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The term structure spread and future changes in long and short rates in the G7 countries

Is there a puzzle?

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According to the expectations hypothesis, when the spread between long and short rates widens, next quarter's long rate should rise. In the United States, however, long rates decline instead of rising. This anomaly is also present in Canada, the UK, Germany, and Japan (four of the additional six G7 countries). Nevertheless, in contrast to the US, where long rates appear to overreact to expected future developments, the anomalous short-run movement of long rates in these other G7 countries is caused by an additive white noise error on long rates that does not materially affect the information in their term structure.

Key words: Term structure; G7 countries; Expectations hypothesis; Overreaction

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1. Introduction

An extensive recent literature on the term structure of interest rates in the United States documents that the spread between long- and short-term interest rates can predict the correct direction of future changes in short rates [Campbell

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and Shiller (1987), Fama (1984, 1990), Fama and Bliss (1987), Hardouvelis (1988), Mankiw and Miron (1985), Mishkin (1988, 1990a, 1991)].[1] When the long rate rises relative to the short rate, future short rates tend to increase. Such predictive power is consistent with the expectations hypothesis of the term structure, which claims that long rates are weighted averages of current and expected future short rates. According to the expectations hypothesis, a rise in the long rate relative to the short rate is due to the expectation of higher short rates in the future. Thus, if the market makes correct predictions on average, future short rates would subsequently tend to rise, generating a positive correlation of the change in short rates with the earlier spread.

On the other hand, Shiller (1979), Shiller, Campbell and Schoenholtz (1983), Campbell and Shiller (1984), Mankiw and Summers (1984), Mankiw (1986), and again more recently Campbell and Shiller (1991) observed that the spread predicts the wrong direction in the subsequent change of the long rate: A rise of the current long rate relative to the current short rate is followed by a subsequent decline, rather than a rise, in the long rate next period. *This behavior is puzzling: How can the movement of future cumulative short rates obey the overall direction predicted by the expectations hypothesis but at the same time the short-run movement of long rates does not?*

Two main alternative explanations to the puzzle have been proposed: The first, and what appears to be the most popular one, claims that movements in current long rates do obey the general direction predicted by the expectations hypothesis, but those movements are sluggish relative to the movements of the current short rates: Long rates underreact to current short rates (or overreact to future short rates).[2] This explanation assumes that risk premia are constant and that the spread between long and short rates correctly incorporates the information about expectations of future interest rates, but that the market's expectations themselves violate the strict definition of rational expectations. Suppose, for example, that a policy announcement by the Federal Reserve increases the market's expectation of future short rates but, since the policy will be implemented in the future, leaves the current short rate intact. The hypothesis claims that markets would overreact to the announcement, raising their expectations of future spot rates by more than is warranted. The current long rate would thus increase by more than warranted, making the spread between long and short rates larger than it should be. During the next month or quarter, long rates

[1] Fama (1990), Mishkin (1990a, 1990b, 1991), and Jorion and Mishkin (1991) further examine the predictability of inflation versus the real rate of interest. Chen (1991) and Estrella and Hardouvelis (1991) show that the spread can predict real economic activity.

[2] Campbell and Shiller (1987, 1991) make this point forcefully, arguing that the spread between long and short rates does track over time very closely the 'theoretical' spread that would obtain if the expectations hypothesis were correct. Nevertheless, the spread moves more than one-for-one with the theoretical spread, i.e., overreacts slightly, resulting in a statistical rejection of the expectations hypothesis.

would fall somewhat correcting the previous overreaction, thus generating a negative correlation between the change in long rates and the previous spread. Short rates, on the other hand, would begin their predicted rise, generating a positive correlation between the change in short rates and the earlier spread.[3]

The second explanation assumes that the market's expectations are rational but the information in the spread is composite information about the variation of both expected future rates and risk premia. This explanation requires, however, that the time-varying risk premium have a very special structure. Consider our earlier example in which the Fed announcement increases the expected level of future short rates. For the risk premium explanation to explain the observed correlations in the data, the risk premium ought not to show a large response at the time of the announcement in order for the expectations hypothesis to approximately hold, yet one to three months later the risk premium ought to decline so that the long-term bond yield also declines. It is hard to think of a simple and plausible economic story that could explain the required behavior of the risk premium. Nevertheless, we explore the plausibility of the risk premium explanation in later sections.

Froot (1989) uses U.S. survey data on short-term and long-term interest rates and is able to distinguish between the two competing hypotheses. Under the maintained hypothesis that survey results reflect the true beliefs of the respondents and that the respondents are representative of the market as whole, the survey expectations can be used to decompose the term structure spread into an expected change in the future spot rate and a risk premium. Froot finds that the negative correlation between changes in long rates and the previous spread is not due to a time-varying risk premium, but is due to a violation of the rational expectations assumption, namely, an overreaction of the spread. When buying long-term bonds, market participants would do better to place more weight on the contemporaneous short rate and less weight on the expected future short rates.[4]

[3]There is some independent evidence in the literature that long rates do overreact to news. Cook and Hahn (1989) examine the reaction of the term structure to discretionary changes in the Fed funds rate and show that forward rates of the distant future appear to react by more than one would expect based on the expectations hypothesis.

Hardouvelis (1987) finds a puzzling overreaction of long-run forward rates to the weekly announcements of bank nonborrowed reserves. Changes in reserves ought to have only a temporary negative influence on interest rates. Yet, one-year forward rates as far as three or four years into the future react negatively to the unanticipated increase in nonborrowed reserves.

A similar puzzle was central to the voluminous money announcements literature of the last decade [Cornell (1983), Hardouvelis (1984), Roley and Walsh (1985)]. For example, five-year forward rates five years into the future respond positively to the weekly surprise in M1. This could be an overreaction, but unlike the case of nonborrowed reserves, the reaction could be also explained by an increase in the inflation premium in bond yields.

[4]Evans and Lewis (1991), who examine the term structure of Treasury bills that mature from one to twelve months, also claim that stationary risk premia cannot explain the same puzzle at the short end of the term structure.

Froot's careful work has not settled the issue in the minds of those economists who place less trust in surveys. Also, Froot's work cannot address the empirical observation of Mankiw (1986) that the puzzle also appears in all the other countries he examined: Canada, Germany, and the United Kingdom. Surveys are not readily available for those countries. Mankiw finds that the spread is negatively correlated with the subsequent change in the long rate and positively correlated with the subsequent change in the short rate. However, Mankiw's analysis of short rates, although suggestive of the general direction of their future change, is incomplete. The spread does not make a prediction of the immediate change in short rates alone; rather, it makes a prediction about a cumulative change in short rates over a long horizon. Thus it is not clear that the expectations hypothesis is consistent with the international evidence on short rates and, therefore, that there is a discrepancy between the behavior of long and short rates, or a puzzle, to be explained in the first place.

This paper takes a fresh look at the puzzle by examining the relation between the spread and the future evolution of long and short rates internationally. It uses post-war data on an approximately ten-year yield and a three-month yield of each country that belongs to the Group of Seven (G7). The paper finds that, curiously, the puzzle is manifested primarily in the United States, the country with the most sophisticated and liquid financial markets. In France and Italy, long rates move in the correct direction. In Canada, the United Kingdom (UK), Germany, and Japan, long rates move in the opposite direction, but this movement is apparently due to a white noise error that does not materially affect the information in the term structure. The use of instrumental variables reverses the negative regression sign. Moreover, a Vector Autoregressive (VAR) methodology similar to Campbell and Shiller (1987, 1991) confirms that, with the exception of the United States, the spread tracks the theoretical spread of the expectations hypothesis very closely. Similarly, multiperiod regressions in these countries show that the cumulative evolution of future short rates corresponds closely to the predictions of the expectations hypothesis.

The rest of the paper is organized as follows: Section 2 discusses the data and conducts a preliminary analysis of their time series properties. Section 3 analyzes the correlation of the term structure spread with the change in long and short rates over the following quarter. It also assesses the success or failure of four different hypotheses – the overreaction, the risk premium, and two additional hypotheses – in explaining the negative correlation of the change in long rates with the earlier spread. Sections 4 and 5 examine whether or not the hypotheses that can explain the short-run changes of the long rate in section 3 can also explain the long-run behavior of interest rates. Section 4 employs the Campbell–Shiller VAR methodology to assess both statistically and economically the deviation in the behavior of the actual term structure spread from the theoretical term structure spread under the expectations hypothesis and, moreover, to explore the most likely alternative hypothesis that is consistent with such a deviation. Section 5 performs a more direct regression analysis over multiperiod horizons of the cumulative change of the short rate on the earlier term structure spread. Section 6 concludes.

2. International yields in historical perspective

The analysis uses post-war end-of-quarter data on a three-month and a ten-year government yield. The data extend as far back as possible and end in the second quarter of 1992. In the US, the sample begins in 1953 in order to exclude data from the period of the Treasury Accord that ended a year earlier. In Canada, the sample begins in 1950. In Japan and the UK, the sample begins in the early 1960s, in Germany and France in the late 1960s, and in Italy in the early 1970s. Table 1 describes the data and their sources in more detail.

Panel A of table 2 provides summary statistics for the level of long and short rates, their first differences, and the term structure spread. The level and standard deviation of those variables does vary from country to country but not by a substantial margin. One exception is the volatile French short rate.[5] Panels B, C, and D of table 2 present the cross-country correlations of those variables. With the exception of the US–Canada correlations, these correlations are small. The behavior of interest rates in foreign countries provides independent information, therefore, not present in the US term structure.

Table 3 presents unit root tests for the two interest rates and their difference, the term structure spread. The levels of the long and short rate have a unit root, but their first difference and the term structure spread do not. There are two minor exceptions to this overall assessment on stationarity: First, in Japan, the null hypothesis of nonstationarity cannot be rejected for the long rate but, surprisingly, is rejected for the short rate. Given the mixed results, I chose to follow the same VAR analysis as in the remaining countries, namely, to use a VAR model of the first difference in short rates and the spread. Campbell and Hamao (1991), who study Japanese interest rates in greater detail for the sample period of the 1970s and 1980s, follow a similar procedure. Second, in the UK, the null hypothesis of nonstationarity of the term structure spread fails to be rejected at the 5% level. However, a closer inspection of the series indicates that the four lags in the Dickey–Fuller regression are too many. When $L = 1$, the Dickey–Fuller t-statistic becomes 2.83, which is significant at the 5% level. I thus treat the UK spread series as a stationary series.

Observe that the stationarity of the spread is consistent with the predictions of the expectations hypothesis. Given that the first difference of short rates is stationary, the stationarity of the spread is a theoretical implication of the

[5]The French short rate is the three-month eurofranc rate. An alternative short rate, the three-month French interbank offer rate, is less volatile but is not available prior to the mid-1970s. Use of the latter short rate provides a similar assessment on the expectations hypothesis.

Table 1

Data: Definitions and sources.

All interest rates are expressed in percentages. Short-term rates are transformed to bond equivalent yields, y, from discount yields, d, according to the formula: $y = (365\,d)/(360 - 0.91\,d)$. Most data series come from the data banks of the Bank for International Settlements (BIS), and to a lesser extent from the International Financial Statistics (IFS) and the Federal Reserve Bank of New York (FRBNY).

USA	The long-term rate is the yield to maturity of a ten-year Treasury bond; source: FRBNY; frequency: last business day of the quarter; begins 1953:2. The short-term rate is the three-month Treasury bill yield; source: FRBNY; frequency: last business day of the quarter; begins 1953:2.
Canada	The long-term rate is the average of government bond yields of maturity over ten years – whenever required, the tables assume a twelve-year maturity; source: BIS, series HHBACA01M; frequency: last Wednesday of the quarter; begins 1950:1. The short-term rate is the discount yield of a three-month T-bill at tender prices; source: BIS, series HEPACA01M; frequency: last Thursday of the quarter; begins 1950:1.
UK	The long-term rate is the yield to maturity of a ten-year government bond; source: BIS, series HHBAGB03M; frequency: last business day of the quarter, but prior to June 1989, last Friday of the quarter; begins 1961:1. The short-term rate is the discount rate on a three-month T-bill at tender prices; source: BIS, series HEPAGB01M; frequency: last Friday of the quarter; begins 1961:1.
Germany	The long-term rate is the yield to maturity on a ten-year government bond; source: BIS, series HHBADE01M; frequency: last business day of the quarter; begins 1967:1. The short-term rate is the three-month euro-deutschmark rate; source: BIS, series JFBADE01M; frequency: last business day of the quarter; begins 1963:2.
Japan	The long-term rate is the yield on interest bearing bonds issued by Nippon Telegraph and Telephone Public Corp. – whenever required, the tables assume a ten-year maturity; source: BIS, series HHEAJP07M; frequency: last business day of the quarter; begins 1961:4. The short-term rate is the one- to three-month call money rate; source: BIS, series HEBAJP01M; frequency: last business day of the quarter; begins 1961:4.
France	The long-term rate is the yield to maturity on public and semi-public bonds – whenever required, the tables assume a ten-year maturity; source: BIS, series HHEAFR01M; frequency: last business day of the quarter; begins 1955:1. The short-term rate is the three-month euro-franc rate; source: BIS, series JFBAFR01M; frequency: last business day of the quarter; begins 1968:1.
Italy	The long-term rate is the yield on bonds issued by Credit Consortium for Public Works – whenever required, the tables assume a maturity of ten years; source: BIS, series HHEAIT06M; frequency: last business day of the quarter; begins 1958:1. The short-term rate is the call money rate – the tables assume a maturity of three months; source: IFS, series $M13660B; frequency: last business day of the quarter; begins 1971:1.

Table 2

Summary statistics and correlation matrix of interest rates across the G7 countries.[a]

Sample: Quarterly, 1971:2–1992:2.

	USA	Canada	UK	Germany	Japan	France	Italy
Panel A: Summary statistics							
Mean(R_t)	9.04	10.26	11.56	7.96	7.43	11.12	12.69
$\sigma(R_t)$	2.28	2.16	2.08	1.25	1.81	2.37	3.81
$\sigma(R_t - R_{t-1})$	0.85	0.86	1.15	0.53	0.73	0.64	0.84
Mean(r_t)	7.74	9.88	10.98	6.45	6.97	11.88	14.08
$\sigma(r_t)$	2.76	3.63	2.90	2.78	2.58	5.04	4.53
$\sigma(r_t - r_{t-1})$	1.48	1.58	1.50	1.29	1.00	4.19	1.92
Mean$(R_t - r_t)$	1.31	0.39	0.58	1.51	0.45	− 0.76	− 1.39
$\sigma(R_t - r_t)$	1.44	2.02	2.29	2.14	1.37	3.68	2.45
Panel B: Cross-country correlations of the change in the long rate, $R_t - R_{t-1}$							
USA	1.00	0.90	0.26	0.60	0.21	0.47	0.17
Canada		1.00	0.34	0.61	0.28	0.51	0.28
UK			1.00	0.35	0.39	0.38	0.16
Germany				1.00	0.41	0.55	0.25
Japan					1.00	0.32	0.10
France						1.00	0.53
Italy							1.00
Panel C: Cross-country correlations of the change in the short rate, $r_t - r_{t-1}$							
USA	1.00	0.71	0.16	0.14	0.10	0.26	0.15
Canada		1.00	0.23	0.25	0.03	0.17	0.26
UK			1.00	0.25	0.33	0.15	0.17
Germany				1.00	0.33	0.14	0.18
Japan					1.00	0.16	0.17
France						1.00	0.35
Italy							1.00
Panel D: Cross-country correlations of the long-short rate spread, $R_t - r_t$							
USA	1.00	0.55	0.18	0.23	0.16	0.26	− 0.15
Canada		1.00	0.49	0.54	0.10	0.22	0.15
UK			1.00	0.53	0.29	− 0.07	0.28
Germany				1.00	0.34	0.32	0.29
Japan					1.00	0.23	0.41
France						1.00	0.34
Italy							1.00

[a] R_t denotes the long-term (ten-year) and r_t the short-term (three-month) end-of-quarter bond equivalent yield in percent. σ denotes standard deviation. The beginning of the sample, 1971:2, is dictated by the availability of Italian short-term interest rate data.

Table 3

Unit root tests.[a]

	USA 53:2–92:2	Canada 50:1–92:2	UK 61:1–92:2	Germany 67:1–92:2	Japan 61:4–92:2	France 68:1–92:2	Italy 71:1–92:2
R_t	− 1.53	− 1.39	− 1.79	− 2.70	− 2.50	− 1.58	− 2.04
$R_t - R_{t-1}$	− 6.09[b]	− 6.56[b]	− 4.16[b]	− 3.92[b]	− 4.09[b]	− 3.86[b]	− 3.33[b]
r_t	− 1.86	− 1.93	− 2.52	− 2.59	− 4.01[b]	− 2.67	− 2.33
$r_t - r_{t-1}$	− 4.45[b]	− 5.45[b]	− 4.91[b]	− 4.37[b]	− 5.33[b]	− 5.61[b]	− 5.49[b]
$R_t - r_t$	− 2.86[b]	− 3.15[b]	− 2.76	− 2.90[b]	− 5.94[b]	− 3.98[b]	− 3.24[b]

[a] The table presents Dickey–Fuller t-statistics of the hypothesis H_0: $\alpha_1 = 0$ in regressions of the form $Y_t = \alpha_0 + \alpha_1 X_{t-1} + \gamma_1 Y_{t-1} + \ldots + \gamma_L Y_{t-L} + u_t$, where $Y_t \equiv X_t - X_{t-1}$ and X_t represents each of the five variables. The sample is quarterly. R_t and r_t denote the long-term (ten-year) and short-term (three-month) end-of-quarter bond equivalent yields and are defined in table 1. The number of lags L equals either 4, 8, or 12 depending on the serial correlation properties of the residuals.

[b] Statistically significant at the 5% level.

expectations hypothesis because the spread is a linear combination of stationary variables (see section 4).

3. The one-period-ahead change in long and short rates

3.1. Theoretical framework

I begin by decomposing the yield to maturity of a bond that matures in N quarters, $R_t^{[N]}$, into two components: a weighted average of expected future yields to maturity of bonds that mature in one quarter, $E_t r_{t+i}$, and a risk premium, $E_t \theta_{N,t}$, the so-called rolling premium:

$$R_t^{[N]} \equiv E_t \theta_{N,t} + \sum_{i=0}^{N-1} w_i E_t r_{t+i}, \tag{1a}$$

$$w_i \equiv g^i (1-g)/(1-g^N), \qquad g \equiv 1/(1 + Rbar),$$

$$E_t \theta_{N,t} \equiv \sum_{i=0}^{N-1} w_i E_t \phi_{N-i,t+i},$$

where E_t denotes expectations conditional on information available at time t; $Rbar$ is the sample mean of $R_t^{[N]}$; and $E_t \theta_{N,t}$, the rolling premium, is a weighted average of the expected one-period holding premia, $E_t \phi_{N-i,t+i}$. Eq. (1a) is a linear approximation to a nonlinear relation between the long rate and short rates and was originally proposed by Shiller (1979), Shiller, Campbell, and

Schoenholtz (1983), and Campbell and Shiller (1984). These authors show that the approximation error is very small and of no consequence in evaluating the expectations hypothesis [see also Campbell (1986) for the analogous argument in continuous time]. Eq. (1a) is an identity and does not impose any restrictions on the data. The expectations hypothesis imposes the restriction that the rolling premium $E_t\theta_{N,t}$ or the sequence of holding premia $E_t\phi_{N-i,t+i}$ are constant.

Observe that the weights w_i sum up to unity and decline monotonically with the time horizon i: Expected short rates or holding premia of the near future carry a larger weight than expected short rates or holding premia of the distant future. This is a consequence of the fact that the long-term bond carries coupons and coupons of the near future have a higher present value than coupons of the distant future. In terms of Macauley's duration, a bond that carries coupons has a duration which is lower than its maturity. Shiller, Campbell, and Schoenholtz (1983) show that the duration of an N-period bond, D_N, approximately equals $(1-g^N)/(1-g)$. Using the definition of duration, eq. (1a) can be rewritten in a more convenient way as follows:

$$R_t^{[N]} \equiv (1/D_N)\sum_{i=0}^{N-1}(D_{i+1}-D_i)\,E_t(r_{t+i}+\phi_{N-i,t+i}), \tag{1b}$$

$$D_i \equiv (1-g^i)/(1-g), \qquad D_{i+1}-D_i = g^i.$$

Eq. (1b) can be used to derive the implications of the expectations hypothesis about the predicted change in the long rate over the quarter. Rewriting eq. (1b) for the yield to maturity of an $(N-1)$-period bond at $t+1$, we get

$$R_{t+1}^{[N-1]} = (1/D_{N-1})\sum_{i=0}^{N-2}(D_{i+1}-D_i)\,E_{t+1}[r_{t+1+i}+\phi_{N-2-i,t+i}]. \tag{2}$$

Multiplying eq. (1b) by the factor D_N, multiplying eq. (2) by the factor $D_N - 1 = g\,D_{N-1}$, and taking the difference between the two sets of products, we get

$$D_N R_t^{[N]} - (D_N-1)\,R_{t+1}^{[N-1]} = r_t + E_t\phi_{N,t} - \varepsilon_{t+1}, \tag{3}$$

$$\begin{aligned}\varepsilon_{t+1} \equiv\ & g(r_{t+1}-E_t r_{t+1}) \\ & + g^2(E_{t+1}r_{t+2}-E_t r_{t+2}) + \ \dots\ + g^{N-1}(E_{t+1}r_{t+N-1}-E_t r_{t+N-1}) \\ & + g(E_{t+1}\phi_{N-1,t+1} - E_t\phi_{N-1,t+1}) \\ & + g^2(E_{t+1}\phi_{N-2,t+2} - E_t\phi_{N-2,t+2}) \\ & + \ \dots\ + g^{N-1}(E_{t+1}\phi_{1,t+N-1} - E_t\phi_{1,t+N-1}).\end{aligned}$$

The left-hand side of eq. (3) is the holding period return of a bond of maturity N held for one period. The right-hand side of eq. (3) is composed of three parts: The first two are the short-term rate, r_t, and the holding premium, $\mathrm{E}_t\phi_{N,t}$, and their sum represents the expected holding period return. The third is an error term, $-\varepsilon_{t+1}$, which represents the unexpected capital loss (gain) from upward (downward) revisions made from period t to period $t+1$ in the market expectations of future short rates and/or holding premia. Under the maintained hypothesis of *rational expectations*, those revisions are unpredictable at time t and, thus, ε_{t+1} can be interpreted as a white noise error term. Eq. (3) represents one of the forms used in testing the expectations hypothesis. The expectations hypothesis claims that the holding premium is constant, and thus a regression of the holding period return on the level of the short rate at the beginning of the holding period should produce a slope coefficient that is equal to unity.

We can now transform eq. (3) to derive the one-period change in the long rate. Subtracting $R_t^{[N]}$ from both sides of eq. (3) and then multiplying both sides by the factor $-1/(D_N-1)$, we derive the following relation:

$$R_{t+1}^{[N-1]} - R_t^{[N]} = (1/(D_N-1))\left[(R_t^{[N]} - r_t) - \mathrm{E}_t\phi_{N,t} + \varepsilon_{t+1}\right]. \tag{4}$$

If the expectations hypothesis holds, the holding premium $\mathrm{E}_t\phi_{N,t}$ is constant and changes in the long rate reflect, apart from noise, previous changes in the term structure spread, $R_t - r_t$.[6]

3.2. OLS results

The empirical counterpart of eq. (4) is

$$R_{t+1}^{[N-1]} - R_t^{[N]} = \alpha + \beta[1/(D_N-1)](R_t^{[N]} - r_t) + u_{t+1}. \tag{5}$$

The expectations hypothesis claims that the estimated slope coefficient, β, equals unity, $\beta = 1$. In practice, holding premia do vary and thus the estimated slope coefficient is different from unity. In fact, investigators have found not only that β is less than unity, but that it is negative. The negative β is a major piece of the puzzle.

[6]Observe that as the maturity N increases towards infinity, the long-term bond becomes a consol and eq. (4) simplifies to the following equation:

$$\text{as}\quad N \to \infty,\quad R_{t+1} - R_t = (1/g)(1-g)(R_t - r_t) - (1/g)(1-g)\mathrm{E}_t\phi_{N,t} + (1/g)(1-g)\varepsilon_{t+1}.$$

This is the equation presented in Mankiw and Summers (1984). Mankiw and Summers do not use consol yields in their empirical work but bonds of finite maturity. Thus their equation is an approximation to the more precise eq. (4).

Eq. (5) requires data on both an N-quarter and an $(N - 1)$-quarter bond. Such fine data are not generally available. We, therefore, treat yields of the two bonds with consecutive maturities $N - 1$ and N as equal: $R_t^{[N-1]} = R_t^{[N]}$. We denote this yield as R_t.

Panel A of table 4a presents the OLS results of eq. (5). Besides the United States, four additional countries show a counterintuitive negative response: Canada, the UK, Germany, and Japan. In three of the five countries with a negative β, the restriction that $\beta = 1$ is rejected at the 5% level, but only in the United States β is statistically different from zero. This evidence is consistent with the findings of Mankiw (1986) who examined the first four countries in the table for a sample period that ended in 1984.

Panel B of table 4a presents the OLS results for the change in the short rate. Observe that in contrast to the negative coefficients of long rates, here the regression coefficients δ are positive and, in five of the seven countries, statistically significant as well. Of course, the size of δ and its distance from unity provide no information on the question of how closely the behavior of short-term rates conforms to the predictions of the expectations hypothesis. To examine this question, we need to trace the evolution of short rates over multiperiod horizons.

3.3. An additive white noise component on long rates?

Can a simple white noise deviation of long rates from their theoretically correct value – the value R_t' predicted by the expectations hypothesis – be responsible for the presence of the estimated negative correlation between changes in long rates and the earlier term structure spread? Such a deviation could be due to temporary mistakes that the market makes or it could be due to a simple econometric measurement error. After all, these yields may be based on thinly traded markets and have undergone some manipulation by the economic agencies responsible for their publication in the different countries. Whatever the source of the error, if a white noise error ε_t is present in the long rate, then $R_{t+1} - R_t = R_{t+1}' - R_t' + (\varepsilon_{t+1} - \varepsilon_t)$ and $R_t - r_t = R_t' - r_t + \varepsilon_t$, where R' denotes the theoretical but unobserved value of the long rate. In this case, the estimated slope coefficient would contain a bias term equal to

$$(D_N - 1)\operatorname{cov}(\varepsilon_{t+1} - \varepsilon_t, \varepsilon_t)/\operatorname{var}(R_t - r_t) = -(D_N - 1)(\sigma_\varepsilon^2/\sigma_{R-r}{}^2). \qquad (6)$$

Panel C of table 4a presents the required size of the standard deviation of the white noise error, $\sigma_1(\varepsilon_t)$, in order to account for the discrepancy of the estimated slope coefficient β from unity. This standard deviation varies from 35 basis points in Japan to 59 basis points in Germany. Mankiw (1986), who first posed this question, views ε_t strictly as an econometric measurement error and argues

Table 4a

The spread as a predictor of the subsequent change in the long rate and the short rate.[a]

OLS estimates

$$R_{t+1} - R_t = \alpha + \beta[1/(D_N - 1)](R_t - r_t) + u_{t+1}, \quad D_N \equiv (1 - g^N)/(1 - g), \quad g \equiv 1/(1 + Rbar)$$

$$r_{t+1} - r_t = \alpha + \delta(R_t - r_t) + v_{t+1}$$

	USA 54:4–92:2	Canada 51:3–92:2	UK 62:3–92:2	Germany 68:2–92:2	Japan 63:2–92:2	France 69:3–92:2	Italy 72:3–92:2
			Panel A: Long rates				
β	-2.901^b	-1.071	-0.576	-1.229	-0.633	0.552	0.067
	(1.308)	(0.955)	(1.142)	(0.683)	(1.144)	(0.411)	(0.922)
$t(\beta=1)$	-2.98^b	-2.17^b	-1.38	-3.26^b	-1.43	-1.09	-1.01
R^2	0.034	0.001	0.002	0.033	0.003	0.020	0.000
			Panel B: Short rates				
δ	0.153	0.134^b	0.126^b	0.073	0.253^b	0.589^b	0.275^b
	(0.082)	(0.058)	(0.060)	(0.057)	(0.069)	(0.098)	(0.090)
R^2	0.023	0.031	0.036	0.025	0.105	0.034	0.108
			Panel C: White noise error on long rates?				
$\sigma(R_t - r_t)$	1.18	1.63	1.96	2.05	1.43	3.71	2.36
$1/(D_N - 1)$	0.0354	0.0323	0.0405	0.0369	0.0371	0.0419	0.0450
$\sigma_1(\varepsilon_t)$	0.44	0.42	0.50	0.59	0.35	0.51	0.48
$\sigma_0(\varepsilon_t)$	0.38	0.31	0.30	0.44	0.22	—	—
$\sigma_{R'-r}/\sigma_{R-r}$	0.93	0.97	0.97	0.96	0.97	0.99	0.98

[a] The sample is quarterly. R_t and r_t denote the end-of-quarter long-term (ten-year) and short-term (three-month) bond equivalent yields in percent. D_N is the duration of the long-term (N-quarter) bond and $Rbar$ is the sample mean of R_t. The factor $1/(D_N - 1)$ adjusts the long-short rate spread so that the Expectations Hypothesis predicts that: $\beta = 1$. σ denotes standard deviation and $t(\beta = 1)$ a t-statistic for the hypothesis $\beta = 1$. R^2 is the coefficient of determination. Numbers in parentheses below the regression coefficients are OLS standard errors.

$\sigma_1(\varepsilon_t) \equiv [(1-\beta)/(D_N - 1)]^{0.5}\sigma_{R-r}$ and $\sigma_0(\varepsilon_t) \equiv [(0-\beta)/(D_N - 1)]^{0.5}\sigma_{R-r}$ are the standard deviations of a hypothetical white noise error ε_t in R_t required to explain the observed discrepancy of the OLS β from unity and zero, respectively. $\sigma_{R'-r}/\sigma_{R-r} \equiv [\sigma^2_{R-r} - \sigma^2_1(\varepsilon_t)]^{0.5}/\sigma_{R-r}$ is the ratio of the standard deviations of the theoretical and actual spreads implied by $\sigma_1(\varepsilon_t)$.

[b] Statistically significant at the 5% level.

that its size is unreasonable. Yet, such an error does not have to be exclusively an econometric measurement error; it can represent a true deviation from the expectations hypothesis caused by a market mistake. Moreover, recall that the puzzle, as posed by Mankiw (1986) or Mankiw and Summers (1984), is not why long rates deviate from the exact predictions of the expectations hypothesis. Rather, the puzzle is why long rates move in the opposite direction from the one predicted. The white noise error required to explain the negative slope

Table 4b

The spread as a predictor of the subsequent change in the long rate and the short rate.[a]
Instrumental variables estimates

$$R_{t+1} - R_t = \alpha + \beta[1/(D_N - 1)](R_t - r_t) + u_{t+1}, \quad D_N \equiv (1 - g^N)/(1 - g), \quad g \equiv 1/(1 + Rbar)$$

$$r_{t+1} - r_t = \alpha + \delta(R_t - r_t) + v_{t+1}$$

	USA 54:4–92:2	Canada 51:3–92:2	UK 62:3–92:2	Germany 68:2–92:2	Japan 63:2–92:2	France 69:3–92:2	Italy 72:3–92:2
			Panel A: Long rates				
β	−2.803	0.098	0.386	0.113	0.658	0.069	0.056
	(1.691)	(1.110)	(1.312)	(0.785)	(1.420)	(0.661)	(1.163)
$t(\beta = 1)$	−2.25[b]	−0.99	−0.47	−1.13	−0.24	−1.41	−0.81
			Panel B: Short rates				
δ	0.141	0.248[b]	0.158[b]	0.165[b]	0.297[b]	0.436[b]	0.389[b]
	(0.106)	(0.068)	(0.069)	(0.072)	(0.085)	(0.158)	(0.114)
			Panel C: Systematic overreaction?				
k	0.96	0.12	0.16	0.20	0.04	0.09	0.11
$\sigma_{R'-r}/\sigma_{R-r}$	0.51	0.89	0.86	0.83	0.96	0.92	0.90

[a] The sample is quarterly. Eight instruments are used: four lags of the spread and four lags of the change in the short rate. R_t and r_t denote the end-of-quarter long-term (ten-year) and short-term (three-month) bond equivalent yields in percent. D_N is the duration of the long-term (N-quarter) bond and *Rbar* is the sample mean of R_t. The factor $1/(D_N - 1)$ adjusts the long-short rate spread so that the Expectations Hypothesis predicts that: $\beta = 1$. σ denotes standard deviation and $t(\beta = 1)$ a t-statistic for the hypothesis $\beta = 1$. Numbers in parentheses below the IV estimates are standard errors.

$k \equiv [(1 - \beta)/\delta]/(D_N - 1)$ denotes the implied degree of market overreaction required to explain the difference of parameter β from unity. $\sigma_{R'-r}/\sigma_{R-r} \equiv 1/(1 + k)$ is the ratio of the standard deviations of the theoretical and actual spreads implied by the overreaction parameter k.

[b] Statistically significant at the 5% level.

coefficients in table 4a in five of the countries, $\sigma_0(\varepsilon_t)$, is smaller. It varies from 13 basis points in Italy to 44 basis points for Germany. Many would still argue, of course, that such an error is too large to be interpreted as a measurement error. In the case of Germany, for example, a 44 basis points error implies that there is a 10 percent probability of an error larger than 73 basis points in either direction. Thus, if a white noise error is responsible for the negative coefficient, it probably reflects both a measurement error and a random deviation of the true yield from the exact predictions of the expectations hypothesis.

A white noise error on long rates, of course, does not necessarily destroy the overall ability of the expectations hypothesis to describe the time series relation between long and short rates. To see this point, observe that the size of $\sigma_1(\varepsilon_t)$ has almost no influence on the size of the ratio of the standard deviation of the theoretical spread to the standard deviation of the actual spread, $\sigma_{R'-r}/\sigma_{R-r} \equiv [\sigma_{R-r}^2 - \sigma_1^2(\varepsilon_t)]^{1/2}/\sigma_{R-r}$. Panel C of table 4a shows that this ratio is very close to unity. Its lowest value of 0.93 occurs in the US data. The ratio $\sigma_{R'-r}/\sigma_{R-r}$ is a metric used by Campbell and Shiller (1987, 1991) to assess the economic relevance of the expectations hypothesis and we will analyze it later in section 4.

To avoid the possible bias that a white noise error on long rates would generate, I reestimate eq. (5) using instrumental variables. The instruments are lags of the spread and lags of the change in the short rate. Table 5 later shows that those lagged variables are highly correlated with the spread and thus are good instruments. The instrumental variables estimates of the slope coefficients β are presented in table 4b. Only in the United States long rates continue to show a negative response and only in the United States β is significantly different from unity. It appears, therefore, that the observed anomaly in four of the six foreign countries is due entirely to an additive white noise error. However, an additive white noise error cannot explain the anomaly in the US data.[7]

Panel B of table 4b presents the instrumental variables estimates of the regression slope δ of short rates. These estimates become slightly stronger than the previous OLS estimates in five of the seven countries. The stronger IV estimates are encouraging for an additional reason: They suggest that a white noise error is not present on short rates. A white noise error on short rates would have biased the OLS estimates of δ in the positive direction and we would have observed a drop in the magnitude of δ in the IV estimates. The fact that the IV estimates of δ are more positive than the OLS estimates of δ suggests that the white noise error is on the long rates.

3.4. *An additive fads component on long rates?*

Since a simple additive white noise error on long rates cannot explain the counterintuitive negative correlation of long rates in the US, the question arises whether a more persistent error component, or a 'fads' component, could account for those results. Let R_t' denote the theoretical long rate that would obtain if the expectations hypothesis were true, and let e_t denote the deviation of the actual long rate R_t from R_t': $R_t \equiv R_t' + e_t$. Suppose next that the error component e_t is persistent, say it has an autoregressive structure of the form: $e_t = \rho e_{t-1} + \varepsilon_t$. Then the bias in the estimated slope coefficient β of eq. (5) would be equal to

$$-(1-\rho)(D_N - 1)(\sigma_\varepsilon^2/\sigma_{R-r}^2). \tag{7}$$

[7]Mankiw (1986) does not present the instrumental variables estimates for the equivalent of eq. (5), so it is hard to make direct comparisons. Nevertheless, the standard errors in his regressions are quite large to be able to reject the expectations hypothesis. Among Canada, Germany, the UK, and the US, Mankiw found significant deviations from the expectations hypothesis only in the US.

It is now much harder to explain the estimated negative coefficients than it was earlier with a simple white noise error. Suppose, for example, that $\rho = 0.9$. The required size of σ_ε^2 (or σ_ε) would then be 10 (or 3.16) times the size required to explain the bias in the earlier eq. (6).

Thus, although at first glance a fads component would appear to be more promising in explaining the puzzle since it represents a much more serious violation of the expectations hypothesis than a simple white noise deviation, in reality a fads component provides a much weaker explanation to the puzzle. The reason is straightforward: To explain the puzzle one has to provide a story of an immediate reversal in long rates. Fads act in the opposite manner: they prolong rather than reverse extraneous shocks to long rates.

3.5. *An overreaction of the term structure spread?*

I now turn to the explanation I mentioned in the introduction, namely an overreaction (underreaction) of long rates to expected future short rates (current short rate). This explanation appeals to a much more serious violation of the expectations hypothesis than the earlier two explanations for it assumes that there is a *systematic* overreaction in the market, or a multiplicative error term on long rates rather than an additive one. It is not accidental, perhaps, that the explanation has been proposed as an explanation of the US data. It is the US data that cannot be explained by a simple white noise error.

To make the spread overreaction hypothesis concrete, set the risk premium to zero in eq. (1b) and write the actual term structure spread, $R_t - r_t$, as follows:

$$R_t^{[N]} - r_t \equiv (1/D_N) \sum_{i=0}^{N-1} g^i (1+k)(\mathrm{E}_t r_{t+i} - r_t) = (1+k)(R_t^{[N]\prime} - r_t), \quad (8)$$

where $R_t^{[N]\prime}$ denotes the theoretical long rate, namely the rate that would prevail if the expectations hypothesis held exactly, and $k > 0$ denotes the degree of overreaction. This is the formulation proposed by Campbell and Shiller (1991) and can be interpreted as a violation of the rational expectations hypothesis: The market expectation of the future total change in the level of short rates, $\mathrm{E}_t^M r_{t+i} - r_t$, equals $1+k$ times the rationally expected change, $(1+k)(\mathrm{E}_t r_{t+i} - r_t)$. Observe that eq. (8) accommodates the notion that long rates either overreact to expected future short rates or underreact to the current short rate.

Eq. (8) implies that the change in the long rate over the quarter is as follows:

$$R_{t+1}^{[N-1]} - R_t^{[N]} = (1/(D_N - 1))[(R_t^{[N]} - r_t) + (1+k)\varepsilon_{t+1}] - k(r_{t+1} - r_t). \quad (9)$$

Thus, under the proposed hypothesis of overreaction, the probability limit of regression coefficient β in eq. (5) equals

$$\begin{aligned}\operatorname{plim}(\beta) &= 1 - k(D_N - 1)\operatorname{cov}(R_t - r_t, r_{t+1} - r_t)/\operatorname{var}(R_t - r_t) \\ &= 1 - [k(D_N - 1)]\operatorname{plim}(\delta),\end{aligned} \tag{10}$$

where δ denotes the regression coefficient in panel B of table 4b of the change in the short rate on the spread. The required overreaction parameter k can, therefore, be identified from the instrumental variable estimates of β and δ of table 4b and the magnitude of duration, D_N, as follows: $k = [1 - \operatorname{plim}(\beta)]/[D_N - 1)\operatorname{plim}(\delta)]$. The presence of k also implies that the ratio of the standard deviation of the theoretical spread to the standard deviation of the actual spread, $\sigma_{R'-r}/\sigma_{R-r}$, equals $1/(1+k)$.

Panel C of table 4b uses the instrumental variable estimates of β and δ to calculate the overreaction parameter k and the ratio of the standard deviation of the theoretical spread to the standard deviation of the actual spread, $\sigma_{R'-r}/\sigma_{R-r}$. These values of k or $\sigma_{R'-r}/\sigma_{R-r}$ are the magnitudes of overreaction required to explain the change in long and short rates that follow prior movements in the term structure spread. Observe that the US results are explained by an overreaction of 96 percent, implying a ratio $\sigma_{R'-r}/\sigma_{R-r}$ of approximately 1/2. Later in section 4, I will estimate this ratio from a VAR model and examine whether or not its implied ratio conforms with the required ratio of the present table. Indeed, I do find a ratio close to 1/2.

3.6. Time-varying risk premia?

I now turn to the second hypothesis we discussed in the introduction, a time-varying risk premium. Previous investigators, including Fama (1984), Mankiw and Miron (1986), and Hardouvelis (1988), show that a time-varying risk premium may destroy the predictive ability of the spread. To see this, observe that under the maintained assumption of rational expectations, the probability limit of the estimated slope coefficient in eq. (5) is

$$\begin{aligned}&\operatorname{plim}\beta \\ &= \frac{\operatorname{var}(\mathrm{E}_t[R_{t+1}^{[N-1]} - R_t^{[N]}]) + \operatorname{cov}(\mathrm{E}_t\phi_{N,t}, \mathrm{E}_t[R_{t+1}^{[N-1]} - R_t^{[N]}])}{\operatorname{var}(\mathrm{E}_t[R_{t+1}^{[N-1]} - R_t^{[N]}]) + 2\operatorname{cov}(\mathrm{E}_t\phi_{N,t}, \mathrm{E}_t[R_{t+1}^{[N-1]} - R_t^{[N]}]) + \operatorname{var}(\mathrm{E}_t\phi_{N,t})} \\ &= (1+\rho q)/(1+2\rho q+q^2) = 1 - q(\rho+q)/(1+2\rho q+q^2),\end{aligned} \tag{11}$$

$$\rho \equiv \operatorname{corr}(\mathrm{E}_t\phi_{N,t}, \mathrm{E}_t[R_{t+1}^{[N-1]} - R_t^{[N]}]),$$

$$q \equiv [\operatorname{var}(\mathrm{E}_t\phi_{N,t})/(\operatorname{var}(\mathrm{E}_t(R_{t+1}^{[N-1]} - R_t^{[N]}))]^{1/2}.$$

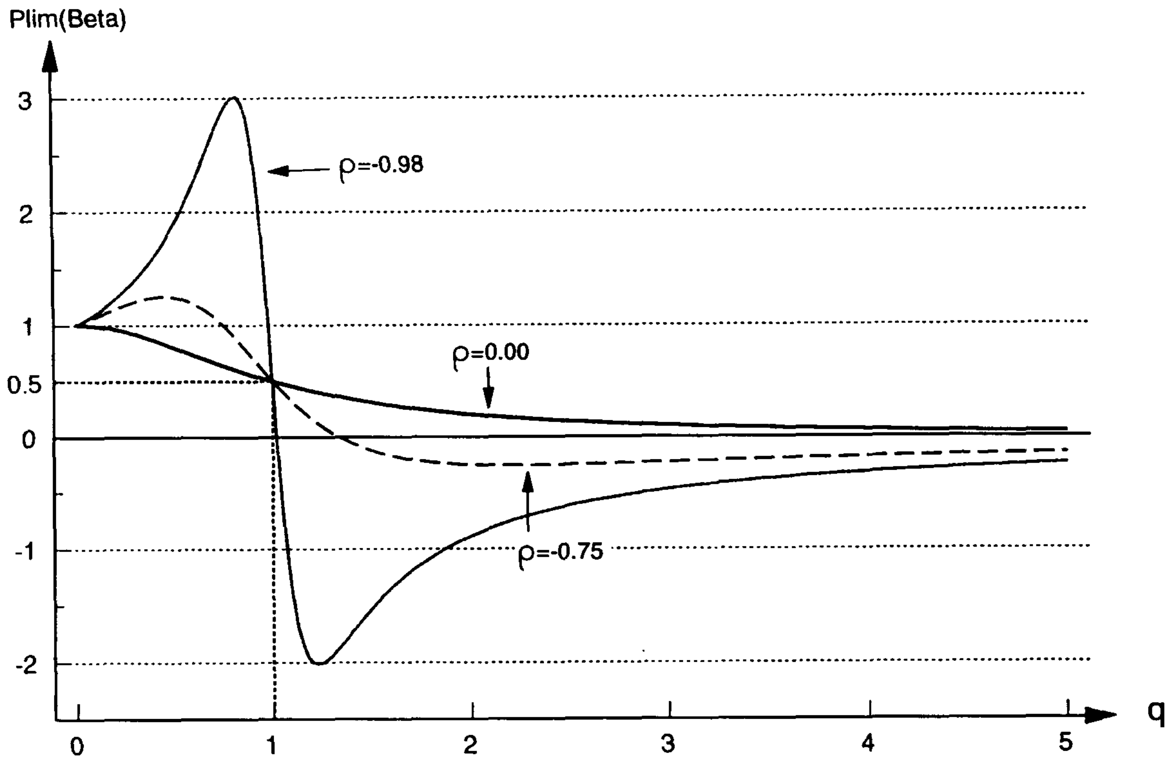

Fig. 1. The regression slope as a function of the relative variability of the risk premium.

$\mathrm{plim}(\beta) = 1 - q(\rho + q)/(1 + 2\rho q + q^2)$, where β is the slope coefficient of the regression equation: $R_{t+1} - R_t = \alpha + \beta[1/(D_N - 1)](R_t - r_t) + u_{t+1}$; q denotes the relative variability of the risk premium and equals the ratio of the standard deviation of the expected risk premium to the standard deviation of the expected change in the long rate; ρ is the correlation between the expected risk premium and the expected change in the long rate.

Fig. 1 plots plim β for various values of ρ, the correlation between the expected change in the long rate and the risk premium, and q, the relative variability of the risk premium. The figure shows that under certain values of q and ρ, β is negative. Observe that values of q larger, say, than 5 cannot accommodate the puzzle because they imply β coefficients that are very close to zero.

It turns out that the volatility of the risk premium relative to the volatility of the expected change in long rates is way too high to accommodate the puzzle. In table 4c, I construct estimates of $E_t R_{t+1} - R_t$ and $E_t \phi_{N,t}$ from regression models that include four lags of the term structure spread and four lags of the change in the short rate as explanatory variables. When the dependent variable is the change in the long rate, $R_{t+1} - R_t$, then the regression fit becomes a proxy for $E_t R_{t+1} - R_t$. When the dependent variable is – from eq. (3) – the realized excess

Table 4c

Can time-varying risk premia explain the puzzle?[a]

Eq. 4c.1: $D_N R_t - (D_N - 1) R_{t+1} - r_t = \alpha_0 + \sum_{j=1}^{4} \alpha_j (R_{t-j+1} - r_{t-j+1}) + \sum_{j=1}^{4} \theta_j (r_{t-j+1} - r_{t-j}) + \varepsilon_t$

Eq. 4c.2: $R_{t+1} - R_t = \gamma_0 + \sum_{j=1}^{4} \gamma_j (R_{t-j+1} - r_{t-j+1}) + \sum_{j=1}^{4} \delta_j (r_{t-j+1} - r_{t-j}) + u_t$

	USA	Canada	UK	Germany	Japan	France	Italy
	54:4–92:2	51:2–92:2	62:2–92:2	68:1–92:2	63:1–92:2	69:2–92:2	72:2–92:2
				Equation 4c.1			
$R^2_{E\phi}$	0.120	0.147	0.090	0.095	0.150	0.128	0.145
$\sigma_{E\phi}$	6.73	7.66	7.33	4.33	6.91	5.25	7.34
				Equation 4c.2			
R^2_{EdR}	0.098	0.128	0.077	0.064	0.136	0.126	0.132
σ_{EdR}	0.212	0.229	0.273	0.129	0.242	0.216	0.312
				Implied parameter values			
q	31.7	33.4	26.8	33.5	28.6	24.3	23.5
ρ	– 0.990	– 0.979	– 0.965	– 0.882	– 0.979	– 0.720	– 0.946
β	– 0.032	– 0.030	– 0.037	– 0.027	– 0.035	– 0.030	– 0.042

[a] The sample is quarterly. R_t and r_t denote the end-of-quarter long-term (ten-year) and short-term (three-month) bond equivalent yields in percent. $D_N \equiv (1 - g^N)/(1 - g)$, with $g \equiv 1/(1 + Rbar)$ the duration of the long-term (N-quarter) bond, and $Rbar$ is the sample mean of R_t. The dependent variable in eq. 4c.1, $D_N R_t - (D_N - 1) R_{t+1} - r_t$, is the excess holding period return of the long-term bond over the quarter.

$R^2_{E\phi}$ and R^2_{EdR} are the R^2s of eqs. 4c.1 and 4c.2, respectively; $\sigma_{E\phi}$ and σ_{EdR} are the standard deviations of the regression fits of eqs. 4c.1 and 4c.2. ρ is the sample correlation between the two regression fits. $q \equiv \sigma_{E\phi}/\sigma_{EdR}$. $\beta \equiv (1 + \rho q)/(1 + 2\rho q + q^2)$ is an implied value for the regression slope coefficient β of tables 4a and 4b.

holding period return, $D_N R_t - (D_N - 1) R_{t+1} - r_t$, then the regression fit becomes our proxy for $E_t\phi_{N,t}$. The ratio of the standard deviations of the regression fits is a proxy for q. Estimates of q are very high and result in implied β estimates of almost zero.[8]

3.7. Summary

Overall, the puzzle in foreign countries appears to be the result of an additive white noise error on the long rate that represents only a minor deviation from the predictions of the expectations hypothesis. The puzzle in the US presents a much more serious serious challenge to the expectations hypothesis: the only plausible alternative that can explain the significant negative response of long rates is the spread overreaction hypothesis.

4. The relation of the term structure spread with its theoretical counterpart under the expectations hypothesis

This section adopts Campbell and Shiller's (1991) VAR methodology to estimate the theoretical term structure spread under the expectations hypothesis and compare it to the actual spread. Given the evidence in section 3, the present section addresses the following major questions of interest: (1) Is the actual term structure spread in the United States twice as volatile as the theoretical spread? (2) Does the actual term structure spread in the remaining G7 countries have the same volatility as the theoretical spread?

Following Campbell and Shiller (1991) and the evidence in table 3, the process generating long- and short-term interest rates is modeled by a vector autoregression with four lags and two variables: The change in the short rate, Δr_t, and the spread between the long and the short rate, $S_t \equiv R_t - r_t$. The VAR system can be written in the companion form: $z_t = Az_{t-1} + u_t$, where $z_t \equiv [\Delta r_t, \dots, \Delta r_{t-3}, S_t, \dots, S_{t-3}]'$ and A is an 8×8 matrix defined appropriately using the coefficients of the original VAR. In the VAR system all variables are in deviations from their respective means, so that constant terms do not appear in the system. Multiperiod VAR forecasts can now be computed easily: $E[z_{t+K} | x_t, x_{t-1}, \dots] = A^K z_t$. The VAR forecast of the change in the short-term rate K periods ahead, $E[\Delta r_{t+K} | x_t, x_{t-1}, \dots]$, is the first element of the vector $A^K z_t$. This first element will be denoted by $h_1' A^K z_t$, where h_1' is the row vector $[1, 0, 0, 0, 0, 0, 0, 0]$. Also, the spread S_t equals $h_5' z_t$, where h_5' denotes the row vector $[0, 0, 0, 0, 1, 0, 0, 0]$.

[8]The standard deviations of $E_t\phi_{N,t}$ are similar to the ones found by Mankiw (1986), who examined the first four countries of the table for a slightly different sample period.

To see the implications of the expectations hypothesis on the parameters of the VAR, let us rewrite eq. (1a) as follows:

$$\sum_{i=0}^{N-1} w_i \mathrm{E}_t r_{t+i} - r_t \equiv R_t^{[N]} - r_t - \mathrm{E}_t \theta_t. \tag{12}$$

Eq. (12) can be written in terms of the stationary expected successive changes in the short-term interest rate, $\mathrm{E}_t \Delta r_{t+i}$, as follows:

$$\sum_{i=1}^{N-1} \{g^i/(1-g^N) - g^N/(1-g^N)\} \, \mathrm{E}_t \Delta r_{t+i} \equiv R_t^{[N]} - r_t - \mathrm{E}_t \theta_t. \tag{13}$$

Imposing the expectations hypothesis $\mathrm{E}_t \theta_t = 0$ on eq. (13) and using the VAR forecasting relation $\mathrm{E}_t \Delta r_{t+i} = h_1' A^i z_t$, we get

$$S_t' \equiv \sum_{i=1}^{N-1} \{g^i/(1-g^N) - g^N/(1-g^N)\} \, h_1' A^i z_t = h_5' z_t \equiv S_t, \tag{14}$$

where S_t' denotes what Campbell and Shiller call the 'theoretical spread', namely, the theoretical size of the spread under the expectations hypothesis and the VAR framework. In our formulation, the theoretical spread is slightly different from the one in Campbell and Shiller (1991) because we deal with coupon carrying bonds.

The restrictions imposed by the expectations hypothesis can be easily seen from examining eq. (14). Since (14) holds for every z_t,

$$\sum_{i=1}^{N-1} \{g^i/(1-g^N) - g^N/(1-g^N)\} \, h_1' A^i = h_5'. \tag{15}$$

This is a set of eight nonlinear restrictions on the coefficients of the VAR.

The purpose of the VAR methodology, of course, is not simply to test the expectations hypothesis, something that can be done in a simpler way through multiperiod regressions. The main purpose is to compare the behavior of S_t' with S_t over time and assess the *economic* relevance of their differences. We examine the correlation $\rho_{s',s}$ between S_t' and S_t, the ratio of their standard deviations, $\sigma_{s'}/\sigma_s$, as well as the regression coefficient $\gamma \equiv \rho_{s',s}\sigma_{s'}/\sigma_s$. Under the expectations hypothesis, all three quantities should equal unity. Under the spread overreaction hypothesis, the ratios $\sigma_{s'}/\sigma_s$ are less than unity.

Table 5 summarizes the results. Observe that the VAR models capture a lot of the variation in the two variables, as evidenced by the regression R^2s. The standard errors which appear in parentheses below the reported parameters do not treat the theoretical spread as data, but allow for the fact that S' is an estimate and thus is measured with error. The reported standard errors are

Table 5

The relation of the spread with its theoretical counterpart under the expectations hypothesis – a VAR methodology.[a]

	USA	Canada	UK	Germany	Japan	France	Italy
	54:3–92:2	51:2–92:2	62:2–92:2	68:1–92:2	63:1–92:2	69:2–92:2	72:2–92:2
$R^2_{\mathrm{VAR},\Delta r}$	0.168	0.210	0.077	0.206	0.463	0.463	0.224
$R^2_{\mathrm{VAR},S}$	0.607	0.744	0.764	0.769	0.656	0.393	0.630
γ	0.249	0.910[b]	0.829[b]	1.078[b]	1.199[b]	1.000[b]	1.041[b]
	(0.521)	(0.294)	(0.255)	(0.151)	(0.178)	(0.086)	(0.252)
ρ	0.750	0.987[b]	0.993[b]	0.991[b]	0.964[b]	0.995[b]	0.966[b]
	(0.569)	(0.131)	(0.111)	(0.024)	(0.059)	(0.014)	(0.057)
$\sigma_{S'}/\sigma_S$	0.332	0.921[b]	0.835[b]	1.087[b]	1.244[b]	1.005[b]	1.077[b]
	(0.281)	(0.267)	(0.241)	(0.151)	(0.176)	(0.083)	(0.287)
$\chi^2(8)$	5.93	11.12	13.30	6.54	9.68	5.49	4.96
Sign. level	0.655	0.195	0.102	0.587	0.288	0.705	0.762

[a] The sample is quarterly. All interest rates are defined in table 1. A fourth-order VAR model which includes the change in the short rate, $\Delta r_t \equiv r_t - r_{t-1}$, and the long–short rate spread, $S_t \equiv R_t - r_t$, is first estimated and used to forecast future changes in short rates, $\mathrm{E}_t\Delta r_{t+j}$, $j = 1,\ldots,N-1$, where N is the maturity of the long rate in quarters. Subsequently, the VAR forecasts are used to construct the theoretical spread, S', according to the formula:

$$S'_t \equiv [1/(1-g^N)] \sum_{j=1}^{N-1} (g^j - g^N)\mathrm{E}_t\Delta r_{t+j}, \quad \text{with} \quad g \equiv 1/(1+Rbar),$$

where $Rbar$ is the sample mean of R_t. ρ denotes the correlation between S_t and S'_t. σ denotes standard deviation. γ is the regression coefficient of S'_t on S_t: $\gamma \equiv \rho\,\sigma_{S'}/\sigma_S$. R^2 is the coefficient of determination. Numbers in parentheses are bootstrap standard errors from a Monte Carlo simulation of 5,000 runs that preserves the original VAR coefficients and draws randomly from the original estimated residuals. In the cases of USA and Canada, the simulation also preserves the structure of the existing conditional heteroskedasticity in the VAR residuals. The $\chi^2(8)$ statistic is a Wald statistic that tests the eight nonlinear restrictions that the Expectations Hypothesis imposes on the coefficients of the VAR model; in the construction of the chi-squared statistic, the required variance–covariance matrix of the VAR coefficients comes from the bootstrap simulation.

[b] Significantly different from zero at the 5% level.

computed from Monte Carlo bootstrap simulations. Each simulation run draws randomly from the original error structure of the VAR.[9] These bootstrap errors together with the original parameter estimates of the VAR and initial values for the two VAR variables are used to construct artificial time series for the change in the short rate and the term structure spread. Subsequently, the artificial series are used to estimate a fourth-order VAR model, construct an artificial

[9] In the United States and Canada, Engle (1982) tests reveal the presence of autoregressive conditional heteroskedasticity. In the simulations of table 5 as well as the subsequent table 6, the bootstrap errors for the United States and Canadian variables follow univariate ARCH processes with two lags.

theoretical spread, and compute artificial parameters ρ, γ, and $\sigma_{s'}/\sigma_s$. This procedure is repeated 5,000 times. The reported standard errors reflect the cross-sectional standard deviations of the artificial parameters across the 5,000 runs.[10]

Table 5 shows that the expectations hypothesis cannot be rejected in any of the seven countries. Even the US data, which showed strong departures from the expectations hypothesis in the one-period regressions of panel A in table 4b, are now unable to detect statistically significant deviations. The χ^2 test statistics are Wald statistics and were computed using the 8×8 variance–covariance matrix of the eight nonlinear parametric restrictions of eq. (15) across the 5,000 simulation runs.

The results in table 5 show that there are no significant economic departures from the expectations hypothesis as well: With the exception of the United States, the correlation coefficient $\rho_{s',s}$ between the theoretical spread and the actual spread is very close to unity in every country. Similarly, the regression slope coefficient γ and the ratio of $\sigma_{s'}/\sigma_s$ are close to unity in these countries. Thus, in countries other than the United States, the expectations hypothesis has an impressive record in describing the data.

In the United States, γ is 0.25, but has a very large standard error, 0.52. This value of γ is very close to zero and is only one and a half standard deviations away from unity. Thus, although the size of γ suggests economically significant departures from the expectations hypothesis, the imprecision of the estimates do not lead to strong statistical rejections.

The ratio $\sigma_{s'}/\sigma_s$ of 0.33 in the United States is estimated with greater precision than the other parameters. This ratio is more than two standard deviations away from unity, thus uncovering both economic and statistical deviations from the expectations hypothesis. This ratio is also consistent with the implied ratio of table 4b, which was estimated to be 0.51. The overreaction hypothesis, therefore, provides a consistent story both for the anomalous US evidence in the one-period regressions and the anomalous US evidence in the present VAR framework.

5. The spread as a predictor of future changes in short rates: Multiperiod regressions

To examine how closely short rate movements conform to the predictions of the expectations hypothesis, and to assess whether or not long rates in the United States show substantial overreaction, I now trace the evolution of short

[10]Campbell and Shiller (1987, 1991) report asymptotic standard errors instead of bootstrap standard errors. The asymptotic standard errors are constructed under the assumption that the original VAR matrix of observations is fixed. This assumption is artificial and is not required in the construction of bootstrap standard errors.

rates over multiperiod horizons. This is an alternative and more straightforward procedure that complements the VAR methodology of section 4. The previous VAR methodology examines the long-run implications of the expectations hypothesis by placing strong emphasis on the short-run dynamics of interest rate movements. A multiperiod regression makes no assumptions about the short-run interest rate dynamics but has its own drawback: At long forecasting horizons a significant chunk of the data is lost.[11]

Under the maintained hypothesis of rational expectations, a future short-term rate equals the market's expectation of the rate plus an unpredictable forecast error: $r_{t+i} = \mathrm{E}_t r_{t+i} + e_{t+i}$, $i = 0, 1, \ldots, N-1$. Substituting this relation in eq. (12), we get

$$\sum_{i=0}^{N-1} w_i r_{t+i} - r_t = R_t^{[N]} - r_t - \mathrm{E}_t \theta_t + v_{t+N-1}, \tag{16}$$

where v_{t+N-1} is a composite error term unforecastable at time t. The expectations hypothesis predicts, therefore, that in a regression of the form:

$$\sum_{i=0}^{N-1} w_i r_{t+i} - r_t = \alpha + \gamma(R_t - r_t) + v_{t+N-1}, \tag{17}$$

the estimated slope coefficient is unity, $\gamma = 1$. On the other hand, the overreaction hypothesis predicts that γ is less than unity.

Table 6 presents the results of estimating equations similar to eq. (17). The term structure spread is used to predict cumulative changes in interest rates over different multiperiod horizons from 3 quarters ahead to 39 quarters ahead. To be able to compare the regression coefficients across the horizons, in each regression it is counterfactually assumed that the long-term bond has a maturity that matches the horizon. Specifically, the first regression assumes a maturity of 4 quarters, the second regression a maturity of 8 quarters, and so forth until we reach the true maturity of 40 quarters. Hence, under the expectations hypothesis, as the forecasting horizon gets closer to the true maturity of the long term-bond, the regression slope coefficient ought to get closer to its theoretical value of unity.

The estimation uses overlapping quarterly data. This overlapping requires the standard by now correction of the standard errors. Here the Newey–West (1987) method is used. Recall that Richardson and Stock (1989) and others have emphasized that as the degree of overlapping increases with the forecasting

[11]The relative merits of the two procedures are an active research topic. Hodrick (1992) claims that in testing present value models of stock prices, the Campbell–Shiller VAR methodology provides more robust results than multiperiod forecasting equations. The Hodrick conclusions, however, cannot be transplanted to the term structure framework.

Table 6

The spread as a predictor of the cumulative change in the short rate – multiperiod regressions.[a]

$$\sum_{i=0}^{N-1} w_i r_{t+i} - r_t = \alpha + \gamma(R_t - r_t) + v_{t+N-1}, \quad w_i \equiv g^i(1-g)/(1-g^N), \quad g \equiv 1/(1+Rbar)$$

	USA		Canada		UK		Germany		Japan		France		Italy	
	53:2–92:2		50:1–92:2		61:1–92:2		67:1–92:2		61:4–92:2		68:1–92:2		71:1–92:2	
# obs.	158 − N		171 − N		127 − N		103 − N		124 − N		99 − N		87 − N	
N − 1	γ	R^2	γ	R^2	γ	R^2	γ	R^2	γ	R^2	γ	R^2	γ	R^2
3	0.126 (0.083) [0.124]	0.026	0.203[b] (0.076) [0.009]	0.089	0.189[b] (0.085) [0.034]	0.084	0.229[b] (0.100) [0.041]	0.122	0.343[b] (0.091) [0.000]	0.197	0.491[b] (0.100) [0.000]	0.402	0.363[b] (0.099) [0.002]	0.235
7	0.259 (0.174) [0.147]	0.046	0.428[b] (0.151) [0.012]	0.178	0.394[b] (0.141) [0.029]	0.161	0.601[b] (0.176) [0.012]	0.297	0.730[b] (0.176) [0.000]	0.340	0.797[b] (0.063) [0.000]	0.599	0.685[b] (0.135) [0.001]	0.436
11	0.374 (0.246) [0.154]	0.070	0.555[b] (0.206) [0.025]	0.208	0.529[b] (0.214) [0.064]	0.192	0.823[b] (0.148) [0.003]	0.405	0.998[b] (0.189) [0.000]	0.469	1.050[b] (0.059) [0.000]	0.704	0.861[b] (0.094) [0.000]	0.510
15	0.439 (0.283) [0.153]	0.082	0.705[b] (0.205) [0.011]	0.254	0.722[b] (0.250) [0.053]	0.267	0.967[b] (0.083) [0.000]	0.496	1.095[b] (0.240) [0.000]	0.539	1.044[b] (0.071) [0.000]	0.749	1.017[b] (0.085) [0.000]	0.559

19	0.513	0.096	0.824[b]	0.318	0.864[b]	0.382	1.002[b]	0.545	1.095[b]	0.576	1.072[b]	0.761	1.101[b]	0.566
	(0.335)		(0.184)		(0.228)		(0.074)		(0.273)		(0.073)		(0.074)	
	[0.174]		[0.005]		[0.031]		[0.000]		[0.002]		[0.000]		[0.000]	
27	0.613	0.113	1.014[b]	0.439	1.018[b]	0.508	1.126[b]	0.704	1.031[b]	0.612	1.083[b]	0.748	1.149[b]	0.507
	(0.403)		(0.164)		(0.148)		(0.098)		(0.262)		(0.079)		(0.114)	
	[0.189]		[0.002]		[0.004]		[0.000]		[0.006]		[0.001]		[0.000]	
39	1.225[b]	0.355	1.229[b]	0.597	1.039[b]	0.458	1.171[b]	0.835	1.088[b]	0.622	1.133[b]	0.841	1.003[b]	0.400
	(0.146)		(0.142)		(0.156)		(0.060)		(0.282)		(0.079)		(0.072)	
	[0.002]		[0.002]		[0.015]		[0.001]		[0.029]		[0.001]		[0.001]	

[a] The sample is quarterly. R_t and r_t denote the end-of-quarter long-term (ten-year) and short-term (three-month) bond-equivalent yields in percent. *Rbar* is the sample mean of R_t. $N-1$ is the forecasting horizon in quarters. In each regression, N denotes the hypothetical maturity of the long-term bond ($N = 4, 8, 12, 16, 20, 28$, and 40, quarters respectively): for each N, the weights w_i sum to unity. Since the actual maturity of the bond is approximately 10 years or 40 quarters, the Expectations Hypothesis claims that, as the forecasting horizon increases, γ gets closer to unity. Numbers in parentheses are Newey–West (1987) standard errors that correct for a moving average of order $N-2$ and for conditional heteroskedasticity. R^2 is the coefficient of determination.

Numbers in brackets are one-sided significance levels of the null hypothesis H_0: $\gamma = 0$, generated from 1,000 simulation runs. Each simulation run constructs artificial independent series $\Delta r_t = r_t - r_{t-1}$ and $S_t = R_t - r_t$ which mimic the fourth-order autoregressive properties of the true series. In the case of U.S. and Canada, the artificial series also mimic the ARCH residuals of the true series. The Δr_t series is subsequently used to construct the dependent variable for each forecasting horizon, the regressions are run, and the Newey–West t-statistics are saved and compared to the true t-statistics.

[b] Significantly different from zero at the 5% level.

horizon, the number of 'truly independent' sample observations diminishes and, therefore, inferences based on asymptotic distribution theory become increasingly tenuous. For this reason, significance levels are reported for the hypothesis that $\gamma = 0$ based on Monte Carlo bootstrap simulations of 1,000 runs. Each simulation run preserves the original fourth-order univariate autoregressive structure of the Δr_t and S_t series, and draws independently from their respective sets of residuals. Subsequently, cumulative averages of future short rates are created, each of the seven regressions of table 6 is run, and the Newey–West t-statistics are saved. The fraction of 1,000 times that the t-statistic of the hypothesis that $\gamma = 0$ exceeded the t-statistic of the actual data is tabulated in brackets.[12] As expected, the simulation significance levels turn out to be larger than the ones implied by the Newey–West standard errors: At multiperiod horizons, correct inferences require standard errors larger than the reported Newey–West ones.

The results of table 6 line up with the expectations hypothesis impressively well. As the forecasting horizon increases, both the size of the slope coefficient and the regression R^2 increase. This indicates that as we match the forecasting horizon more closely with the true maturity of the bond, the informative content of the spread rises. Consistent with the expectations hypothesis, at $N = 40$ the slope coefficients γ reach and slightly exceed unity. The hypothesis that there is no predictive power in the term structure spread, namely $\gamma = 0$, is overwhelmingly rejected. For example, at $N = 40$, the weakest rejection occurs in Japan with a 2.9 percent significance level.

In the US, the evidence appears somewhat anomalous relative to the earlier findings. At forecasting horizons up to seven years, the evidence is similar to the evidence of tables 4 and 5: The R^2s in the US are the lowest among the G7 countries and the estimated γ's are insignificantly different from zero. However, at the correct horizon of ten years the γ parameter is 1.23 and the R^2 is 0.35. Given the earlier VAR evidence that in the US the spread is more volatile than the theoretical spread, the present evidence at the ten-year horizon is surprising. The discrepancy between the results of the two methodologies may be due to the truncation of the sample in table 6 (39 observations, or the last ten years of term structure spreads, are lost out of a total of 158). Campbell and Shiller (1991), who find a similar discrepancy between the two methodologies, argue that it is due exactly to this truncation in table 6. However, the discrepancy may also be due to the inability of the VAR to capture the true market expectations of future short rates beyond, say, five years into the future. Thus, to investigate this issue, I re-estimated the single-period equations of tables 4a and 4b and the multi-period equations of table 6 by eliminating the last ten years of the sample. All

[12] In the United States and Canada, whose series show evidence of conditional heteroskedasticity, the simulations create artificial Δr_t and S_t series which are conditionally heteroskedastic, using the ARCH parameters of the true series.

regressions use the same time series of term-structure spreads, with the first spread observed in 1953:2 and the last spread observed in 1982:3, as in the ten-year horizon regression of table 6. Consistent with the truncated sample explanation, the puzzle is now weaker over the sample 1953:2–1982:3: The multi-period regressions at horizons one to seven years have larger δ coefficients and higher R^2s, and the single-period regressions have smaller negative β coefficients. Thus, even in the United States, the evidence in table 6 is consistent with the evidence in table 5.

6. Conclusions

The negative correlation between the term structure spread and the one-quarter-ahead change in long rates may strike many economists as a curious but, nevertheless, minor anomaly. After all, the spread between long rates and short rates contains composite information about the level of rates over a very long period of time. To see how the expectations hypothesis fares, one would have to trace the evolution of rates over a long period of time and compare them with the predictions of the spread. One should not, however, expect to find that changes in long rates over the very short run fall exactly within the predictions of the expectations hypothesis: A minor white noise deviation from the expectations hypothesis would receive unduly large weight in those short-run regressions and would obscure the information in the term structure.

Our evidence in all of the G7 countries except the United States supports the above intuition. In France and Italy, the long rate does move in the correct direction in the first place. In Canada, Japan, Germany, and the UK, the long rate moves in the opposite direction from the one predicted by the expectations hypothesis, but this counterintuitive movement is due to a simple additive white noise deviation of the long rate from the level predicted by the expectations hypothesis. The use of instrumental variables reverses the negative correlation and makes the regression slope statistically indistinguishable from the predictions of the expectations hypothesis. Moreover, the size of the white noise discrepancy does not materially affect the informative content of the term structure in these countries, namely, it is not *economically* important: the term structure spread does move almost one-for-one with its theoretical counterpart under the expectations hypothesis, and with the subsequent cumulative change in future short rates.

In the United States, a white noise error on long rates cannot explain the puzzle. The use of instrumental variables results in equally sharp rejections of the expectations hypothesis. Furthermore, time-varying risk premia cannot provide an adequate explanation to the puzzle: Holding premia vary way too much relative to the variability of expected changes in long rates to be able to accommodate regression estimates that are different from zero. In addition, a risk premium explanation lacks plausibility, as became evident with the earlier example in the introduction. Can an extraneous fads component on long rates explain the puzzle in the United States? Fads are often proposed to explain irrational price behavior in the stock market and are based on the notion that mispricing errors on asset prices are persistent. Unfortunately, the fads hypothesis cannot provide a plausible explanation. To explain the puzzle one has to provide a story of an immediate reversal in long rates. Fads act in the opposite manner: they prolong rather than reverse extraneous shocks to long rates. The only alternative hypothesis that comes close to explaining the evidence in the United States is the overreaction hypothesis of Froot (1989) and Campbell and Shiller (1991). The hypothesis is consistent with both the short-run response of long rates and the larger volatility of the spread relative to theoretical spread of the VAR analysis. It is also supported by the survey evidence of Froot.

The contrast between the evidence in the United States and all the other G7 countries is an interesting phenomenon, particularly because the US financial markets have been the most liquid markets in the world during the post-war period. If markets are dominated by rational traders, then it is the US markets where one would least expect to find phenomena of overreaction. If, however, markets are dominated by irrational – often called noise – traders, the higher trading volumes in the US would signify a stronger presence of noise traders, and the observed bond price overreaction would be less surprising. Although very interesting, a clear-cut resolution of this issue is beyond the scope of the present paper and requires further research.

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