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JEL Classification: E43, G15, E42

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Time-Variation in Term Premia: International Survey-Based Evidence

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Abstract

Using a large, previously unexplored international dataset of market expectations that covers a broad range of deposits, this paper presents a wealth of empirical evidence on the behavior of the term structure of interest rates in an international perspective. We find that our survey forecasts are of quite good quality, outperforming a relevant naive benchmark in most cases. We also find considerable international evidence in favor of rejecting the 'pure' version of the expectations hypothesis. We also find some evidence that the behavior of market participants, when making predictions about the future level of interest rates, is not entirely in line with rational behavior. There is strong evidence of time-variation in term premia. Furthermore, while this variation in term premia can be captured adequately by low-order members of the ARMA class models, there is clear evidence that conditional heteroskedasticity in the movement of term premia plays an important role in explaining the time-variation for a number of countries.

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1. Introduction

The term structure of interest rates describes the relationship between short- and long-term interest rates, and has become an important concept both for academics and policy makers. One of the theories explaining the shape of the term structure is the expectations hypothesis, which states that the entire term structure at a given point in time reflects the market's expectations of future rates. In one version of the expectations hypothesis (see, for instance, the seminal paper by Fama, 1984) the forward rate implicit in the term structure is an unbiased predictor of the future realized spot interest rate, up to a constant liquidity premium. In other versions the expected holding return is equal for all bond maturities except for a maturity-specific constant term premium, or long-term interest rates are equal to the (weighted) average of expected future short rates, up to a constant difference.

Shiller (1990) and Campbell (1995) review some of the literature on term structure studies and find that the expectations hypothesis for the U.S. term structure is rejected. The general result is that although the forward rate contains some information about future spot rates, it is by no means an unbiased predictor. In fact, it has become one of the well-established empirical regularities in the financial economics literature that the implicit forward rate is a biased predictor of the future spot interest rate. It must be stressed that results are sample-dependent. Christiansen (2003) finds that, when investigating two sample periods, the expectations hypothesis is rejected in one, though not in the other. Rejection of the expectations hypothesis has traditionally been attributed to the existence of a time-varying term premia, as manifested by Fama (1984) or more recently by Harris (2001) and Dai and Singleton (2002), the existence of irrational behavior on the part of market participants, or some combination thereof.

However, international evidence in favor of the expectations hypothesis is by far not as strong as for the U.S. Dahlquist and Jonsson (1995), Gerlach and Smets (1997), Jondeau and Ricart (1999), and Domínguez and Novales (2000) find that while for the U.S. the expectations hypothesis is firmly rejected, this is by no means the case in a broader international, non-U.S. context. Their results are irrespective of the various forms in which the expectations hypothesis has been presented in the literature and are focused both on forward rate unbiasedness regressions or cointegration techniques, and on spread regressions. In the current paper we address this issue by relying on a large international dataset that includes simultaneous American, European, and Asian interest rate deposits, thus allowing an interesting comparison of the strength of the expectations hypothesis in a broader setting.

Moreover, even when the expectations hypothesis cannot be rejected in a wider international context, it may still be the case that time-varying term premia exist in the term structure. Dahlquist and Jonsson (1995) show that while the expectations hypothesis cannot be rejected for the Swedish term structure over the sample under investigation, in that forward rates are not biased estimates of future spot rates, the parameter estimates are nevertheless unstable over time and this instability is amplified with maturity. The instability in turn may then well be the result of time-variation in the term premia. So even though the expectations hypothesis may not be rejected for the usual data sources, time-variation in term premia may still exist.

However, the inherently unobserved nature of the term premium never allows a clear assessment of how the results from previous expectations hypothesis studies are biased by the assumption of rational expectations and whether rejection of the expectations hypothesis is in fact attributed to the existence of a time-varying term premium. In the current paper we try to address this issue by relying on survey measures of interest rate expectations. Using survey measures of expectations allows us to verify whether commonly-used assumptions about the underlying process that characterizes the expectations hypothesis are correct. Not only does the use of survey measures of expectations allow us to assess whether time-varying term premia exist in the term structure of interest rates, it also allows to model the term premia, in terms of time series models or economic fundamentals.

The use of survey data is not uncommon in the literature. Frankel and Froot (1987) and Cavaglia, Verschoor and Wolff (1993, 1994) use survey data on foreign exchange to examine whether exchange rate expectations are formed rationally. Dokko and Edelstein (1989) review the usefulness of the Livingston forecasts of stock market rates of returns and Branch (2004) re-examines rational inflation expectations. Although Friedman (1979) and Froot (1989) briefly examine the expectations hypothesis of the U.S. term structure using survey data, these studies are hampered by small sample problems and do not allow an assessment of the expectations hypothesis in a wider international context.

While the failure of the expectations hypothesis of the term structure of interest rates in the U.S. is often attributed to the existence of a time-varying term premium, it remains ambiguous as to what the sources of this time-variability are. Most of the recent literature on term premium determinants has focused on models of its second moments. Engle, Lilien and Robins (1987), for instance, suggest that the conditional variance of the excess holding yield on a long bond is a determinant of the current term premium, above the information contained in the often-used long-short yield spread. Yet Tzavalis and Wickens (1995) and Henry (1999) argue that a structural break in the unconditional variance can give rise to a spurious persistence in volatility. This suggests that conditional heteroskedasticity models may not adequately capture the time-variation in the term premium. Henry (1999), for instance, accounts for regime shifts in the U.S. monetary policy in tests of time-varying term premia and finds that conditional heteroskedasticity models appear less than adequate. On the other hand, Lee (1995) and Bekdache (2001) find that term premia should be modeled as a function of macroeconomic variables instead of as a function of mere asset covariances. In particular production, money supply, and maturity composition of federal debt appear important determinants.

Little if anything is documented about the behavior and the determinants of foreign term premia, while the behavior of international term premia can be substantially different. Hejazi, Lai, and Yang (2000) report that, in contrast to the U.S., the Canadian term premium is not related to the conditional variances of Canadian macroeconomic variables, like industrial production and money supply. Yet, an in-depth comparative study on the presence of term premia in international interest rate deposits is still lacking in the literature. We address these issues by explicitly modeling the survey-based term premia using various techniques.

Finally, the rejection of the expectations hypothesis is particularly pronounced for longer maturity bonds, while for the shorter maturity bonds the evidence is mixed. We therefore focus in this paper on the short end of the term structure, in an effort to obtain more clarity for this segment.

Our results are easily summarized. Although we find strong evidence for the rejection of the 'pure' version of the expectations hypothesis, we only find the null hypothesis that forward rates are biased estimates of future interest rates to be rejected for a limited number of international interest rate deposits. Nevertheless, we find some evidence that the behavior of market participants, when making predictions about the future level of interest rates, is not entirely in line with what rational behavior would suggest. We find that there is strong evidence of time-variation in the term premia. Furthermore, while the dynamics of term premia can be accounted by low-order members of the ARMA class models, there is clear evidence to conclude that the conditional heteroskedasticity in term premium movement plays an important role.

The remainder of the paper is presented as follows. In section 2 we introduce the basic methodology. Section 3 presents the data and examines several important standard propositions regarding interest rate expectations. In section 4 we examine the validity of the expectations hypothesis and we look at the extent to which time-varying term premia are present in the term structure of interest rates and whether the behavior of market participants can be labeled rational. In section 5 the time-series behavior of term premia is examined and section 6 models the time-variation in the term premia. Finally, section 7 concludes.

2. The Expectations Hypothesis of the Term Structure

The expectations hypothesis of the term structure of interest rates is presented in the literature in various basic forms. In one form, interest rates on long-term contracts are represented as an average of expected future short-term rates, plus a constant premium. A constant is allowed by the expectations hypothesis, for instance to account for the liquidity preference theory that states that investors are risk-averse and hence require an (monotonically) increasing return for bonds with longer maturities, since the latter are less liquid. An alternative form of the expectations hypothesis asserts that investors try to equalize the expected holding returns of investments strategies with various horizons, again up to a constant (see, e.g. Campbell, 1995). A final version of the expectations hypothesis asserts that forward rates should equal the expectation of the corresponding future realized interest rates, and hence forward rates can be used to forecast future rates. An advantage of this version is that it depends only on contemporaneous information. All three versions allow for a constant term premium. Shiller, Campbell and Schoenholtz (1983) give a clear overview of the implications of these three versions. In this article we focus on the third version of the expectations hypothesis.

The continuously compounded yield to maturity at time t on an n-period zero coupon (or discount) bond with a normalized face value of 1 currency unit and a market price of $P_t^{(n)}$ equals

$$r_t^{(n)} = -\frac{1}{n}\ln(P_t^{(n)}) \tag{1}$$

for each period when held to maturity. The forward rate implied by the term structure of interest rates at time t for an $(n - k)$-period contract to be delivered at time $t + k$ is the interest rate at which either holding an n-period bond until maturity or investing in a k-period bond at time t and subsequently in a $(n - k)$-period bond at time $t + k$, yields the same return, defined as

$$f_t^{(n,k)} = \frac{1}{n-k}[\ln(P_t^{(n)}) - \ln(P_t^{(k)})] = \frac{n}{n-k} r_t^{(n)} - \frac{k}{n-k} r_t^{(k)}, \tag{2}$$

where $k < n$. The expectations hypothesis states that the implicit forward rate should equate the expected future interest rate, up to a term premium, which is defined as

$$\varphi_t^{(n,k)} \equiv f_t^{(n,k)} - E_t[r_{t+k}^{(n-k)}]. \tag{3}$$

From this specification it can be seen that although the forward rate may contain information about future spot interest rates, it also contains a premium component. Variation through time of the term premium can therefore reduce the predictive power of forward rates as predictors of future spot rates. Time-variation in term premia might arise because of changes in market participants' preferred investment horizon, so that term premia evolve over time with investors' presumed risk aversion. More general equilibrium asset pricing models, like Cox, Ingersoll and Ross (1985), imply that the term premium is a function of both investors' attitudes towards risk, so that the term premium can be interpreted as a risk premium, and the covariance of long and short rates with consumption or wealth.

A technical problem that has plagued model builders on the expectations hypothesis is that market expectations are inherently unobservable. A commonly used assumption is that expectations are formed in a rational way, such that the market forms unbiased expectations of future interest rates and that the forecast error, defined as $\varepsilon_{t+k}^{(n-k)} \equiv E_t[r_{t+k}^{(n-k)}] - r_{t+k}^{(n-k)}$, is orthogonal to all information available at time t, and in fact is distributed around a zero mean and constant variance. Conditional on this assumption of rational expectations, the term premium is approximated by the ex-post forward bias

$$\phi_t^{(n,k)} \equiv f_t^{(n,k)} - r_{t+k}^{(n-k)} = \varphi_t^{(n,k)} + \varepsilon_{t+k}^{(n-k)}. \tag{4}$$

Fama (1984) builds on this assumption by developing a technique based on forward rates that are implicit in the term structure of interest rates, while allowing for time-varying term premia. He assumes that premia are time-variant when their covariance with other elements from the investor's information set, such as the forward premium, is nonnegative. Therefore, the expectations hypothesis of the term structure of interest rates can be tested by means of the following equation:

$$r_{t+k}^{(n-k)} - r_t^{(n-k)} = \alpha_1 + \beta_1[f_t^{(n,k)} - r_t^{(n-k)}] + \eta_{t+k}^{(n-k)}. \tag{5}$$

The null hypothesis that the implicit forward rate is an unbiased predictor of the future spot interest rate is presented by the joint hypothesis that $\alpha_1 = 0$ and $\beta_1 = 1$. In fact, this specification tests the pure version of the expectations hypothesis, where no constant premium is allowed. In a weaker version the constant is allowed to be nonzero. In the next section we introduce a survey data set that allows us to deviate from the commonly used rationality assumption. We denote $r_{t+k}^{e(n-k)}$ as the survey-based proxy for the (unknown) market expectation $E_t[r_{t+k}^{(n-k)}]$.

3. The Survey Data and Standard Propositions Regarding Expectations

Every second Monday of each calendar month Consensus Economics of London publishes results from a survey among up to 150 professional market participants and forecasting agencies (per deposit) for their 3-months-ahead expectations for a large number of international interest rate deposits All contracts have a maturity of 3 months. To keep the analysis tractable we take the cross-sectional arithmetic average of the individual market participants and use this in the analyses. Replications of the analyses using other aggregate measures of expectations, like the median, did not alter the general message from the results.

For our panel, we obtain expectations on 20 interest rate series from developed economies for the period of January 1995 through December 2004 (see Table 2 for a listing of the various deposits included). This period is of particular interest since it contains several financial crises, the introduction of a single monetary currency unit, and several dramatic changes in the level of interest rates for some deposits. The panel is unbalanced since for some series the expectations were discontinued in January 1999, whilst for others the first expectations were published somewhat later than January 1995. We categorize the deposits in four areas based on their geographic origin.

Although survey participants have a few days time to return their expectations, we learned that the vast majority send their responses by e-mail on the Friday before the publication day (second Monday of the month). We consider this Friday to be the day on which the expectations are formed. On this Friday, we obtain spot interest rate series with different maturities to match with the survey data. All spot rate series are obtained through Datastream. To verify that the information sets of market participants are not too diverse, all of the analyses throughout this study were re-estimated using spot data from various days surrounding this Friday, yet the overall results remained virtually unchanged. The sample primarily consists of expectations for eurodeposit rates, interbank rates, and treasury bill rates. The use of eurodeposit rates has the added advantage of not being affected by possible government regulations such as capital controls and are sold in a number of national markets simultaneously, thus making the rates more comparable than national rates.

Typical concerns when using survey data, in any setting, are whether these data reflect the true market expectations, whether the expectations are biased because of strategic behavior from the panelists, or whether forecasts from surveys are of any use in an out-of-sample forecast setting—a criterion that has often been put forward to evaluate the quality of survey expectations. It should be noted that for survey data in the present setting it is of primary concern that the survey expectations reflect the market's sentiment at the time they are formed, that is, the survey data should reflect expectations, nothing more than that. While it is not the primary concern that the expectations outperform other forecasting techniques, there is general consensus that expectations from surveys in most fields perform no worse than any other forecast technique and in this respect we can also learn much from related fields.

Ang et al. (2005), for instance, provide recent evidence that aggregate expectations from various surveys on inflation consistently deliver better forecasts than time-series models, models based on the yield curve, and forecasts based on the Phillips curve, which highlights the usefulness of survey measures of expectations. Elliott and Ito (1999) find that in the foreign exchange market portfolio's based on survey expectations produce small, but positive, profits. It is therefore surprising that Hafer et al. (1992), the only study that investigates the forecast performance of survey-based *interest* rate expectations, find that survey forecasts for the 3-month U.S. treasury bill rate have larger root mean squared errors (RMSE) and mean absolute errors (MAE) than simple random walk (no change) forecasts.

Insert table 1 here

To ascertain the quality of the survey-based forecasts, table 1 presents MAE's and RMSE's of the survey-based and forward rate forecasts in comparison to the benchmark forecast performance of a random walk model. In addition, the table presents Theil's *U*-statistic, in our case defined as the ratio of the RMSE of the survey forecast divided by the RMSE of the random walk forecast. For about half of the deposits the errors for the survey forecasts are lower that those of the forward forecasts, indicating that the forecast performance from survey data is about equal to that of forward rate forecast. Of more significance is that both in terms of MAE and RMSE the survey-based forecasts outperform the benchmark random walk model. In fact, except for U.S., Australian, Swiss and Belgium interest deposits, the *U*-statistics are all below unity, which stresses the outperformance of the survey forecasts. It is interesting to compare these findings to those of Hafer et al. (1992). Generalizing their finding on the forecast performance of the U.S. treasury bill would erroneously have led to the conclusion that survey-based forecasts cannot outperform a simple random walk forecasts, where in fact we document that the failure to 'beat' the random walk for U.S. forecasts is only an exception to a much larger sample where survey forecasts do outperform their benchmark.

Another issue is that consensus measures of expectations are likely to perform better than the individual expectations that together make the consensus. Although some individuals' forecast performance may be better than others' in terms of criteria like root mean squared errors, it is difficult to identify a priori who these individuals are, in particular since forecast performance is not constant. Since surveys aggregate the expectations from many market participants, the information in consensus measures is usually superior, which may be due to an effect similar to Bayesian Model Averaging or due to implicitly filtering out the common components of the forecasts.

Insert table 2 here

We continue the analysis by providing summary statistics for the annualized percentage realized spot interest rate, its first difference, and the survey-based expected future rate, in Table 2. Several findings are worth noting. Over the period investigated, January 1995 through December 2004, the general level of all interest rate deposits decreased, given the uniformly negative mean change in the interest rates. For some deposits, like the Hong Kong interbank rate, the decrease is over 15 percent over a period of 10 years. We must therefore bear in mind that all results in this paper should be interpreted against the background of a period of decreasing interest rates.

Furthermore, there does not seem to be a consistent difference in the variability in expectations for the various blocks of deposits in terms of their standard deviation. For instance, the standard deviation for the so-called EMS deposits (most continental Europe rates) is by no means lower than the variability of expected rates from other blocks of countries. This is an interesting finding, for it could reflect that the market did not believe in the stabilizing role of the EMS, and expected larger swings in the EMS' interest rates than actually occurred.

Insert table 3 here

Table 3 presents summary statistics from the survey-based term premia, φ, following the definition in equation (3). The majority of average term premia are positive, indicating that forward rates are on average larger than the expected future levels of the interest rates. Interpreting the results, for instance for the Hong Kong interbank rate, suggests that the yield curve is priced in such a way that the total return of a longer-term contract is higher than the return participants expect when short rates are rolled over. In other words, market participants add a premium for longer-term instruments. Results from a sign test corroborate that for most interest rate deposits the term premium is more often positive than negative, such that the finding of positive average premia is not the result of outliers. Note that premia are typically skewed to the left and show leptokurtic behavior, which results in a regular rejection of the normal distribution.

First-order autocorrelation coefficients are reported for the premia and their squares, accompanied by significance levels from Ljung-Box Q-statistics for first-order serial correlation. The level premia show strong evidence of serial correlation. Modern equilibrium theory can explain this behavior. If term premia are functions of both investors' attitudes towards risk and the covariance of long and short rates with consumption or wealth, as in the Cox et al. (1985) framework, and these parameters evolve only slowly over time as is often assumed, then term premia should be partially predictable from past observations. In this framework the serial correlation in the levels of the term premia makes sense from an economic theory point of view. Serial correlation in the squares of the premia furthermore suggests the presence of at least first order autoregressive conditional heteroskedasticity.

We subsequently analyze the (non)stationarity of expected, realized, and forward interest rates. When the forward rate, realized spot rate, and expected future spot rate all contain a unit root, a linear relation between any of these variables can be spurious and hence may lead to an erroneous non-rejection of the expectations hypothesis. If, say, the forward rate contains a unit root, but the expected interest rate is stationary, the resulting term premium is not stationary, since a linear combination of a stationary and nonstationary variable is nonstationary.

The (non)stationarity of nominal interest rates remains an issue of central concern in the literature. In a seminal paper Rose (1988) investigates the (non)stationarity of various international nominal short- and long-term interest rate measured at yearly, quarterly, and monthly data frequencies and finds that the null hypothesis of a single unit root in the levels of the series cannot be rejected. Rapach and Weber (2004) re-examine the stationarity condition of long-term government bonds as the nominal interest rates for about the same sample of international deposits as Rose, using new unit root tests with improved size and power. They find that, except for Austrian, German, and Swiss interest rates, the results conform to those in Rose (1988) in that nominal rates contain a unit root.

It remains ambiguous as to whether *expected* interest rates inherit the same (non)stationarity chracteristics as their nominal realized counterparts. To the best of our knowledge, no previous study has attempted to investigate the stationarity of expected interest rates (or expectations series in general), by means of survey data or other. When the order of integration of expected interest rate series differs than that of realized interest rate series, any linear combination of the two is nonstationary, and results of a test of time-varying term premia may then be spurious.

We examine the stationarity of realized, expected, and implicit forward rate series using two nonparametric unit root tests. First, we use a traditional Phillips-Perron (PP) unit root test that is robust to a variety of serial correlation and heteroskedasticity, where the null hypothesis is that of a unit root. A criticism of such traditional unit root tests is that they cannot distinguish between a unit root and a near-unit root process, since the classical way of hypothesis testing ensures that the null hypothesis cannot be rejected unless there is strong evidence against it. Therefore, the test due to Kwiatkowski, Phillips, Schmidt and Shin (1992) (henceforth KPSS) is also considered, where the null hypothesis is that of a stationary series.

In both tests a linear trend does not seem relevant for the data at question and a constant term is included for generality; thus we check for level stationarity. Spectral estimation is performed using the Bartlett kernel and the optimal lag length is selected automatically using the Newey-West bandwidth.[1]

Insert table 4 here

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[1] Other bandwidth estimators, like the Andrews bandwidth, were considered but did not alter the general conclusions in any way.

Table 4 presents first-order autocorrelation coefficients, together with test statistics from the unit root/stationarity tests for the forward rate, realized future spot rate, expected future spot rate, and some linear combinations hereof. The general finding that emerges from the Phillips-Perron unit root test is that nominal realized interest rates, $r_t^{(3)}$, appear to be $I(1)$ in that the null hypothesis of a unit root in the levels cannot be rejected, whilst for their first differences the null of a unit root is strongly rejected.[2] Three exceptions are the German, Dutch, and Japanese interest rate deposits, for which a unit root in the level cannot be rejected. Interestingly is that for survey-based *expected* interest rates the null hypothesis of a unit root also cannot be rejected for a vast majority of the countries. This would imply that *expected* interest rates seem to share the stationarity properties with their realized counterparts. However, we must be careful with such an interpretation, for results on Austrian interest rates show that although realized rates can be $I(1)$, expected rates can nevertheless be $I(0)$. Finally, most forward rates appear to contain a unit root as well.

However, for the expected and implicit forward rates in excess of the contemporaneous nominal spot rate, the null hypothesis of a unit root is generally rejected, suggesting that any test the expectations hypothesis in not influenced by a spurious relationship between $f_t^{(6,3)}$ and $r_{t+3}^{(3)}$ or $f_t^{(6,3)}$ and $r_{t+3}^{e(3)}$. KPSS test results support these results, although for continental European deposits the results remain somewhat non-uniform. At minimum, for all non-continental European, and nearly all continental European deposits the null hypothesis of stationarity can be rejected for the levels of the realized, expected, and forward interest rates.

Although it is possible that for some interest rate deposits the levels of the realized, expected, and forward interest rates contain a unit root, whereas for other deposits a unit root is absent, it is also possible that our sample size simply does not give us enough testing power to reject the unit root in the levels for a few countries.

Based on results from both tests, the general picture that emerges is that while nominal realized, implicit forward, and expected interest series seem to contain a unit root, linear combinations thereof are stationary. In addition, Bekaert, Hodrick, and Marshall (1997) show that traditional single-equation regression models are characterized by a positive bias in the slope coefficients, even with relatively large samples. Tests of the expectations hypothesis, rational expectations, or time-variation in the term premia should therefore be expressed in terms of these linear combinations. In the next section we examine the expectations

[2] Statistics on the differenced series are not reported to conserve space, but are available on request.

hypothesis of the term structure of interest rates empirically and investigate to what extent expectations are rational and whether term premia are time-varying.

4. Rational Expectations and Time-Varying Term Premia

The unbiasedness specification in (5) has been used regularly as a tool to verify the validity of the expectations hypothesis, and in a comparable version to explain the forward premium behavior in the exchange rate literature (see, e.g. Engel, 1996, and Hodrick, 1987, for an overview). Failure of the hypothesis that $\alpha_1 = 0$ and $\beta_1 = 1$ is interpreted as failure of the expectations hypothesis, and conditional on the assumption of rational expectations the failure is generally attributed to the existence of a time-varying term premium.

However, from the decomposition in (4) it follows that the forward premium consists of both an expectational error and term premium component and therefore the above regression-based test is in fact a joint test of time-varying term premia *and* the existence of errors in expectations. Whenever the assumption of rational expectations is invalid, rejection of the null hypothesis that $\alpha_1 = 0$ and $\beta_1 = 1$ does not automatically imply the existence of time-varying term premia, but instead can be attributed to the existence of constant or time-varying expectations errors, constant or time-varying term premia, or a combination of these. Moreover, even when one fails to reject the null hypothesis that $\alpha_1 = 0$ and $\beta_1 = 1$, this does not imply that no time-varying term premia exist, whenever the variability of term premia and expectations errors move in opposite directions, thereby dampening any movement in the ex-post forward error.

Insert table 5 here

Table 5 reports GMM estimates for the unbiasedness test in equation (5), with standard errors corrected to allow for a k-order moving average.[3] Most slope coefficients are positive and significant, indicating that the forward rates contain information about the future spot rates. Conditional on the assumption of rational expectations, the null hypothesis that the

[3] When the forecast horizon is longer than the observational frequency, the forecast error will be serially correlated up to a moving average process of order $k - 1$. Since the survey forecasts are for 'approximately 3 months ahead' projections, and since panellists are known to often make their projections towards the end of a calendar month, it is expected that the k-month-ahead forecast is in fact a k-month-plus-a-few-days-ahead

forward rate is an unbiased estimate of the future spot interest rate is presented by the joint hypothesis that $\alpha_1 = 0$ and $\beta_1 = 1$.

The null hypothesis is rejected for about 75 percent of the countries, with varying degrees of significance. This implies that there is considerable evidence against at least the pure version of the expectations hypothesis. Rejection of this null is not predominantly caused by either a constant or a time-varying premia, but must be attributed to both. Estimates for β_1 are furthermore mostly less than unity and occasionally even negative, although when relaxing the zero constant premium, we find that the null hypothesis of $\beta_1 = 1$ can only be rejected for five interest rate deposits.

Focusing on the role of the estimate for β_1 alone, one could erroneously be tempted to interpret the findings in Table 5 as evidence against time-variability of term premia. However, since the test in (5) is a joint test of rational expectations and the existence of constant or time-varying term premia, results in Table 5 remain inconclusive as to the existence of time-varying term premia. In fact, even when time-varying term premia exist in the term structure of interest rates, β_1 may still be equal to unity provided that expectations are irrational. Equation (5) hence cannot capture the presence of time-varying term premia, nor provide evidence of irrationality of expectations in the term structure of interest rates.

Survey data can be used to decompose the forward premium into a time-varying term premium component and an irrational expectations component. It can be shown that the probability limit of (5) reduces to

$$\text{plim}\,\beta_1 = 1 - \frac{\text{cov}(\phi_t^{(n,k)}, f_t^{(n,k)} - r_t^{(n-k)})}{\text{var}(f_t^{(n,k)} - r_t^{(n-k)})}, \tag{6}$$

which exactly measures time-variability in term premia in the Fama sense. The probability limit in (6) can be decomposed into an expectational error and time-varying term premium component as $\beta_1 = \beta_2 + \beta_3$, where

$$\beta_2 \equiv \frac{\text{cov}(\varepsilon_{t+k}^{(n-k)}, f_t^{(n,k)} - r_t^{(n-k)})}{\text{var}(f_t^{(n,k)} - r_t^{(n-k)})} \tag{7}$$

and

$$\beta_3 \equiv 1 - \frac{\text{cov}(\varphi_t^{(n,k)}, f_t^{(n,k)} - r_t^{(n-k)})}{\text{var}(f_t^{(n,k)} - r_t^{(n-k)})}. \tag{8}$$

forecast. We have replicated all results assuming a moving average process of order $k - 1$, yet these results differ only marginally from the ones reported and do not alter any of our conclusions.

To test the rationality of survey-based interest rate expectations we employ an orthogonality test (see, e.g. Pesaran, 1987). The orthogonality test aims to assess whether economic agents use information that is available to them efficiently to forecast future interest rates. The null hypothesis of rational expectations implies that $\alpha_2 = 0$ and $\beta_2 = 0$ in regressions of the following form

$$\varepsilon_{t+k}^{(n-k)} = \alpha_2 + \beta_2^{'} X_t + \upsilon_{t+k}^{(n-k)}, \tag{9}$$

where X_t is a vector of elements from the investor's information set, a subset from her complete, yet unobserved information set Ω_t. When expectations are formed in a rational way, the survey-based forecast error should be orthogonal to all elements form the investor's information set at the time she forms her forecasts. Although the information set may be infinite and unobserved, we choose to use the forward premium, $f_t^{(n,k)} - r_t^{(n-k)}$, as the sole element for X_t. We choose this specification since under the null hypothesis of rational expectations and under the assumption that any measurement error in the survey is orthogonal to the forward discount, the β_2 coefficient is precisely equal to β_2 in equation (7). Equation (9) was fitted using GMM for each deposit.

Insert table 6 here

Table 6 reports parameter estimates of this orthogonality test. The null hypothesis of rational expectations is rejected for about half of the deposits. It is interesting to note that the vast majority of estimates for β_2 are positive of sign, indicating that an increase in interest rate implied by the term structure (through the implicit forward rate) leads to an underestimation of future interest rates.

These results should be interpreted with care. If conditional forecasts are formed rationally, whilst allowing for a small probability of a large interest rate movement, then forecasts will appear biased when judged from *ex-post* forecast errors. This is the familiar 'peso problem' due to Krasker (1980). Bekaert, Hodrick, and Marshall (2001) find that the failure of the expectations hypothesis in the U.S. can be explained at least partly by the existence of such a peso problem effect. An alternative explanation would be that the time series process which describes the expected interest rate movement, is not ergodic as is implied in the application of the GMM procedure.

In order to discover time-variation in the term premia and to see to what extent the existence of a time-varying term premium is an economically important reason for rejection of the expectations hypothesis, the following regression test is used:

$$r_{t+k}^{e(n-k)} - r_t^{(n-k)} = \alpha_3 + \beta_3[f_t^{(n,k)} - r_t^{(n-k)}] + v_{t+k}^{(n-k)}, \qquad (10)$$

where the mathematical expectations are replaced by their survey-based counterparts. The null hypothesis of no time-varying term premia is represented by the hypothesis that $\beta_2 = 1$, where the correlation of the risk premium with the forward discount is zero. By inspection, the β_3 coefficient is precisely equal to β_3 in equation (8). Similarly, the hypothesis of a zero mean term premium can be tested by examining whether the α_3 coefficient is significantly different from zero.

Insert table 7 here

Table 7 reports GMM estimates for the test of time-varying term premia in equation (10). There is strong evidence of time-variation in the term premia, given that for all but a few deposits the null hypothesis of no time-variation, $\beta_2 = 1$, is resoundingly rejected. Although for the remaining deposits, Belgium and France, and Denmark and Hong Kong the null hypothesis of no time-variation in the term premium cannot be rejected, slope coefficients are below unity and the joint null hypothesis of no constant or time-varying premia is nevertheless rejected. All slope estimates are between zero and unity, a finding which corroborates the early finding of Froot (1989) and MacDonald and Macmillan (1994) for US and UK interest rate deposits, respectively. Whilst strong evidence of time-varying term premia, all estimates of β_3 are significantly positive, indicating that not all of the variation in the expectations hypothesis can be attributed to the existence of a time-varying term premium.

Lagrange multiplier tests for the presence of serial correlation and conditional heteroskedasticity in the residuals are reported. For most deposits, significant serial correlation is present up to one lag, indicating that term premia might be characterized by ARMA class models of low order. In general, evidence of serial correlation is more evident for non-continental European rates than for continental European rates. Volatility clustering furthermore appears predominantly in non-continental European term premia, indicating that the conditional variance of term premia varies over time. One noteworthy case is the Hong

Kong interbank rate, for which autocorrelations up to six months is reported, and strong ARCH effects. Hong Kong interbank rates have experienced dramatic fluctuations in 1997 and 1998 and have decreased considerably.

We can question what the implications are of these results. To do so, a closer examination of time-varying term premia in the financial economics literature is required. In most financial models, time-varying term premia are essentially exogenous, like in the square root model of Cox et al. (1985), the affine term structure model of Dai and Singleton (2002), and the macro models of Ang and Piazessi (2003) and Rudebusch and Wu (2005). A way to endogenize time-varying term premia is by arguing that transaction costs of entering the long-run bond market are higher and vary over time due to this segment's smaller liquidity, which can obviously explain why the yield curve is upward sloping most of the time. Alternatively, a segmented market approach may also explain why term premia vary over time when realizing that the preferred contract length of certain type of investors may change with their hedging requirements. In any case, an identification of the determinants of the time-variation in the term premia is desired, so that they can be controlled for in traditional tests of the expectations hypothesis.

On the other hand, the present findings may well present yet another challenge to the rational expectations hypothesis. Several ways to salvage the rational expectations hypothesis in the present setting are when we consider the existence of peso problems, learning behavior as introduced by Lewis (1989), or the presence of regime shifts by the monetary authority. Yet, it seems more plausible to depart, or at least redefine, the expectations hypothesis. Several theories that postulate deviations from it have recently drawn the attention. A common argument is that individual agents have limited capacity for processing information, which is in line with ordinary experience and basic behavior theory. A recent example of this line of argumentation is the rational inattention model due to Sims (2003).

In the next section we analyze various time-series properties of the term premia and look at potential explanations for the time-variation.

5. Time Series Behavior of the Term Premia

In most modern, intertemporal general equilibrium asset pricing models that describe the term structure of interest rates, for instance those of Vasicek (1977) or Cox, Ingersoll, and Ross (1985), the term premium is a function of variables like expectations, investors' risk and intertemporal consumption preferences, investment alternatives, and interest rates' covariance with consumption. Although these determinants are extremely difficult to observe, they are commonly believed to evolve only gradually over time. This is an important assumption, for if determinants evolve gradually, so does the term premium and historical information about the term premium should therefore have at least some predictive power for future levels thereof. This argument has commonly resulted in treating it as a state variable, or unobserved component, and estimating it by means of a Kalman filter. Such state variables are commonly assumed to follow low-order versions of the ARMA class models.

The existence of low-order serial correlation in the residuals of the time-variability equation (10) furthermore supports the assumption that term premia evolve only gradually and that past values have predictive power. In order to examine the statistical time-series behavior of the risk premia implied by the survey-based expectations, we estimate AR(1), MA(1), and ARMA(1,1) models for the term premia and present the model with the best fit in terms of the Bayesian Information Criterion (BIC) in Table 8. Even though in a rare occasion a higher lag seems to improve the model in terms of the selection criterion, we do not report these results to preserve model parsimony and comparability and we select the most appropriate model from the three variations mentioned earlier. Lagrange Multiplier statistics for the presence of serial correlation that is not picked up by the models are presented, as well as a measure of ARCH presence.

Insert table 8 here

The low-order versions of the ARMA class models are quite capable of accounting for variation in the term premia. For all AR or MA models, parameter estimates for γ_1 and/or γ_2 are significantly positive. These findings are consistent with the theories laid out above, in that knowledge about last month's premium provides considerable information about next month's premium. It is noteworthy that non-continental European deposits are better characterized by these low-order ARMA models than other deposits, given the higher significance of the terms and higher adjusted R^2 statistics.

The ARMA models account for serial correlation in the term premia adequately, except for Hong Kong and Norway. We do note, however, that conditional heteroskedasticy is present in a number of cases. In the next section we look at this issue more closely

6. Determinants of the Time-Variation in the Term Premia

In this section we examine the determinants of term premia implicit in the term structure. Although the presence of time-variation in the premia is firmly documented in the previous section, it remains a question as to what determines this time-variation. In this section we examine three models of the term premium.

There is evidence in the literature that term premia are forecastable using variables such as the spread between the *n*-period and the *k*-period rates and the level of the 1-period interest rate (see, e.g. Bekdache, 2001, or Campbell, 1995). This relationship is partly explainable by the preferred habitat theory of the term structure that states that risk-averse investors require a premium when investing in assets with a maturity that does not match their habitat. We verify this preferred habitat argumentation by linking the term premium to

$$\varphi_t^{(n,k)} = \delta_0 + \delta_1 r_t^{(n-k)} + \delta_2 spread_t + \delta_3 risk_t + \vartheta_t^{(n-k)}. \tag{11}$$

where *spread* is defined as $r_t^{(n)} - r_t^{(k)}$. Although no particular measure of risk has been consistently been found to be significant for term premia, we use the crude proxy *risk*, defined as the absolute percentage change in the yield on the *n*-period bond, or $|(r_t^{(n)} - r_{t-1}^n)/r_{t-1}^n|$. Least squares estimates of equation (11) are presented in Table 9. Consistent with other studies, such as Bekdache (2001) or Hejazi et al. (2000) we find that the yield spread is highly significant for all but two interest rate deposits, and the relationship between the spread and the premium has the expected sign. The level of the 3-month spot interest rate also appears significantly in a number of cases, yet the risk measure seems to be irrelevant.

Insert table 9 here

It can be questioned whether the finding that risk measures are irrelevant is plausible or whether the risk proxy actually measures risk. The presence of strong ARCH effects in the tests of time-variation in the term premia, as well as in some of the low-order ARMA models, suggests that at least the inclusion of the second moments are relevant for explaining the time-variation of the premia. We therefore use a GARCH-M model, originally due to Engle et al. (1987), where the risk premia are heteroskedastic and the standard deviation of each observation directly affects the mean of that observation, or

$$\varphi_t^{(n,k)} = \alpha_0 + \alpha_1 h_t + \varsigma_t^{(n-k)}$$
$$h_t^2 = \beta_0 + \sum_{i=1}^{p} \beta_i \varepsilon_{t-i}^2 + \sum_{i=1}^{q} \gamma_i h_{t-i}^2 . \tag{12}$$

where $\varsigma_t^{(n-k)} \sim N(0, h_t^2)$. To preserve model parsimony and keeping the summary statistics in Table 3 in mind that show ARCH effects usually up to only lag, we select an GARCH(1,1)-M specification.

Insert table 10 here

The parameters in equation (12) are estimated by maximum likelihood using the BHHH algorithm and the standard errors are heteroskedasticity-consistent. Estimates are reported in Table 10. The estimates for β_1 and γ_1 are both highly significant for most countries, indicating that that the GARCH specification is relevant. The inclusion of the conditional standard deviation in the mean equation results in significant parameter estimates for about half of the deposits, indicating that the variation in the term premia can be explained by a risk component. Furthermore, the conditional standard deviation has the correct sign: an upward change in the conditional standard deviation leads to an upward change in the level of the term premium.

Finally, we document that there is considerable evidence of a high degree of persistence in the volatility of term premia, given large estimates for β_1 and γ_1. Tzavalis and Wickens (1995) show that when $\beta_1 + \gamma_1 < 1$, the conditional variance of innovations in interest rates has a direct, yet fading, effect on the slope of the yield curve, while if $\beta_1 + \gamma_1 > 1$ the conditional variance is potentially explosive. For the majority of the interest rate deposits we find that the sum of β_1 and γ_1 is smaller than unity. However, for the remaining models, we cannot reject the null hypothesis that the model is integrated in variance.

7. Conclusions

In the current paper we examine the expectations hypothesis using survey data for a large set of international, short-maturity interest rate deposits. In contrast to commonly-cited results from U.S. studies we find that, although the ‘pure’ version of the expectations hypothesis can be firmly rejected for all deposits, forward rates are not biased estimates of future interest rates. Nevertheless we find some evidence of irrationality on the part of market participants, and strong evidence of the existence of a time-varying term premium in the short end of the term structure of our deposits.

Term premia furthermore are commonly positive, suggesting that the premia could be liquidity premia as proclaimed by the liquidity preference theory. A closer look at the time-varying term premia reveals a large degree of persistence, which is in line with intertemporal general equilibrium asset pricing models, such as for instance those of Vasicek (1977) or Cox et al. (1985). We find that low-order versions of ARMA models describe the term premia well, although low-order ARCH effects are also present in several interest rate deposits. Finally, we find that although continental European term premia show lower degrees of persistence than other premia, the former nevertheless are time-variant and the survey respondents did make irrational forecasts, in that not all available information to market participants was used efficiently.

References

Ang, A, Bekaert, G., and Wei, M. (2005). Do Macro Variables, Asset Markets or Surveys Forecast Inflation Better?, NBER Working Paper 11538.

Ang, A. and Piazzesi, M. (2003). A No-Arbitrage Vector Autoregression of Term Structure Dynamics with Macroeconomic and Latent Variables, *Journal of Monetary Economics*, 50(4): 745-787.

Bekaert, G., Hodrick, R.J., and Marshall, D.A. (1997). On Biases in Tests of the Expectations Hypothesis of the Term Structure of Interest Rates, *Journal of Financial Economics*, 44(3): 309-348.

Bekaert, G., Hodrick, R.J., and Marshall, D.A. (2001). Peso Problem Explanations for Term Structure Anomalies, *Journal of Monetary Economics*, 48(2): 241-270.

Bekdache, B. (2001). Term Premia and the Maturity Composition of the Federal Debt: New Evidence from the Term Structure of Interest Rates, *Journal of Forecasting*, 20(7): 519-539.

Branch, W.A. (2004). The Theory of Rationally Heterogeneous Expectations: Evidence from Survey Data on Inflation Expectations, *Economic Journal*, 114(497): 592-621.

Campbell, J.Y. (1995). Some Lessons from the Yield Curve, *Journal of Economic Perspective*, 9(3): 129-152.

Cavaglia, S., Verschoor, W.F.C., and Wolff, C.C.P. (1993). Further Evidence of Exchange Rate Expectations, *Journal of International Money and Finance*, 12(1): 78-98.

Cavaglia, S.M.F.G., Verschoor, W.F.C., and Wolff, C.C.P. (1994). On the Biasedness of Forward Foreign Exchange Rates: Irrationality or Risk Premia?, *Journal of Business*, 67(3): 321-343.

Christiansen, C. (2003). Testing the Expectations Hypothesis Using Long-Maturity Forward Rates, *Economics Letters*, 78(2): 175-180.

Cox, J.C., Ingersoll, J.E., and Ross, S.A. (1985). A Theory of the Term Structure of Interest Rates, *Econometrica* 53(2): 385-407.

Dahlquist, M. and Jonsson, G. (1995). The Information in Swedish Short-Maturity Forward Rates, *European Economic Review*, 39(6): 1115-1131.

Dai, Q. and Singleton, K.J. (2002). Expectations Puzzles, Time-Varying Risk Premia, and Affine Models of the Term Structure, *Journal of Financial Economics*, 63(3): 415-441.

Deaves, R. (1998). Term Premium Determinants, Return Enhancement and Interest Rate Predictability, *Journal of Business Finance and Accounting*, 25(3-4): 485-499.

Dokko, Y. and Edelstein, R.H. (1989). How Well do Economists Forecast Stock Market Prices? A Study of the Livingston Surveys, *American Economic Review*, 79(4): 865-871.

Domínguez, E. and Novales, A. (2000). Testing the Expectations Hypothesis in Eurodeposits, *Journal of International Money and Finance*, 19(5): 713-736.

Elliott, G. and Ito, T. (1999). Heterogeneous Expectations and Tests of Efficiency in the Yen/Dollar Forward Exchange Rate Market, *Journal of Monetary Economics*, 43(2): 435-456.

Engel, C. (1996). The Forward Discount Anomaly and the Risk Premium: A Survey of Recent Evidence, *Journal of Empirical Finance*, 3(2): 123-192.

Engle, R.F., Lilien, D.M. and Robins, R.P. (1987). Estimating Time Varying Risk Premia in the Term Structure: The ARCH-M Model, *Econometrica*, 55(2): 391-407.

Fama, E.F. (1984). The Information in the Term Structure, *Journal of Financial Economics*, 13(4): 509-528.

Frankel, J.A. and Froot, K.A. (1987). Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations, *American Economic Review*, 77(1): 133-153.

Friedman, B.M. (1979). Interest Rate Expectations Versus Forward rates: Evidence from an Expectations Survey, *Journal of Finance*, 34(4): 965-973.

Froot, K.A. (1989). New Hope for the Expectations Hypothesis of the Term Structure of Interest Rates, *Journal of Finance*, 44(2): 283-305.

Gerlach, S. and Smets, F. (1997). The Term Structure of Euro-Rates: Some Evidence in Support of the Expectations Hypothesis, *Journal of International Money and Finance*, 16(2): 305-321.

Hafer, R.W., Hein, S.E. and MacDonald, S.S. (1992). Market and Survey Forecast of the Three-Month Treasury-Bill Rate, *Journal of Business*, 65(1): 123-138.

Harris, R.D.F. (2001). The Expectations Hypothesis of the Term Structure and Time-Varying Risk Premia: A Panel Data Approach, *Oxford Bulletin of Economics and Statistics*, 63(2): 233-245.

Hejazi, W., Lai, H. and Yang, X. (2000). The Expectations Hypothesis, Term Premia, and the Canadian Term Structure of Interest Rates, *Canadian Journal of Economics*, 33(1): 133-148.

Henry, O.T. (1999). The Volatility of US Term Structure Term Premia 1952-1991, *Applied Financial Economics*, 9(3): 263-271.

Hodrick, R.J. (1987). *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets*, Chur, Switzerland: Harwood Academic Publishers.

Jondeau, E. and Ricart, R. (1999). The Expectations Hypothesis of the Term Structure: Tests on US, German, French, and UK Euro-Rates, *Journal of International Money and Finance*, 18(5): 725-750.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P. and Shin, Y. (1992). Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root, *Journal of Econometrics*, 54(1-3): 159-178.

Lee, S.-S. (1995). Macroeconomic Source of Time-Varying Risk Premia in the Term Structure of Interest Rates, *Journal of Money, Credit, and Banking*, 27(2): 549-569.

Lewis, K.K. (1989). Changing Beliefs and Systematic Rational Forecast Errors with Evidence from Foreign Exchange, *American Economic Review*, 79(4): 621-636.

MacDonald, R. and Macmillan, P. (1994). On the Expectations View of the Term Structure, Term Premia and Survey-Based Expectations, *Economic Journal*, 104(425): 1070-1086.

Rapach, D.E. and Weber, C.E. (2004). Are real Interest Rates Really Nonstationary? New Evidence from Tests with Good Size and Power, *Journal of Macroeconomics*, 26(3): 409-430.

Rudebusch, G.D. and Wu, T. (2004). A Macro-Finance Model of the Term Structure, Monetary Policy, and the Economy, Federal Reserve Bank of San Francisco Working Paper 2003-17.

Rose, A.K. (1988). Is the Real Interest Rate Stable?, *Journal of Finance*, 43(5): 1095-1112.

Shiller, R.J. (1990). The Term Structure of Interest Rates, in: B.M. Friedman and F.H. Hahn, eds., *Handbook of Monetary Economics*, 1, Amsterdam: North-Holland, 627-722.

Shiller, R.J., Campbell, J.Y. and Schoenholtz, K.L. (1983). Forward Rates and Future Policy: Interpreting the Term Structure of Interest Rates, *Brookings Papers on Economic Activity*, 1: 173-223.

Sims, C.A. (2003). Implications of Rational Inattention, *Journal of Monetary Economics*, 50(3): 665-690.

Tzavalis, E. and Wickens, M.R. (1995). The Persistence in Volatility of the US Term Premium 1970-1986, *Economics Letters*, 49(4): 381-389.

Vasicek, O. (1977). An Equilibrium Characterization of the Term Structure, *Journal of Financial Economics*, 5(2): 177-188.

Table 1. Forecast Performance of Survey-Based, Forward, and Random Walk Forecasts

	MAE			RMSE			*U*-statistic
	forward	survey	no-change	forward	survey	no-change	
a) Continental Europe							
Austria	0.182	0.163	0.148	0.280	0.236	0.259	0.912
Belgium	0.208	0.190	0.178	0.297	0.277	0.259	1.071
Europe	0.175	0.228	0.248	0.242	0.308	0.363	0.847
France	0.219	0.225	0.251	0.322	0.306	0.381	0.802
Germany	0.173	0.204	0.209	0.244	0.278	0.319	0.870
Ireland	0.339	0.357	0.350	0.498	0.530	0.630	0.841
Italy	0.338	0.257	0.389	0.494	0.343	0.578	0.594
Netherlands	0.196	0.216	0.225	0.267	0.291	0.324	0.900
Spain	0.199	0.254	0.310	0.281	0.345	0.436	0.791
Switzerland	0.307	0.315	0.298	0.411	0.430	0.424	1.015
b) Asia-Pacific							
Australia	0.267	0.295	0.276	0.349	0.393	0.386	1.016
Hong Kong	1.055	0.809	0.931	1.870	1.526	1.827	0.835
Japan	0.095	0.089	0.087	0.145	0.142	0.152	0.933
New Zealand	0.493	0.531	0.550	0.743	0.738	0.794	0.929
c) Anglo-Saxon Countries							
Canada	0.403	0.333	0.350	0.559	0.479	0.503	0.953
U.K.	0.254	0.267	0.280	0.334	0.369	0.400	0.924
U.S.	0.243	0.305	0.276	0.372	0.458	0.444	1.031
d) Scandinavia							
Denmark	0.277	0.201	0.263	0.391	0.303	0.401	0.755
Norway	0.426	0.513	0.517	0.669	0.756	0.786	0.961
Sweden	0.333	0.317	0.326	0.461	0.414	0.512	0.810

Notes: MAE and RMSE are the mean absolute error and the root mean squared error, respectively. The *U*-statistic is Theil's *U*-statistic, which is expressed as the ratio of the RMSE of a survey-based forecast divided by the RMSE of a no-change (random walk) forecast.

Table 2. Summary Statistics of Realized and Expected Interest Rate Series

	$r_t^{(3)}$				$\Delta r_t^{(3)}$				$r_{t+k}^{e(3)}$			
	Mean	Stdev.	High	Low	Mean	Stdev.	High	Low	Mean	Stdev.	High	Low
a) Continental Europe												
Austria (VIBOR)	3.772	0.535	5.208	3.242	-0.037	0.138	0.250	-0.570	3.691	0.626	5.300	3.000
Belgium (euro-Franc)	3.805	0.775	6.875	3.060	-0.041	0.390	1.500	-1.687	3.804	0.848	6.200	3.000
Europe (Euribor)	3.232	0.966	5.112	2.025	-0.007	0.167	0.571	-0.706	3.202	0.973	5.100	1.900
France (euro-Franc)	4.376	1.404	8.750	3.234	-0.056	0.596	3.125	-1.312	4.102	1.074	6.600	3.100
Germany (euro-Mark)	3.612	0.551	5.062	2.984	-0.036	0.135	0.312	-0.375	3.647	0.548	5.100	3.100
Ireland (interbank)	5.847	0.678	7.375	3.406	-0.061	0.344	1.250	-1.093	5.568	0.889	7.100	3.100
Italy (treasury bill)	7.842	2.279	11.840	4.160	-0.109	0.521	1.920	-1.020	7.247	2.441	11.300	3.300
Netherlands (euro-Guilder)	3.523	0.616	5.187	2.625	-0.040	0.173	0.343	-0.390	3.608	0.646	5.400	2.700
Spain (euro-Peseta)	6.622	2.078	9.875	3.297	-0.126	0.302	1.250	-0.781	6.497	2.228	10.000	3.200
Switzerland (euro-Franc)	1.729	1.047	4.187	0.171	-0.029	0.255	1.156	-0.843	1.832	1.006	4.200	0.300
b) Asia-Pacific												
Australia (dealer bill)	5.453	0.863	7.620	4.180	-0.019	0.174	0.300	-0.700	5.491	0.881	7.700	4.100
Hong Kong (interbank)	4.573	3.012	15.500	0.105	-0.049	1.321	6.485	-6.218	4.473	2.548	10.000	0.300
Japan (certificate of deposit)	0.410	0.435	2.350	0.050	-0.019	0.105	0.400	-0.550	0.410	0.424	2.400	0.100
New Zealand (bank bill)	6.535	1.532	9.970	4.040	-0.014	0.386	1.380	-1.340	6.488	1.341	9.900	4.300
c) Anglo-Saxon Countries												
Canada (treasury bill)	4.079	1.483	8.170	1.760	-0.034	0.273	1.270	-0.870	4.209	1.456	8.200	1.800
United Kingdom (interbank)	5.569	1.228	7.690	3.375	-0.014	0.184	0.562	-0.656	5.645	1.229	7.700	3.500
United States (treasury bill)	3.815	1.800	6.190	0.850	-0.030	0.225	0.730	-1.050	4.008	1.813	6.300	1.000
d) Scandinavia												
Denmark (euro-Krone)	4.467	1.023	6.875	3.560	-0.041	0.312	0.968	-0.718	4.445	1.037	7.000	3.500
Norway (euro-Krone)	5.602	1.993	8.050	1.625	-0.034	0.402	2.485	-0.859	5.543	1.897	7.600	1.900
Sweden (euro-Krona)	4.437	1.790	9.100	2.090	-0.050	0.251	0.796	-1.425	4.481	1.779	9.600	2.100

Notes: Sample is from January 1995 through December 2004. Stdev. is the standard deviation. The first column gives a description of the 3-month spot rate series used for the particular countries. The choice for a particular series is driven by the availability of survey data. For some continental European deposits the sample ranges until the introduction of the Euribor rate.

Table 3. Summary Statistics of the Term Premia

	Mean	Stdev.	Skewness	Kurtosis	BJ	% > 0	$\tau_1(\varphi)$	$\tau_1(\varphi^2)$
a) Continental Europe								
Austria	0.194	0.150	0.56	-0.19	2.58	89	0.178	0.320**
Belgium	0.101	0.240	0.27	0.42	0.94	68	0.177	0.099
Europe	0.072	0.162	0.57	1.67	11.84***	67	0.708***	0.612***
France	0.154	0.215	-0.15	2.43	12.03***	95	0.141	0.031
Germany	0.022	0.143	0.43	0.44	1.75	56	0.416***	0.106
Ireland	0.092	0.419	0.04	-0.54	0.59	54	0.577***	0.424***
Italy	0.203	0.416	-0.75	3.33	26.69***	75	0.136	-0.034
Netherlands	0.056	0.136	0.58	-0.39	3.05	60	0.184	-0.001
Spain	-0.009	0.235	2.40	8.35	186.07***	33	0.044	0.032
Switzerland	0.051	0.233	0.90	1.47	27.14***	53	0.492***	0.411***
b) Asia-Pacific								
Australia	-0.045	0.177	0.36	-0.31	2.60	37	0.599***	0.379***
Hong Kong	0.474	0.869	2.87	11.97	530.91***	82	0.668***	0.417***
Japan	-0.016	0.097	0.08	1.22	7.60**	39	0.528***	0.237***
New Zealand	0.066	0.234	-0.15	0.20	0.64	62	0.575***	0.453***
c) Anglo-Saxon Countries								
Canada	0.146	0.234	0.17	1.64	14.18***	74	0.523***	0.284***
United Kingdom	-0.014	0.180	0.22	4.11	85.51***	43	0.412***	0.037
United States	-0.080	0.209	0.20	1.74	16.05***	30	0.397***	0.232***
d) Scandinavia								
Denmark	0.135	0.194	0.15	-0.29	0.34	75	0.272**	0.081
Norway	-0.153	0.240	0.12	0.14	0.27	23	0.764***	0.594***
Sweden	0.020	0.222	1.43	4.72	153.00***	48	0.423***	0.500***

Notes: Sample is from January 1995 through December 2004. Stdev. is the standard deviation. Kurtosis is in fact *excess* kurtosis. BJ is the Bera-Jarque statistic for normality of the distribution. % > 0 is a sign test statistic that gives the percentage of months the term premium is positive over the sample period. $\tau_1(\varphi)$ is the first order autocorrelation of the premium, and $\tau_1(\varphi^2)$ for it's squared; significance levels come from Ljung-Box *Q*-statistics. A *, **, and *** means rejection of the null hypothesis at a 1, 5, and 10 percent significance level, respectively.

Table 4: Unit Root Test Coefficients and First Order Auto-Correlation Coefficients

	PP					KPSS				
	$r_t^{(3)}$	$r_{t+3}^{e(3)}$	$f_t^{(6,3)}$	$r_{t+3}^{e(3)} - r_t^{(3)}$	$f_t^{(6,3)} - r_t^{(3)}$	$r_t^{(3)}$	$r_{t+3}^{e(3)}$	$f_t^{(6,3)}$	$r_{t+3}^{e(3)} - r_t^{(3)}$	$f_t^{(6,3)} - r_t^{(3)}$
a) Continental Europe										
Austria	-3.12**	-2.59	-2.58	-3.50**	-1.82	0.47**	0.39*	0.45*	0.13	0.22
	(0.884)	(0.861)	(0.877)	(0.527)	(0.691)					
Belgium	-2.03	-2.01	-2.19	-6.21***	-6.24***	0.43*	0.42*	0.43*	0.22	0.37*
	(0.833)	(0.907)	(0.864)	(0.087)	(0.127)					
Europe	-0.81	-0.99	-0.96	-3.14**	-2.54	0.60**	0.61**	0.68**	0.15	0.27
	(0.956)	(0.960)	(0.954)	(0.734)	(0.831)					
France	-1.41	-1.37	-1.89	-4.26***	-4.37***	0.61**	0.61**	0.67**	0.66**	0.33
	(0.891)	(0.944)	(0.902)	(0.443)	(0.480)					
Germany	-3.09**	-2.84*	-3.34**	-2.86*	-3.74***	0.45*	0.38*	0.38*	0.35*	0.16
	(0.892)	(0.895)	(0.846)	(0.611)	(0.542)					
Ireland	-0.49	0.03	-0.55	-2.34	-2.14	0.21	0.44*	0.61**	0.49**	0.79***
	(0.727)	(0.836)	(0.828)	(0.773)	(0.811)					
Italy	0.13	0.40	0.39	-4.32***	-3.25**	0.85***	0.86***	0.87***	0.50**	0.64**
	(0.941)	(0.954)	(0.927)	(0.409)	(0.606)					
Netherlands	-3.21**	-3.04**	-3.16**	-3.51**	-3.68***	0.32	0.33	0.29	0.07	0.14
	(0.866)	(0.875)	(0.846)	(0.381)	(0.507)					
Spain	0.15	0.11	-0.80	-3.63***	-4.17***	0.87***	0.87***	0.88***	0.55**	0.53**
	(0.944)	(0.956)	(0.936)	(0.565)	(0.716)					
Switzerland	-2.26	-2.31	-2.62*	-6.80***	-5.73***	0.44*	0.45*	0.47**	0.09	0.09
	(0.935)	(0.947)	(0.927)	(0.416)	(0.561)					
b) Asia-Pacific										
Australia	-2.41	-2.55	-2.39	-4.21***	-3.41**	0.45*	0.39*	0.38*	0.24	0.13
	(0.954)	(0.947)	(0.940)	(0.726)	(0.816)					
Hong Kong	-1.62	-0.62	-1.18	-8.48***	-9.12***	0.91***	0.92***	0.88***	0.46*	0.17
	(0.877)	(0.963)	(0.927)	(0.232)	(0.269)					
Japan	-5.76***	-5.80***	-5.90***	-4.44***	-4.81***	1.14***	1.14***	1.16***	0.11	0.08
	(0.871)	(0.879)	(0.855)	(0.716)	(0.702)					
New Zealand	-2.02	-2.03	-2.02	-3.79***	-3.21**	0.68**	0.65**	0.69**	0.48**	0.38*
	(0.962)	(0.969)	(0.958)	(0.756)	(0.839)					
c) Anglo-Saxon Countries										
Canada	-1.69	-2.08	-2.46	-5.50***	-6.29***	0.67**	0.70**	0.80***	0.05	0.41*
	(0.957)	(0.954)	(0.932)	(0.691)	(0.669)					
U.K.	-1.31	-1.49	-1.55	-3.65***	-4.79***	0.94***	0.93***	0.96***	0.09	0.06
	(0.982)	(0.980)	(0.968)	(0.799)	(0.692)					
U.S.	-1.03	-1.21	-1.49	-5.74***	-5.78***	0.97***	0.96***	0.97***	0.05	0.11
	(0.976)	(0.979)	(0.970)	(0.549)	(0.664)					
d) Scandinavia										
Denmark	-1.53	-1.77	-1.96	-3.05**	-3.86***	0.51**	0.51**	0.55**	0.10	0.18
	(0.928)	(0.937)	(0.911)	(0.648)	(0.534)					
Norway	-0.50	-0.49	-0.33	-2.39	-2.72*	0.74***	0.65**	0.71**	0.57**	0.21
	(0.911)	(0.918)	(0.919)	(0.839)	(0.842)					
Sweden	-1.67	-1.99	-2.06	-4.36***	-4.14***	0.83***	0.78***	0.83***	0.19	0.09
	(0.950)	(0.951)	(0.943)	(0.714)	(0.756)					

Notes: Sample is from January 1995 through December 2004. Given are the test statistic for the respective unit root/stationarity test. First order autocorrelation coefficient of the series are given between parentheses. A *, **, and *** means rejection of the null hypothesis at a 1, 5, and 10 percent significance level, respectively. For the Phillips-Perron (PP) test the null hypothesis is that of a unit root, while for the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test the null hypothesis is that of a stationary series.

Table 5. Forward Rate Unbiasedness

	$\hat{\alpha}_1$	$\hat{\beta}_1$	t: $\beta_1 = 1$	χ^2
a) Continental Europe				
Austria	-0.1110	-0.0810	2.59***	37.05***
	(0.0908)	(0.4161)		(0.0000)
Belgium	-0.1794	-0.0154	3.47***	47.77***
	(0.1105)	(0.1915)		(0.000)
Europe	-0.0707*	1.1624***	0.96	3.45
	(0.0415)	(0.1677)		(0.1777)
France	-0.2155**	1.1733***	1.11	6.81**
	(0.0901)	(0.1552)		(0.0331)
Germany	-0.1262**	0.2135	-2.44**	15.67***
	(0.0588)	(0.3213)		(0.0004)
Ireland	-0.0959	0.6337*	-1.03	1.96
	(0.1110)	(0.3529)		(0.3746)
Italy	-0.1046	0.8293**	-0.50	0.36
	(0.1749)	(0.3363)		(0.8350)
Netherlands	-0.1146	-0.0779	2.35**	21.85***
	(0.1019)	(0.4567)		(0.0000)
Spain	-0.2904***	0.6461***	-2.25**	26.36***
	(0.0628)	(0.1570)		(0.0000)
Switzerland	-0.2199***	0.8837***	-0.60	19.38***
	(0.0585)	(0.1923)		(0.0000)
b) Asia-Pacific				
Australia	-0.0464	0.7793***	-1.20	2.94
	(0.0511)	(0.1831)		(0.2290)
Hong Kong	-0.3793	0.6336	-0.80	6.42**
	(0.3023)	(0.4523)		(0.0403)
Japan	-0.0489	0.4890***	-3.24***	11.53***
	(0.0330)	(0.1574)		(0.0031)
New Zealand	-0.0512	0.7825***	-0.86	1.56
	(0.1248)	(0.2516)		(0.4581)
c) Anglo-Saxon Countries				
Canada	-0.3892***	0.9439***	0.35	33.33***
	(0.0784)	(0.1572)		(0.0000)
United Kingdom	-0.0929	0.7819***	-1.04	7.25**
	(0.0561)	(0.2091)		(0.0265)
United States	-0.2092***	1.0845***	0.30	20.84***
	(0.0644)	(0.2764)		(0.0000)
d) Scandinavia				
Denmark	-0.2827***	1.1771***	0.72	9.34***
	(0.1038)	(0.2458)		(0.0094)
Norway	0.0975	1.0075***	0.00	0.70
	(0.1271)	(0.2380)		(0.7068)
Sweden	-0.2308***	1.1469***	0.53	18.52***
	(0.0667)	(0.2732)		(0.0000)

Notes: Sample is from January 1995 through December 2004. Reported are GMM regression results. The standard errors of the coefficients are given in parentheses. A *, **, *** denotes rejection at a 10, 5, and 1 percent significance level. t: $\beta = 1$ reports the t-statistic and significance levels for the hypothesis that $\beta = 1$. The χ^2 statistic pertains to the *joint* hypothesis that $\alpha = 0$ and $\beta = 1$ (p-values are given in parentheses).

Table 6. Error Orthogonality to Forward Premia

	$\hat{\alpha}_2$	$\hat{\beta}_2$	χ^2
a) Continental Europe			
Austria	0.0518	-0.8051*	4.37
	(0.0798)	(0.4284)	(0.1124)
Belgium	-0.1138	-0.6548*	7.44**
	(0.1115)	(0.3367)	(0.0242)
Europe	-0.0176	0.7066***	15.43***
	(0.0494)	(0.1802)	(0.0004)
France	0.0414	0.3027	2.32
	(0.0867)	(0.2223)	(0.3124)
Germany	-0.1394***	-0.1698	9.95***
	(0.0491)	(0.2653)	(0.0069)
Ireland	0.0525	-0.0654	0.21
	(0.1248)	(0.3693)	(0.8996)
Italy	0.3068*	0.3612	3.63
	(0.1640)	(0.2766)	(0.1628)
Netherlands	-0.1458	-0.4571	9.07**
	(0.0947)	(0.4587)	(0.0107)
Spain	-0.2380***	0.1080	14.42***
	(0.0628)	(0.1274)	(0.0007)
Switzerland	-0.2705***	0.5669**	18.97***
	(0.0642)	(0.2245)	(0.0001)
b) Asia-Pacific			
Australia	-0.0845	0.3728*	4.89*
	(0.0583)	(0.2076)	(0.0867)
Hong Kong	-0.0473	0.0692	0.03
	(0.2852)	(0.4488)	(0.9846)
Japan	-0.0563	0.0150	2.78
	(0.0338)	(0.1841)	(0.2489)
New Zealand	0.0102	0.1126	0.27
	(0.1273)	(0.2688)	(0.8707)
c) Anglo-Saxon Countries			
Canada	-0.3868***	0.4880***	25.55***
	(0.0774)	(0.1560)	(0.0000)
United Kingdom	-0.1357**	0.2701	6.02**
	(0.0578)	(0.2056)	(0.0492)
United States	-0.3505***	0.6436**	30.41***
	(0.0679)	(0.2785)	(0.0000)
d) Scandinavia			
Denmark	-0.1552*	0.2463***	3.22
	(0.0865)	(0.2478)	(0.1993)
Norway	0.0554	0.3722	1.79
	(0.1434)	(0.2837)	(0.4067)
Sweden	-0.2281***	0.4785**	15.33***
	(0.0582)	(0.2344)	(0.0005)

Notes: Sample is from January 1995 through December 2004. A *, **, *** denotes rejection at a 10, 5, and 1 percent significance level. The χ^2 statistic pertains to the *joint* hypothesis that $\alpha = 0$ and $\beta = 0$ (p-values are given in parentheses).

Table 7. Time-Variation in Term Premia

	$\hat{\alpha}_3$	$\hat{\beta}_3$	t: $\beta_3 = 1$	χ^2	BG (opt. # lags)	ARCH (opt. # lags)
a) Continental Europe						
Austria	-0.1629***	0.7241***	-2.70***	66.88***	2.62	1.68
	(0.0232)	(0.1019)		(0.0000)	(1)	(1)
Belgium	-0.0656*	0.6393***	-1.54	8.75**	0.99	0.99
	(0.0363)	(0.2333)		(0.0126)	(1)	(1)
Europe	-0.0493***	0.4585***	-8.15***	80.68***	20.71***	0.59
	(0.0171)	(0.0664)		(0.0000)	(1)	(1)
France	-0.2570***	0.8706***	-0.96	57.75***	7.56**	0.24
	(0.0360)	(0.1345)		0.0000)	(2)	(1)
Germany	0.0132	0.3833***	-7.99***	81.98***	12.13***	7.79***
	(0.0211)	(0.0771)		(0.0000)	(1)	(1)
Ireland	-0.1485**	0.6991***	-2.71***	13.25***	17.24***	11.04***
	(0.0756)	(0.1108)		(0.0013)	(1)	(1)
Italy	-0.4115***	0.4680***	-5.30***	52.30***	0.03	1.64
	(0.0572)	(0.1006)		(0.0000)	(1)	(3)
Netherlands	0.0311	0.3792***	-7.42***	70.21***	1.91	0.00
	(0.0189)	(0.0836)		(0.0000)	(1)	(1)
Spain	-0.0524	0.5381***	-3.21***	19.94***	0.46	11.45***
	(0.0423)	(0.1435)		(0.0000)	(1)	(1)
Switzerland	0.0536**	0.3217***	-8.86***	85.86***	20.36***	5.54**
	(0.0260)	(0.0765)		(0.0000)	(1)	(1)
b) Asia-Pacific						
Australia	0.0410**	0.4087***	-9.04***	89.45***	32.93***	12.80***
	(0.0175)	(0.0654)		(0.0000)	(1)	(1)
Hong Kong	-0.3127	0.5683**	-1.54	28.30***	66.82***	18.33***
	0.1961	0.2798		(0.0000)	(6)	(2)
Japan	0.0080	0.4735***	-5.94***	35.55***	32.74***	26.44***
	(0.0078)	(0.0885)		(0.0000)	(1)	(1)
New Zealand	-0.0604**	0.6696***	-3.86***	43.67***	29.63***	30.25***
	(0.0281)	(0.0855)		(0.0000)	(1)	(1)
c) Anglo-Saxon Countries						
Canada	0.0044	0.4532***	-10.03***	228.28***	38.17***	5.70***
	(0.0299)	(0.0544)		(0.0000)	(1)	(2)
United Kingdom	0.0442***	0.5116***	-7.17***	54.95***	30.55***	1.59
	(0.0163)	(0.0681)		(0.0000)	(1)	(1)
United States	0.1421***	0.4530***	-9.36***	145.78***	12.32***	14.26***
	(0.0157)	(0.0584)		(0.0000)	(1)	(1)
d) Scandinavia						
Denmark	-0.1275***	0.9308***	-0.77	15.39***	4.67**	1.78
	(0.0329)	(0.0889)		(0.0005)	(1)	(1)
Norway	0.0772**	0.6415***	-5.77***	38.83***	48.15***	42.08***
	(0.0354)	(0.0620)		(0.0000)	(2)	(1)
Sweden	0.0004	0.6738***	-3.54***	22.48***	27.46***	36.35***
	(0.0297)	(0.0919)		(0.0000)	(1)	(1)

Notes: Sample is from January 1995 through December 2004. The standard errors of the coefficients are given in parentheses. A *, **, *** denotes rejection at a 10, 5, and 1 percent significance level. t: $\beta = 1$ reports the t-statistic and significance levels for the hypothesis that $\beta = 1$. The χ^2 statistic pertains to the *joint* hypothesis that $\alpha = 0$ and $\beta = 1$. BG is the Breusch-Godfrey Lagrange Multiplier statistics for residual serial correlation (optimal lag-length in parentheses). ARCH is a Lagrange Multiplier statistic for the presence of autoregressive conditional heteroskedasticity in the residuals (optimal lag-length in parentheses).

Table 8. ARMA Specifications for Term Premia

	$\hat{\gamma}_0$	$\hat{\gamma}_1$	$\hat{\gamma}_2$	R^2_{adj}	LM	ARCH
a) Continental Europe						
Austria	0.1940***	-	0.2243	0.02	0.19	0.45
	(0.0262)		(0.1442)		(1)	(1)
Belgium	0.1017**	-	0.1789	0.01	0.03	0.14
	(0.0405)		(0.1451)		(1)	(1)
Europe	0.0768	0.7087***	-	0.49	0.28	2.83
	(0.0482)	(0.0863)			(1)	(1)
France	0.2373***	0.1445	-	0.00	0.58	0.24
	(0.0332)	(0.1326)			(1)	(1)
Germany	0.0082	0.4163***	-	0.18	0.54	0.42
	(0.0305)	(0.1232)			(1)	(1)
Ireland	0.0692	0.5773***	-	0.32	2.15	0.62
	(0.1185)	(0.1193)			(1)	(1)
Italy	0.1666	0.7979***	0.9627	0.05	1.61	0.10
	(0.0272)	(0.1025)	(0.0295)		(1)	(1)
Netherlands	0.0568**	-	0.2896**	0.03	1.89	0.00
	(0.0249)		(0.1409)		(1)	(1)
Spain	-0.0331	0.0443	-	-0.01	0.87	15.36***
	(0.0274)	(0.1106)			(1)	(3)
Switzerland	0.0446	0.4926***	-	0.24	1.15	1.42
	(0.0365)	(0.0792)			(1)	(1)
b) Asia-Pacific						
Australia	-0.0411	0.6045***	-	0.36	1.49	0.13
	(0.0341)	(0.0760)			(1)	(1)
Hong Kong	0.4497	0.8547***	-0.3297**	0.48	19.22***	18.55***
	(0.2776)	(0.0703)	(0.1262)		(2)	(2)
Japan	-0.0206	0.5292***	-	0.28	0.88	15.99**
	(0.0157)	(0.0761)			(1)	(6)
New Zealand	0.0708	0.5753***	-	0.32	0.10	3.91**
	(0.0434)	(0.0786)			(1)	(1)
c) Anglo-Saxon Countries						
Canada	0.1279***	0.5410***	-	0.31	0.00	0.14
	(0.0371)	(0.0731)			(1)	(1)
United Kingdom	-0.0195	0.4134***	-	0.16	0.44	0.96
	(0.0254)	(0.0872)			(1)	(1)
United States	-0.0906***	0.3953***	-	0.16	0.05	8.44***
	(0.0249)	(0.0792)			(1)	(1)
d) Scandinavia						
Denmark	0.1354***	-	0.3019**	0.06	0.02	3.33*
	(0.0351)		(0.1407)		(1)	(1)
Norway	-0.1646**	0.7631***	-	0.58	6.75**	18.55***
	(0.0747)	(0.0735)			(2)	(5)
Sweden	0.0208	-	0.4665***	0.17	0.06	20.93***
	(0.0269)		(0.0810)		(1)	(1)

Notes: Sample is from January 1995 through December 2004. Reported are estimates for an AR(1), MA(1), or ARMA(1,1) model, where $\varphi_t^{(n,k)} = \gamma_0 + \gamma_1 \varphi_{t-1}^{(n,k)} + \eta_t^{(n,k)} + \gamma_2 \eta_{t-1}^{(n,k)}$. The standard errors of the coefficients are given in parentheses. A *, **, *** denotes rejection at a 10, 5, and 1 percent significance level. ARCH is a Lagrange Multiplier statistic for the presence of autoregressive conditional heteroskedasticity in the residuals (optimal lag-length in parentheses) and BG is the Breusch-Godfrey Lagrange Multiplier statistic for residual serial correlation (optimal lag-length in parentheses).

Table 9. Linear Model of the Term Premium

	$\hat{\delta}_0$	$\hat{\delta}_1$	$\hat{\delta}_2$	$\hat{\delta}_3$
a) Continental Europe				
Austria	0.2698	-0.0414	0.6946*	1.1403
	(0.1970)	(0.0570)	(0.3958)	(0.7227)
Belgium	-0.1693	0.0785	0.7039	-1.5659**
	(0.1970)	(0.0577)	(0.4390)	(0.6598)
Europe	0.0858	-0.0185	1.0771***	0.3131
	(0.0416)	(0.0122)	(0.0965)	(0.3041)
France	-0.2470	0.1058***	0.8271***	0.3459
	(0.1208)	(0.0285)	(0.1973)	(0.4751)
Germany	-0.1451	0.0368	1.0732***	-0.6058
	(0.0994)	(0.0270)	(0.1653)	(0.5102)
Ireland	-0.1114	0.0340	0.5741**	0.7484
	(0.5760)	(0.0945)	(0.2182)	(1.1850)
Italy	0.7731***	-0.0426*	1.1114***	-2.0638**
	(0.2148)	(0.0234)	(0.1906)	(0.9803)
Netherlands	-0.0903	0.0033	1.2140***	0.7426
	(0.1043)	(0.0298)	(0.2320)	(0.4577)
Spain	0.2360**	-0.0348**	0.7949***	-0.6653
	(0.1126)	(0.0150)	(0.1764)	(0.7284)
Switzerland	-0.1494***	0.0416**	1.4012***	0.0959
	(0.0410)	(0.0164)	(0.1452)	(0.1119)
b) Asia-Pacific				
Australia	0.0351	-0.0226	1.1617***	0.1785
	(0.0775)	(0.0139)	(0.1043)	(0.4962)
Hong Kong	-0.6008***	0.1914***	0.7378***	0.1745
	(0.1142)	(0.0201)	(0.2256)	(0.1165)
Japan	-0.0200**	0.0253*	1.0637***	0.0062
	(0.0091)	(0.0146)	(0.0821)	(0.0228)
New Zealand	-0.2414**	0.0326**	0.9018***	0.8896**
	(0.1031)	(0.0146)	(0.1277)	(0.4196)
c) Anglo-Saxon Countries				
Canada	-0.0791	0.0091	1.1227***	0.1085
	(0.0503)	(0.0097)	(0.0951)	(0.2660)
United Kingdom	-0.0509	-0.0070	0.9702***	0.2270
	(0.0561)	(0.0093)	(0.0869)	(0.4762)
United States	-0.0440	-0.0263***	0.9672***	-0.3202
	(0.0354)	(0.0072)	(0.1084)	(0.1915)
d) Scandinavia				
Denmark	-0.0137	0.0325	0.0499	-0.0835
	(0.1313)	(0.0298)	(0.2667)	(0.6774)
Norway	-0.1393*	0.0078	0.7223***	-0.3873
	(0.0757)	(0.0118)	(0.1168)	(0.3800)
Sweden	-0.0079	-0.0006	0.6412***	-0.2093
	(0.0545)	(0.0110)	(0.1549)	(0.3404)

Notes: Sample is from January 1995 through December 2004. The standard errors of the coefficients are given in parentheses. A *, **, *** denotes rejection at a 10, 5, and 1 percent significance level.

Table 10. A GARCH-in-Mean Model for Term Premia

	α_0	α_1	β_0	β_1	γ_1
a) Continental Europe					
Austria	0.1646*	0.1342	0.0010	0.1393	0.8054***
	(0.1000)	(0.7146)	(0.0032)	(0.1565)	(0.2631)
Belgium	-0.6782***	3.4940***	0.0293***	-0.1765***	0.5996***
	(0.1183)	(0.5166)	(0.0045)	(0.0347)	(0.1091)
Europe	-0.0283***	0.5430***	0.0410***	0.4284**	-0.5632***
	(0.0040)	(0.0878)	(0.0067)	(0.1929)	(0.0980)
France	-0.0235	1.6794***	0.0024***	-0.0577**	0.9701***
	(0.0422)	(0.3856)	(0.0000)	(0.0250)	(0.0537)
Germany	-0.0532	0.4313	0.0071	0.4628	0.2272
	(0.0557)	(0.4001)	(0.0053)	(0.3043)	(0.2273)
Ireland	-0.6308**	1.7558**	0.0414**	0.4410**	0.3388**
	(0.2882)	(0.7524)	(0.0201)	(0.2046)	(0.1633)
Italy	-0.1665***	0.9686***	-0.0009*	0.0470**	0.9496***
	(0.0496)	(0.2000)	(0.0005)	(0.0196)	(0.0238)
Netherlands	0.0560	0.0056	0.0038	0.2323*	0.5663
	(0.1012)	(0.7620)	(0.0054)	(0.1266)	(0.3785)
Spain	-0.0599***	0.2266***	0.0010	1.7925**	0.1152
	(0.0090)	(0.0790)	(0.0007)	(0.8892)	(0.0745)
Switzerland	-1.8249	9.8762	0.0098***	0.0576	0.6687***
	(1.4028)	(7.2771)	(0.0031)	(0.0431)	(0.0563)
b) Asia-Pacific					
Australia	-0.1012*	0.2005	0.0054	0.2812***	0.5415***
	(0.0561)	(0.3305)	(0.0046)	(0.1054)	(0.2048)
Hong Kong	-0.1611***	0.9832***	0.0129	1.0806	0.3301***
	(0.0524)	(0.2283)	(0.0084)	(0.6873)	(0.1066)
Japan	-0.0628***	0.5980***	0.0000	0.2840**	0.7255***
	(0.0017)	(0.1258)	(0.0000)	(0.1263)	(0.0916)
New Zealand	-0.1061***	0.6637***	0.0060**	0.5438***	0.4035***
	(0.0342)	(0.1739)	(0.0026)	(0.1446)	(0.0919)
c) Anglo-Saxon Countries					
Canada	-0.2613*	2.2863**	-0.0002	0.0078	0.9851***
	(0.1436)	(1.0076)	(0.0002)	(0.0115)	(0.0115)
United Kingdom	-0.3522*	1.8109	0.0105**	0.1641	0.4945***
	(0.2012)	(1.2894)	(0.0049)	(0.1256)	(0.1552)
United States	-0.1523***	0.3191	0.0193***	0.5959**	-0.0007
	(0.0494)	(0.2652)	(0.0051)	(0.2345)	(0.0462)
d) Scandinavia					
Denmark	0.2914**	-0.9476	0.0044	-0.1083***	0.9883***
	(0.1441)	(0.8924)	(0.0035)	(0.0269)	(0.1170)
Norway	-0.1272***	-0.0726	0.0051*	0.5566***	0.3996**
	(0.0325)	(0.1537)	(0.0027)	(0.1814)	(0.1588)
Sweden	-0.0444**	0.3143*	0.0251***	0.4729***	-0.1183**
	(0.0267)	(0.1562)	(0.0061)	(0.1395)	(0.0500)

Notes: Sample is from January 1995 through December 2004. Reported are maximum likelihood estimates of a GARCH(1,1)-in-mean model. The quasi-maximum likelihood heteroskedasticity-consistent standard errors of the coefficients are given in parentheses. A *, **, *** denotes rejection at a 10, 5, and 1 percent significance level.