What do asset price movements in Germany tell monetary policy makers?

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Introduction

Asset prices can play a twofold role in monetary policy. First, they may be seen as important elements in the chain along which monetary policy stimuli are transmitted to the real economy. From this perspective, asset price movements cause changes in aggregate demand or the price level through substitution, income and wealth effects. If these structural relationships were stable and could be estimated reliably, asset prices could be used as indicators of, or even target variables for, monetary policy. Second, they may be seen as predictors of the future course of the economy, independently of their active role in the transmission process. This view does not depend on the causal influence of asset prices on the macroeconomic variables to be predicted. Instead, it takes due account of the fact that the price of rationally valued assets should reflect the expected path of the asset's income components and the equilibrium returns used for discounting the future stream of income. If these expectations were influenced by the anticipated development of certain macroeconomic fundamental factors, and if, furthermore, market expectations were not systematically biased, asset prices could be used by the central bank as predictors of real activity and inflation.

The monetary policy implications of both roles depend crucially on the informational efficiency of asset markets. Market inefficiencies would cause asset prices to deviate from their fundamental values, distorting their informational content and their indicator quality. Furthermore, if asset prices play an important role in the transmission process, mispricing may adversely affect economic activity and price stability. The main body of this paper is devoted to assessing the predictive power or the informational content, respectively, of dividend yields and the term structure spread to draw some preliminary conclusions about the efficiency of the stock and government bond markets in Germany.

The theoretical framework is provided by the rational valuation approach. Applied to the bond market and the stock market, this approach leads to the expectations hypothesis and the dividend discount model, respectively, both on the assumption of rational expectations. The informational content is judged by metrics from univariate regression techniques using short and long-horizon measures for future inflation, stock returns, dividend growth, and interest rate changes as dependent variables and the spread or the dividend yield as regressors. The paper closes with some implications of the results for monetary policy.

1. Pricing stocks and bonds with the rational valuation approach

The value of financial assets generally depends on the future stream of payments the holder is entitled to receive. Hence, it is economically reasonable to calculate an asset's fundamental value as the discounted present value of the expected stream of income. The discount rate used can be interpreted as the required (expected) rate of return which attracts investors to hold the asset in their portfolios. In an informationally efficient market, an asset's actual market price should then equal its fundamental value as calculated by all or the marginal investor depending on whether expectations are assumed to be homogeneous or not. Thus, testing the informational efficiency of asset prices requires

* The views expressed in the paper are those of the authors and not necessarily those of the Deutsche Bundesbank.

an assumption about the behaviour of equilibrium returns and a hypothesis as to how market agents form expectations.

1.1 Stock pricing

Applied to the stock market, this general valuation approach is the dividend discount model. We can derive it starting with the approximation formula for the continuously compounded one-period return h_{t+1} on stocks as suggested by Campbell and Shiller:[1]

$$h_{t+1} = k + \rho p_{t+1} + (1-\rho)d_{t+1} - p_t \tag{1.1}$$

with h_{t+1} = approximate continuously compounded (or logarithmic) one-period return on stocks over the holding period t+1; p_t = log of stock price measured at the end of period t; d_{t+1} = log of dividend paid out before the end of period t+1; $\rho \equiv 1/(1+\exp(\overline{d-p}))$, where $\overline{d-p}$ = average log dividend yield; and $k \equiv -\log(\rho) - (1-\rho)\log(1/\rho - 1)$.

Equation (1.1) provides a loglinear relation between stock prices, returns and dividends, which is more convenient for calculation purposes if equilibrium returns are allowed to be time-varying. It is a first-order linear difference equation in the stock price. Solving forward and imposing the terminal condition $\lim_{j\to\infty} \rho^j p_{t+j} = 0$, yields:[2]

$$p_t = \frac{k}{1-\rho} + \sum_{j=0}^{\infty} \rho^j \left[(1-\rho)d_{t+1+j} - h_{t+1+j}\right] \tag{1.2}$$

Equation (1.2) is a mere identity, which says that today's stock price is high if future dividends are high and/or future returns are low. By applying the conditional expectations operator $E_t x_{t+1} = E[x_{t+1} | \Omega_t]$ (with Ω_t the market-wide information set available at the end of period t) and the law of iterated expectations, equation (1.2) can be changed to an ex ante relationship: [3]

$$p_t = \frac{k}{1-\rho} + \sum_{j=0}^{\infty} \rho^j \left[(1-\rho)E_t d_{t+1+j} - E_t h_{t+1+j}\right] \tag{1.3}$$

Further assuming homogeneous expectations on the part of all market participants and instantaneous market clearing, the log stock price always equals its single fundamental value, which in turn is the specifically weighted, infinite sum of expected log dividends discounted by principally time-varying expected equilibrium returns. Thus, equation (1.3) just represents the dividend discount model. Combined with rational expectations, it is also a valid representation of the "rational valuation formula" (RVF) for stocks.[4]

The loglinear approximation framework has two important advantages: first, it allows a linear and thus rather simple, analysis of the stock price behaviour. Second, it conforms with the empirically plausible assumption that dividends and stock returns follow loglinear stochastic

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1 See Campbell et al. (1997), pp. 260-2.

2 This terminal condition rules out rational bubbles that would cause the log stock price to grow exponentially forever at rate $1/\rho$ or faster (Campbell et al. (1997), pp. 262 f.).

3 In technical terms, the law of iterated expectations can be expressed as $E_t[E_{t+i-1}h_{t+i}] = E_t h_{t+i}$ which may be interpreted as a consistency condition under rational expectations.

4 See Cuthbertson (1996), who applies the RVF to various financial instruments (stocks, bonds, foreign exchange).

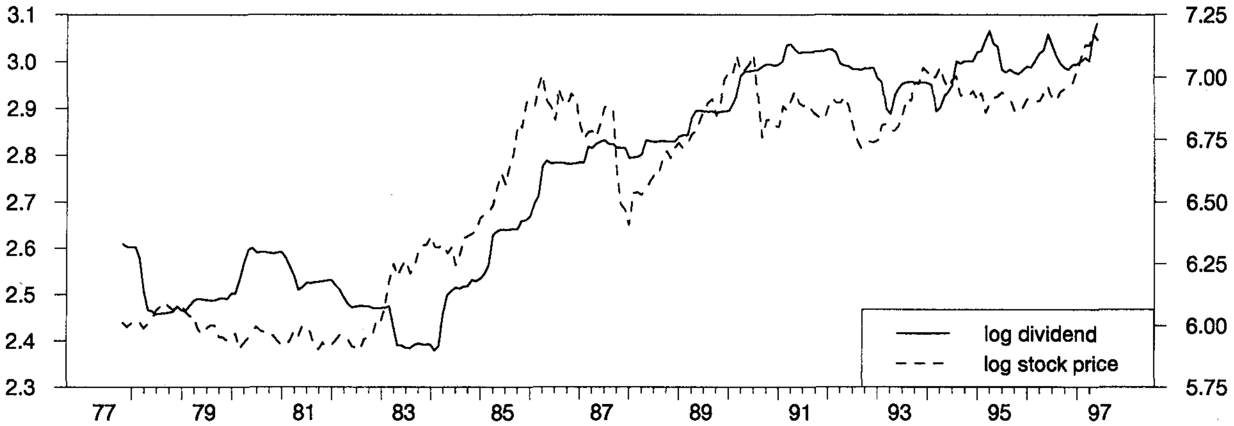
processes. For the empirical analysis it turns out to be advantageous to rearrange equation (1.3) such that the log dividend yield (or log dividend-price ratio) is singled out as the left-hand variable:

$$d_t - p_t = -\frac{k}{1-\rho} + \sum_{j=0}^{\infty} \rho^j (-E_t \Delta d_{t+1+j} + E_t h_{t+1+j}) \qquad (1.4)$$

Figure 1

Dividends (left-hand scale) and stock prices (right-hand scale)

Nominal values (DM) in logs; November 1977 to June 1997



The empirical evidence generally suggests that the logs of dividends and stock prices follow non-stationary I(1) processes (see Figure 1). Dividend changes (the first differences) are therefore I(0) or stationary, as are the one-period stock returns. Thus, the right-hand side of equation (1.4) – a weighted sum of (expected) dividend changes and stock returns – should also be stationary. Dividends and stock prices must then cointegrate so that the (log) dividend yield can form a stationary process, too. If these stationarity assumptions were true, equation (1.4) would only consist of stationary variables and could be used for regression analysis without any further data transformations or use of non-standard distribution theory.

1.2 Bond pricing

Now we turn to the RVF for bonds. Since our analysis of the German bond market is based upon estimated spot rates (zero coupon rates), we start with the definition of the one-period return on a pure discount bond:

$$h_{t+1}^{(n)} = \ln(1 + H_{t+1}^{(n)}) = \ln P_{t+1}^{(n-1)} - \ln P_t^{(n)} \qquad (1.5)$$

with $h_{t+1}^{(n)}$ = continuously compounded (or log) one-period return on a pure discount bond over the holding period $t+1$; $P_t^{(n)}$ = price of an n-period pure discount bond measured at the end of period t.

To cast equation (1.5) in terms of continuously compounded spot yields $z_t^{(n)}$, we substitute out bond prices by using the relation $\ln P_t^{(n)} = \ln M - n\ln(1 + Z_t^{(n)}) = \ln M - nz_t^{(n)}$. M is the redemption price of the n-period bond and $Z_t^{(n)}$ is the simple spot rate. Equation (1.5) then becomes:

$$h_{t+1}^{(n)} = nz_t^{(n)} - (n-1)z_{t+1}^{(n-1)} \qquad (1.6)$$

The different theories of the term structure of interest rates are now based on different assumptions about the required or expected one-period return that attracts investors to hold an n-period bond over one period. We assume that investors require a rate of return which exceeds the one-period risk-free rate r_t by a term premium $T_t^{(n)}$:[5]

$$E_t h_{t+1}^{(n)} = E_t\left[nz_t^{(n)} - (n-1)z_{t+1}^{(n-1)}\right] = r_t + T_t^{(n)} \tag{1.7}$$

or $$nz_t^{(n)} = (n-1)E_t z_{t+1}^{(n-1)} + r_t + T_t^{(n)} \tag{1.8}$$

Now leading (1.8) one period, applying the law of iterated expectations and substituting the result into equation (1.8) gives:

$$nz_t^{(n)} = (n-2)E_t z_{t+2}^{(n-2)} + r_t + E_t r_{t+1} + T_t^{(n)} + E_t T_{t+1}^{(n-1)} \tag{1.9}$$

Further substituting and noting that $(n-j)E_t z_{t+j}^{(n-j)} = 0$ for $j = n$, we finally obtain a familiar term structure relationship which also represents the RVF for bonds:[6]

$$z_t^{(n)} = E_t\left[\frac{1}{n}\sum_{i=0}^{n-1} r_{t+i}\right] + E_t\left[\frac{1}{n}\sum_{i=0}^{n-1} T_{t+i}^{(n-i)}\right] = E_t\left[\frac{1}{n}\sum_{i=0}^{n-1} r_{t+i}\right] + E_t\phi_t^{(n)} \tag{1.10}$$

with $\phi_t^{(n)}$ = the average risk premium on the *n*-period bond until it matures. The *n*-period long-rate equals a weighted average of expected future short rates plus the expected average risk premium. But this equation is non-operational unless we assume a specific form of the term premium.[7] Different assumptions about the term premia also characterise the different term structure theories. For example, the pure expectations hypothesis (*PEH*) rests on the assumption of zero term premia for all maturities, while the expectations hypothesis (*EH*) only requires constant term premia which are the same for all maturities.[8]

Under empirically plausible assumptions about the time-series characteristics of interest rates, the following rearrangement of equation (1.10) leads to a stationary transformation, which is now widely used for regression purposes:[9]

$$S_t^{(n,1)} = z_t^{(n)} - r_t = \sum_{i=1}^{n-1}(1 - i/n)E_t\Delta r_{t+i} + E_t\phi_t^{(n)} \tag{1.11}$$

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5 The expected excess return may generally be called a risk premium. But since the yield data we use are for government bonds only which carry little or no default risk, the remaining risk of such bonds mainly arises from different *terms to maturity*. The expression "term premium" draws on this fact (see Cuthbertson (1996), p. 214).

6 The RVF for coupon-paying bonds is very similar to the formula for stocks. Uncertain dividend streams in the latter case are replaced by known coupon payments over a limited period of time, and, at maturity, the also known nominal value will be redeemed. This certain stream of (nominal) income has to be discounted using consecutive expected one-period returns required by the investors to hold the bond over its time to maturity, just as in the case of stocks. For pure discount bonds, only the redemption price has to be discounted to get the fundamental bond value and thus the RVF.

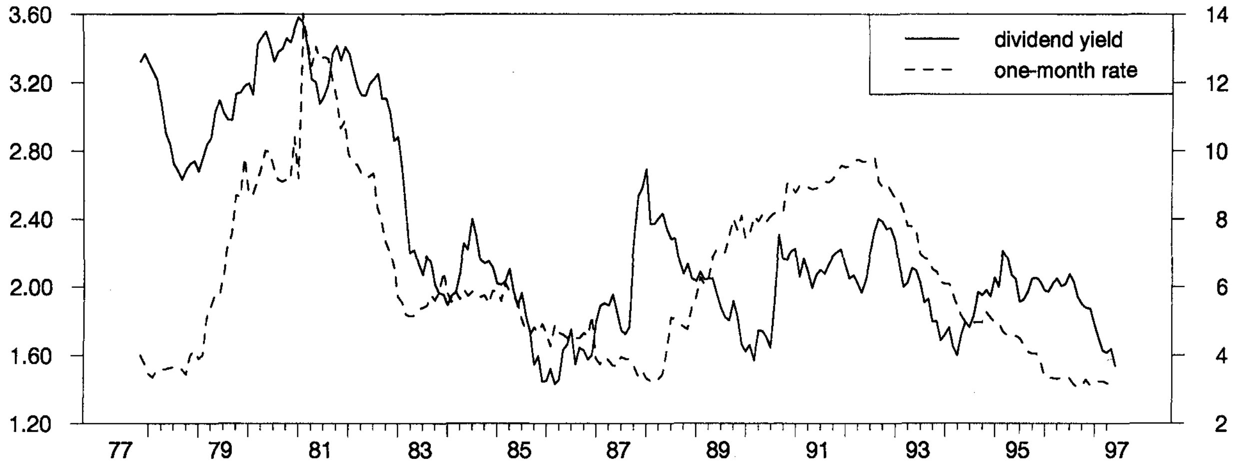
7 See Cuthbertson (1996), p. 225.

8 For a short survey of different term structure theories see, e.g., Cuthbertson (1996), pp. 218-23.

9 Although there are theoretically strong reasons for regarding interest rates as stationary variables, conventional integration tests most often suggest interest rates to be near-integrated variables whose time-series behavior may better be represented by non-stationary I(1) processes, at least in finite samples of typical size.

Figure 2

Dividend yield (left-hand scale) and one-month interest rate (right-hand scale)



Hence the spread between a long rate and a short rate should reflect the agents' expectations about future changes in the short rate and, under the expectations hypothesis, a constant term premium $\phi^{(n)}$ (see Figure 2). This is essentially an arbitrage condition saying that the investment in the long bond should earn the same return as successive short-term investments plus a risk premium that compensates for the capital risk incurred by holding the long bond.

2. Econometric evidence on the informational content and efficiency of German stock and bond market prices

The study of prices of long-term assets is intimately related to the study of long-horizon asset returns.[10] As equation (1.3) or, analogously, (1.4) shows, an infinite sum of future dividends enters into the calculation of the fundamental share value. Thus, the dividend of a single period can only be a small fraction of the stock price. Persistent changes in dividends therefore have a much larger influence on the stock price than do temporary dividend movements. A similar insight applies to changes in the discount rate used to value any financial asset.

This general conclusion provides the basis for the econometric analysis of this section. If dividend growth and discount rates follow predictable patterns, and if agents' expectations are not systematically biased, then the actual prices of longer-term assets like stocks and bonds should on average give useful information about the future course of asset returns or other variables correlated with the return process. It is intuitively plausible from the *RVF*s that in this case the forecast performance of current asset prices should generally be better for longer-term return measures (average returns), since these make up a larger part of the asset's calculated equilibrium price, and are, moreover, presumably less susceptible to large one-time shocks and peso effects than highly volatile short-term returns.[11]

In the following, long-horizon regressions are employed to determine the informational content of stock and bond market indicators regarding future stock returns, dividend growth, and short-term interest rate changes, respectively. Future ex post returns or short-rate changes measured over varying horizons are regressed on the current dividend yield or interest rate spread. The forecast

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[10] See Campbell et al. (1997), p. 253.

[11] See Kaul (1996), p. 284.

performance (predictive content) of each regression then helps to evaluate whether dividend yields or spreads better reflect *correctly anticipated* developments over longer or shorter horizons. This regression framework does not presuppose any specific model of the equilibrium return process. Thus, partial forecastability of returns (or excess returns) given lagged information, may indicate that conditional expected (equilibrium) returns are not constant, but vary over time, perhaps driven by time-variation in risk premia.[12] In addition, the predictive content of the same financial market indicators for future developments of macroeconomic variables like inflation or output generally provides some stylised facts about which fundamental factors are likely to determine equilibrium asset returns. In the present case, we ask about the informational content regarding inflation, since this is the most important variable from a monetary policy point of view.[13]

With respect to market efficiency, the long-horizon regressions for stock returns can be used to test the null hypothesis of constant equilibrium returns. Under this "traditional" hypothesis, future returns in excess of a constant should be unpredictable regardless of the return horizon and the information variables used.[14] In this single-equation setting, the unpredictability of stock returns can easily be tested by zero coefficient restrictions. However, in line with modern economic theory and the overall empirical evidence, it is now commonly believed that equilibrium returns vary over time. In this case, only returns in excess of the time-varying equilibrium component should be unpredictable. Efficiency tests under this assumption thus require a proxy for expected equilibrium returns. A short-term interest rate (the risk-free rate corresponding to the time-horizon over which returns are measured) is sometimes used for that purpose. As demonstrated above, this idea of constant equilibrium excess returns over a short-rate, applied to the bond market, leads to the expectation hypothesis of the term structure. Testing this hypothesis, which will be done below, is tantamount to testing bond market efficiency within the present framework.

Finally, a few comments on the data.[15] The RVF will not be applied to individual instruments but to broad portfolios of German stocks and bonds. While it is rather uncontroversial to refer to "average" bond yields calculated from a basket of homogeneous bonds (with comparable terms to maturity), it is more questionable using aggregate stock market data instead of data on single shares, since companies are likely to pursue very different dividend policies. But as Marsh and Merton have shown, *"it is (...) possible for aggregate dividends to exhibit stable and consistent time-series properties even if no such stability were found for individual firms."* Since, for theoretical and empirical reasons, the opposite is much less likely, it is advisable to use aggregate data if the empirical testing methodology strongly depends on capturing any systematic and stable element of dividend (policies) behaviour.[16]

2.1 The informational content of the dividend yield

Dividend yields, stock returns and dividend growth

We will begin with regressions that should reveal the information contained in the dividend yield for future stock returns and dividend growth. Equation (1.4) shows that the current

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[12] See Campbell (1987), p. 373.

[13] See Cuthbertson (1996), p. 129.

[14] Under risk-neutrality, asset returns should behave like martingales or random walks, respectively, which are unforecastable by definition. The neglect of time-variation in rational risk premia in a risk-averse world thus led to the long-held view that return predictability is synonymous to market inefficiency. See Kaul (1996), pp. 270-2.

[15] A more detailed description of the data is provided in the Appendix.

[16] See Marsh and Merton (1987), pp. 4 f.

dividend yield should predict future returns if the discount rates used by forward-looking investors actually depend on expected holding period returns for subsequent periods, and if these expectations do not deviate systematically, and too much, from realised returns. Since stock prices also depend on expected dividends, the dividend yield can only provide noisy measures of variation in expected returns, though, as Keim and Stambough put it, " *(...) whether this low signal-to-noise ratio destroys any ability of prices to predict returns is an empirical question.*"[17] The regressions for dividend growth are subject to the same omitted-variables problem because, in that case, expected stock returns introduce noise. To circumvent this problem, we also use the difference between returns and dividend growth as a single dependent variable.

Table 1

Long-horizon regressions of stock return measures on the log dividend yield

Regression equation: $\frac{1}{K}(x_{t+1}+...+x_{t+K}) = \alpha(K) + \beta(K)(d_t - p_t) + \varepsilon_{t+K,K}$

Estimation period with monthly data: December 1977 to June 1997

	Forecast horizon (*K*)					
	1	**3**	**12**	**24**	**36**	**48**
			$x_t = h_t$			
$R^2(K)$	0.001	0.013	0.052	0.102	0.120	0.352
$\beta(K)$	8.461	13.317	18.566	17.6113	16.240	17.389
t-value Newey and West	0.559	0.982	1.498	1.629	2.243	2.715
			$x_t = \Delta d_t$			
$R^2(K)$	0.046	0.108	0.229	0.166	0.143	0.107
$\beta(K)$	-18.770	-20.616	-15.058	-9.863	-7.217	-5.168
t-value Newey and West	-3.360	-3.327	-3.592	-2.761	-2.263	-1.659
			$x_t = h_t - \Delta d_t$			
$R^2(K)$	0.013	0.048	0.187	0.318	0.462	0.629
$\beta(K)$	27.231	33.933	33.624	27.474	23.457	22.557
t-value Newey and West	1.722	2.321	2.551	2.641	3.677	4.816

Notes: *h* is the annualised one-month continuously compounded stock return in per cent. Δd is the annualised one-month dividend growth rate in per cent. $(d - p)$ is the log dividend yield. $\alpha(K)$ (not shown) and $\beta(K)$ are the coefficients for the regression constant and the dividend yield, respectively, estimated by OLS. $\varepsilon_{t+K,K}$ are the error terms which are autocorrelated owing to data overlap for $K > 1$ under the null hypothesis of no predictability. Standard errors and t-values are corrected for serial correlation and heteroskedasticity in the equation error using the method of Newey and West (1987). Number of observations: 235 - (*K*-1).

Table 1 shows the regression results for each of the three dependent variables measured over a holding period (*K* months), ranging from one month to four years.[18] The regressions use monthly data, which means that data-overlap for the forecast horizons exceeding one month, induces

[17] See Keim and Stambaugh (1986), pp. 360 f.

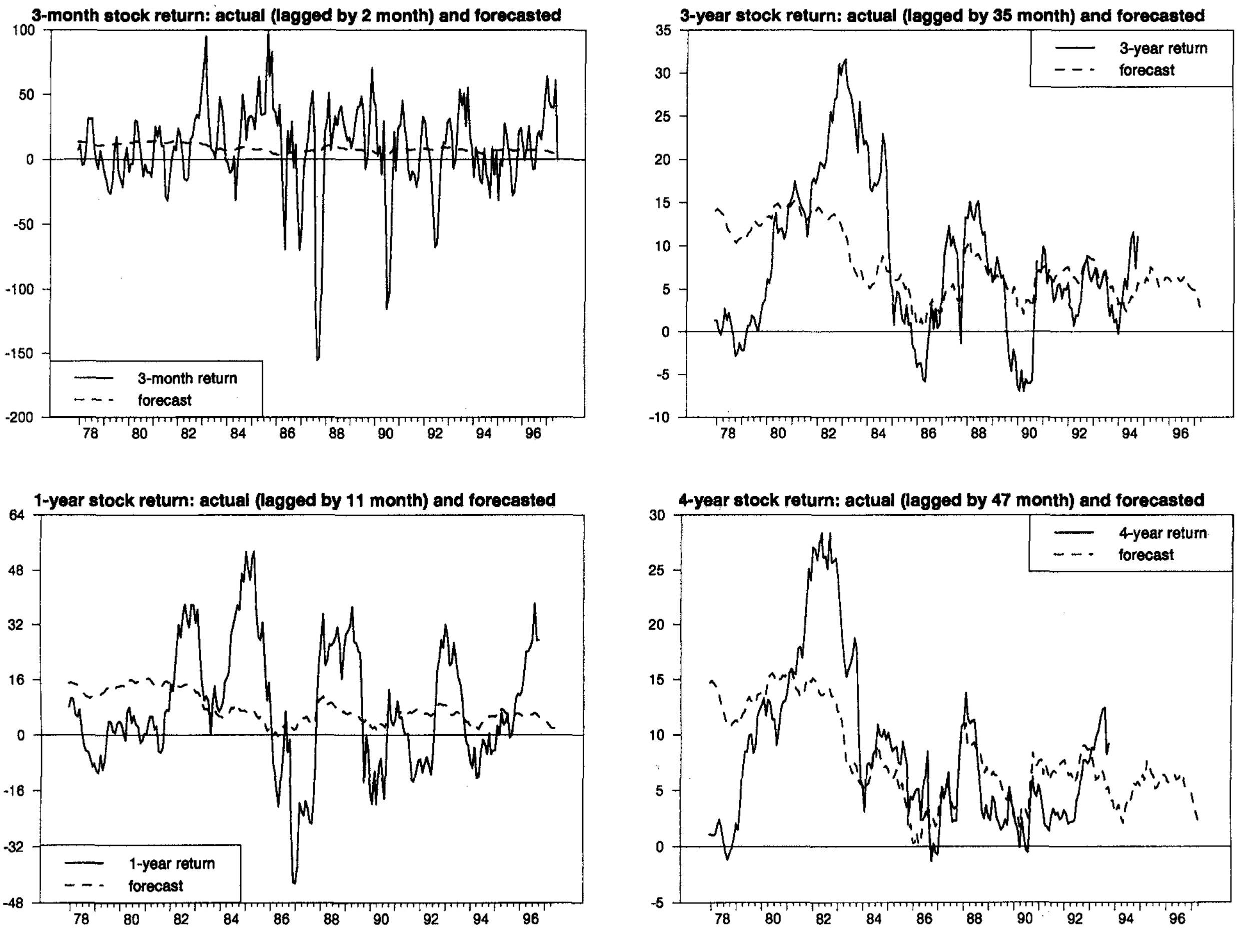
[18] The forecast horizons are chosen rather arbitrarily and follow the influential work of Fama and French (1988, 1989).

serial correlation of the error terms even under the null hypothesis of no return predictability (zero coefficient on the dividend yield). In this case, errors are correlated with K-1 previous error terms. But under alternative hypotheses, in which returns have a variable conditional mean, the serial correlation can in fact be arbitrary if dividend yields do not capture all of the variation in the conditional mean.[19] Additionally, since the regressor is only predetermined and not strictly exogenous, asymptotic distribution theory must be used to generate standard errors. The alternative t-statistics shown in the table for the null hypothesis of a zero coefficient are corrected for serial correlation and possible heteroskedasticity as suggested by Newey and West (1987) using a lag length of K-1.[20]

The upper part of Table 1 (see also Figure 3) summarises the main results for the stock returns regressions. The coefficient of determination (the $R^2(K)$ statistic) increases continuously with the forecast horizon, as do the t-values. The slope coefficients also increase from the one-month to the twelve-month horizon and remain roughly at that level for the longer forecast horizons. But statistical significance can only be attached to the 3-year and the 4-year return periods.[21]

Figure 3

Long-horizon regressions: stock returns and dividend yield



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19 See Hodrick (1992), p. 360.

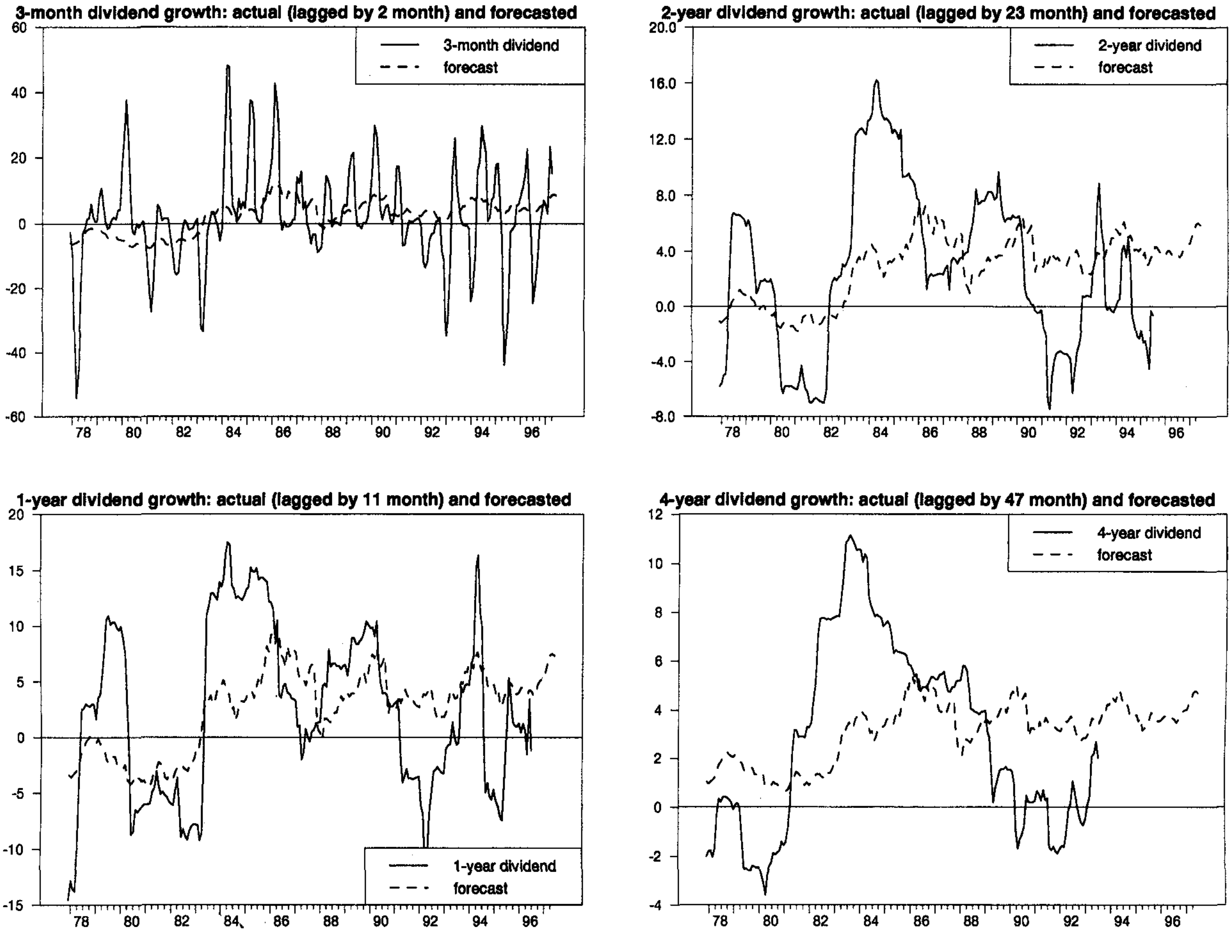
20 These issues are discussed in more detail in Campbell et al. (1997), pp. 534-6.

21 The results broadly conform to those for the US stock market, although our sample is much smaller, which weakens the comparability of results; see Fama and French (1988), p. 13 or Campbell et al. (1997), p. 269. The results for nominal returns are very similar to the results for real stock returns.

The R^2 statistics for the dividend growth regressions show, instead, a hump-shaped pattern and are much higher than in the stock returns case except for the 4-year horizon. They peak at the 1-year horizon with more than 20% of explanatory power (see Figure 4). What is more important is the high statistical significance of the slope parameters, particularly for the short to medium forecast horizons.

Figure 4

Long-horizon regressions: dividend growth [*ddiv*(*k*)] and dividend yield



The results for the combined returns variable (returns less dividend growth) are even more impressive. Although, by mere visual inspection, the time-series of this variable shows a very similar and volatile pattern as stock returns alone, the predictive power and the statistical significance of the slope coefficients are much higher for every forecast horizon (compare the results of the upper and the lower parts of Table 1, and see also Figure 5). The R^2 statistic increases to a remarkable 46 and 63% for the 3 and 4-year horizon, respectively. This comparison indicates that the noise introduced by dividend growth to the stock returns regressions is not negligible.

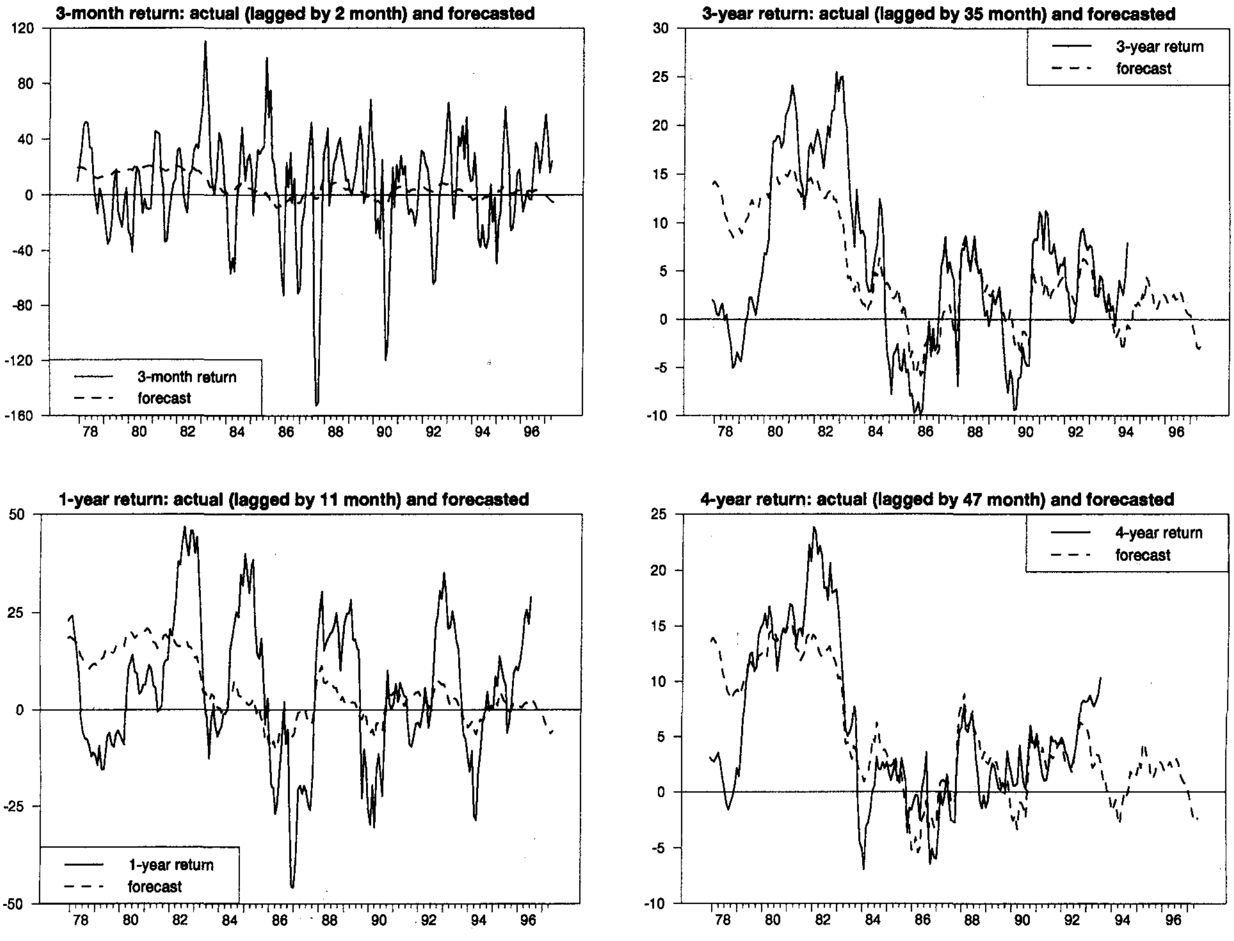
Although there are some serious doubts about the statistical reliability of long-horizon regressions, we interpret the results as providing sufficient preliminary evidence that future stock returns, and especially future dividend growth, contain predictable components which are reflected in the current dividend yield.[22] The fact that return predictability increases with the length of the holding period

[22] There are some general problems with long-horizon regressions in small samples. If the data are sampled more finely than the forecast interval, the error terms are autocorrelated at least of the order at which the data overlap. This autocorrelation is usually corrected for by use of some asymptotic distribution theory, in most cases with additional

considered may result from a better forecastability of the medium to long-term prospects of the economy (interest rates, business cycle positions).

Figure 5

Long-horizon regressions: combined returns [$r(k) - ddiv(k)$] and dividend yield



From the point of view of market efficiency in terms of rational expectations (summarised by the RVF) the results of Table 1 indicate that dividend yields signal persistent time-variation in expected equilibrium returns, rejecting the long-held hypothesis that equilibrium returns are constant. The efficient markets hypothesis only postulates that abnormal returns are unpredictable,

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correction for heteroskedasticity. But if the time span of data overlap is not small relative to the sample size, this approach is also flawed because there are not enough data points to reliably estimate the variance-covariance matrix. Monte Carlo simulations indicate that asymptotic standard errors can be very misleading in small samples (see Hodrick (1992), and Gerlach (1997), p. 164). An alternative is to use empirical standard errors using a bootstrapping procedure. But this method can only deal with biased standard errors. Another finite sample problem that puts into question the statistical reliability of long-horizon regressions derives from the fact that the independent variable, although predetermined with respect to the dependent variable, is stochastic and most likely correlated with past regression disturbances. This phenomenon leads to a finite-sample bias in the regression coefficients and the standard errors, *"(...) and the bias can be non-trivial even in samples of several hundred observations if the independent variable has both high autocorrelation and a high correlation with the past regression disturbance"* (Keim and Stambaugh (1986), p. 370). The dividend yield and term structure spreads – the independent variables used in this study – share at least the first property of being rather highly autocorrelated (i.e., highly persistent). But there are also some more theoretical problems caused by the very strong restrictions which rather simplistic models of the equilibrium returns process impose on the data. Modern theory suggests that the behavior of asset prices has much to do with the fundamental forces driving risk premia on the different assets. The assumption of constant risk premia provides a suitable starting point, but if risk premia actually play a significant role in asset pricing the econometrician most probably faces a serious omitted variables problem which biases coefficient estimates.

not that actual returns are unpredictable. High stock price volatility, as usually observed, is therefore compatible with persistent movements in rationally expected returns and need not indicate irrational investor behaviour.[23] But since return predictability could also result from irrational bubbles in stock prices, the question of whether the forecastability of stock returns is driven by rational economic behaviour or by animal spirits is still unresolved.[24] Further efficiency tests cannot fundamentally change this general conclusion but can only add evidence on the empirical plausibility of the rational valuation approach.[25]

Dividend yields and inflation

If we accept the view that stock prices are driven by expected equilibrium returns, it seems reasonable to ask whether the required rate of return includes a premium that compensates for inflation as expected over the holding period.[26]In that case, one could argue that the dividend yield should also have predictive power for future inflation.[27] But it has to be recognised that any empirical relationship between the two variables does not necessarily arise owing to an inflation premium in the dividend yield itself. If expected nominal dividend growth adjusts to inflation expectations in exactly the same way as the nominal discount rate does, the two effects on the dividend yield cancel out. The dividend yield can then be regarded as a real measure of stock returns and should not have any predictive power for future inflation unless expected real returns (including various risk premia) vary systematically with inflation expectations.

However, the regression results show for all forecast horizons high and significant slope coefficients which decrease with the horizon (see Table 2). The R^2 statistic is also always high, ranging from a minimum of 15% for the one-month period to a maximum of 54% for the 1-year horizon. The hump-shaped pattern of the R^2 statistic indicates that the forecast performance is best in a medium-term perspective (see also Figure 6).

How can this finding be interpreted in the light of the real nature of the dividend yield as explained above? We provide the following ad hoc explanation: First, assume dividend growth adjusts sluggishly to changes in the inflation environment. The expected dividend growth then falls short of the change in expected inflation. Second, if investors furthermore expect the central bank to raise (lower) short-term interest rates above (below) the upward (downward) shifts in expected or forecasted inflation, market participants will correspondingly require holding period returns which

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23 See Cuthbertson (1996), p. 129 and Campbell et al. (1997), p. 254.

24 In the case of bubbles, *"(...) dividend yields and expected returns are high when prices are temporarily irrationally low (and vice versa)"* (Fama and French (1989), p. 26).

25 To improve our understanding of the regression results in light of the rational valuation model we provide an illustrative example. When the log dividend yield decreases by 0.05 units from its long-term average (2.35% in logs) – which means a fall in the dividend yield of about twelve basis points – the average stock return tends to decrease by roughly 90 basis points over the next 4 years. This may be interpreted as follows: if investors require and expect a 90 basis points lower return on stocks, the log dividend yield will fall by 0.05 units. This in turn equals a 5% increase in the current stock price if dividends remain constant. The 25% increase from December 1996 until June 1997 (as measured with the price index used in this study) went along with a fall in the dividend yield of about 34 basis points. As predicted with the regression equation for 4-year returns, this fall is tantamount to a decrease in expected 4-year returns from 4.6 to 2.2%. This is a very low figure compared with average annualized stock returns of 8.5% over the past 18 years or so, but also relative to the level of short-term interest rates. Hence, if the forecast equation is not too biased, either rational investors are currently very risk prone regarding stock market investments, or economic agents behave irrationally, believing that the capital gains accrued over the recent months will continue or will at least not be reversed.

26 This does not preclude time-variation in real returns, which can be analysed separately, but is not the question of interest here.

27 The Fisher-effect can be analysed separately by running regressions between nominal stock returns and inflation or various proxies for inflation expectations. For some cross-country evidence see Solnik (1983).

increase (decrease) in excess of the inflation premium change. The net effect of the two offsetting channels through which changes in expected inflation influence share prices is to raise (lower) the current dividend yield, thus inducing a positive correlation between the dividend yield and future inflation.

Table 2

Long-horizon regressions of inflation on the log dividend yield

Regression equation: $\frac{1}{K}(\pi_{t+1}+\ldots+\pi_{t+K}) = \alpha(K)+\beta(K)(d_t-p_t)+\varepsilon_{t+K,K}$

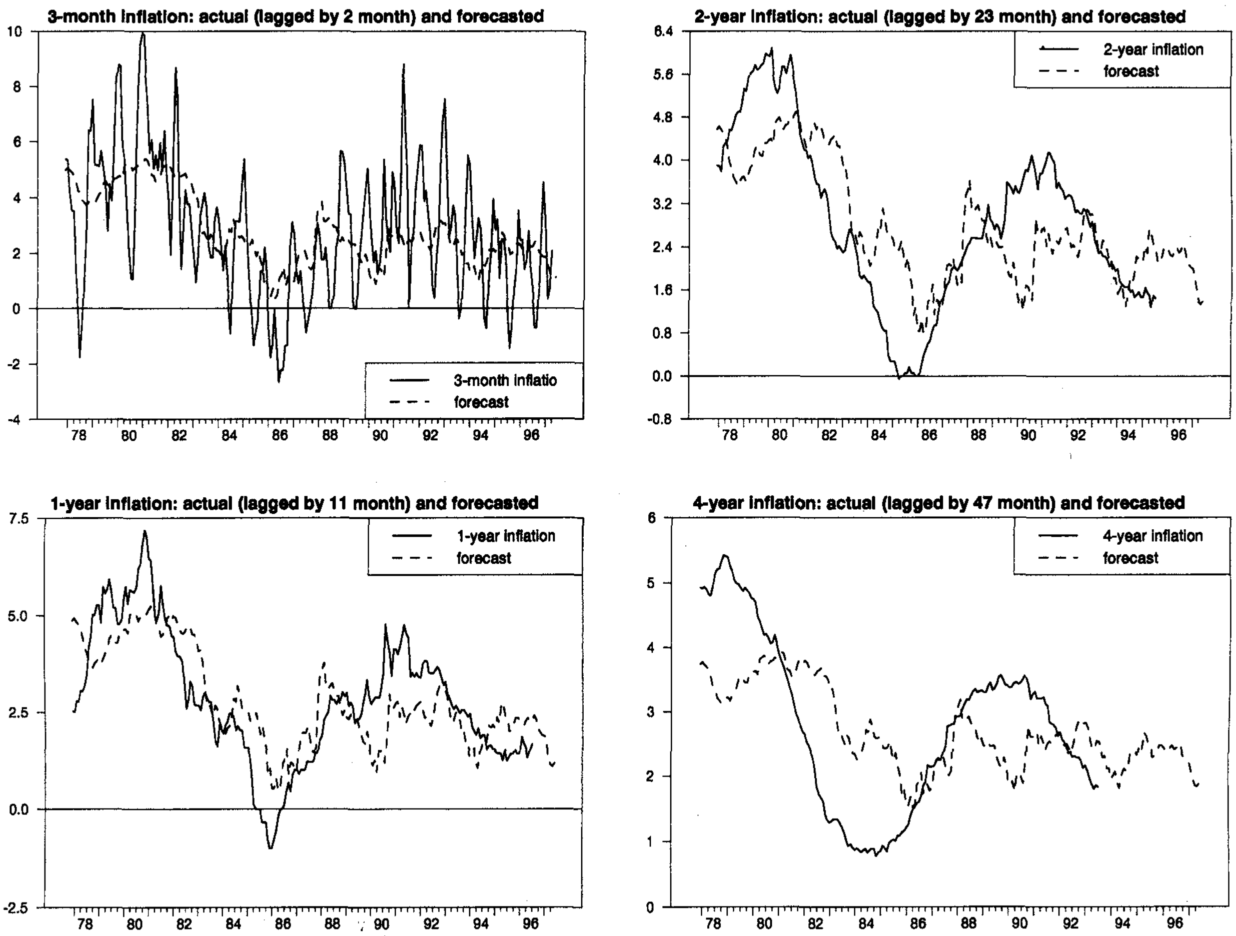
Estimation period with monthly data: December 1977 to June 1997

	Forecast horizon (*K*)					
	1	**3**	**12**	**24**	**36**	**48**
$R^2(K)$	0.151	0.277	0.542	0.470	0.367	0.241
$\beta(K)$	5.471	5.459	5.207	4.476	3.600	2.622
t-value Newey and West	6.436	5.857	5.342	3.847	3.117	2.366

Notes: π is the one-month continuously compounded rate of consumer price inflation. $(d - p)$ is the log dividend yield. $\alpha(K)$ (not shown) and $\beta(K)$ are the coefficients for the regression constant and the dividend yield, respectively, estimated by OLS. $E_{t+K,K}$ are the error terms which are autocorrelated owing to data overlap for $K > 1$ under the null hypothesis of no predictability. Standard errors and t-values are corrected for serial correlation and heteroskedasticity in the equation error using the method of Newey and West (1987). Number of observations: 235 - (*K*-1).

Figure 6

Long-horizon regressions: inflation and dividend yield



But some words of caution have to be added. Inflation and the dividend yield are highly persistent variables. According to standard unit-root tests, both variables can be regarded only as borderline stationary or near-integrated. From a mere statistical point of view, it is thus possible that the high R^2 statistics result from stochastic trends in the data and are thus spurious.

2.2 The information content of the term structure spread

The term structure spread and short-term interest rate changes

According to the expectation hypothesis with rational expectations (*EH-RE*), the spread is an optimal predictor for future changes in short-term interest rates. The spread should equal a weighted average of expected short-rate changes over the life of the long bond plus a constant risk premium. Referring to the long-horizon regression methodology, one can test the forecast accuracy by constructing the perfect foresight spread, $S_t^{n(pf)}$, for each bond maturity n from ex post values of short-rate changes as:

$$S_t^{n(pf)} = \sum_{i=1}^{n-1}(1 - i/n)E_t\Delta r_{t+i} + \phi^{(n)} \qquad (2.1)$$

Table 3

Long-horizon regressions of the perfect foresight spread on the actual spread

Regression equation: $S_t^{n(pf)} = \alpha(n) + \beta(n)S_t^n + \varepsilon_t^n$

Estimation period with monthly data: September 1972 to June 1997

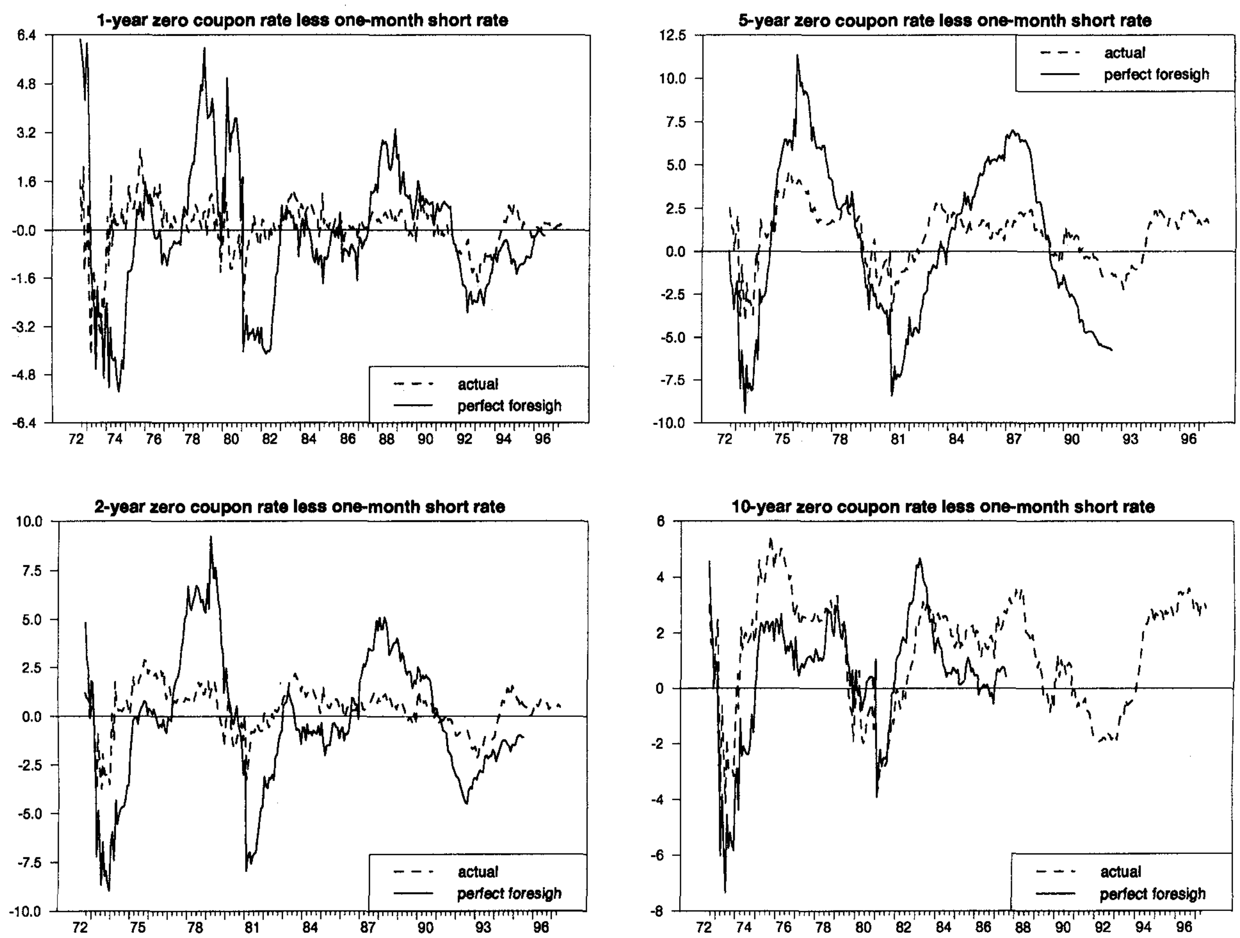
	Long-bond maturity in years (n/12)									
	1	**2**	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**
Term premium $\phi(n)$	0.07	0.34	0.58	0.78	0.93	1.05	1.14	1.21	1.26	1.31
R^2	0.13	0.33	0.40	0.51	0.70	0.71	0.59	0.43	0.47	0.50
$\alpha(n)$	-0.21	-0.76	-1.02	-1.30	-1.75	-1.56	-0.76	-0.13	-0.34	-0.56
	(0.36)	(0.64)	(0.77)	(0.72)	(0.53)	(0.62)	(0.88)	(0.79)	(0.27)	(0.65)
$\beta(n)$	0.89	1.69	1.86	2.08	2.36	2.11	1.51	0.94	0.84	0.73
	(0.19)	(0.38)	(0.37)	(0.23)	(0.17)	(0.22)	(0.39)	(0.40)	(0.08)	(0.14)
H0: $\beta(n) = 1$	0.54	0.07	0.02	0.00	0.00	0.00	0.19	0.88	0.05	0.06
H0: $\alpha(n) = 0$, $\beta(n) = 1$	0.72	0.03	0.00	0.00	0.00	0.00	0.40	0.85	0.14	0.00
Variance ratio (*VR*)	0.40	0.34	0.34	0.34	0.35	0.40	0.51	0.71	0.80	0.93

Notes: $S_t^{n(pf)}$ is the perfect foresight spread as defined in equation (2.1) using the respective term premium as given in the first line of the table. S_t^n is the actual spread between the n-period (in months) bond and the one-month interest rate. $\alpha(n)$ and $\beta(n)$ are the coefficients (standard errors in brackets) for the constant term and the actual spread, estimated by OLS. ε_t^n are the error terms which are autocorrelated of order n-1 due to data overlap. Standard errors are corrected for serial correlation and heteroskedasticity in the equation error using the Newey and West (1987) method. The values shown for the hypothesis tests are p-values; the test statistic for the Wald-test is distributed as $\chi^2(df)$ with df = 1 and 2 degrees of freedom. The variance ratio is defined as the sample standard deviation of the actual spread, divided by the standard deviation of the perfect foresight spread. Number of observations: 298 - (n-1).

and regress it on the actual spread and a constant. We do this for spreads between long-bond zero-coupon rates and the one-month interest rate on the interbank money market in Frankfurt. The long-bond maturities tested range from 1 year (n = 12 months) to 10 years (n = 120 months). In constructing the perfect foresight spread we face the problem of how to get an estimate of the term premium. We use a common but rather crude method and estimate the term premium for each maturity by the difference in the sample means of the respective long rate and the short-term interest rate.[28] As can be seen from the first line of Table 3, the estimated term premia increase with bond maturity. This is not compatible with the conventional interpretation of the *EH* which assumes *constant and equal* term premia for all maturities. Instead, the relevant hypothesis to be tested is the liquidity preference hypothesis, which exactly adds to the *EH* the assumption of term premia increasing with bond maturity. For the sake of simplicity, we subsume the liquidity preference hypothesis under the notion *EH*.

Figure 7

Long-horizon regressions: perfect foresight spread and actual spread



The R^2 statistic is rather high for all maturities but the one-year horizon. It peaks at the medium-term maturities of 5 and 6 years, at about 70%. The slope coefficients show a more pronounced hump-shaped pattern with the highest value of 2.36 for the 5-year maturity. Thus, high (low) R^2 statistics tend to be associated with high (low) slope coefficients. Taken together, this suggests that investors can reliably predict only medium-term, but not very near-term, developments of future short rates, which may be based on better medium-term forecastability of real activity and

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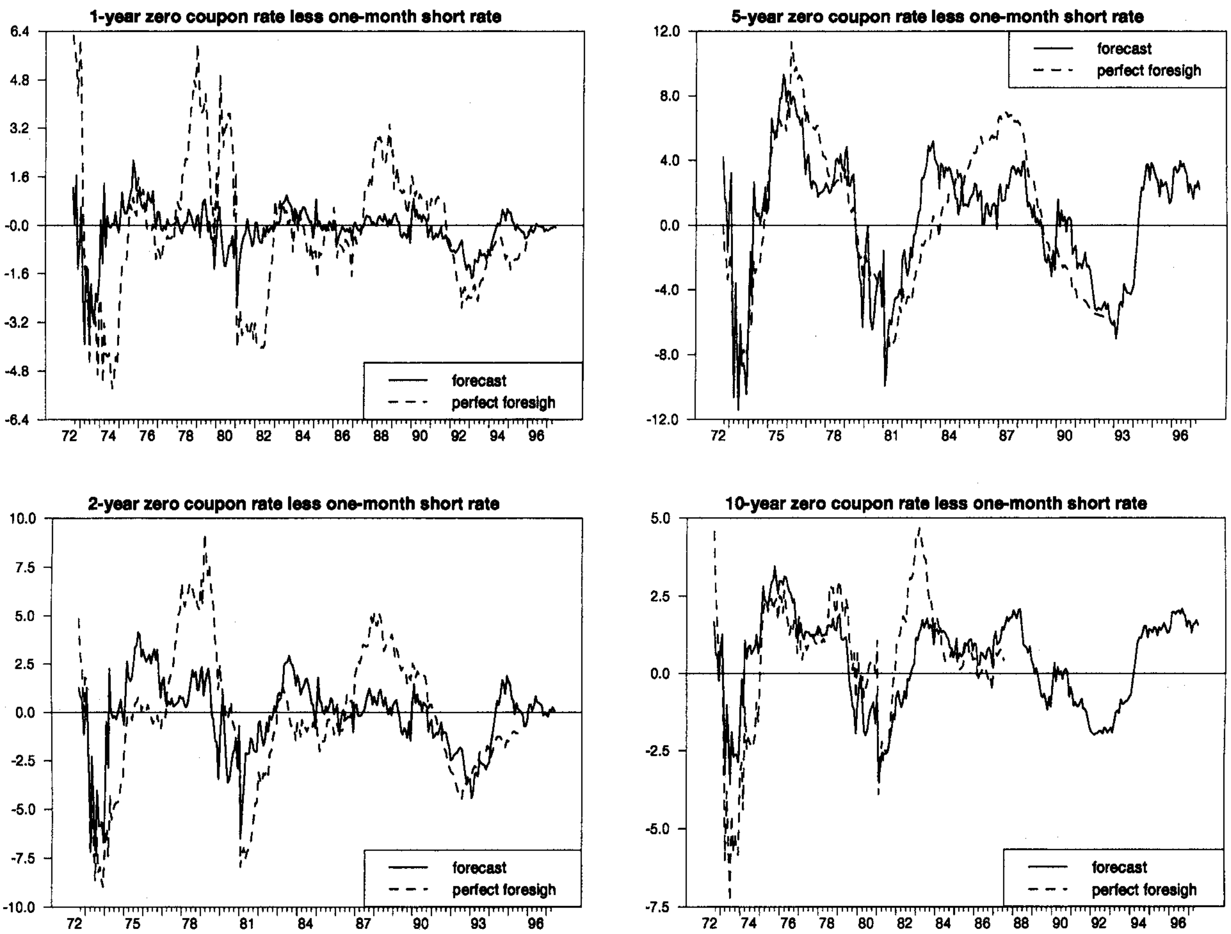
28 See, for example, Shiller (1989), p. 225.

inflation. But although the direction of change has been anticipated quite well, the magnitudes of the medium term interest rate changes have been underestimated, causing the slope coefficients to get significantly larger than one to improve the regression fit (compare Figures 7 and 8). This explanation may hold at least within the given sample, which includes highly volatile periods, such as the oil-price shocks and German reunification. However, the interest rate shocks associated with these exceptional phases were only temporary and vanished or cancelled out after the medium term. Thus, for times to maturity of 8 to 10 years the accumulated short-rate changes are much lower, the slope coefficients are around one, and the relative standard deviations of the actual and the perfect foresight spreads (the variance ratio) approach unity. But it has to be conceded that differences in the slope coefficients may also arise from the influence of omitted variables, especially those factors which may introduce time-variation in the term premia.

Figure 8

Long-horizon regressions: perfect foresight spread and its forecast



This regression framework also forms a basis for testing market efficiency or the *EH* using a rather strong definition of rational expectations (*EH-RE*). It assumes that investors can forecast future short-rate changes perfectly save a pure white noise error which is orthogonal to all information at time t (the forecast origin):

$$\Delta r_{t+i} = E_t \Delta r_{t+i} + \eta_{t+i} \tag{2.2}$$

with i = 1, ..., n-1. Substitution into (2.1) leads to the testable hypothesis that the perfect foresight spread should equal the actual spread (its optimal predictor); differences between the two should be purely random and uncorrelated with all information available at time t or earlier (to which the actual spread itself belongs, too):

$$S_t^{n(pf)} = S_t^n + \varepsilon_t^n \tag{2.3}$$

The regression equation in Table 3 represents the appropriate testing framework. Under hypothesis (2.2) the regression error is a moving average process of order (n-1) for monthly data:

$$\varepsilon_t^n = \sum_{i=1}^{n-1} (1 - i/n)\eta_{t+i} \tag{2.4}$$

The expected value of the compound forecast error is still zero, but successive errors are autocorrelated and possibly heteroskedastic. The standard errors for the regression coefficients are therefore again corrected for serial correlation and heteroskedasticity, using the Newey-West method.[29] The *EH-RE* or efficient market hypothesis implies the restrictions $\beta(n) = 1$ and $\alpha(n) = 0$. Table 3 shows the p-values for Wald-tests of the first restriction (fourth line) and of both restrictions together (fifth line). The best results from the efficient market view are for the 1-year and the 7-year to 9-year maturities with sufficiently high p-values for both restriction sets. Particularly for the medium-term maturities, the spread is a biased (with slope coefficients much above one, the value implied by the efficient markets hypothesis), although a better predictor of future short-rate changes.

This model-consistent performance of the longer maturities also shows up in the variance ratios, which are much higher than for the shorter maturities and approach unity for the 10-year maturity. As can be derived from equation (2.3) and the null hypothesis of *RE*, the variance (or the standard deviation) of the perfect foresight spread must always be higher than the variance of the actual spread.[30] This is actually the case for all maturities, but since the variance ratio (actual to perfect foresight spread) approaches one with decreasing variance of forecast errors, a high (low) variance ratio indicates low (high) forecast error variances. Hence, the accumulated long-run forecast errors tend to be significantly lower than errors summed over shorter time periods. This in turn confirms our conjecture, above, that the cancelling-out of temporary strong interest-rate movements over the longer periods reduces the bias in the slope coefficient and hence weakens evidence against the efficient market hypothesis.

However, there are still some more fundamental doubts about the appropriateness of using perfect foresight measures of expectations as the basis for testing market efficiency. This very strong hypothesis of *RE* assumes that agents can forecast with 100% accuracy, regardless of any unforeseeable special events that occur during the sample. An alternative, ex ante oriented, approach tries to find a suitable (multivariate) time-series representation of the data and expectations generating process and to draw inferences about market efficiency from forecasts based on such models.[31]

Term structure spread and inflation changes

The Fisher theorem states that the current nominal interest rate of a bond in equilibrium equals the expected real interest rate plus the (annualised) expected rate of inflation over the life of the bond. The real rate also contains any risk premium required by investors. If this relation holds and

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29 See Cuthbertson (1996), p. 325.

30 See Cuthbertson (1996), p. 138.

31 The so-called Campbell and Shiller (1987) approach provides some metrics to test market efficiency in this context. For some exemplary evidence on the German bond market see Gerlach (1996). Domanski and Kremer (1997) apply this approach to the German stock market.

if the real interest rate is constant, then the spread between the interest rates of an m-year and a j-year bond should exactly correspond to the (annualised) difference in expected inflation m years and j years ahead, respectively. Hence it makes sense to use term structure spreads as indicators of changes in inflation expectations held by market participants. In a recent study, Schich (1996) analyses the predictive content of spreads regarding future inflation changes by using zero-coupon rates for the German government bond market. We refer to this study for the details and show slightly updated results for the long-horizon regressions in Table 4 (see also Figure 9).[32]

Table 4

Long-horizon regressions of inflation changes on spreads

Regression equation: $\Delta\pi_t^{(m,1)} = \alpha(m,1) + \beta(m,1)S_{t-1}^{(m,1)} + \varepsilon_t^{(m,1)}$

Estimation period with monthly data: September 1972 to June 1997

	Longer-bond maturity in years (m)								
	2	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**
R^2	0.040	0.113	0.203	0.305	0.362	0.357	0.348	0.330	0.273
$\alpha(m,1)$	-0.188	-0.441	-0.724	-1.022	-1.267	-1.322	-1.468	-1.540	-1.528
	(0.120)	(0.210)	(0.282)	(0.349)	(0.389)	(0.466)	(0.400)	(0.373)	(0.388)
$\beta(m,1)$	0.268	0.450	0.633	0.801	0.893	0.851	0.816	0.746	0.646
	(0.164)	(0.180)	(0.135)	(0.120)	(0.125)	(0.143)	(0.125)	(0.116)	(0.132)

Notes: $\Delta\pi_t^{(m,1)} = \pi_t^{(m)} - \pi_t^{(1)}$ is the difference between the annualised one-year and m-year-ahead rate of inflation defined as $\pi_t^{(j)} = 100/j(p_{t+12j} - p_t)$ with $j = 1, \ldots m$ and p the log of the German consumer price index, for monthly data. The spread variable is simply defined as the difference between the zero-coupon rate for an m-year and the one-year bond, i.e., as $S_t^{(m,1)} = z_t^m - z_t^1$. $\alpha(m,1)$ and $\beta(m,1)$ are the coefficients (standard errors in brackets) for the constant term and the spread variable, estimated by OLS. $\varepsilon_t^{(m,1)}$ are the error terms which are autocorrelated of order $(12m - 1)$ due to data overlap. Standard errors are corrected for serial correlation and heteroskedasticity in the equation error using the Newey and West (1987) method. Number of observations: 298 - $(12m + 1)$.

The empirical evidence suggests that especially the medium to long-term segments of the German term structure of interest rates have significant forecast power for future inflation changes over the respective horizons with R^2 statistics of up to about 40%. This, in turn, favours the interpretation that market agents can anticipate longer-term developments of inflation better than short-term changes. The values of the slope coefficients show the hump-shaped pattern sometimes observed in the other long-horizon regressions above. For the longer maturities, they are in the neighbourhood of one, the value which would obtain if the strong *RE* assumption were imposed on inflation expectations. These general results are robust to the use of either zero-coupon rates or yields to maturity.[33]

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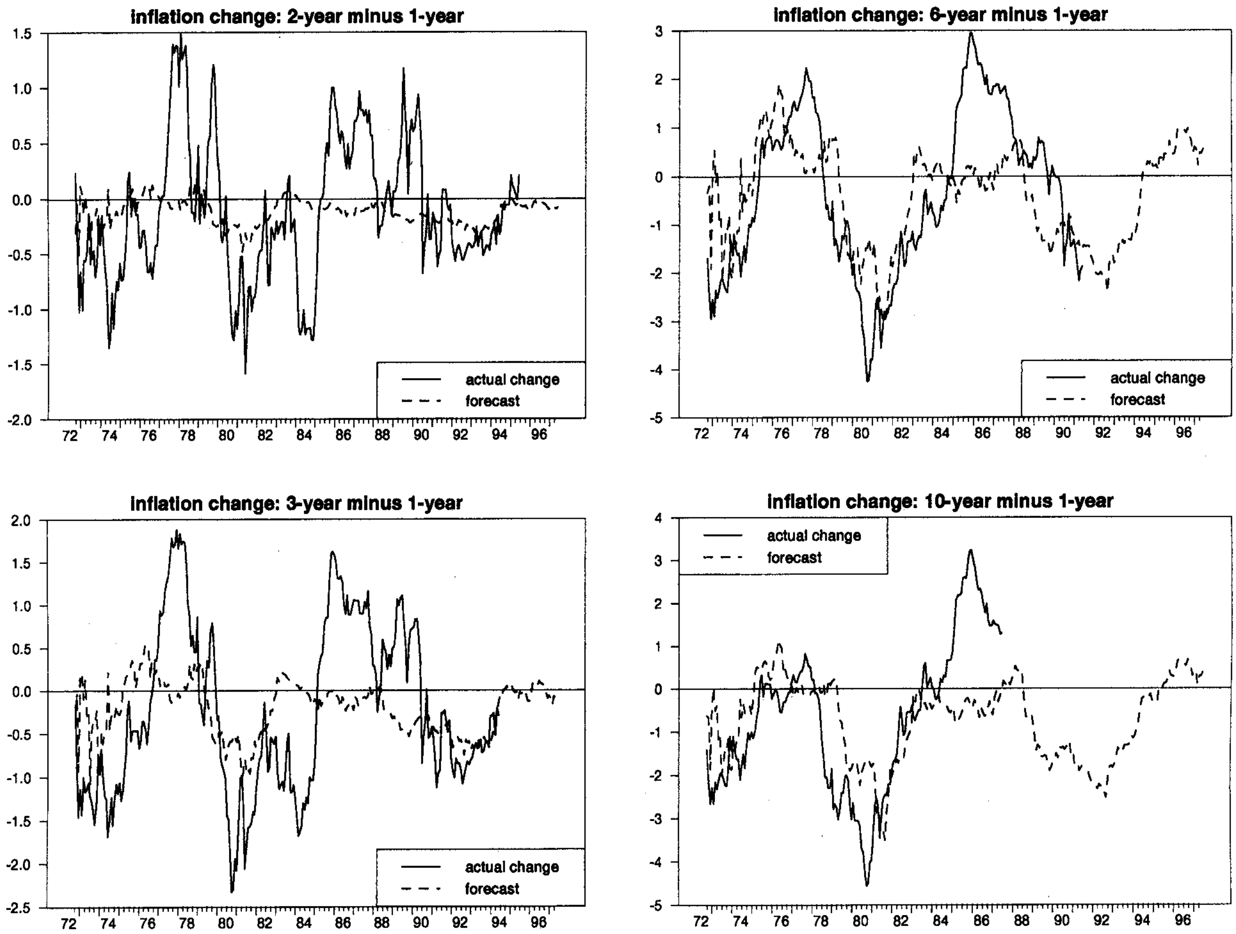
32 See Schich (1996), pp. 39-50. See also Gerlach (1997) for a brief survey of related studies and the origins of the testing methodology.

33 See again Schich (1996), who compares the results obtained for both interest rate measures.

Figure 9

Long-horizon regressions: inflation and spread forecast



3. Implications for monetary policy

3.1 Impact of monetary policy on asset prices

According to the rational valuation approach, monetary policy can influence asset prices by two channels. First, the price of long-term assets like stocks and bonds reflects agents' expectations about the course of the short-term interest rate which may serve as a benchmark for equilibrium one-period returns. These short-term returns are used for discounting the assets' future streams of income. As monopolistic supplier of base money, the central bank can – at the operational level – determine short-term interest rates and thus affect asset prices via agents' expectations about the future path of money market rates. Second, since in general the nominal long-term returns which investors require to hold an asset in their portfolios should contain an inflation premium, it is the long-horizon perspective about future inflation that influences today's prices of long-term assets. At the strategic level, however, monetary policy controls inflation in the longer run. Hence, monetary policy has a strong impact on asset prices by affecting agents' inflation expectations over longer horizons.

But both channels are merely two sides of the same coin since in equilibrium successive short-run returns simply have to add up to long-run expected returns. The long and short-term perspectives are interlinked by the central bank's reaction function as perceived by economic agents.

A change in inflation expectations, for example, should cause a shift in the path of expected short-term rates and vice versa. This link has clear implications for the way monetary policy decisions affect the level of and changes in asset prices. By reducing short-term rates below equilibrium level, the central bank may increase stock prices and the term spread if long-term expectations remain unaffected. But this only occurs if the central bank measure conforms to the monetary policy regime implied by the reaction function which agents use to determine their expectations about future inflation and short-term interest rates. If the measure comes as a surprise and does not fit to previous experiences with monetary policy, there always exists the danger that asset prices react in a way which counteracts the intentions of central bankers. Short-term fluctuations of asset prices – their volatility – in this view depend on how often and to what extent expectations have to be revised by market participants.

A first conclusion from this is that a predictable monetary policy makes it easier for economic agents to form expectations. First, through an unambiguous obligation to the goal of price stability monetary policy provides a nominal anchor for inflation expectations over longer horizons. Second, a transparent strategy establishes a link between this strategic level and the operational level reflected in short-term interest rates. Under this conditions, it is reasonable for market participants to assume that short-term (policy-determined) rates might fluctuate significantly in the short run (in order to contain inflationary pressures and to make real "monetary" rates conformable to expected changes in real "capital" rates), but should return to a "normal" level in the medium run. Third, monetary policy should be able to smooth market volatility by reducing uncertainty of future rate changes. This, again, is a facette of a transparent strategy, but is also related to the implementation of monetary policy. If money market rates fluctuate by chance or in an undesired manner owing to unexpected changes in banks' liquidity, there can be volatility spill-overs to other financial markets.[34]

From this point of view, the empirical results presented above can be seen as an indication that monetary policy in Germany has been able to provide a relatively reliable medium-term orientation, thus facilitating the process of expectation formation regarding inflation and short-term interest rates. The fact that the forecast performance of the dividend yield with respect to future inflation is better at shorter horizons than in the case of the term structure spread may indicate that other factors which determine stock returns dominate the influence of inflation, especially over longer horizons. That is, the noise introduced by the omitted variables in the forecast equations for inflation (changes) is probably stronger for the dividend yield regressions. Furthermore, the results support the view that short-term expectations about stock returns and money market rates are often subject to disappointments reflecting unprecedented macroeconomic shocks. In the short run, these shocks can have a very strong and unexpected impact on inflation rates and the path of short-term interest rates which renders econometric analysis – using either ex post data or ex ante measures of the variables to be forecasted – more difficult.

3.2 The use of asset prices as monetary policy indicators

The empirical evidence presented in this paper shows that the dividend yield and the term structure spread contain useful information about future stock returns, dividend growth, short-term interest rate changes and inflation (changes) as expected by market participants, at least over medium-term horizons. At a first glance, this seems to support an outstanding role for financial market prices as indicators for monetary policy. However, although the regression fit is in most cases impressive according to standard metrics, the forecast errors are generally rather high from an operational point of view. Thus, policy makers face a lot of uncertainty if they try to evaluate whether any change in the indicator variable reflects shifts in agents' expectations or, instead, the influence of other factors omitted from the forecasting equation. Moreover, from a strategic perspective, it is crucial that monetary policy still relies on an "external" anchor and not on market expectations themselves.

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[34] See Schmid and Asche (1997).

The anchoring of expectations about monetary policy can probably best be achieved by a strong and credible commitment to long-term price stability. The respective long-term inflation goal is then given a heavy weight in any reaction function which economic agents use in forming their expectations about the future course of short-term interest rates.

By instead linking monetary policy decisions to market expectations, the form of expectations about inflation and, connected to that, the future path of short-term interest rates becomes self-fulfilling and could lead to policy instability and hence inflation instability.[35] This makes room for speculative attacks in financial markets and jeopardises the credibility of the central bank.

Independently of the danger of sliding into a vicious circle, putting more weight on market expectations could be interpreted as a shift in the monetary policy regime by market participants. This makes it difficult for the central bank to assess the stance of monetary policy because market indicators become less reliable (which should show up in coefficient changes in the forecasting equations) and other indicators (as, for example, the money stock) may lose their indicator properties owing to changes in the behaviour of market participants. Finally, the central bank could end up in a situation in which it is impossible, or at least rendered more difficult, to stabilise expectations just because monetary policy has been geared to market expectations. All this suggests, as Woodford convincingly argued, that modelling structural relationships, including the monetary policy reaction function, is unavoidable in order to make more reliable inferences about the indicator quality of a financial market variable and to assess its usefulness for monetary policy purposes.[36]

Appendix: Data description

The monthly stock price and dividend series used in this study are calculated by the Federal Statistical Office up to June 1995. The computations are based on a fictitious share having the face value of DM 100. The stock price series is the arithmetic mean of the end-of-month prices of all the shares of public limited companies officially listed on German stock exchanges (stock prices of each company are previously multiplied by a factor which raises or lowers its face value to DM 100). The series is thus equivalent to an equally-weighted stock price index. The dividend series is calculated correspondingly. However, the monthly dividend (excluding tax credit) of each share is the dividend as last paid out. The dividend yield (in per cent per annum) is defined as the ratio of dividends to stock prices multiplied by one hundred. While the stock price series is available for the period from January 1960 to June 1995, the dividend series only begins in November 1977. Both series are published in Deutsche Bundesbank, Capital Market Statistics, Statistical Supplement to the Monthly Report 2, Table IV.2. Complementary series for the period from July 1995 to the present are calculated by the Deutsche Börse AG. But as the number of stocks included in the calculation is reduced (only ordinary and preference shares officially listed on the Frankfurt stock exchange of companies domiciled in Germany are included) a statistical break occurs which is accounted for in the empirical analysis.

The interest rates representing the German term structure are estimated zero-coupon rates. They are estimated from the prices of listed coupon bonds issued by the Federal Government. For a detailed description of the estimation procedure see Deutsche Bundesbank (1997). The monthly series comprise end-of-month data as published in Deutsche Bundesbank, Capital Market Statistics, Statistical Supplement to the Monthly Report 2, Table II.7e).

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35 See Woodford (1994), p. 104.

36 See Woodford (1994), p. 112 f.

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