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Working Paper

Macroeconomic sources of risk in the term structure

CESifo Working Paper, No. 1329

Provided in Cooperation with:

Ifo Institute – Leibniz Institute for Economic Research at the University of Munich

Suggested Citation: Balfoussia, Chiona; Wickens, Mike (2004) : Macroeconomic sources of risk in the term structure, CESifo Working Paper, No. 1329, Center for Economic Studies and ifo Institute (CESifo), Munich

This Version is available at:

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MACROECONOMIC SOURCES OF RISK IN THE TERM STRUCTURE

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CESIFO WORKING PAPER No. 1329
CATEGORY 5: FISCAL POLICY, MACROECONOMICS AND GROWTH
NOVEMBER 2004

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MACROECONOMIC SOURCES OF RISK IN THE TERM STRUCTURE

Abstract

In this paper we develop a new way of modelling time variation in term premia. This is based on the stochastic discount factor model of asset pricing with observable macroeconomic factors. The joint distribution of excess holding period US bond returns of different maturity and the fundamental macroeconomic factors is modelled using multivariate GARCH with conditional covariances in the mean to capture the term premia. We show how by testing the assumption of no arbitrage we can derive a specification test of our model. We estimate the contribution made to the term premia at different maturities by real and nominal macroeconomic sources of risk. From the estimated term premia we recover the term structure of interest rates and examine how it varies through time. Finally, we examine whether the large number of reported failures of the rational expectations hypothesis of the term structure can be attributed to an omitted time-varying term premium.

JEL Code: C5, E4, G1.

Keywords: term structure, the stochastic discount factor model, term premia, GARCH.

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Financial support for this research was provided by the ESRC, Grant no. R000237862. We would like to express particular thanks to Peter Smith and Steffen Sorensen for their comments and help in the preparation of this paper.

1 Introduction

Historically, the bond market has been of greater concern to macroeconomists than macroeconomics has been to bond market analysts. In macroeconomic explanations of bond prices the emphasis is on the fundamentals whilst the principle of no-arbitrage is often neglected. In contrast, in finance the emphasis is on relative asset pricing rather than fundamentals, but is usually within a no-arbitrage framework. Typically this consists of a time series analysis based on affine latent factor models constrained to satisfy a no-arbitrage condition. Consideration of the fundamentals is usually confined to giving the factors an ex-post ad hoc macroeconomic interpretation. In this paper we attempt to integrate macroeconomics and finance by undertaking a fundamentals analysis of bond prices for US data based on the stochastic discount factor (SDF) model with observable macroeconomic factors. This includes consumption CAPM (C-CAPM), a general equilibrium model of bond pricing, as a special case. The SDF model we use is constrained to be arbitrage free across the term structure.

Our analysis has a number of novel features. We use an SDF model with observable macroeconomic factors instead of the more restrictive, but popular, Vasicek model. We consider three factors: consumption, inflation and output. We carry out the analysis using nominal returns rather than real returns on the grounds that an observable real risk-free rate does not exist. This alters the pricing kernel by including the inflation rate, in addition to consumption. A key feature of our analysis is the use of a multi-variate GARCH model with conditional covariances in the mean of the excess holding period returns. We show that neither univariate GARCH models, nor multi-variate GARCH models without conditional covariances in the mean, are capable of imposing no-arbitrage, and hence, although widely used, are inappropriate.

Apart from estimating the SDF and C-CAPM models, by relaxing some of the assumptions we are able to carry out tests of the no-arbitrage condition. Assuming that the no-arbitrage condition is satisfied in practice, these tests become specification tests of our models. We obtain estimates

of the term premia for each maturity in each period of time, and we estimate the contribution of each factor in determining the term premia. Having estimates of the term premia enables us to derive point estimates of the yield curve. We are also able to re-examine whether the failure of traditional tests of the rational expectations hypothesis of the term structure are due to omitting the term premia.

The paper is set out as follows. In Section 2 we provide a selective review of the related literature. In Section 3 we provide the theoretical framework for our analysis based on the SDF model. Our econometric methodology modifies the multi-variate GARCH model so that it satisfies a no-arbitrage condition and permits us to obtain point estimates of the time-varying term premia. This is set out in Section 4. The data are described in Section 5. In Section 6 we report our estimates, compare the SDF model with C-CAPM and carry out tests of the no-arbitrage condition. In Section 7 we reconstruct the yield curve using our estimates of the term premia. In this way we can estimate the term structure. In Section 8 we revisit tests of the rational expectations hypothesis of the term structure. Finally, in Section 9 we present our conclusions.

2 Selective review of related literature

There is a large body of literature of the empirical evidence on the term structure. Much of the early work was based on the rational expectations hypothesis of the term structure (REHTS) and came to the conclusion that it was rejected by the data. Several possible explanations of this failure have been proposed over the years.¹ The most fundamental criticism of the REHTS is that it is theoretically incorrect as it assumes that investors are risk neutral, whereas they are well-known to be risk averse. This version of the REHTS (also known as the expectations

¹ Campbell and Shiller (1991) and Hardouvelis (1994) concluded that long rates over-react to short rates. An alternative opinion, first emphasized by Mankiw and Miron (1986), followed by Hardouvelis (1988) and Rudebusch (1995), and formalized in a theoretical model by McCallum (1994), is that short rates respond to changes in monetary policy. Another strand of the literature relates the failure of the REHTS to “noise traders” and various other forms of irrational expectations of at least some market participants, see Froot (1989), Bekaert, Hodrick and Marshall (1997b).

hypothesis of the term structure) is therefore misspecified due to the omission of a time-varying risk premium. The early studies of Fama (1984), Keim and Stambaugh (1986), Mankiw (1986) and Hardouvelis (1988) supported the significance of time-varying risk premia in bond returns. More recent evidence is by Froot (1989), Fama (1990), Mishkin (1990) and Tzavalis and Wickens (1997). This suggests that omitted variables bias may be the cause of the empirical failure of the REHTS. Unfortunately, the omitted variable - the term premium - is not directly observable, so the solution is not as simple as just including an additional observable variable.

Looking beyond the traditional REHTS literature, there is broad consensus that, in any model of asset pricing, unless risk neutrality is explicitly assumed, a risk premium should be included. From an investor's perspective, risk premia also have substantial informational value. Hence, efforts to correctly model and estimate the risk premium are an integral part of financial research.

Two approaches have been adopted to account for the term premium. One is to find an observable proxy. Modigliani and Shiller (1973), Fama (1976), Mishkin (1982) and Jones and Roley (1983) employed various measures of the variability of interest rates. Shiller, Campbell and Schoenholtz (1983) used a measure of the volume of trade in bonds. Campbell (1987) used a latent variable model of the returns on bills, bonds and common stocks to infer time-varying risk premia in all three markets. Engle, Lilien and Robins (1987) used a time-varying conditional variance of interest rates derived from an ARCH-M model. Simon (1989) used the square of the excess holding-period return to proxy the term premium. Tzavalis and Wickens (1997) found the ex-post holding-period return of one maturity provided a good proxy for the risk premia of other maturities, thereby allowing the yield curve to have a single factor representation. Cochrane and Piazzesi (2001) used a single factor, a tent-shaped function of forward rates to predict one-year excess holding-period returns.

A second, and recently very popular, approach is to specify the term premium based on an appropriate model of the term structure. The usual way to do this is to use the SDF model with latent affine factors, see Campbell (1987), Knez, Litterman and Scheinkman (1994), Gong and

Remolona (1997), Cochrane and Piazzesi (2001) and Dai and Singleton (2002). Piazzesi (2002) surveys this literature. For a detailed treatment of the SDF model see Cochrane (2001) and Smith and Wickens (2003).

The attraction of affine term structure models is that they provide linear models of yields and term premia. This also means that they are a restricted form of the SDF model. In general, the risk premia implied by the SDF model are captured by the conditional covariances of the excess return over the risk-free rate. In, for example, the Vasicek (1977) and Cox, Ingersoll and Ross (1985) models, two widely-used affine models of bond prices, these conditional covariances are linear functions of the factors. In single factor affine models in general, all bond yields are just a function of the short rate and the shape of the yield curve is fixed through time. Multi-factor affine factor models are more flexible but, arguably, are still over-constrained. Ghysels and Ng (1998) use a semi-parametric framework to test the restrictions imposed by affine term structure models and find that, despite their common use, the empirical evidence does not support them. Duffee (2002) shows that all "completely affine" models of the term structure restrict the variability of the risk premium because they do not allow it to vary independently of interest rate volatility. He proposes a more general class of term structure models he calls "essentially affine", which offer greater flexibility in producing time-varying risk premia. The ability of these models to fit the time variation in conditional variances of yields is however poorer.

Another drawback of latent affine factor models is the difficulty in interpreting the latent factors extracted from the data. Pearson and Sun (1994), for example, labelled their two factors "short rate" and "inflation". In general, such labels are only loose ex-post descriptions. They give only rough guidance on the causes of changes in the shape of the yield curve. On grounds of interpretation, it would be preferable, in principle, to use observations on the short rate and inflation.

Although it is a growing field, to date little work has been done on observable factor models of the term structure. Ang and Piazzesi (2003) estimate a multifactor Vasicek model, allowing

the SDF model to depend on shocks to both observable and unobservable factors. This model has a VAR structure. They use 3 latent and 2 observable variables (measures of inflation and real activity), and draw on the literature on monetary policy rules to justify their choice of macroeconomic variables. They find that time variation in bond risk premia depend significantly on the macroeconomic factors. Using a New Keynesian macroeconomic model to support the specification of their macroeconomic variables, Rudebusch and Wu (2003) develop and estimate an affine term structure model where the factors are explicit functions of inflation and output. They contrast their macroeconomic factors with the latent factors extracted from an affine latent factor term structure model and find that they are highly correlated. This suggests that the latent factors have macroeconomic underpinnings. Dewachter and Lyrio (2004) draw similar conclusions on the basis of macroeconomic factors and their long run expectations. They find that their model can compete with the three factor latent model, despite the latter's increased flexibility. Finally, Diebold, Rudebusch and Aruoba (2003) find strong evidence of macroeconomic effects on the yield curve. These studies provide considerable support for the use of observable macroeconomic factors in modelling the term structure, but they are not sufficiently informative about precisely what role the factors play in determining bond risk premia.

The advantage of our methodology is that it enables us to obtain estimates of term premia at each point of time, within a no-arbitrage framework, without having to make the restrictive assumptions of previous studies. According to the SDF model, the cause of this time variation is a changing conditional covariation between the factors and the excess holding period returns on bonds. Furthermore, in order to satisfy no-arbitrage, these condition covariances must be present in the conditional mean of the distribution of the excess returns.

A number of implications follow from this. First, the analysis of term premia must be conducted in a multi-variate context as it is necessary to model the joint distribution of the excess holding period returns and the factors. Second, in order to ensure there are no arbitrage opportunities across the term structure, this joint distribution should involve the excess returns at all

maturities. Third, the conditional covariances must be included in the conditional mean of the joint distribution of excess returns as these are the risk premia. This implies that multi-variate GARCH alone is not adequate as it does not satisfy no arbitrage. Fourth, according to the SDF model, the coefficients of the conditional covariances in the mean are the same for all maturities and are constant across time; time variation in the risk premia is due solely to changing conditional covariances.

This approach was first proposed by Wickens and Smith (2000) for FOREX, and by Smith, Sorensen and Wickens (2003) for equity. For a general survey of the approach see Smith and Wickens (2002). Here, we develop this methodology for application to bond pricing and the term structure. We explicitly price US bonds with reference to their exposure to fundamental sources of macroeconomic risk. Our choice of factors is motivated from the consumption-based intertemporal capital asset pricing model of Rubinstein (1976) and Lucas (1978). We estimate the contribution made to the term premia at different maturities by real and nominal macroeconomic sources of risk, and explicitly test the assumption of no-arbitrage in the US bond market. The time-variation in our term premia has an intuitive interpretation which can be related to the macroeconomic events and policies of the period. From the estimated term premia we recover the term structure of interest rates. Finally, we re-examine whether, by including our estimates of the term premia, it is possible to account for the lack of empirical support for the REHTS.

3 Theoretical framework

3.1 Basic concepts

We use the following notation. $P_{n,t}$ = price of an n -period, zero-coupon (pure discount), default-free, bond at t , where $P_{0,t} = 1$ as the pay-off at maturity is 1. $R_{n,t}$ = yield to maturity of this bond, where the one-period, risk-free rate $R_{1,t} = s_t$. $h_{n,t+1}$ = return to holding an n -period bond for one period from t to $t + 1$. It follows that

$$P_{n,t} = \frac{1}{[1 + R_{n,t}]^n}$$

and

$$R_{n,t} \simeq -\frac{1}{n} \ln P_{n,t}$$

Thus

$$\begin{aligned} 1 + h_{n,t+1} &= \frac{P_{n-1,t+1}}{P_{n,t}} \\ &= \frac{(1 + R_{n-1,t})^{-(n-1)}}{(1 + R_{n,t})^{-n}} \end{aligned}$$

And if $p_{n,t} = \ln P_{n,t}$ then taking logs

$$h_{n,t+1} \simeq p_{n-1,t+1} - p_{n,t} = nR_{n,t} - (n-1)R_{n-1,t+1}$$

In the absence of arbitrage opportunities, after adjusting for risk, investors are indifferent between holding an n -period bond for one period and holding a risk-free 1-period bond. In the absence of default, the risk is due to the price of the bond next period being unknown this period.

$$E_t[h_{n,t+1}] = s_t + \rho_{n,t} \quad (1)$$

where $\rho_{n,t}$ is the risk, or term, premium on an n -period bond at time t .

3.2 SDF model of the term structure

The SDF model relates the price of an n -period zero-coupon bond in period t to its discounted price in period $t+1$ when it has $n-1$ periods to maturity. Thus

$$P_{n,t} = E_t[M_{t+1}P_{n-1,t+1}]$$

where M_{t+1} is a stochastic discount factor, or pricing kernel. It follows that

$$E_t[M_{t+1}(1 + h_{n,t+1})] = 1$$

and for $n = 1$,

$$(1 + s_t)E_t[M_{t+1}] = 1$$

If $P_{n,t}$ and M_{t+1} are jointly log-normally distributed and $m_{t+1} = \ln M_{t+1}$ then

$$p_{n,t} = E_t(m_{t+1}) + E_t(p_{n-1,t+1}) + \frac{1}{2}V_t(m_{t+1}) + \frac{1}{2}V_t(p_{n-1,t+1}) + Cov_t(m_{t+1}, p_{n-1,t+1}) \quad (2)$$

and as $p_{o,t} = 0$,

$$p_{1,t} = E_t(m_{t+1}) + \frac{1}{2}V_t(m_{t+1}) \quad (3)$$

Subtracting (3) from (2) and re-arranging gives the no-arbitrage equation

$$E_t(p_{n-1,t+1}) - p_{n,t} + p_{1,t} + \frac{1}{2}V_t(p_{n-1,t+1}) = -Cov_t(m_{t+1}, p_{n-1,t+1}) \quad (4)$$

This can be re-written in terms of yields as

$$-(n-1)E_t(R_{n-1,t+1}) + nR_{n,t} - s_t + \frac{(n-1)^2}{2}V_t(R_{n-1,t+1}) = (n-1)Cov_t(m_{t+1}, R_{n-1,t+1})$$

and, since $h_{n,t} \simeq nR_{n,t} - (n-1)R_{n-1,t+1}$, as

$$E_t(h_{n,t+1} - s_t) + \frac{1}{2}V_t(h_{n,t+1}) = -Cov_t(m_{t+1}, h_{n,t+1}) \quad (5)$$

This is the fundamental no-arbitrage condition for an n -period bond². It must be satisfied simultaneously for all maturities to ensure that there are no arbitrage opportunities across the term structure. The term on the right-hand side is the term premium and $\frac{1}{2}V_t(h_{n,t+1})$ is the Jensen effect. Comparing equations (1) and (5), we note the SDF model implies that

$$\rho_{n,t} = -\frac{1}{2}V_t(h_{n,t+1}) - Cov_t(m_{t+1}, h_{n,t+1})$$

Empirical work on the term structure can be distinguished by the choice of $\rho_{n,t}$ and the discount factor m_t . The expectations hypothesis, which is the basis of most tests of the REHTS, assumes that $\rho_{n,t} = 0$ but the evidence cited earlier rejects this assumption.

To complete the specification, it is necessary to specify m_t , the logarithm of the pricing kernel. Assuming joint log-normality of the excess returns and the factors, m_t is a linear function of the stochastic discount factors z_{it}

$$m_{t+1} = a + \sum_i \beta_i z_{i,t+1}.$$

² Arbitrage opportunities are excluded if and only if there exists a *positive* stochastic discount factor M_{t+1} that prices *all* assets (Cochrane 2001). In the models described in this section, a positive discount factor is used to price bonds of all maturities.

The resulting no-arbitrage condition for the SDF model is

$$E_t(h_{n,t+1} - s_t) + \frac{1}{2}V_t(h_{n,t+1}) = - \sum_i \beta_i Cov_t(h_{n,t+1}, z_{i,t+1}) \quad (6)$$

Although the SDF model places no restrictions on the coefficients of the conditional covariances, it does restrict the coefficient on the conditional variance.

3.3 General equilibrium models of the term structure

The SDF model does not specify which factors should be used. They can be latent or observable. An advantage of general equilibrium models of asset pricing is that they do specify the factors. The most widely-used general equilibrium asset pricing model is the C-CAPM based on power utility. This gives a single factor SDF model with real consumption as the factor, and it implies a constant coefficient of relative risk aversion σ .

C-CAPM is usually expressed in real terms. This involves the use of the excess return and hence assumes the existence of a real risk-free rate. Since, in practice, only a nominal risk-free rate exists (assuming no default risk), we carry out our analysis in nominal terms. This implies two factors: real consumption and inflation. It can be shown that for *nominal* returns³

$$m_{t+1} = \theta - \sigma \Delta c_{t+1} - \pi_{t+1}$$

where m_t = the log stochastic discount factor, c_t = log real consumption and π_t = inflation.

The resulting no-arbitrage condition is, see Smith and Wickens (2002) and Smith, Sorensen and Wickens (2003),

$$E_t(h_{n,t+1} - s_t) + \frac{1}{2}V_t(h_{n,t+1}) = \sigma Cov_t(h_{n,t+1}, \Delta c_{t+1}) + Cov_t(h_{n,t+1}, \pi_{t+1}) \quad (7)$$

We note again that the coefficients in the no-arbitrage condition are the same for all maturities.

³ The familiar solution of this intertemporal optimisation problem, in a nominal setup, is the Euler equation $E_t[\frac{\beta U'_{t+1}}{U'_t} \frac{P_t}{P_{t+1}} (1 + r_{t+1})] = 1$, where U_t is the single-period utility function derived from consumption and P_t is the price level at time t . The nominal discount factor implied by the C-CAPM is hence: $M_{t+1} = \frac{\beta U'_{t+1}}{U'_t} \frac{P_t}{P_{t+1}}$.

A generalisation of C-CAPM is to allow utility to be time non-separable. Smith, Sorensen and Wickens (2003) show that for the Epstein-Zin (1979) utility function the no-arbitrage condition can be written as

$$E_t(h_{n,t+1} - s_t) = \beta_1 V_t(h_{n,t+1}) + \beta_2 Cov_t(h_{n,t+1}, \Delta c_{t+1}) + \beta_3 Cov_t(h_{n,t+1}, \pi_{t+1}) \quad (8)$$

The β_i , which are the same for all maturities, are functions of the deep structural parameters of the general equilibrium model; β_2 is now no longer equal to σ . This formulation removes a number of the restrictions of C-CAPM with power utility.

4 Econometric methodology

4.1 Estimation

Our aim is to estimate the no-arbitrage condition either jointly for all maturities, in order to impose no-arbitrage across the term structure, or separately for each maturity, in order to test the no-arbitrage condition. In other words, we must model the joint distribution of the excess holding-period returns jointly with the macroeconomic factors in such a way that the mean of the conditional distribution of each excess holding-period return in period $t + 1$, given information available at time t , satisfies the no-arbitrage condition. The conditional mean of the excess holding-period return involves selected time-varying second moments of the joint distribution. We therefore require a specification of the joint distribution that admits a time-varying variance-covariance matrix. A convenient choice is the multivariate GARCH-in-mean (MGM) model. For a review of multivariate GARCH models see Bollerslev, Chou and Kroner (1997), and see Flavin and Wickens (1998) for a discussion of specification issues for their use in financial models.

Let $\mathbf{x}_{t+1} = (h_{n_1,t+1} - s_t, h_{n_2,t+1} - s_t, \dots, z_{1,t+1}, z_{2,t+1}, \dots)'$, where $z_{1,t+1}, z_{2,t+1}, \dots$ include the macroeconomic variables that give rise to the factors in the SDF model through their conditional covariances with the excess holding-period returns. In principle, they may also include additional variables that are jointly distributed with these macroeconomic variables as this may improve the

estimate of the joint distribution. The MGM model can then be written

$$\mathbf{x}_{t+1} = \boldsymbol{\alpha} + \boldsymbol{\Gamma}\mathbf{x}_t + \mathbf{B}\mathbf{g}_t + \boldsymbol{\varepsilon}_{t+1}$$

where

$$\boldsymbol{\varepsilon}_{t+1} \mid I_t \sim D[0, \mathbf{H}_{t+1}]$$

$$\mathbf{g}_t = vech\{\mathbf{H}_{t+1}\}$$

The *vech* operator converts the lower triangle of a symmetric matrix into a vector. The distribution is the multivariate log-normal distribution. If there are K maturities (excluding the short rate) then the first K equations of the model are restricted to satisfy the no-arbitrage condition.

It will be noted that the theory requires the macroeconomic variables to display conditional heteroskedasticity. This is not something traditionally assumed in macroeconomics, but seems to be present in our data. Ideally, we would like to use high frequency data for asset returns, but very little macroeconomic data are published for frequencies higher than a month, and then only a few variables are available. Although more macroeconomic variables are published at lower frequencies, they tend not to display conditional heteroskedasticity.

Whilst the MGM model is convenient, it is not ideal. First, it is heavily parameterised which can create numerical problems in finding the maximum of the likelihood function due to the likelihood surface being relatively flat, and hence uninformative. Second, asset returns tend to be excessively volatile. Assuming a non-normal distribution such as a t -distribution can sometimes help in this regard by dealing with thicker tails. The main problem, however, is not thick tails, but a small number of extreme values. The coefficients of the variance process of the MGM model have a tendency to produce a near unstable variance process in their attempt to fit these extreme values. In principle, a stochastic volatility model, which includes an extra random term in the variance, could capture these extreme values. Unfortunately, as far as we are aware, no multivariate stochastic model with in-mean effects in the conditional covariances has been proposed in the literature.

The choice of specification of the conditional covariance matrix is a compromise between generality and feasibility. The latter requires that we limit the number of parameters to estimate. Accordingly, our specification of \mathbf{H}_t is that of Ding and Engle (1994), the vector-diagonal multivariate GARCH-in-mean

$$\mathbf{H}_t = \mathbf{H}_0(\mathbf{i}\mathbf{i}' - \mathbf{a}\mathbf{a}' - \mathbf{b}\mathbf{b}') + \mathbf{a}\mathbf{a}' * \boldsymbol{\Sigma}_{t-1} + \mathbf{b}\mathbf{b}' * \mathbf{H}_{t-1}$$

where \mathbf{i} is a vector of ones, $*$ denotes element by element multiplication (the Hadamard product) and $\boldsymbol{\Sigma}_{t-1} = \boldsymbol{\varepsilon}_{t-1}\boldsymbol{\varepsilon}_{t-1}'$. This is a special case of the diagonal Vech model, in which each conditional covariance depends only on its own past values and on surprises. The restrictions implicitly imposed by this parameterisation of the multivariate GARCH process guarantee positive-definiteness and also substantially reduce the number of parameters to be estimated, thus facilitating computation and convergence. Stationarity conditions are imposed.

First we estimate the whole system as a standard homoskedastic VAR. This gives a consistent estimator of the long-run variance covariance matrix \mathbf{H}_0 . In the subsequent estimation step we constrain \mathbf{H}_0 to this value. Our sample has 348 observations. The estimation is performed using quasi-maximum likelihood.

5 Data

The complete sample is monthly, from January 1970 to December 1998. Until 1991, the term structure data are those of McCulloch and Kwon.⁴ ⁵ We then use the data extended by Bliss which is based on the same technique. The maturities available are a 6 month, 1 year, 2 year, 5 year and 10 year bond returns. Excess holding-period returns are taken in excess of the one-

⁴ McCulloch used tax-adjusted cubic splines to interpolate the entire term structure from data on most of the outstanding Treasury bills, bonds and notes, see McCulloch (1975) and McCulloch and Kwon (1993). These data have been used extensively in empirical work and are considered very “clean”, in that they are based on a broad spectrum of government bond prices and are corrected for coupon and special tax effects. Furthermore, spline-based techniques allow for a high degree of flexibility, since individual curve segments can move almost independently of each other (subject to continuity constraints). Hence they are able to accommodate large variations in the shape of the yield curve.

⁵ We are not using the complete term structure dataseries available by McCulloch and Kwon because no real personal consumption data was available for earlier dates.

month risk-free rate provided by K. French⁶ ⁷. Essentially, this is the one-month Treasury bill rate. Selected yields are plotted in Figure 9a.

The macroeconomic variables are the month-on-month growth rates of the consumer price index for all urban consumers (CPI) and real total personal consumption (TPC).⁸

Table 1 presents some descriptive statistics for these data which are expressed in annualised percentages. The average yield curve is upward sloping. Average excess holding-period returns are positive for all maturities and increase with time to maturity. Like most financial data, excess holding-period returns exhibit excess skewness and kurtosis, particularly for short maturities. Fitting a univariate GARCH(1,1) to them indicates there is also significant heteroskedasticity.⁹ Their unconditional variances increase with maturity, as do the absolute unconditional covariances of the excess returns with the macroeconomic variables.

Table 2 reports the unconditional sample correlations. The unconditional correlations of all excess returns with the macroeconomic variables are negative, and for the CPI their absolute values increase monotonically, with maturity. The excess returns are highly correlated, especially at the short end of the yield-curve. For example, the 6-month and 1-year excess returns have a correlation of 0.95.

6 Empirical results

It is important to note that the models are non-linear and, as a result, the coefficient estimates are not independent of the units of measurement. The theory for the SDF model and C-CAPM were developed in absolute terms, but our data are in annualised percentages. Consequently, the

⁶ In constructing holding-period returns we use the change in n -period yields $\Delta R_{n,t+1}$ instead of $R_{n-1,t+1} - R_{n,t}$, since $R_{n-1,t+1}$ is not available. This approximation is very commonplace in the literature, and is likely to be good for medium and long yields. For a discussion on the effects of this approximation see Bekaert, Hodrick and Marshall (1997a).

⁷ <http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/>

⁸ Data on CPI were obtained from the U.S. Department of Labor's Bureau of Labor Statistics, while consumption data were obtained from the Federal Reserve Bank of St Louis.

⁹ Estimation results are available upon request.

coefficient on the own variance of the holding-period return is constrained to be $\frac{1}{2400}$ and not $\frac{1}{2}$; the coefficient of the inflation covariance in the SDF CAPM is constrained to be $\frac{1}{1200}$ and not 1; and the coefficient on the consumption covariance should be interpreted as $\frac{1}{1200}\sigma$ and not σ .

For the sake of brevity, estimates of the covariance structure are not reported.¹⁰ For all models the ARCH and GARCH coefficients are highly significant for the excess holding-period return equations and for the macroeconomic variables; typically, the *t*-statistics have more than two significant places. The GARCH coefficients are always much larger than the ARCH coefficients, with typical estimates close to 0.9 and 0.3, respectively. Further, the GARCH parameter shows a tendency to increase with maturity. As a result, the conditional covariance structure of excess holding-period returns depends largely on the lagged conditional covariance matrix, and much less on lagged innovations. In the equations for the macroeconomic variables, the constant and own lagged values are always highly significant.

6.1 Coefficient estimates

The coefficient estimates for each model, and for excess returns on bonds of 6 months, 1, 2, 5 and 10 years to maturity, are reported in Table 3a, and the likelihood ratio tests are reported in Table 3b.¹¹

6.1.1 C-CAPM

The estimates in the first row of Table 3a are for C-CAPM, equation (7), constrained to satisfy no arbitrage across the term structure, and hence are the same for all maturities. In practice, as the coefficient for inflation is constrained to unity, this implies that only one parameter is estimated in the entire system of equations, namely the coefficient of relative risk aversion σ . The estimate of the coefficient on the consumption covariance is -0.072 and is significant at conventional levels.

¹⁰ These are available upon request.

¹¹ The same estimations have been performed for systems of smaller dimensions, using a number of different combinations of holding period returns. The statistical significance and the explanatory power of the model vary very little, while the signs of the estimated coefficients and the trends along the yield curve are almost identical for all other combinations we examined.

It implies an estimate of σ of 86.4. Like most estimates of the coefficient of relative risk aversion in the literature, this is implausibly large. Based on a likelihood ratio test, from Table 3b we note that the consumption covariances of all maturities are also jointly significant. This implies that the term premia are significant. The explanatory power of the model is very low (see Table 3a, columns 11 to 15). This suggests that the dominant components of the excess holding-period returns are innovations in the price of bonds.

6.1.2 SDF model

The SDF model, equation (6), is less restrictive than C-CAPM. It is obtained by relaxing the constraint on the coefficient of inflation and dropping the interpretation of the coefficient of the conditional covariance of consumption as the coefficient of relative risk aversion. We continue to impose the no-arbitrage condition across the term structure by constraining the estimates to be the same for all maturities. Estimates of the SDF model are in the second row of Table 3a.

The estimate of the coefficient on the consumption covariance is -0.057 , but is not significant at the 5% level. Were we still able to interpret the coefficient as for C-CAPM, this would imply an estimate of σ of 68.7. The estimate of the coefficient of the inflation covariance has switched sign; it is -0.153 and is highly significant. This is very far from its theoretical value of unity under C-CAPM, to which it was previously constrained. From Table 3b, the conditional covariances are jointly significant in the conditional mean of the excess returns. Comparing C-CAPM with the SDF model, a likelihood ratio test of the C-CAPM restrictions strongly rejects them in favour of the more general SDF model. Nevertheless, the explanatory power of the SDF model is also very low.

6.1.3 Arbitrage models

We now relax the restrictions that impose no arbitrage on the system. We do this in three ways. First we allow the coefficient on the inflation covariance to differ across maturities; second we allow the coefficient on the consumption covariance to differ across maturities; third we allow

both coefficients to differ across maturities. The exact specifications of these models are given in the footnote to Table 3a.

The estimates for the first of these models are reported in the third row of Table 3a. An entirely different picture now emerges. The coefficient on the inflation covariance has changed considerably. For the shortest maturity it is 0.424, even further from the theoretical value implied by C-CAPM. The coefficient estimates decline in absolute size and significance as the time to maturity of the holding-period returns increases, but all are highly significant. The consumption covariance coefficient has now switched sign compared with the two no-arbitrage models. The absolute value of the coefficient is greater, but is insignificant. The explanatory power of the model has increased dramatically compared with the no-arbitrage models. 14% of the variance of the excess holding-period return is explained for the shortest bond, 7% is explained for the medium-term bond and 2% for the longest bond.

The estimates for the second model are reported in the fourth row of Table 3a. The explanatory power of this model is also higher than that of the no-arbitrage models. The results for relaxing the coefficient on the consumption covariance are similar to those for relaxing the inflation covariance. 12% of the variance of the excess holding-period return is explained for the shortest bond and 5% is explained for the longest bond. Both the inflation and the consumption coefficients are significant.

The estimates after removing both sets of restrictions are reported in row 5 of Table 3a. The coefficients on the inflation covariance are once again highly significant for all maturities, whereas those for consumption are now significant only at the short end of the yield curve.

Likelihood ratio test statistics for the hypothesis of no-arbitrage against that of arbitrage are reported in Table 3b. Both C-CAPM and the no-arbitrage version of the SDF model are rejected against each of the arbitrage models. These results can be interpreted to mean either that there are in fact arbitrage opportunities across the term structure, or that the models are misspecified as in fact there are no arbitrage opportunities. The latter seems more plausible. These findings suggest

that testing whether or not the no-arbitrage condition holds provides a valuable new specification test for asset pricing models that involve more than one risky asset. Certainly, just imposing no-arbitrage, as is common in empirical work on the term structure, may hide the inadequacies of a model.

Although we reject the no-arbitrage condition, our results show that macroeconomic variables are significant sources of risk in the term structure. Inflation risk appears to be very strongly priced across the yield curve. These findings support those of Ang and Piazzesi (2004) and Dewachter, Lyrio and Maes (2001) who also find that macroeconomic factors are an important source of variation in bond risk premia.

Another finding is that the coefficient estimates tend to be larger and much more significant at the short end of the term structure than for the longest maturities. These results suggest that investors are better informed about risk at shorter than longer time horizons. Consequently, we would expect term premia to be adjusted more often, and hence to be more volatile, at short horizons than at longer horizons. This conclusion is similar to those of Dewachter and Lyrio (2004) who find that their observable macroeconomic factors are important for modelling the shorter end of the yield curve, while long-run inflation expectations are the prime determinant of the risk premium for longer maturities.

6.2 Estimated Term Premia

We now consider the implications of our estimates for the term premia of excess holding period returns. We are interested in the variation in the term premia over time and across maturities, and in the relative contributions of inflation and consumption. We use the most general model in these calculations even though it does not satisfy the no-arbitrage condition. This corresponds to the last row of Table 3a.

As the coefficient estimates for all maturities are negative for the inflation covariance and positive for the consumption covariance, and as the unconditional (i.e. average long-run) correlations

between the excess returns at all maturities and inflation and consumption are negative in Table 2, we may infer that in general inflation makes a positive contribution to the term premia and consumption makes a negative contribution. We also note that whereas the coefficients for inflation are highly significant, those for consumption are not significant for longer maturities. This suggests that inflation will be the dominant factor determining the term premia and consumption will contribute little. This is precisely what we observe in Figure 1 where we plot the estimated time-varying term premia (plus Jensen effect) for each maturity. Table 4 gives summary statistics for the estimated risk premia. The total term premium is positive in nearly every period and this is due almost entirely to inflation risk. The contribution of consumption risk is very small and the “incorrect” sign of the consumption coefficient has hardly any effect.

An idea of the importance of the term premia relative to the (scaled) excess holding-period return may be obtained from Figure 2. We note that investors required an abnormally large term premium during the 1979-82 period when the Fed was conducting its “monetary experiment” of strict money-base targeting. This led to high and volatile short rates and was accompanied by high and volatile inflation and a US recession. It might therefore be reasonable to expect that both real and nominal factors would contribute to the term premium during this period. It is however the inflation risk premium that dominates. Term premia are also relatively high in the early 1970’s, during the first oil crisis, and throughout the 1980’s. In both periods inflation volatility was still relatively high. The 1990’s was a decade of macroeconomic stability and low inflation (the Greenspan era). Term premia are much lower and more stable as a result.

A comparison of the term premia for different maturities may be made from Figure 3. We find that the term premia increase in magnitude with maturity. Since the coefficients of the conditional covariances decline in absolute value as the time to maturity increases, the size of the term premia may be attributable to the magnitude of the conditional covariances. This finding supports the conclusions of McCulloch (1987) and Wickens and Tzavalis (1997) who found that risk premia increase with maturity, but at a decreasing rate. Heston’s (1992) conjecture that risk premia

for excess holding-period returns of different maturities are related to each other also finds some support here, as the estimated risk premia seem to follow a similar pattern for all maturities and they are highly correlated with each other, see Table 5.

We also find that the explanatory power of the model decreases with maturity. This result is consistent with Ang and Piazzesi (2001) who find that their two macroeconomic factors primarily explain movements at the short end and middle of the yield curve, while their unobservable factors account for most of the movement at the long end. This suggests that additional factors might be needed in our model. Earlier, using survey data, Froot (1989) had found that the failure of the REHTS at the short end of the yield curve is due primarily to a time-varying risk premium, while at the very long end the principal reason is expectations error.

Although the results are not reported, we experimented with different definitions of consumption and with additional variables in the SDF model. We found that using retail sales instead of total personal consumption allowed us to explain a maximum of 13% of the variance of the excess holding-period returns on the shortest bond; using personal consumption on non-durables and services gave similar results. We conclude from this that there is little to choose between the alternative measures of consumption. Using industrial production as the real factor, gave poorer results than consumption.

7 Reconstructing the yield curve

In this section we consider the risk premia embedded in the yields to maturity. We call these the rolling risk premia. This enables us to reconstruct the whole yield curve. One advantage of obtaining point estimates of time-varying term premia (rather than estimating the model using GMM as many studies of the term structure have done) is that we can estimate the yield curve. First we discuss how the risk premia in the yields can be estimated from the term premia.

7.1 Estimating the rolling risk premia

The no-arbitrage equation for an n -period bond can be re-expressed as

$$E_t[h_{n,t+1}] = nR_{n,t} - (n-1)E_t[R_{n-1,t+1}] = s_t + \rho_{n,t}$$

Hence,

$$(n-1)[E_t[R_{n-1,t+1}] - R_{n,t}] = (R_{n,t} - s_t) - \rho_{n,t}$$

where $R_{n,t} - s_t$ is the term spread and

$$\rho_{n,t} = -\frac{1}{2}(n-1)^2 V_t(R_{n-1,t+1}) - (n-1)Cov_t(m_{t+1}, R_{n-1,t+1})$$

It follows that

$$\begin{aligned} R_{n,t} &= \frac{n-1}{n} E_t[R_{n-1,t+1}] + \frac{1}{n}(s_t + \rho_{n,t}) \\ &= \frac{1}{n} \sum_{i=0}^{n-1} E_t[s_{t+i} + \rho_{n-i,t+i}] \\ &= \frac{1}{n} \sum_{i=0}^{n-1} E_t s_{t+i} + \omega_{n,t} \end{aligned} \tag{9}$$

where

$$\omega_{n,t} = \frac{1}{n} \sum_{i=0}^{n-1} E_t \rho_{n-i,t+i} \tag{10}$$

is the rolling risk premium. Thus the yield to maturity is the average of expected future short rates plus the average risk premium on the bond over the rest of its life, $\omega_{n,t}$.

In order to re-construct yields it is necessary to estimate $\omega_{n,t}$. In principle, this requires having term premia at all maturities. In practice, as only a few yields are observable at any time, we calculate the missing term premia using a linear interpolation of our estimated term premia. We also need estimates of future term premia. We consider two ways of obtaining these. The first uses estimates of the current term premia plus the assumption of static expectations for future term premia. The second uses estimated future term premia plus the assumption of perfect foresight. All calculations are based, as before, on the most general model estimated.

We consider time-series plots of these two ways of estimating the rolling risk premium for each maturity. Figure 4 is based on static expectations, and Figure 5 is based on perfect foresight. Figure 6 expresses the rolling premia based on static expectations as a percentage of the total yield, and Figure 7 gives the corresponding information based on perfect foresight. Being averages, the rolling premia show much more persistence than the term premia. Otherwise, they reflect previous findings, being highest during the late seventies, the early 1980's, and during the recession of the early 1990's. Tables 6a and 6b present some descriptive statistics for the two sets of rolling premia.

The average estimated rolling risk premium based on static expectations for the 6-month yield during the 1970's is 0.27 (which is 4.3% of the yield); in the period from 1979-1982 the corresponding figure is 0.92 (7.6%); the average until the end of the 1980's is 0.20 (2.5%). During the 1990's, a decade associated with low and very stable inflation as well as positive growth rates, estimated rolling risk premia are very low. For 6-month bonds the average is 0.09 (1.8%). The rolling risk premia are larger the longer the time to maturity. On a few occasions our estimates become negative. Sometimes, particularly during the period 1979-1982, they are clearly excessively large. An example is April 1980 when our estimate for the 10-year bond is 114% of the actual yield. However, in general, these estimates seem plausible, at least for short and medium term maturities, and are comparable to those found in other studies.

The rolling premia based on perfect foresight are clearly very different from those based on static expectations, particularly in the 1970's. The reason is that they assume that the market anticipates the high term premia of the period of the Fed's "monetary experiment". These estimates are clearly not plausible. This becomes obvious if we reconstruct the yields implied by the estimated rolling premia.

7.2 Yield curves

To estimate the yields we combine the estimated rolling term premia with estimates of expected future short rates as in equation (9). Again we consider the use of static expectations and perfect

foresight, this time for the expected future short rates. Under static expectations, expected future short rates are set equal to the current short rate s_t . This implies that the shape of the yield curve and of the term structure reflects nothing more than changes in the rolling risk premia at different maturities.

Figures 8 and 9 give actual and estimated yields under static expectations for short rates and term premia, and under perfect foresight, for three maturities: 1-year, 2-year and 10-year bonds. The panels in Figure 8 are for each maturity, while those in Figure 9 are for each type of estimate. The two figures show that yields are better estimated at shorter than longer maturities, and that the assumption of static expectations provides better estimates than that of perfect foresight. The estimation errors for 6-month and 2-year yields based on static expectations are remarkably small. Even for 10-year bonds they are small from 1984 onwards. Perfect foresight does not introduce large errors for 6-month bonds, but for 2-year bonds they are larger. And for 10-year bonds prior to 1984 the errors are huge.

Inflation is well-known to be an important factor affecting short rates. And we have shown inflation volatility is the major factor in term premia. When inflation is low it appears to be both more predictable and to have lower and more predictable volatility. It is in this case that we are able to estimate yields best. The larger estimation errors under perfect foresight suggest that markets are more likely to mis-price bonds when inflation is more volatile.

8 Testing the REHTS

We now return to consider the question raised at the beginning of whether the empirical failures of the REHTS are due to the omission of a time-varying risk premium. We can include our estimate of the term premium $\rho_{n,t}$ as an additional observable variable in the model. The model based on returns can be written

$$h_{n,t+1} = \alpha + \beta s_t + \gamma \rho_{n,t} + \varepsilon_{n,t+1} \quad (11)$$

where $\varepsilon_{n,t+1}$ is a zero-mean disturbance. According to the REHTS $\alpha = 0$, $\beta = 1$ and $\gamma = 0$. But if a risk premium is in fact present and we have estimated it accurately, then we would expect $\gamma = 1$. Alternatively, we can express the model in terms of yields as

$$(n - 1)(R_{n-1,t+1} - R_{n,t}) = \delta + \mu(R_{n,t} - s_t) + \nu\rho_{n,t} + \epsilon_{n,t+1} \quad (12)$$

where we expect $\delta = 0$, $\mu = 1$, $\nu = -1$ and $\epsilon_{n,t+1} = -\varepsilon_{n,t+1}$.

OLS estimates of these equations are reported in Table 7 for each of the five maturities. Columns *a* and *c* are estimates of the standard tests of the REHTS without a risk premium, and columns *b* and *d* include the estimated risk premium derived from the most general model (the last row of Table 3a). In the models without the risk premium the estimates of β are close to unity for each maturity, but those of μ are not. This is the usual result. The difference is due to the holding-period returns and the short rate in equation (11) being I(1), or close to I(1), variables. This results in super-consistent estimates of β even if there is a stationary omitted term premium. In contrast, in equation (12), the change in the yield and the term spread are I(0) variables, and so an omitted stationary term premium is likely to give biased estimates of μ .

Including the risk premium is expected to have little effect on the estimate of β but ought in principle to remove the bias in the estimate of μ , if the risk premium is well estimated and provided it is the principal reason for the empirical failure of the REHTS. The results in column *b* confirm this for equation (11), which is based on holding period returns. The estimate of γ is very close to 1, its theoretical value, and the coefficient of the short rate is close to its theoretical value. These findings provide considerable support for including a risk premium in tests of the REHTS. They also suggest that we can use this test of the REHTS to provide a further diagnostic check on models of the term premia.

The estimates based on equation (12), the term spread equation, are reported in column *d*. The estimate of ν is not far from its theoretical value of -1 , especially for short and medium maturities. However, the bias in the estimate of μ appears to have increased for most maturities.

This suggests that the estimates of the term premia are less accurate at long maturities and as a result the inclusion of time-varying risk premia is not sufficient to provide strong evidence in support of the REHTS based on an equation using the term spread.

9 Conclusion

We have shown how it is possible to directly estimate time-varying term premia derived from observable factors, whilst satisfying the condition of no-arbitrage along the yield curve. Based on the stochastic discount factor model, this requires the specification of the joint distribution of excess holding-period returns and observable factors in which the conditional mean of the excess returns is the no-arbitrage condition. Although multivariate GARCH models are widely used in empirical finance, they are rarely specified with conditional covariances in the mean. We show that, in general, this is essential if a no-arbitrage condition is to be satisfied.¹²

This paper is the first application of this methodology to the term structure. We find that the main source of risk at all maturities is nominal risk due to inflation. We also find that consumption is a significant but less important source of risk. Term premia at shorter maturities are better explained than at longer maturities. The time-variation in our term premia can be related in an intuitive way to the macroeconomic events of the period. The magnitude of the term premia tends to increase with maturity due to the size of the conditional covariance of excess holding period returns with inflation, rather than differences in the coefficient of the covariance terms.

We have found that both the SDF model and C-CAPM constrained to satisfy the no-arbitrage condition across the term structure are rejected in favour of separate SDF models for each maturity. An alternative, but less plausible, interpretation is that the data themselves do not satisfy no-arbitrage.

We show how it is possible to reconstruct yields, and hence estimate the term structure, by estimating the risk premia embedded in yields from the term premia of excess holding period

¹² An exception is when a Vasicek model is used, but this has the drawback of not having a time-varying risk premium.

returns. Our estimates seem to be more accurate for the shorter maturities. When inflation is predictable and has low volatility they are also accurate for longer maturities. The larger errors when estimating yields under perfect foresight suggest that markets may be more prone to misprice bonds when inflation is hard to predict and is volatile.

We re-examine whether or not the main cause of the familiar finding of the failure of the rational expectations hypothesis of the term structure is an omitted term premium. Using our estimates of the term premia as proxies we find that for the holding period return model of the REHTS there is strong evidence in favour of the REHTS when a term premium is included. The evidence is weaker when the test is based on a term spread model of the REHTS.

Our conclusion is that this way of modelling the term structure provides a viable alternative to the methods proposed in the literature. It has produced new ways to test theories of the term structure and has cast new light on several familiar issues. It has also thrown up a new puzzle: why the C-CAPM and SDF models do not appear to satisfy the condition of no-arbitrage. The question of whether to emphasise the fundamental macroeconomic sources of risk, possibly at the expense of no arbitrage, or to impose no arbitrage, possibly at the expense of using macroeconomic fundamentals, remains and takes a new twist.

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Table 1: Descriptive Statistics

		Mean	Median	Maximum	Minimum	Std. Dev.	Skewness	Kurtosis
Risk Free	1 month	6.80	6.04	17.46	2.55	2.85	1.27	4.85
Bond returns	6 months	7.10	6.43	16.51	2.88	2.74	1.14	4.22
	1 year	7.34	6.85	16.35	3.09	2.68	1.06	3.94
	2 years	7.64	7.07	16.15	3.80	2.52	1.07	3.85
	5 years	8.04	7.53	15.70	4.35	2.31	1.09	3.65
	10 years	8.32	7.80	15.07	4.51	2.14	1.03	3.40
Excess holding period bond returns	6 months	0.35	0.19	24.81	-14.40	3.19	1.23	15.79
	1 year	0.65	0.74	47.24	-33.23	6.72	0.56	12.50
	2 years	1.07	1.02	82.48	-71.01	12.60	0.20	11.30
	5 years	1.77	2.17	111.49	-128.11	26.02	-0.22	6.08
	10 years	2.36	2.55	152.26	-163.35	42.95	0.02	4.62
Growth rates	CPI	5.29	4.18	23.96	-5.35	4.34	1.00	4.25
	TPC	3.52	3.35	29.87	-25.59	7.16	0.14	4.76

Notes

1. All series are in annualised percentages.
2. The macroeconomic variables' growth rates are month-on-month.

Table 2: Sample Correlations

		6 months	1 year	2 years	5 years	10 years	CPI	TPC
Excess holding period bond returns	6 months	1.00						
	1 year	0.95	1.00					
	2 years	0.88	0.96	1.00				
	5 years	0.76	0.87	0.94	1.00			
	10 years	0.67	0.76	0.84	0.94	1.00		
Growth rates	CPI	-0.10	-0.12	-0.15	-0.17	-0.22	1.00	
	TPC	-0.05	-0.07	-0.06	-0.09	-0.07	-0.26	1.00

Table 3a: Estimation results
Bond maturities: 6 months, 1, 2, 5 & 10 years
Macroeconomic variables: CPI Inflation & Total Personal Consumption

Estimation	In mean coefficients										Explanatory power				
	6 months		1 year		2 year		5 year		10 year		6 months	1 year	2 year	5 year	10 year
	CPI	TPC	CPI	TPC	CPI	TPC	CPI	TPC	CPI	TPC					
	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15
No-arbitrage C-CAPM	0.001	-0.072	0.001	-0.072	0.001	-0.072	0.001	-0.072	0.001	-0.072	0.00	0.00	0.00	0.00	0.00
No-arbitrage SDF	-0.153	-0.057	-0.153	-0.057	-0.153	-0.057	-0.153	-0.057	-0.153	-0.057	0.00	0.00	0.00	0.00	0.00
Arbitrage 1	-0.424	0.003	-0.415	0.003	-0.387	0.003	-0.384	0.003	-0.298	0.003	0.14	0.10	0.07	0.04	0.02
Arbitrage 2	-0.426	0.109	-0.426	0.040	-0.426	0.071	-0.426	0.067	-0.426	0.124	0.12	0.09	0.08	0.05	0.05
Arbitrage General	-0.460	0.146	-0.373	0.033	-0.342	0.043	-0.268	0.014	-0.182	-0.020	0.06	0.03	0.02	0.01	0.01

Notes

1. The excess return equations of the models estimated are the following:
 - No-arbitrage C-CAPM: $E_t(h_{n,MH} - s_t) = 0.5V_t(h_{n,MH}) + \beta_1 \text{Cov}_t(h_{n,MH}, \Delta c_{n,MH}) + \text{Cov}_t(h_{n,MH}, \pi_{n,MH})$
 - No-arbitrage SDF: $E_t(h_{n,MH} - s_t) = 0.5V_t(h_{n,MH}) + \beta_1 \text{Cov}_t(h_{n,MH}, \Delta c_{n,MH}) + \beta_2 \text{Cov}_t(h_{n,MH}, \pi_{n,MH})$
 - Arbitrage 1: $E_t(h_{n,MH} - s_t) = 0.5V_t(h_{n,MH}) + \beta_1 \text{Cov}_t(h_{n,MH}, \Delta c_{n,MH}) + \beta_2 \text{Cov}_t(h_{n,MH}, \pi_{n,MH})$
 - Arbitrage 2: $E_t(h_{n,MH} - s_t) = 0.5V_t(h_{n,MH}) + \beta_{1,n} \text{Cov}_t(h_{n,MH}, \Delta c_{n,MH}) + \beta_{2,n} \text{Cov}_t(h_{n,MH}, \pi_{n,MH})$
 - Arbitrage General: $E_t(h_{n,MH} - s_t) = 0.5V_t(h_{n,MH}) + \beta_{1,n} \text{Cov}_t(h_{n,MH}, \Delta c_{n,MH}) + \beta_{2,n} \text{Cov}_t(h_{n,MH}, \pi_{n,MH})$
2. Columns TPC and CPI report the estimated β_1 and β_2 respectively, for different maturities and models.
3. Since the model is highly non-linear, the estimates are not unit free. For example, for the C-CAPM estimation, the theory implies that the coefficient of the inflation covariances should be constrained to 1. However, all our data are in annualised percentages, for numerical ease. Subsequently we must instead constrain the coefficient of the inflation covariances to $1/(1/1200) = 0.001$ as appears above. Conversely, in the subsequent estimation which does not restrict the coefficient of the inflation covariances, one can recover the implied value for the raw data by inverting the above calculation: $-0.153 \times 1200 = -183.6$.
4. *t*-statistics are below the estimated parameters in italics.
5. The explanatory power of the model is measured as the share of the variance of the excess holding-period return explained for each maturity.

Table 3b: Likelihood ratio tests

Estimation	Joint Significance	No-arbitrage C-CAPM	No-arbitrage SDF	Arbitrage 1	Arbitrage 2
No-arbitrage C-CAPM	4.4 # (1)				
No-arbitrage SDF	12.1 # (2)	7.8 # (1)			
Arbitrage 1	18.9 # (6)	14.6 # (5)	6.8 (4)		
Arbitrage 2	25.4 # (6)	21.1 # (5)	13.3 # (4)		
Arbitrage General	53.2 # (10)	48.9 # (9)	41.1 # (8)	34.3 # (4)	27.8 # (4)

Notes

1. The column title corresponds to the null hypothesis tested (the restricted model). The row title corresponds to the alternative hypothesis (the unrestricted model).
2. # denotes rejection of the null at the 5% significance level.
3. The first column tests the joint significance of all covariance terms in the estimation.
4. The number of parameter constraints tested is in parentheses, to the right of the corresponding test statistic.

Table 4: Descriptive statistics of estimated risk premia

	6 months	1 year	2 years	5 years	10 years
Mean	0.57	1.09	1.81	2.89	4.68
Median	0.32	0.78	1.34	2.21	3.74
Maximum	5.17	6.24	11.89	13.50	15.14
Minimum	-1.43	-0.36	-2.57	-4.07	-0.18
Std. Dev.	0.77	1.11	1.84	2.72	2.79
Skewness	2.27	2.28	2.28	1.65	1.48
Kurtosis	10.24	8.10	8.79	5.64	4.70

Notes

1. Estimated risk premia from estimation 10 have been used.

Table 5: Correlations of estimated risk premia

	6 months	1 year	2 years	5 years	10 years
6 months	1.00				
1 year	0.90	1.00			
2 years	0.87	0.99	1.00		
5 years	0.72	0.91	0.94	1.00	
10 years	0.62	0.85	0.88	0.96	1.00

Notes

1. Estimated risk premia from estimation 10 have been used.

**Table 6a: Reconstructed rolling risk premia and yields
(static expectations)**

Maturities	Observed Yields (A)	Reconstructed Yields (B)	Rolling Risk Premium	Rolling Risk Premium (% of A)	Error (A-B)	Error (as a % of A)
Average of period 1970-1998						
6 months	7.10	7.09	0.29	4.04	0.01	0.17
1 year	7.34	7.38	0.58	7.93	-0.04	-0.53
2 years	7.64	7.83	1.03	13.52	-0.19	-2.51
5 years	8.04	8.63	1.83	22.80	-0.59	-7.32
10 years	8.32	9.88	3.08	37.02	-1.56	-18.72
Average of period 1970-1978						
6 months	6.28	6.12	0.27	4.34	0.16	2.61
1 year	6.51	6.36	0.51	7.83	0.15	2.36
2 years	6.72	6.66	0.82	12.14	0.05	0.81
5 years	7.05	7.00	1.16	16.39	0.05	0.70
10 years	7.28	7.65	1.80	24.71	-0.36	-5.01
Average of period 1979-1982						
6 months	12.06	12.66	0.92	7.59	-0.60	-4.98
1 year	12.11	13.51	1.77	14.59	-1.40	-11.55
2 years	11.98	14.79	3.04	25.36	-2.80	-23.39
5 years	11.77	16.86	5.12	43.47	-5.09	-43.24
10 years	11.65	19.17	7.42	63.65	-7.51	-64.45
Average of period 1983-1989						
6 months	7.86	7.73	0.20	2.48	0.13	1.60
1 year	8.24	8.00	0.47	5.67	0.24	2.86
2 years	8.73	8.47	0.94	10.74	0.25	2.91
5 years	9.30	9.58	2.04	21.98	-0.28	-3.02
10 years	9.63	11.43	3.90	40.46	-1.80	-18.72
Average of period 1990-1998						
6 months	5.10	5.06	0.09	1.84	0.05	0.92
1 year	5.34	5.18	0.22	4.05	0.16	3.05
2 years	5.77	5.39	0.43	7.45	0.38	6.61
5 years	6.39	5.84	0.88	13.73	0.55	8.55
10 years	6.85	6.74	1.78	25.97	0.11	1.57

Notes

1. Estimated risk premia from estimation 10 have been used.

**Table 6b: Reconstructed rolling risk premia and yields
(perfect foresight of short rates and risk premia)**

Maturities	Observed Yields (A)	Reconstructed Yields (B)	Rolling Risk Premium	Rolling Risk Premium (% of A)	Error (A-B)	Error (as a % of A)
Average of period 1970-1988						
6 months	7.99	8.00	0.39	4.50	-0.01	0.48
1 years	8.24	8.41	0.77	8.72	-0.17	-1.70
2 years	8.50	9.06	1.34	14.89	-0.56	-7.35
5 years	8.82	10.17	2.37	27.52	-1.35	-19.87
10 years	9.03	11.50	3.90	46.87	-2.47	-36.68
Average of period 1970-1978						
6 months	6.28	6.20	0.28	4.32	0.09	1.63
1 years	6.51	6.56	0.54	8.13	-0.05	-0.90
2 years	6.72	7.34	0.97	14.06	-0.62	-9.47
5 years	7.05	9.94	2.39	32.70	-2.89	-39.77
10 years	7.28	13.46	4.98	68.18	-6.18	-85.18
Average of period 1979-1982						
6 months	12.06	12.62	0.95	7.91	-0.56	-5.56
1 years	12.11	13.46	1.87	15.57	-1.35	-13.32
2 years	11.98	14.45	3.06	26.19	-2.46	-24.73
5 years	11.77	13.48	3.61	33.11	-1.71	-20.75
10 years	11.65	12.53	4.19	37.82	-0.87	-11.81
Average of period 1983-1988						
6 months	7.78	7.56	0.18	2.50	0.23	2.82
1 years	8.21	7.76	0.39	5.03	0.45	4.87
2 years	8.79	8.00	0.72	8.59	0.79	7.35
5 years	9.46	8.28	1.52	16.17	1.18	10.02
10 years	9.84	7.93	2.12	21.51	1.91	18.16

Notes

1. Estimated risk premia from estimation 10 have been used.

Table 7: Tests of the REHTS with and without risk premia

	Holding-period equation		Term spread equation	
	a	b	c	d
6-month bond				
Constant	-0.183 #	0.415 #	-0.173 #	0.332 #
	-0.24	0.69	-0.83	1.41
Short rate	1.076 #	0.904 #	0.471 #	0.402
	8.47	8.89	1.44	1.33
Risk Premium		0.995 #		-0.842 #
		2.08		-1.06
Adjusted R^2	0.482	0.498	0.017	0.056
1-year bond				
Constant	-0.726 #	0.052 #	-0.411 #	0.639 #
	-0.49	0.04	-0.78	1.06
Short rate	1.197 #	0.917 #	0.632 #	0.546 #
	4.79	4.47	1.10	1.03
Risk Premium		1.029 #		-0.917 #
		1.36		-1.27
Adjusted R^2	0.205	0.214	0.007	0.027
2-year bond				
Constant	-0.838 #	0.458 #	-0.537 #	1.315 #
	-0.32	0.21	-0.48	1.07
Short rate	1.269 #	0.787 #	0.457 #	0.269 #
	2.89	2.00	0.53	0.36
Risk Premium		1.095 #		-0.933 #
		1.19		-1.14
Adjusted R^2	0.074	0.084	-0.001	0.014
5-year bond				
Constant	1.115 #	2.647 #	0.477 #	2.915 #
	0.23	0.68	0.20	1.30
Short rate	1.082 #	0.396 #	-0.721 #	-0.878 #
	1.37	0.60	-0.55	-0.73
Risk Premium		1.083 #		-0.775 #
		1.24		-0.95
Adjusted R^2	0.011	0.015	-0.001	0.002
10-year bond				
Constant	4.647 #	4.223 #	2.430 #	4.838 #
	0.62	0.56	0.61	0.95
Short rate	0.643 #	0.167 #	-2.044 #	-2.142 #
	0.53	0.13	-1.14	-1.25
Risk Premium		0.784 #		-0.483 #
		0.61		-0.41
Adjusted R^2	-0.001	-0.002	0.003	0.002

Notes

1. Estimated risk premia from estimation 10 have been used.
2. # denotes that the hypothesis of equality to the theoretical value can not be rejected at the 5% significance level.
3. t-statistics using White heteroskedasticity-consistent standard errors are below the estimated parameters in italics.

**Figure 1: Estimated Risk Premia (including Jensen effect)
and their macroeconomic components**

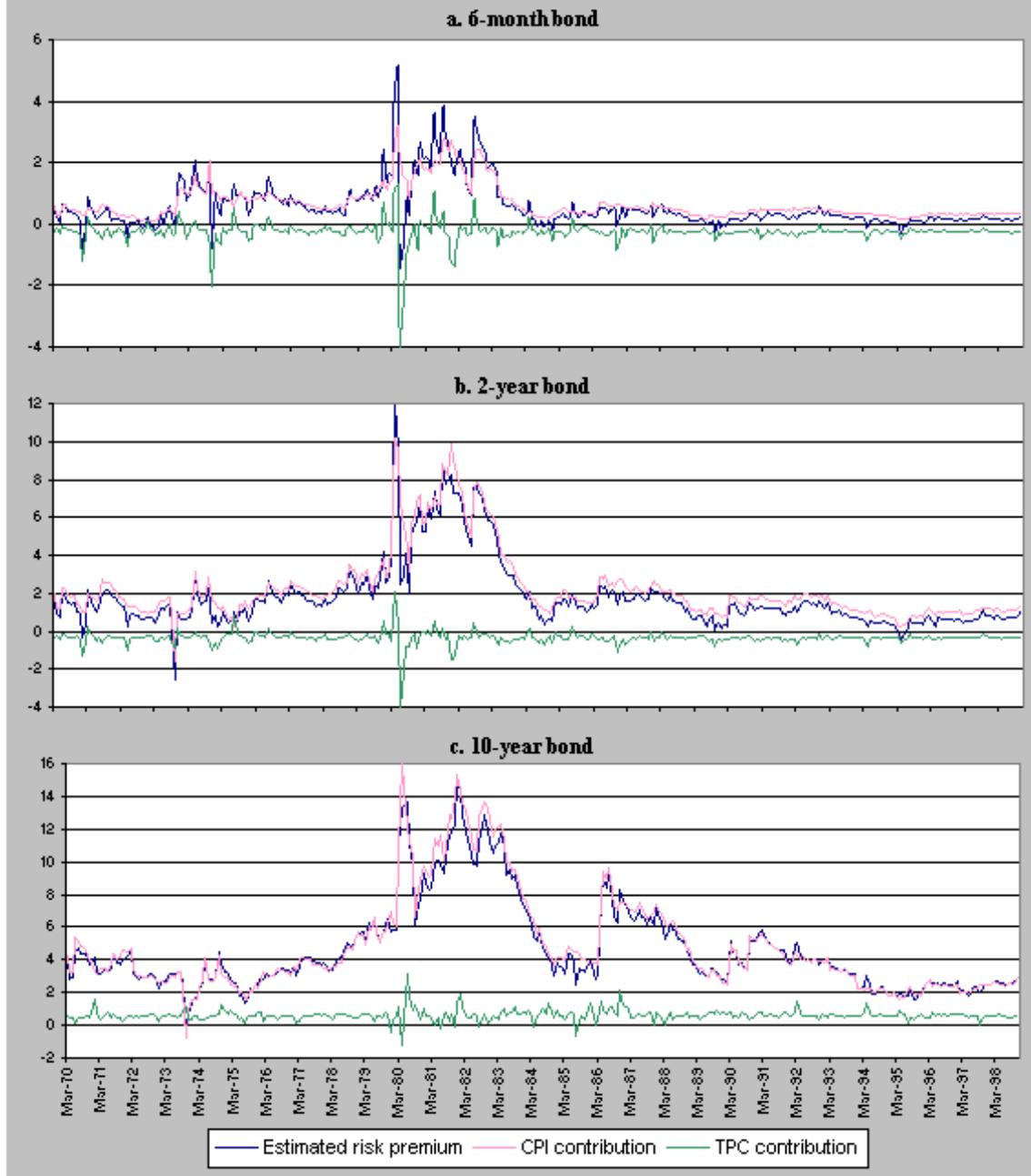


Figure 2: Estimated Risk Premia (including Jensen effect) and Excess Returns

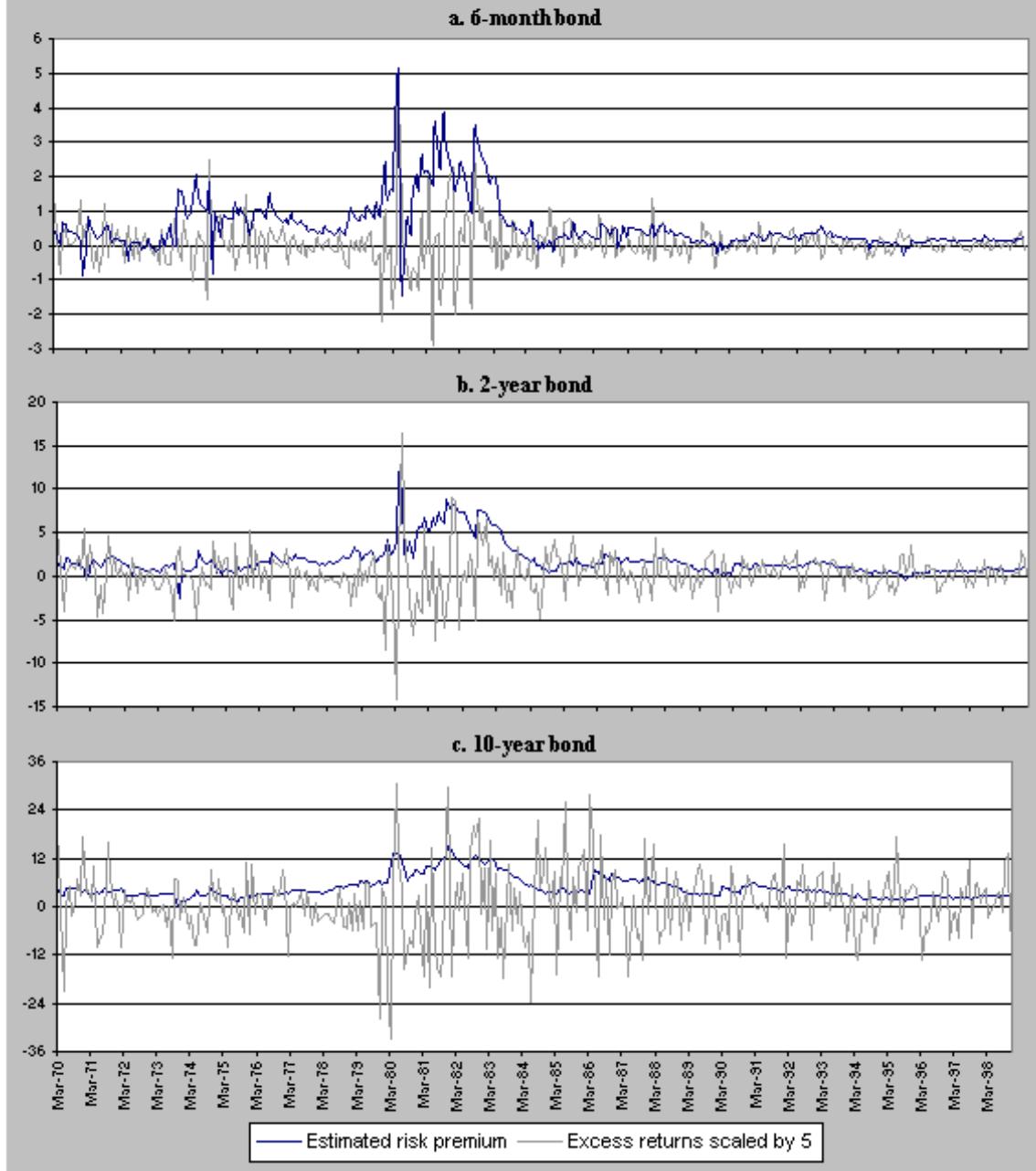


Figure 3: Estimated Risk Premia (including Jensen effect)

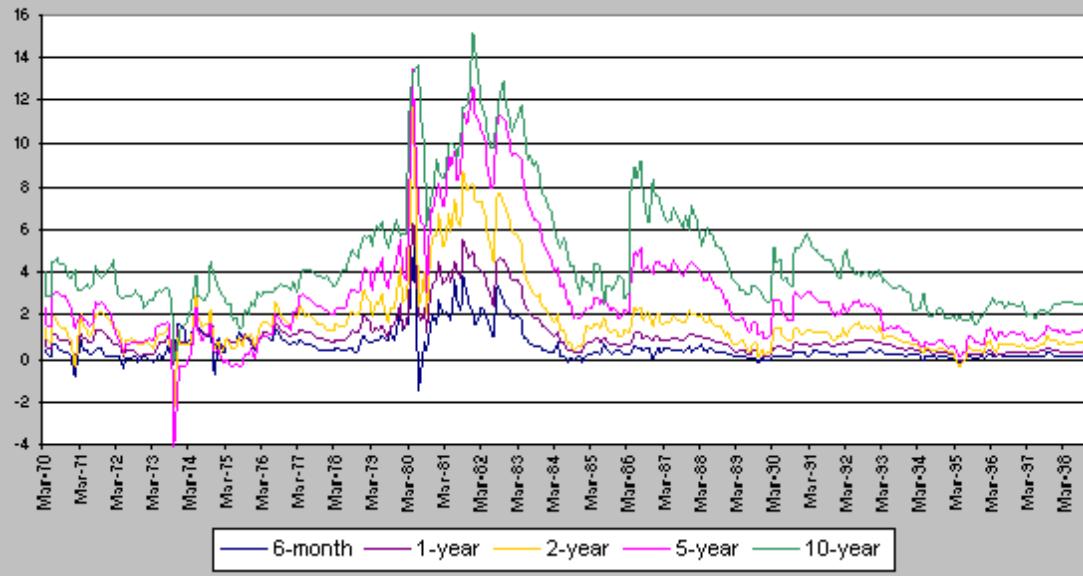


Figure 4: Rolling risk premia in bonds of different maturities
Static expectations

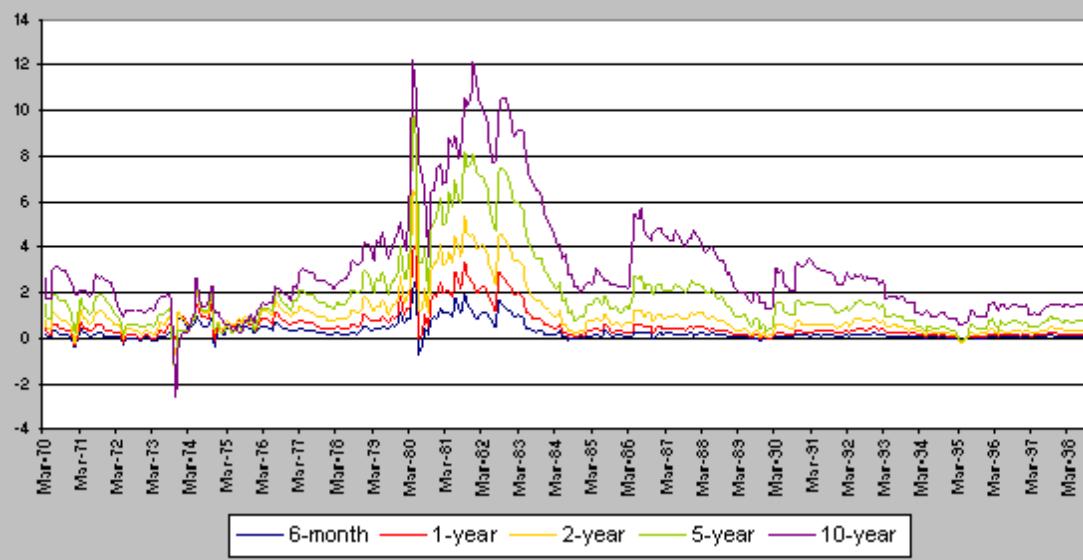


Figure 5: Rolling risk premia in bonds of different maturities
Perfect foresight

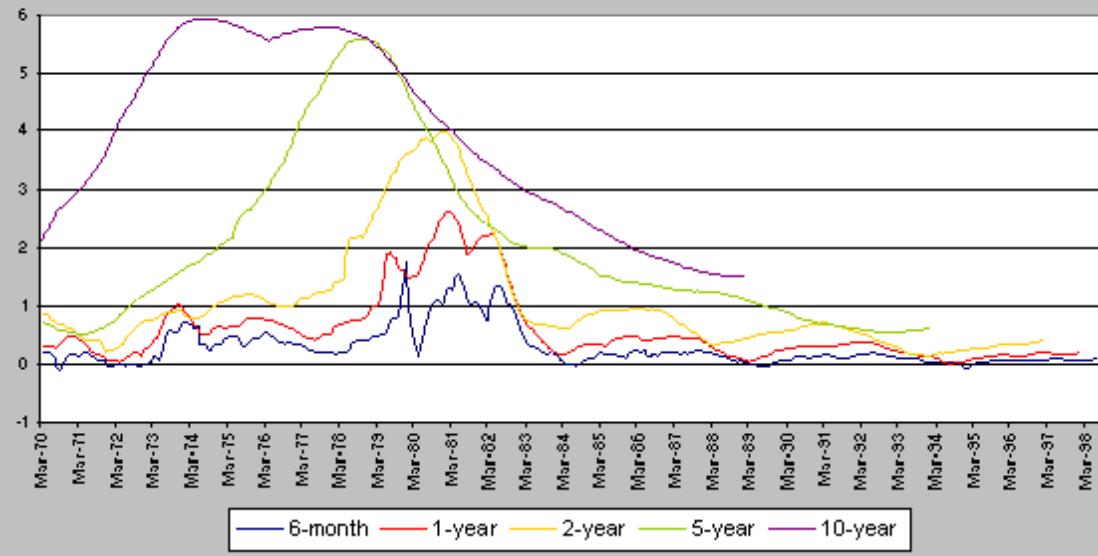


Figure 6: Rolling risk premia in bonds of different maturities as % of yield
Static expectations

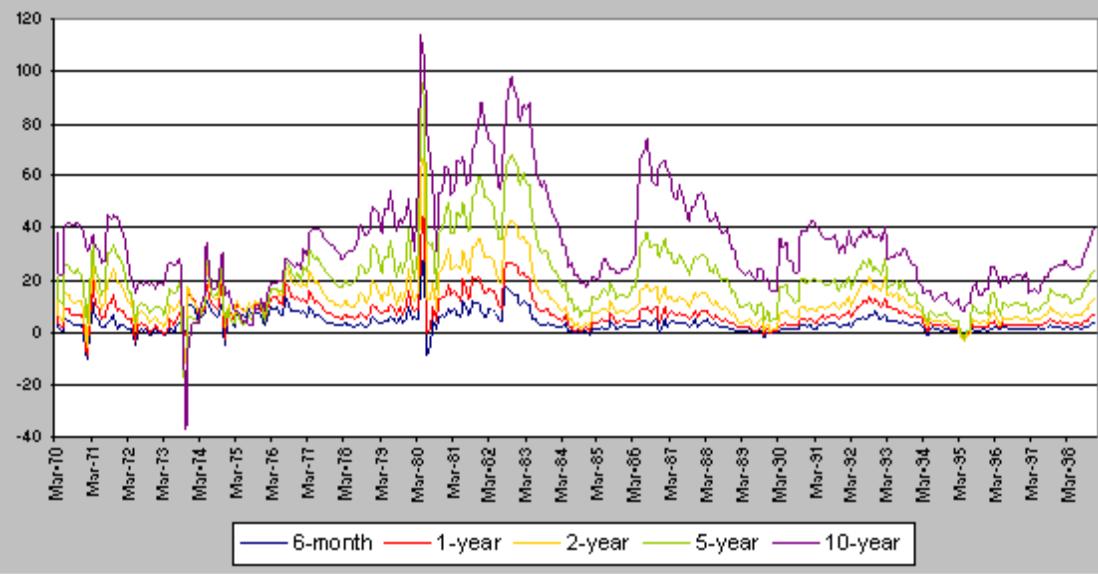


Figure 7: Rolling risk premia in bonds of different maturities as % of yield
Perfect foresight

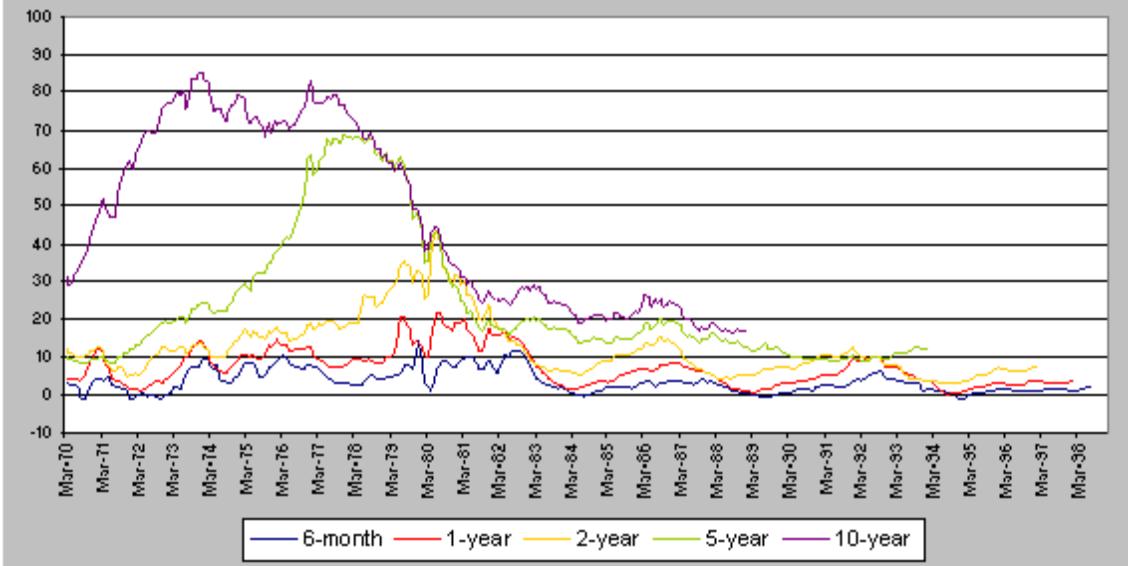
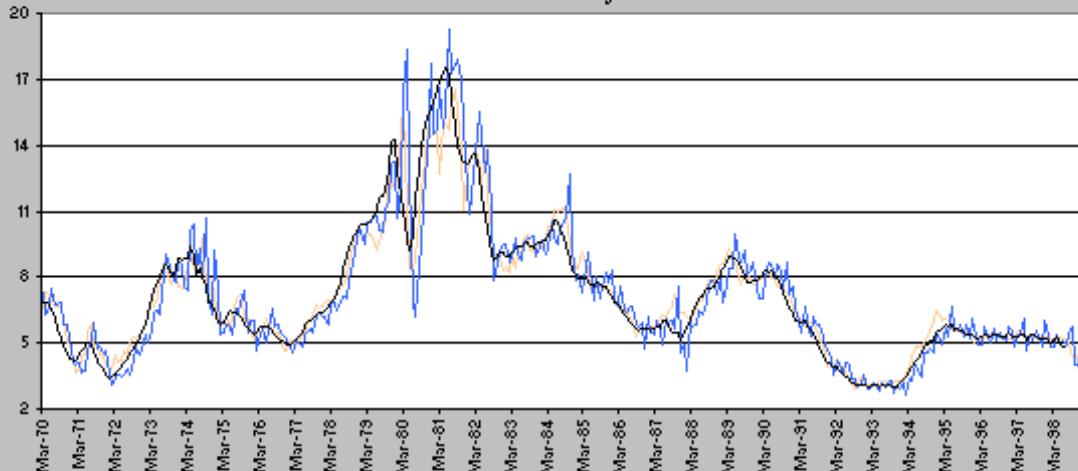
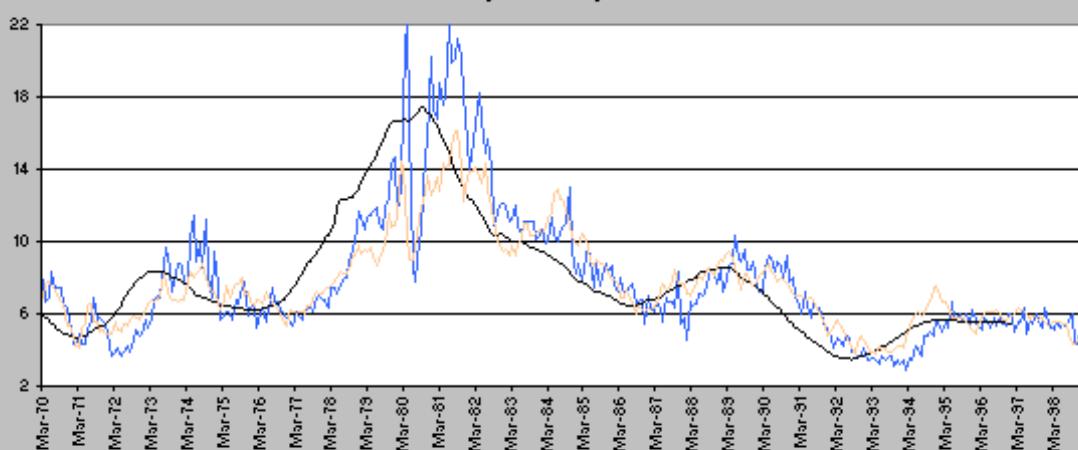


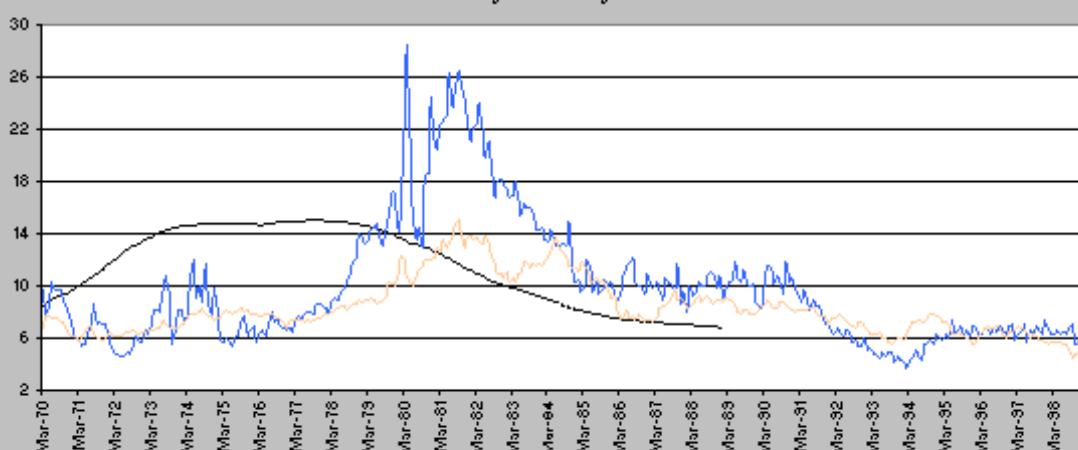
Figure 8: Yields
a. 6-month bond yields



b. 2-year bond yields



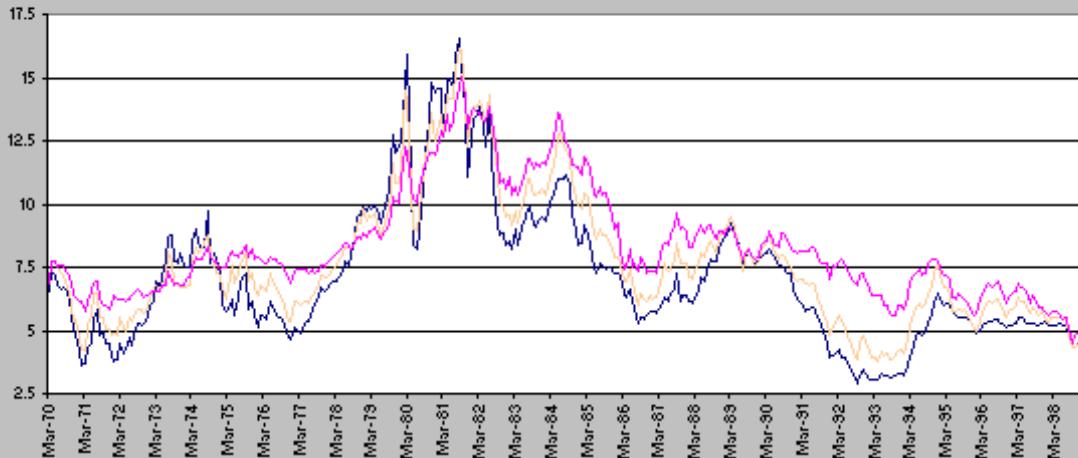
c. 10-year bond yields



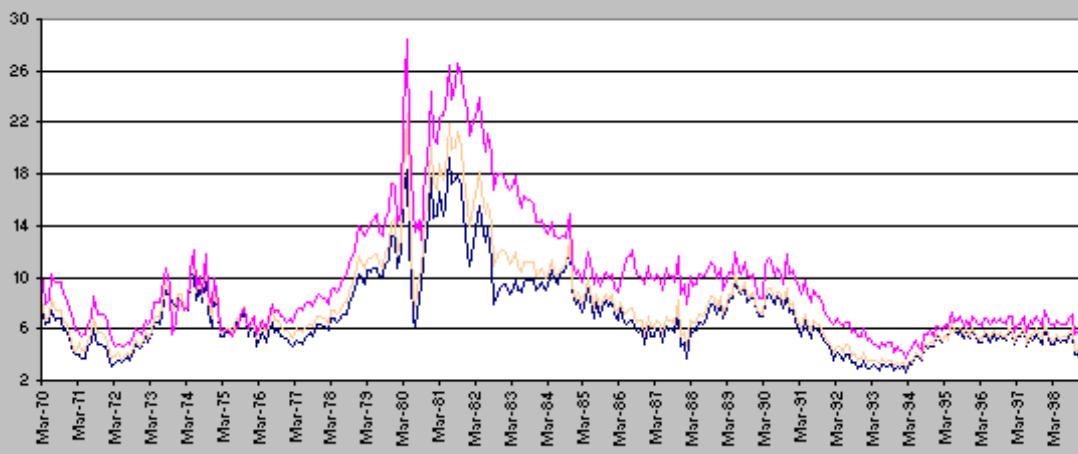
— Perfect foresight of risk premia and short rates — Static Expectations — Actual Yields

Figure 9: Term Structures

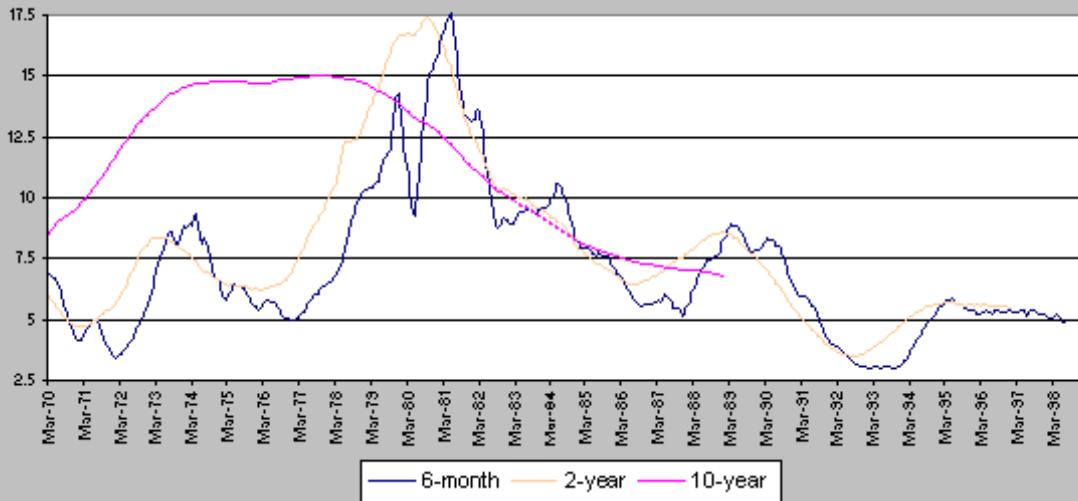
a. Actual Yields



b. Reconstructed Yields - Static expectations of risk premia and short rates



c. Reconstructed Yields - Perfect foresight of risk premia and short rates



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