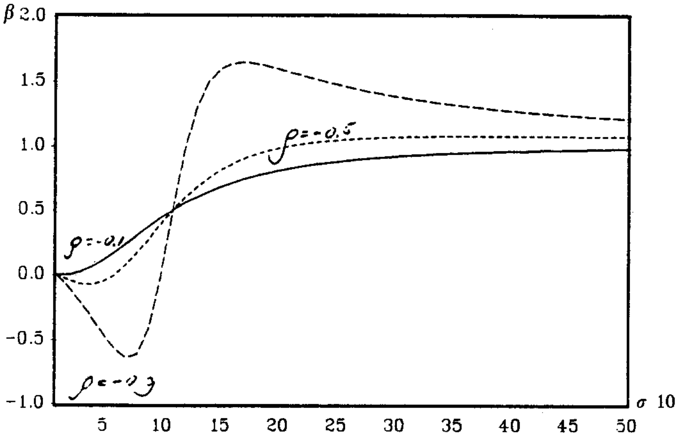

Fig. 3. Responsiveness of β in equation (1) to σ

To shed further light on this issue, Blough (1994) explicitly includes estimates of the expected real term structure of interest rates as an additional explanatory variable in (a variant of) equation (l) for the U.S. Under the null hypothesis its coefficient should be minus unity. Blough is unable to empirically reject this null hypothesis. Moreover, the coefficient of the nominal term structure is found to be insignificantly different from both zero and one and has no additional explanatory power. Blough, therefore, concludes that the simple correlation between the nominal term structure and future inflation is an artifact of the omission of the expected term structure of real rates from the equation.[6]

4.3 Levels Versus Spreads

In an attempt to further analyze the underlying factors of the observed variability of the real interest rate term structure, we turn to a level specification of the link between inflation and interest rates. From equations (2) and (3), the following regression equation can be derived:

$$\pi_t^m = \alpha^m + \beta^m(i_t^m) + v_t^m \qquad (6)$$

Under the assumption that expectations are rational and the level of real interest rates is constant, the null hypothesis is that β^m equals one. Part A of table 4 contains the results for each country over the sample period starting in 1982, for maturities m of one to five years. Part B presents similar results for the U.S. and Germany over the longer sample. Again, standard errors are corrected for moving average error terms.

A few things stand out. First, for all European countries the coefficient β^1 is significantly positive and, except for Germany, never significantly different from unity. This provides evidence in support of the null hypothesis to some extent. For Japan, the explanatory power of the regression is low and the β^1-estimate only marginally different from zero, while for the US the one year nominal interest rate appears to have no forecasting power at all for the one year future inflation. This holds both for the short and the long sample period.

Second, the longer the maturity (m), the stronger the null hypothesis can be rejected. This holds for all countries. With longer maturity, the estimates of β^m decline towards zero. Correspondingly, the explanatory power of the regressions declines. In some cases, the estimated β^m coefficients even become significantly negative, resulting in higher explanatory power.

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[6] Blough, moreover, argues that the nominal term structure can only be useful in forecasting the term structure of inflation if it is able to forecast changes in interest rates. He refers to Shiller (1990) for an overview of – generally disappointing – empirical work and concludes that this is sufficient to understand the failure of the term structure of interest rates to predict future inflation.

Table 4. Information about future inflation in the interest rate level, $\pi_t^m = \alpha^m + \beta^m(i_t^m) + \varepsilon_t^m$

PART A Period	Series (m)	α^m (s.e.)	β^m (s.e.)	$\bar{R}^2$	SE
U.S.					
1982:1–1991:9	1 year	4.184 (0.771)	−0.050 (0.070)	−0.002	1.065
1982:1–1990:9	2 year	5.428 (0.741)	−0.178 (0.078)	0.172	0.731
1982:1–1989:9	3 year	5.757 (0.517)	−0.208 (0.049)	0.418	0.521
1982:1–1988:9	4 year	5.893 (0.385)	−0.220 (0.034)	0.711	0.317
1982:1–1987:9	5 year	5.426 (0.108)	−0.165 (0.008)	0.797	0.188
Germany					
1982:1–1991:9	1 year	−1.324 (1.144)	0.549 (0.143)	0.502	0.957
1982:1–1990:9	2 year	−0.421 (1.235)	0.373 (0.153)	0.228	1.013
1982:1–1989:9	3 year	1.753 (0.862)	−0.014 (0.103)	−0.011	0.990
1982:1–1988:9	4 year	3.293 (1.190)	−0.230 (0.163)	0.148	0.717
1982:1–1987:9	5 year	3.480 (0.934)	−0.252 (0.124)	0.331	0.449
Netherlands					
1982:1–1991:9	1 year	−2.952 (1.344)	0.707 (0.163)	0.462	1.148
1982:1–1990:9	2 year	−2.641 (1.269)	0.619 (0.141)	0.370	1.095
1982:1–1989:9	3 year	0.008 (1.220)	0.211 (0.128)	0.053	1.098
1982:1–1988:9	4 year	1.394 (1.504)	−0.007 (0.188)	−0.013	0.839
1982:1–1987:9	5 year	1.808 (1.048)	−0.072 (0.129)	0.018	0.544
Belgium					
1982:1–1991:9	1 year	−4.255 (1.432)	0.825 (0.134)	0.651	1.233
1982:1–1990:9	2 year	−2.740 (1.966)	0.616 (0.198)	0.479	1.284
1982:1–1989:9	3 year	−1.027 (1.848)	0.409 (0.193)	0.329	1.194
1982:1–1988:9	4 year	0.172 (1.578)	0.258 (0.157)	0.252	0.899
1982:1–1987:9	5 year	0.135 (1.081)	0.242 (0.102)	0.388	0.593
France					
1982:1–1991:9	1 year	−4.542 (0.867)	0.873 (0.082)	0.753	1.102
1982:1–1990:9	2 year	−3.452 (0.919)	0.718 (0.089)	0.821	0.790
1982:1–1989:9	3 year	−1.822 (0.925)	0.530 (0.084)	0.799	0.682
1982:1–1988:9	4 year	−0.708 (0.716)	0.396 (0.061)	0.823	0.482
1982:1–1987:9	5 year	−0.141 (0.373)	0.329 (0.031)	0.877	0.335
Switzerland					
1982:1–1991:9	1 year	−1.098 (0.783)	0.875 (0.132)	0.490	1.174
1982:1–1990:9	2 year	−0.170 (0.885)	0.704 (0.148)	0.240	1.274
1982:1–1989:9	3 year	3.273 (1.922)	−0.083 (0.532)	−0.010	1.263
1982:1–1988:9	4 year	8.332 (3.590)	−1.311 (0.793)	0.323	0.810
1982:1–1987:9	5 year	5.340 (2.759)	−0.608 (0.582)	0.090	0.694
Japan					
1982:1–1991:9	1 year	0.134 (1.189)	0.281 (0.174)	0.104	1.048
1982:1–1990:9	2 year	1.716 (1.077)	0.009 (0.134)	−0.010	0.984
1982:1–1989:9	3 year	3.037 (0.899)	−0.230 (0.142)	0.136	0.773
1982:1–1988:9	4 year	3.337 (0.636)	−0.286 (0.104)	0.408	0.479
1982:1–1987:9	5 year	3.177 (0.263)	−0.256 (0.041)	0.699	0.214

Table 4 Continued

PART B Period	Series (m)	α^m (s.e.)	β^m (s.e.)	$\bar{R}^2$	SE
U.S.					
1976:4–1991:9	1 year	3.875 (2.017)	0.208 (0.249)	0.018	3.049
1976:4–1990:9	2 year	7.430 (2.539)	−0.184 (0.238)	0.016	2.906
1976:4–1989:9	3 year	9.564 (3.286)	−0.405 (0.256)	0.117	2.618
1976:4–1988:9	4 year	10.672 (3.559)	−0.520 (0.260)	0.233	2.259
1976:4–1987:9	5 year	11.199 (3.498)	−1.580 (0.258)	0.341	1.861
Germany					
1976:4–1991:9	1 year	−0.360 (1.105)	0.513 (0.137)	0.355	1.411
1976:4–1990:9	2 year	0.564 (1.421)	0.356 (0.158)	0.150	1.529
1976:4–1989:9	3 year	2.021 (1.472)	0.135 (0.144)	0.014	1.581
1976:4–1988:9	4 year	3.368 (1.542)	−0.061 (0.119)	−0.003	1.513
1976:4–1987:9	5 year	4.595 (1.852)	−0.230 (0.151)	0.050	1.370

Overall, the information in table 4 suggests that short-term (one year) interest rates do contain valuable information about inflation over the corresponding period, but that longer-term interest rates do not. The failure of the term structure of nominal interest rates to predict the term structure of future inflation, as documented in table 3, therefore, mainly comes from the long end of the maturity specter. In other words, for most countries the assumption of a constant one-year real interest rate is not a too bad approximation over the period since 1982. For longer-term real interest rates, on the other hand, the assumption of constancy appears to be rejected. As a consequence, the term structure of real interest rates has not been constant either.

The case of the U.S. deserves special attention. Huizinga and Mishkin (1986) document significant changes in the relation between short-term interest rates and inflation in 1979 and 1982, using monthly data, while Antoncic (1986) provides corresponding evidence about persistent real interest changes in the period 1979–82. Barsky (1987) shows theoretically that any relation may be found between nominal interest rates and inflation empirically, depending on both the degree of persistence and forecastability of inflation.

Figure 4 shows the recursively estimated coefficient β^1 (plus and minus 2 standard deviations[7]) of equation (6) for the US, while figure 5 presents the recursive coefficient β^5 corresponding to the OLS regression for the 5-1 year US spread in table 3. Finally, figure 6 displays the time path of the US 1-year nominal interest rate and inflation.

The conclusion from figures 4 to 6 is quite straightforward. For the US, the relation between inflation and the one-year interest rate breaks down in early

[7] Here, the standard errors have not been adjusted for MA-terms induced by the overlapping data and, therefore. must be interpreted with care.

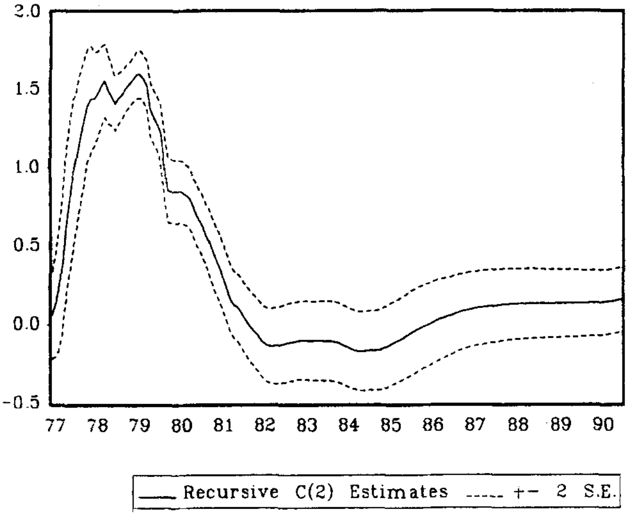

Fig. 4. Recursive β-coefficient equation (6), $m = 1$, United States

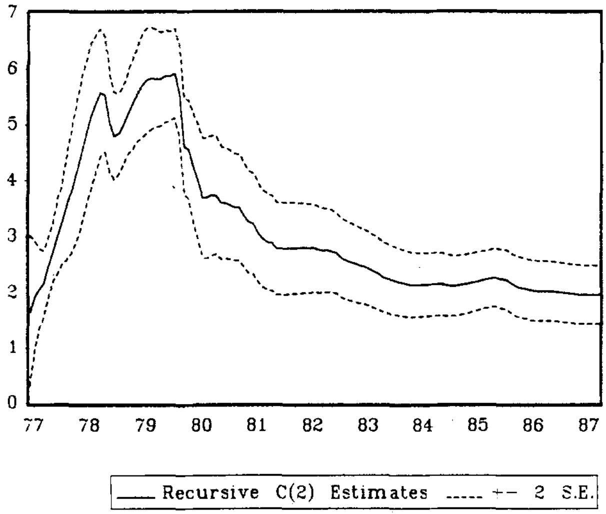

Fig. 5. Recursive β-coefficient equation (1), United States

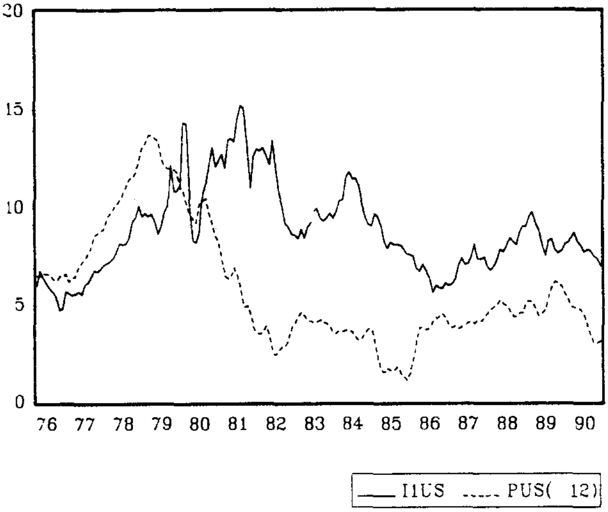

Fig. 6. 1-Year inflation and interest rate level, United States

1980, probably due to the change in monetary policy and the ensuing fast disinflation. At the same time, nominal interest rates remain high, resulting in a persistent rise in the ex post real interest rates in the period 1980–1982. For an explanation, Barro and Sala-i-Martin (1990) point to the economic boom as reflected in rising stock prices from the early eighties onward. In a comment, Lucas (1991) suggests the possibility of persistently biased inflation expectations. The spread equation collapses at approximately the same time, as shown in figure 5. Apparently, the shift in the level of the one-year real interest rate is at the heart of our rejection of the one-to-one relation between nominal interest rates and future inflation for the US.

Overall, the above discussion points to the structural weakness of predicting inflation (spreads) through interest rate (spreads). Over the sample(s) considered here, movements in either the short-term real interest rate level, the term structure of real interest rates or both have been too large to neglect.

4.4. *Are Nominal Interest Rates Driven by Inflation?*

The evidence in the previous paragraph shows long-term nominal interest rates have little explanatory power in explaining future inflation. One reason for this phenomenon may be that longer-term inflation forecasts are hard to make. As a corollary, it may be argued that short-term interest rates are driven more by changes in expected inflation than are long-term rates.[8] Here, we further analyze this issue in the following way. We regress the change in the interest rate on an asset with remaining maturity of m years on the last realized change in the annual inflation rate:

$$\Delta i_t^m = \gamma_0^m + \gamma_1^m \Delta \pi_{t-12} , \qquad m = 1, \ldots, 5 \tag{7}$$

The coefficient γ_1^m measures to what extent the new information about inflation[9] is reflected in forward-looking interest rates of different maturities. Table 5 contains the estimated values and associated standard deviations of γ_1^m across countries and maturities for the period 1982:1 to 1992:12. For Germany and the U.S. results for the period 1976:4–1992:12 are presented as well. For these latter two countries stability of the results over the longer period cannot be

Table 5. Sensitivity of interest rates to changes in inflation: γ_1^m
$\Delta i_t^m = \gamma_0^m + \gamma_1^m \Delta \pi_{t-12}$, $m = 1, \ldots, 5$

	m				
Period 1982:1–1992:12	1	2	3	4	5
U.S.	0.298	0.281	0.273	0.260	0.247
	(0.110)	(0.111)	(0.111)	(0.110)	(0.106)
Germany	0.264	0.229	0.202	0.183	0.176
	(0.097)	(0.086)	(0.080)	(0.078)	(0.077)
Netherlands	0.192	0.190	0.189	0.186	0.182
	(0.111)	(0.086)	(0.079)	(0.077)	(0.074)
Belgium	0.050	0.053	0.067	0.087	0.117
	(0.118)	(0.090)	(0.081)	(0.082)	(0.088)
France	0.349	0.256	0.208	0.202	0.242
	(0.204)	(0.142)	(0.109)	(0.100)	(0.125)
Switzerland	−0.020	−0.000	0.014	0.023	0.031
	(0.105)	(0.080)	(0.063)	(0.052)	(0.047)
Japan	0.101	0.105	0.102	0.095	0.088
	(0.064)	(0.058)	(0.056)	(0.054)	(0.052)
	m				
Period 1976:4–1992:12	1	2	3	4	5
U.S.	0.339	0.312	0.290	0.267	0.245
	(0.134)	(0.124)	(0.115)	(0.107)	(0.099)
Germany	0.239	0.197	0.162	0.135	0.118
	(0.091)	(0.079)	(0.075)	(0.074)	(0.074)

[8] According to Blough (1994), a necessary condition for the term structure of interest rates to be useful as a predictor of the term structure of future inflation, is that changes in interest rates are mainly driven by changes in inflation.

[9] Alternatively, the residual from an autoregression of the change in inflation on its own past was used as an explanatory variable in equation (7). Results were only marginally different, indicating that it is the innovation in inflation that is relevant in revising inflationary expectations and interest rates.

rejected. The most important conclusion to be drawn from the table is that for no country the equality of coefficient γ_1^m across maturities can be rejected. That is, innovations in current inflation lead to interest rate revisions of equal magnitude across maturities. This suggests that financial markets have a hard time in predicting a correct term structure of expected inflation. It corresponds to the stylized fact that (changes in) interest rates of different maturities have a very high correlation. In combination with the evidence in table 4, we may conclude that currently observed changes in inflation may be helpful in predicting next year's inflation, but contribute little to the prediction of inflation over much longer periods. Consequently, movements in the term structure of nominal interest rates mainly reflect movements in the real term structure of interest rates.

4.5 *Monetary Policy Considered*

Looking at longer horizons of four and five years, both tables 3 and 4 suggest a perverse negative relation between interest rates and future inflation for a number of countries. That is, as opposed to the theory developed in section 2, high long-term interest rates predict low future inflation, and steep positive term structures predict declining inflation over time.

A potential explanation for this phenomenon is the interaction of inflation and nominal interest rates through the conduct of monetary policy. Suppose the monetary authorities have a reaction function of the form:

$$\Delta^2 m_t = \delta_1 i_t^1 + \delta_2 (i_t^m - i_t^1) \ , \qquad \delta_1, \delta_2 < 0 \tag{8}$$

where m is the logarithmic level of the money supply. Equation (8) indicates that both a steep yield curve or a high nominal interest rate level may be indications of high current and future inflation and, therefore, may lead the monetary authorities to decelerate the growth of the money stock (Δm_t). According to the quantity theory, this would lead to lower inflationary expectations in the following way:

$$\begin{aligned} E_t(\pi_t^k - \pi_t^1) &= 100 * E_t[(p_{t+k} - p_t)/k - (p_{t+1} - p_t)] \\ &= 100 * E_t[(m_{t+k} - m_t)/k - (m_{t+1} - m_t)] \\ &\quad + 100 * E_t[(v_{t+k} - v_t)/k - (v_{t+1} - v_t)] \\ &\quad - 100 * E_t[(y_{t+k} - y_t)/k - (y_{t+1} - y_t)] \end{aligned} \tag{9}$$

where p, v, and y are the logarithmic levels op prices, velocity and real income respectively. The combination of equations (8) and (9) imply that a steep yield curve and/or high current interest rate levels will trigger contractionary monetary policy, leading in the end to lower future expected inflation.

The main objection against the above argument is that under rational expectations the reaction function of the monetary authorities and its consequences for future inflation should already be incorporated into the market's inflation expectations, and, thus, in nominal interest rates as well: if a steep yield curve leads to lower future inflation, the slope of the yield curve now should decline to capture that information. Moreover, the existence of a liquidity effect could raise short-term nominal interest rates and, thereby also contribute to a lower slope of the yield curve.

Counter arguments do exist, however. First, the liquidity effect is generally not found to be particularly significant, especially not for interest rates on assets with a maturity of one year or more. If a liquidity effect exists, it is generally thought to be relevant for weekly or monthly interest rates. Even then, it is weak and time dependent[10]. Second, a consensus has emerged in the literature that monetary policy affects inflation with long and variable lags in the order of two to three years. Moreover, the income velocity of money has been shown to exhibit nonstationary behavior, so that its future time path provides only limited information for inflation predictions over longer horizons, see Goldfeld and Sichel (1990). Finally, the determinedness with which the monetary authorities stick to their reaction function may vary through time as a function of unpredictable variables related to the state of the economy.

In the following, we explicitly test for the information in the current interest rate level about the future term structure of inflation. Figures 7 and 8 are complements to figures 1 and 2 and graphically show the strong negative linkages between one-year interest rates and (5-1)-year inflation spreads for the United States and Germany over the period 1976–1992. Table 6 contains the results of regressing (m-1)-year inflation spreads on both the corresponding interest rate spread and the current one-year interest rate level. Again, standard errors are corrected for moving average patterns in the residuals due to overlapping data.

The information in part A of table 6 concerns the period starting in 1982. The results are quite similar across countries: both the term structure and the level term have negative signs and are often significant. The explanatory power is quite high, especially when compared with table 3, and increases with maturity. Exceptions are Switzerland where explanatory power and significance drop when moving from (4-1) to (5-1) slopes, and the United States, where there is no explanatory power at all across maturities. The longer sample results for the United States in part B suddenly show an increase in explanatory power and significance of the coefficients. This instability of the US results corresponds to the instabilities documented earlier both for the spread and level equations. For Germany, similar results are obtained as in the shorter sample.

[10] See Wasserfallen and Kursteiner (1994) for an overview of recent work on the liquidity effect.

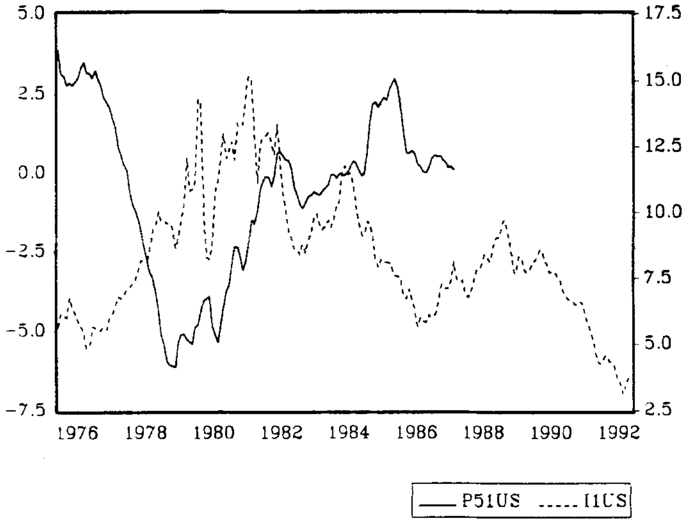

Fig. 7. (5-1)-year inflation spread (drawn) and one-year interest rate level (dotted): United States

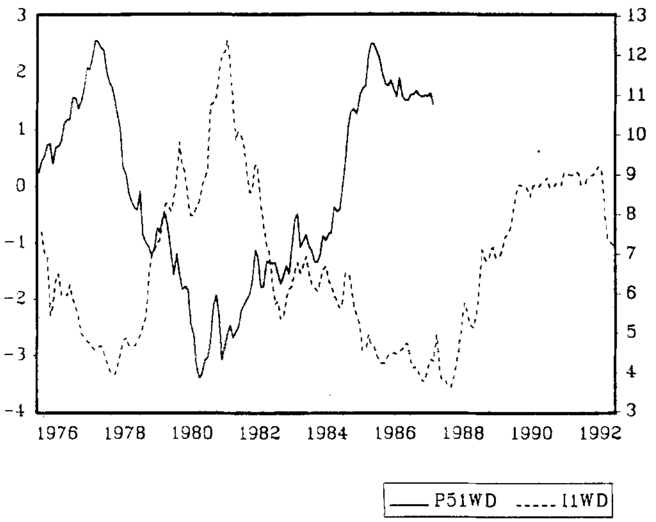

Fig. 8. (5-1)-year inflation spread (drawn) and one-year interest rate level (dotted): Germany

Table 6. Information about the term structure of inflation in the short-term interest rate level, $\pi_t^m - \pi_t^1 = \alpha^m + \beta^m(i_t^m - i_t^1) + \gamma^m i_t^1 + \varepsilon_t^m$

PART A Period	m	α^m (s.e.)	β^m (s.e.)	γ^m (s.e.)	$\bar{R}^2$	SE
U.S.						
82:1–90:9	2-1 yr	0.479 (0.924)	0.194 (0.475)	−0.064 (0.111)	−0.004	0.735
82:1–89:9	3-1 yr	1.111 (0.966)	0.048 (0.500)	−0.120 (0.126)	0.037	0.839
82:1–88:9	4-1 yr	1.347 (0.412)	−0.321 (0.406)	−0.092 (0.057)	0.090	0.881
82:1–87:9	5-1 yr	1.730 (0.422)	−0.251 (0.214)	−0.115 (0.038)	0.065	0.977
Germany						
82:1–90:9	2-1 yr	1.825 (0.361)	−1.145 (0.140)	−0.219 (0.066)	0.222	0.560
82:1–89:9	3-1 yr	4.328 (0.480)	−1.400 (0.188)	−0.546 (0.081)	0.572	0.628
82:1–88:9	4-1 yr	6.241 (0.585)	−1.285 (0.266)	−0.819 (0.064)	0.669	0.739
82:1–87:9	5-1 yr	8.408 (0.210)	−1.343 (0.163)	−1.092 (0.017)	0.760	0.751
Netherlands						
82:1–90:9	2-1 yr	1.068 (0.682)	−0.815 (0.361)	−0.096 (0.090)	0.176	0.618
82:1–89:9	3-1 yr	3.174 (1.379)	−1.005 (0.247)	−0.354 (0.019)	0.396	0.873
82:1–88:9	4-1 yr	5.207 (1.751)	−1.242 (0.139)	−0.598 (0.245)	0.566	0.977
82:1–87:9	5-1 yr	6.676 (1.613)	−1.588 (0.168)	−0.744 (0.207)	0.652	1.007
Belgium						
82:1–90:9	2-1 yr	2.078 (0.704)	−0.336 (0.345)	−0.237 (0.074)	0.342	0.633
82:1–89:9	3-1 yr	3.531 (0.841)	−0.150 (0.148)	−0.420 (0.083)	0.515	0.882
82:1–88:9	4-1 yr	5.670 (0.859)	−0.214 (0.132)	−0.649 (0.084)	0.652	1.049
82:1–87:9	5-1 yr	7.890 (0.659)	−0.264 (0.083)	−0.862 (0.062)	0.735	1.077
France						
82:1–90:9	2-1 yr	1.255 (0.360)	−0.652 (0.351)	−0.135 (0.029)	0.402	0.516
82:1–89:9	3-1 yr	2.716 (0.565)	−0.387 (0.191)	−0.305 (0.046)	0.553	0.715
82:1–88:9	4-1 yr	4.312 (0.365)	−0.384 (0.207)	−0.472 (0.024)	0.678	0.835
82:1–87:9	5-1 yr	5.160 (0.546)	−0.279 (0.129)	−0.571 (0.038)	0.700	0.952
Switzerland						
82:1–90:9	2-1 yr	2.362 (0.649)	−1.172 (0.707)	−0.497 (0.118)	0.275	0.684
82:1–89:9	3-1 yr	4.771 (1.228)	−0.496 (0.703)	−1.042 (0.214)	0.314	0.922
82:1–88:9	4-1 yr	7.242 (1.807)	−0.846 (0.561)	−1.611 (0.349)	0.302	1.177
82:1–87:9	5-1 yr	6.041 (1.435)	−0.527 (0.395)	−1.299 (0.266)	0.105	1.479
Japan						
82:1–90:9	2-1 yr	1.356 (0.399)	0.357 (0.481)	−0.239 (0.073)	0.264	0.526
82:1–89:9	3-1 yr	2.333 (0.702)	−0.457 (0.797)	−0.365 (0.136)	0.333	0.711
82:1–88:9	4-1 yr	3.982 (0.507)	−1.603 (0.393)	−0.517 (0.091)	0.565	0.745
82:1–87:9	5-1 yr	5.672 (0.258)	−1.956 (0.215)	−0.702 (0.043)	0.721	0.659

Table 6 Continued

PART B Period	m	α^m (s.e.)	β^m (s.e.)	γ^m (s.e.)	$\bar{R}^2$	SE
U.S.						
76:4–90:9	2-1 yr	2.266 (0.631)	0.786 (0.338)	−0.298 (0.065)	0.367	0.889
76:4–89:9	3-1 yr	3.617 (1.192)	0.864 (0.272)	−0.484 (0.138)	0.439	1.358
76:4–88:9	4-1 yr	3.288 (1.443)	1.199 (0.450)	−0.518 (0.164)	0.445	1.706
76:4–87:9	5-1 yr	1.861 (1.445)	1.398 (0.363)	−0.413 (0.135)	0.423	1.967
Germany						
76:4–90:9	2-1 yr	2.497 (0.419)	−1.379 (0.254)	−0.312 (0.050)	0.386	0.555
76:4–89:9	3-1 yr	4.630 (0.491)	−1.246 (0.201)	−0.603 (0.057)	0.639	0.646
76:4–88:9	4-1 yr	5.881 (0.539)	−1.021 (0.193)	−0.791 (0.063)	0.675	0.807
76:4–87:9	5-1 yr	6.695 (0.754)	−0.796 (0.222)	−0.932 (0.077)	0.712	0.907

5 Summary and Conclusions

This paper investigates the information contents of the term structure of interest rates in the United States, Japan, Switzerland, Germany, France, Belgium and the Netherlands with respect to future inflation using a new data set of yield to durations for 1 to 5 year fixed duration bonds.

Following JM (1991), we regress the ex post term structure of inflation on the term structure of nominal interest rates for various maturities. Our results differ from theirs, though. We conclude that apparently the choice of sample period is crucial to the results.

Overall, our results suggest that there is more information about future inflation in the current level of interest rates than there is in the term structure. In both cases, though, the information is in the opposite direction from that predicted by the Fisher equation. High interest rates and a steep yield curve are followed by declining and low inflation.

On the basis of our additional experiments, we postulate that this perverse effect may be due to a combination of two factors. On the one hand, the expected inflation level incorporated in long interest rates is close to that incorporated in short-term interest rates, that is, financial markets have only limited power to forecast a true term structure of inflation. An alternative way to make the same point is to argue that short-term – say quarterly – inflation is nonstationary and close to a random walk, for which there is ample evidence.

On the other hand, however, the behavior of monetary authorities that are concerned about both inflation and real activity may be characterized as follows. If inflation gets high, at some point the monetary authorities will start fighting inflation to bring it down, while if inflation is low monetary authorities may allow inflation to rise to higher levels. Neither the point at which the monetary authorities take action and with what determinedness, nor the lag with which this action will affect future inflation is known, however. Consequently, the relevance of this information for financial markets is limited.

Ex post, however, the two factors mentioned above will cause observed long-term inflation and observed long-term real interest rates to be strongly negatively correlated, leading to a negative relation between the term structure of nominal interest rates and the term structure of observed inflation.

For the United States, an additional problem is present. Over the longer sample, the level of the short-term real interest rate is observed to change considerably too, thereby disturbing the relation between the level of nominal interest and inflation even in the shorter run. Obviously, this affects term structure analyses as well. Again, monetary policy may be at the heart of the shift in the real interest rate level, see Lucas (1990), and Huizinga and Mishkin (1986).

Overall, our results strongly suggest that the term structure of interest rates is an inappropriate indicator of future inflationary developments regardless of the precise horizon.

Appendix Data Construction

Yield to duration time series for (artificial) constant duration bonds have been calculated as follows:

1. For each country k and each period t a cross section regression was performed on the $J(k, t)$ available bond quotations[11]:

$$i_j = c_0 + c_1 D_j + c_2 D_j^2 + c_3(P_j - 100) + u_j \qquad j = 1, J \tag{A1}$$

where i_j, D_j and P_j are the yield to duration, duration and price of bond j respectively. The quadratic polynomial in D_j as an explanation for i_j is reminiscent of Nelson and Siegel (1987) and Jorion and Mishkin (1991), although the functional specification is somewhat different. The inclusion of the term $(P_j - 100)$ takes account of possible differences in tax treatment of coupon payments and capital gains respectively.[12]

2. A sensitivity check on the above regressions was performed to avoid a too large influence of outliers on the results. If necessary, individual quotations were removed from the sample and the corresponding regressions were re-estimated.

[11] Bonds containing option-like elements (like callable bonds) have been removed from the sample.

[12] Using only those bonds with prices close to par, would have resulted in too little observations for many countries and periods.

As especially the results for the European countries were unduly influenced by quotations on bonds with a very short duration, the sample was confined to those bonds which had a remaining duration of at least 1 year for all countries.
3. Constant-duration time series of yield-to-durations for 1, 2, 3, 4, and 5 year par bonds were constructed using the final coefficient estimates of equation (A1). Note that the part $c_3(P_j - 100)$ automatically falls out now. These series are used in the remaining part of this paper.

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