NEW YORK UNIVERSITY
STERN SCHOOL OF BUSINESS
FINANCE DEPARTMENT

Working Paper Series, 1996

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FIN-96-17

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March 29, 1996

Abstract

We provide a test of the liquidity preference hypothesis (i.e., the monotonicity of *ex ante* term premiums), conditioning on the shape of the yield curve. The approach we use is general, and does not require a structural model for conditional expected returns. Using nonparametric estimates, the evidence supports previous conclusions in the literature regarding time-varying negative term premiums. For example, in periods in which the term structure is downward sloping, we find that the premiums can be significantly negative and are often monotonically decreasing in maturity. Interestingly, in these periods the volatility of the term premium is still increasing in maturity, indicating that bond return volatility is not a priced risk.

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1 Introduction

The liquidity preference hypothesis (LPH) (see, for example, Hicks (1939) and Kessel (1965)) states that the *ex ante* return on government securities is a monotonically increasing function of time to maturity. That is, conditional on all available information, the expected 3-month return on a T-bill with 9 months to maturity should exceed the expected 3-month return on a 6-month T-bill, which should be greater than the certain yield on a 3-month T-bill, and so forth. The LPH implies this condition, regardless of the shape of the term structure or any other economic variables contained in the agent's information set. The underlying intuition behind the LPH is that longer term bonds are more *risky*. With flat term structures and parallel shifts of the yield curve, longer term bonds are more sensitive to interest rate changes than shorter term bonds. Individuals need to be compensated for the risk of holding these bonds, hence the higher expected return.

Using post 1963 data on T-bills, direct tests of the unconditional version of the liquidity preference hypothesis have been performed by Fama (1984b), McCulloch (1987) and Richardson, Richardson and Smith (1992). While there is some disagreement concerning the reliability of the data in the 1964-1972 period, the evidence suggests that expected returns are monotonic (throughout, monotonic is taken to mean monotonically increasing). For example, Richardson, Richardson and Smith (1992) find that, when one correctly tests for monotonicity using inequality constraints, there is little evidence against the LPH.

This evidence, however, comprises only unconditional tests of the theory. These tests are expected to have low power because the econometrician is ignoring the information available to economic agents. The LPH relates conditional expected returns across maturities; thus, unconditional tests provide very weak tests of the underlying theory. In fact, asset pricing theory suggests that the current term structure contains important information for expected returns on bonds of different maturities. For example, suppose that expected returns are monotonically increasing in maturity when the term structure is upward sloping, yet decreasing in maturity when the term structure is downward sloping. Since upward sloping term structures occur more often, unconditional tests will not be able to reject the monotonicity of returns because the tests average over all term structure shapes.

A strand of the literature recognizes this problem, and documents time-varying expected returns on bonds, e.g., Fama (1986), Fama and Bliss (1987), Stambaugh (1988), Fama and French (1989) and Klemkosky and Pilotte (1992), among others. All of these papers suggest that the *fitted* values of *ex ante* bond returns are not always increasing with the maturity of the bond. In order to correctly interpret these results, however, it is necessary to consider the joint statistical properties of these estimates of *ex ante* bond returns across maturities. Consequently, it may not be surprising then that no formal test of the LPH (using conditioning information) has been performed. The difficulty is that the LPH implies a set of inequality restrictions on the *ex ante* returns on bonds of different maturities. Since these *ex ante* returns are unobservable, and statistical methods for testing inequality restrictions have only recently been developed, only anecdotal evidence regarding the LPH appears in the finance literature.

This paper provides several contributions to the literature on bond returns. First, we present some theoretical results on *ex ante* bond returns which provide some additional insights on the validity of the LPH. These results complement existing theory in the literature and motivate the choice of the shape of the term structure as a conditioning variable. Second, we conduct a formal *ex ante* test of the LPH. Using information contained in the yield curve, we estimate conditional mean returns on bonds of different maturities. These means are then compared cross-sectionally using recently developed techniques from the econometric literature for testing inequality constraints (see, for example, Wolak (1989)). Third, while the results are mixed due to sampling variation of the term premium estimates, we provide evidence to support previous conclusions in the literature regarding time-varying negative term premiums. In particular, we document important states of the world in which the LPH may be violated. For example, in periods in which the term structure is downward sloping, we find that the premiums can be negative (as low as -7%, annualized) and are often monotonically decreasing in maturity. Interestingly, in these periods the volatility of the term premium is still increasing in maturity, indicating that bond return volatility is not a priced risk.

The paper is organized as follows. Section 2 reviews the relevant economics and econometric literature on the LPH. We focus on identifying useful conditioning information for testing the LPH. In Section 3, we illustrate the inequality testing methodology by replicating and then reinterpreting Fama's (1986) original work on time-varying premiums. In Section 4, we provide the main results of the paper. We first describe the data used in the study, and some of the salient characteristics of the relation between bond returns and the yield curve. We then provide formal tests of the LPH. In particular, we analyze term premiums on bonds with long maturities. Section 5 provides some concluding remarks including an economic interpretation of the empirical results.

2 Preliminaries

2.1 Theory

Forming more powerful tests of the LPH through the use of conditioning information does not require a structural model of *ex ante* bond returns. However, for the test to be powerful, the conditioning set must provide useful information about alternative theories, i.e., states in which the LPH may not be valid. Although the LPH is consistent with a variety of term structure models, bond pricing theory does not imply the LPH as a condition for equilibrium.[1] Below, we investigate how the LPH relates to the term structure of interest rates in a general setting. We focus on information which may point to violations of the LPH.

Let $X_{t,t+j}$ be the n-dimensional vector of asset returns from t to $t+j$, and define E_t as the expectation operator conditional on the information available at time t. In the absence of arbitrage opportunities, there exists a positive stochastic discount factor M which satisfies the following condition (Harrison and Kreps (1979)):

$$E_t[X_{t,t+j}M_{t,t+j}] = 1. \tag{1}$$

Thus, the price at time t of a pure discount bond which pays off 1 unit at time $t+j$, irrespective of the state, is just

$$P_{t,j} = \frac{1}{(Y_{t,t+j})^j} = E_t[M_{t,t+j}], \tag{2}$$

[1] In an economic environment in which future production possibilities are independent of the current economic state, Benninga and Protopapadakis (1986) examine a general equilibrium economy for general specifications of utility and production functions. They find that, in complete markets, the LPH's main conclusions are valid. The intuition is that longer bonds are a poor hedge for the representative agent's short-term consumption, thus requiring a premium to hold these bonds over shorter horizons. Alternatively, in the Cox, Ingersoll and Ross (1985) one-factor model of interest rates, term premiums on instantaneous holding periods are monotonic in the maturity of the bond.

where $Y_{t,t+j}$ is the yield-to-maturity on a j-period zero coupon bond. The holding period return on this j-period bond is then equal to

$$R_{t,t+1}(j) = \frac{P_{t+1,j-1}}{P_{t,j}} = \frac{E_{t+1}[M_{t+1,t+j}]}{E_t[M_{t,t+j}]}. \tag{3}$$

Since it is common to define yields in terms of logarithms, let $y_{t,t+j} = \ln(Y_{t,t+j})$ and $r_{t,t+1}(j) = \ln(R_{t,t+1}(j))$ (see, for example, Fama (1984), Campbell and Shiller (1991) and Engle and Ng (1993)). Then using equation (2), the fact that $E_t[M_{t,t+j}]$ equals $E_t[M_{t,t+1}]E_t[M_{t+1,t+j}] + cov_t(M_{t,t+1}, M_{t+1,t+j})$, and the linearization $z \approx \ln(1+z)$ for small z, the expected holding period log return on a j-period bond $E_t[r_{t,t+1}(j)]$ can be written as

$$E_t[r_{t,t+1}(j)] \approx y_{t,t+1} - cov_t(M_{t,t+1}, M_{t+1,t+j})Y_{t,t+1}Y^j_{t,t+j}, \tag{4}$$

where cov_t is the covariance operator conditional on information at time t.

It is common to define the risk premium on long bonds (i.e., $j > 1$) as the expected return on the bond in excess of the risk-free rate, that is, $E_t[r_{t,t+1}(j) - y_{t,t+1}]$.[2] Further, denote $cov_t(M_{t,t+1}, M_{t+1,t+j})Y_{t,t+1}Y^j_{t,t+j}$ as $cov_t(M^*_{t,t+1}, M^*_{t+1,t+j})$, which can be interpreted as the covariance between scaled discount factors.[3] Equation (4) states that the one-period risk premium on long bonds is approximately equal to minus the conditional covariance between the scaled stochastic discount factor next period and the scaled stochastic discount factor over the following $j-1$ periods, i.e.,

$$E_t[r_{t,t+1}(j) - y_{t,t+1}] \approx -cov_t(M^*_{t,t+1}, M^*_{t+1,t+j}) .$$

The intuition behind this relation is clear if we interpret stochastic discount factors in terms of the consumption based asset pricing model of Lucas (1978), among others. In this model, the discount factor is the marginal rate of substitution between consumption flows at different times, which, for standard parameterizations, is inversely related to the growth rate of consumption. If $cov_t(M_{t,t+1}, M_{t+1,t+j}) > 0$, then low consumption growth over the next period means low expected consumption growth in the future. In these states of the world, future consumption levels are expected to

[2]Without the linear approximation, $E_t[r_{t,t+1}(j) - y_{t,t+1}]$ should be replaced by $E_t[R_{t,t+1}(j) - Y_{t,t+1}]$. Since the literature focuses on log returns, we maintain the approximation throughout the rest of the paper.

[3]Recall that $\frac{1}{(Y_{t,t+j})^j} = E_t[M_{t,t+j}]$, so the scaled stochastic discount factors are weighted by values close to their conditional means.

be very low and therefore interest rates are low and long bonds are relatively valuable. Consequently, the high return on long bonds effectively hedges consumption risk and they command a smaller risk premium than shorter term bonds. On the other hand, if $cov_t(M_{t,t+1}, M_{t+1,t+j}) < 0$, then low consumption growth over the next period means higher expected future consumption. Thus, the relation between the return on long bonds and consumption is reversed. The premium on long bonds needs to be sufficiently high to keep investors holding these bonds.

The above interpretation immediately provides clues for the identification of states in which the risk premium on bonds may not be monotonically increasing. The key variable is $cov_t(M^*_{t,t+1}, M^*_{t+1,t+j})$; however, autocovariances of the stochastic discount factor are one of the cornerstones for pricing long-term assets. Thus, the term structure of interest rates will yield information about these autocovariances.

In particular, the slope of the term structure, defined here as the yield spread between j-period and 1-period bonds can be written as

$$\begin{aligned} y_{t,t+j} - y_{t,t+1} &= E_t\left[\frac{1}{j}\sum_{i=1}^{j-1}(y_{t+i,t+i+1} - y_{t,t+1})\right] \\ &\quad - E_t\left[\frac{1}{j}\sum_{i=1}^{j-1} cov_{t+i-1}(M^*_{t+i-1,t+i}, M^*_{t+i,t+j})\right], \end{aligned} \tag{5}$$

where the same linear approximation is used as before. The yield spread is made up of two components: (i) expected changes in future short-term rates, and (ii) expectations of future conditional autocovariances of the scaled stochastic discount factor.

Equation (5) implies that the term structure will be downward sloping under two circumstances. On the one hand, if expected future short rates are expected to decrease relative to the current short rate, then the first component of the term structure slope will be negative. If the autocovariances of the stochastic discount factors are sufficiently small, this effect will dominate. Note that when the autocovariances are all zero, the risk premium on bonds is zero, and the Expectations Hypothesis (EH) holds. On the other hand, the larger and more positive we expect the future autocovariances to be, the more likely the term structure is to be flat or negatively sloped. That is, positive autocovariances of stochastic discount factors tend to be associated with downward sloping term structures.

The second component of the term structure slope is simply the sum of current and expected future risk premiums on a j-period bond. Ceteris paribus, equation (5)

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implies that negative risk premiums on bonds occur during periods in which the term structure is relatively flat or inverted. For example, if future expected short-term rates are equal to the current short-term rate, then the sign of the risk premium coincides with the slope of the term structure.[4] This relation is to be expected — the risk premium on long bonds is reflected in their yields which also determine the slope of the term structure.

Of course, this analysis depends on the relation between the yield spread between long- and short-term bonds and expected movements in future short-term yields. Existing evidence suggests that this relation is positive, yet far from one-to-one (e.g., see Fama and Bliss (1987) and Campbell and Shiller (1991)). Thus, the shape of the yield curve may have important information about the conditional covariances between future stochastic discount factors, suggesting a possible conditioning set which will provide a high hurdle for tests of the LPH. It is an empirical question whether these conditional covariances are highly correlated with expected movements in short-term yields, driving a wedge between *ex ante* bond returns and the shape of the yield curve.

2.2 Existing Empirical Evidence

Existing results on the LPH's validity are usually embedded in broader investigations of time-varying risk premia on bonds. For example, Shiller, Campbell and Schoenholtz (1983), Fama (1984b, 1986), Keim and Stambaugh (1986), Fama and Bliss (1987), Stambaugh (1988), Campbell and Shiller (1991), Klemkosky and Pilotte (1992) and Engle and Ng (1993) all report evidence that the risk premiums on bonds of various maturities are predictable. In a seminal piece, Fama (1986) documents time-varying movements in term premiums which depend on the business cycle. With respect to the liquidity preference hypothesis, Fama (1986, p.176) states

> *term premiums are generally interpreted as rewards for risk. In this view, the changes from upward sloping term structures of expected returns during good times to humped and inverted term structures of expected returns*

[4]While this intuition is explained for a comparison between a j-period and 1-period bond, it carries through for any maturity. Ceteris paribus, if $cov_t(M^*_{t+1}, M^*_{t+j})$ is greater than $cov_t(M^*_{t+1}, M^*_{t+k})$ $(j > k)$, then "flatter/more inverted" term structures from maturities k to j will also correspond to decreasing risk premiums between periods j and k — a violation of the LPH.

during recessions imply that the ordering of risks and rewards across maturities changes with the business cycle and is not always monotonic.

While Fama's (1986) findings are consistent with the theory in Section 2.1 above, it is difficult to assess the statistical significance of these results. First, although the individual mean estimates of the premiums suggest expected returns are not monotonically increasing, these results are not interpreted jointly across maturities. Given the high correlation across the premiums, the need for a joint test seems especially clear. Second, Fama (1986) uses term structure shapes as his conditioning variable for the state of the economy. Given that these shapes may be correlated from month to month, the relevant test statistics need to be adjusted for serial correlation in the series.

In a related, and somewhat more formal, setting, Stambaugh (1988) adds to Fama's (1986) evidence by showing that a two latent variable model of expected returns on T-bills produces similar results. In particular, he shows that expected returns exhibit variation with business cycles which is non-monotonic. However, while his paper certainly suggests non-monotonicity of term premiums, the evidence is based on *ex post* fitted estimates. Stambaugh (1988) does not perform an *ex ante* test of the LPH. This requires tests of multiple inequality constraints and a model for conditional expected returns.

Most of the existing studies generally investigate term premiums on shorter-term bonds. An exception is Fama and Bliss (1987) who extend Fama (1984a, 1986) to investigate variation in *ex ante* term premiums on U.S. Treasury bonds of maturities greater than one year (see also Fama and French (1989)).[5] Using a methodology similar to Fama (1984a), Fama and Bliss (1987) document non-monotonic term premiums for longer maturity bonds. The signs of these *fitted* premiums depend on the shape of the term structure as determined by information in long-term forward rates. In terms of the LPH, however, it is difficult to interpret the statistical significance of these results without looking at the estimates jointly across maturities.

[5] More recently, Klemkosky and Pilotte (1992) investigate variation in *ex ante* term premiums on U.S. Treasury bills and bonds of a variety of maturities. As predicted by a popular class of term-structure models (e.g., Vasicek (1977)), they find that the premiums are monotonically related to the conditional volatility of the short-rate of interest, as measured by information contained in the short-term yield curve. The Klemkosky and Pilotte (1992) findings are especially interesting because they represent some of the first results displaying time variation in short-term expected returns on bonds of long maturities.

In our analysis, we explore how various yield curve shapes interact with premiums on bonds of different maturities, and examine the implication for tests of the LPH. Similar to Klemkosky and Pilotte (1992), these tests are performed using short-horizon bond returns, and thus also represent some new evidence on term premium magnitudes.

2.3 Term Premiums and the LPH

Using the above notation, the term premium for a bond with maturity τ is defined as

$$P_{\tau,t+1} \equiv E_t[r_{t,t+1}(\tau)] - y_{t,t+1}.$$

The liquidity preference hypothesis implies that term premiums increase with τ, i.e.,

$$P_{\tau,t+1} \geq P_{\tau-k_1,t+1} \geq \cdots \geq P_{\tau-k_1-\ldots-k_{i-1},t+1}, \quad k_i > 0. \tag{6}$$

The model in (6) implies that, conditional on all information available to the market at time t, expected returns are larger for longer maturity bonds. Of particular interest, the available information contains the entire term structure and thus the market's expectations about future rates.

The problem with the formulation in (6) is twofold. First, $P_{\tau,t+1}$ is unobservable. That is, pinning down the premium, $P_{\tau,t+1}$, requires an equilibrium model for expected returns, which is not known by the econometrician. Compounding this problem is the fact that the researcher has a smaller information set than investors did at time t. Second, even given a model for $P_{\tau,t+1}$, the restriction in (6) suggests a multiple one-sided test procedure, which is not covered by standard econometric theory. Below, we outline a methodology for evaluating the LPH stated in equation (6).

Consider information available to the researcher at time t. To coincide with the discussion in section 2.1, let us condition on monotonic and non-monotonic yield curves. As noted above, existing theory *suggests* that expected bond returns move with the shape of the term structure.

To generate testable restrictions implied by the LPH from information in the term structure, first define

$$I_t = \begin{cases} 1 & \text{if the term structure is inverted or humped} \\ 0 & \text{if this term structure is upward sloping} \end{cases}$$

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For normalization purposes, we let the instrument z_t be defined as

$$z_t \equiv \frac{I_t}{E[I_t]}.$$

Note that equation (6) implies that the difference between two term premiums (the first having a longer maturity) is non-negative. For example, for 1-period returns from t to $t+1$,

$$P_{\tau,t+1} - P_{\tau-1,t+1} \equiv E_t[r_{t,t+1}(\tau) - r_{t,t+1}(\tau-1)] \geq 0 \qquad \forall \tau \geq 2. \tag{7}$$

Since z_t is a nonnegative random variable and in the information set at time t, equation (7) can be rewritten as

$$E_t[(r_{t,t+1}(\tau) - r_{t,t+1}(\tau-1)) \times z_t] \geq 0 \times z_t = 0 \qquad \forall \tau \geq 2. \tag{8}$$

Rearranging equation (8) and applying the law of iterated expectations,

$$E[(r_{t,t+1}(\tau) - r_{t,t+1}(\tau-1)) \times z_t - \theta] = 0 \qquad \forall \tau \geq 3, \tag{9}$$

where under the null model of the LPH, $\theta \geq 0$.

Equation (9) provides a set of moment conditions that identify the vector θ in terms of observables — the *ex post* return on bonds, $r_{t,t+1}(\tau)$, and the shape of the term structure, z_t. The vector θ has a very clear economic interpretation; it equals the average term premiums, *conditional on non-monotonic term structures.* With respect to the LPH, the restrictions on these conditional means (i.e., $\theta \geq 0$) are testable using results from the recent econometric literature on testing inequality constraints. The method for this particular problem is outlined in the appendix of the paper (though the general method for conditional asset pricing can be found in Boudoukh, Richardson and Smith (1993)). The approach provides a joint test of monotonicity of the *ex ante* term premiums across maturities and does not require a specific model for conditional expected returns. Of particular interest, the power of these tests can be substantially enhanced by focusing on information which is most likely to provide evidence against the LPH null, such as inverted term structures.

The description so far conditions on whether a state occurs or does not occur, and may ignore other relevant information. For example, it may be the case that term premiums are negative only in periods of sharply inverted term structures. Thus, it may be important to put more weight on these periods in the empirical analysis. As an illustration, suppose we want to condition not only on downward sloping term structures, but also on the magnitude of the slope. In this case, we choose I_t^* such that

$$I_t^* = \begin{cases} \max(y_{t,t+k} - y_{t,t+l}), \; k < l \leq j & \text{if the term structure is inverted or humped} \\ 0 & \text{if the term structure is upward sloping} \end{cases}$$

Here we define

$$z_t^* \equiv \frac{I_t^*}{E[I_t^*]}.$$

Using these "informative" instruments, z_t^*, equation (9) still provides a set of moment conditions that identify the vector θ in terms of observables — the *ex post* return on bonds, $r_{t,t+1}(\tau)$, and now both the shape and *magnitude* of the slope of the term structure, z_t^*. The vector of parameters θ has a new economic interpretation; it now equals the weighted average term premium, where the weights correspond to the steepness of the yield curve (adjusted by the probability of such events).

3 Example: Replication of Fama (1986)

Fama (1986) documents time-varying term premiums for bonds of short maturity. He finds that these term premiums are not monotonic in the maturity of the bond; instead, the expected returns can be humped, inverted or decreasing in maturity depending on different stages of the business cycle. Moreover, these stages tend to coincide with humped, inverted or downward sloping forward-rate term structures, which is consistent with the theory in Section 2.1. While this evidence contradicts the LPH, it is important to provide a formal test of this hypothesis.

We collected data from the Fama bond files for one to twelve month bills (in yearly percent) over the period November 1971 to July 1984.[6] Both the period and the maturities are chosen to coincide with Fama (1986). These bills are used for two main purposes: (i) to construct holding period returns on bills of different maturities, and (ii) to separate relevant economic states into either upward sloping forward-rate term structures or inverted/humped term structures. Since stylized facts regarding

[6]We did not include the bill with twelve months to maturity for two reasons. First, there are a substantial number of missing observations at these maturities. Second, the twelve month bill is actually defined in the data to be bills of at least 11 months and ten days. As such, we considered its definition too unreliable for our analysis.

the term premiums on these bills and their behavior with different term structure shapes have been well documented, we refer the interested reader to the relevant literature (e.g., Fama (1986)).

Using the inequality testing methodology of Section 2, we perform two tests of the LPH: (i) a formal test of Fama's (1986) analysis given in his Table 1, which documents average term premiums in different term structure environments, and (ii) a test of Fama's (1986) analysis given in his Table 2 which relates term premiums over particular maturities to forward rates at these same maturities.

With respect to both cases (i) and (ii), Fama (1986) looks at four annualized holding period returns on bills — $y_{t,t+1}$, $r_{t,t+2}(3)$, $r_{t,t+3}(6)$ and $r_{t,t+6}(12)$, where $r_{t,t+j}(\tau)$ means buying a τ-period bill and holding it for j periods. Our estimation uses data from the Fama files for one to eleven month bills (in yearly percent) over the period 1974-1985, so that we replace $r_{t,t+6}(12)$ with $r_{t,t+6}(11)$ (see footnote 6). Under the null of the LPH, the conditional expectations of the annualized $r_{t,t+j}(\tau)$ should be monotonic in maturity. Using the methodology of Section 2.3, we consider the following cross-section of moment conditions implied by the various maturities,

$$E[(r_{t,t+6}(11) - r_{t,t+3}(6))z_t - \theta_1] = 0$$
$$E[(r_{t,t+3}(6) - r_{t,t+2}(3))z_t - \theta_2] = 0$$
$$E[(r_{t,t+2}(3) - y_{t,t+1})z_t - \theta_3] = 0,$$

with the restriction $\theta_i \geq 0 \; \forall i$ under the liquidity preference hypothesis.

To replicate Fama (1986), we choose two different sets of instruments (i.e., case (i) and (ii) above). Fama chooses the instruments in order to capture business cycle effects as described by the term structure of interest rates. For case (i), we consider all periods in which the forward-rate term structure is not monotonic. For case (ii), we choose a different instrument for each moment condition; in particular, this second case investigates term premiums over a particular maturity and holding period which exactly coincide with the non-monotonicity of the forward-rate term structure at the corresponding maturity.[7]

Table 1A reports results for a conditional test of the monotonicity of the term premium based on case (i). With four holding period returns and one instrument,

[7] All tests are repeated using a more informative instrument which conditions on the magnitude of the decline in forward rates (see Section 2.3).

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the system imposes three inequality restrictions. Note that the variance-covariance matrix of the estimators takes into account the correlation across the sample moment estimators and also adjusts for serial correlation due to both time-varying behavior of the premiums and autocorrelation of the term structure shapes through time.

The difference in term premiums, θ_i $(i = 1, \ldots, 3)$, is only negative at the longer-end of the yield curve, that is, $\hat{\theta}_3$ equals -0.614% (annualized). When we condition on the magnitude of the non-monotonicity present, $\hat{\theta}_3$ declines to -1.558%, but the standard error also increases proportionately. Using the 1/0 instrument and the informative instrument, the one-sided joint test statistics for a test of conditional monotonicity of the term premium equal 1.066 (with a P-value of 0.382) and 2.170 (with P-value 0.301), respectively.[8] These tests illustrate that, even though some premiums are individually negative, it is important to perform joint tests in a cross-sectional analysis across maturities.

These results are in contrast to conclusions drawn from Fama's (1986) Table 1; that is, we cannot reject the null hypothesis that the conditional term premiums are monotonic. It is interesting to note that the majority of the unconstrained estimators are negative (using the informative instruments). Similar to Fama's results for humped and inverted forward-rate term structure shapes, the term structure instrument chosen produces many of these negative means. However, these estimated means provide little statistical evidence against the null hypothesis.

The reasons are threefold. First, the test takes into consideration the joint nature of the hypothesis and, in particular, the high cross correlation patterns across the premiums. Second, autocorrelation in the data induced by serial correlation in the forward-rate term structure shapes through time is explicitly accounted for. Third, the test is formal and therefore adjusts for the special distribution of the statistic under the null.

Of course, the sample size of Fama's (1986) study is small, and this may explain why monotonicity is not rejected. However, it is still not appropriate to consider the term premium estimates individually given their joint correlation properties across maturities. At the very least, these results show the different types of conclusions

[8] The P-value has a slightly different interpretation than under tests of equality constraints. Here, we calculate the distribution of the one-sided Wald test statistic for the least favorable value of the null hypothesis (i.e., $\theta \geq 0$ in (9)) and thus of any size test. This can, but does not necessarily, lead to complications in determining the least favorable value of the null if Ω depends on θ in equation (9) (see Boudoukh, Richardson and Smith (1993) and Wolak (1991)). In any event, the test can always be interpreted locally.

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which can be reached by using tests for inequality restrictions. In particular, the commonly held belief that the liquidity preference hypothesis is violated once we take into account available information may be statistically unreliable. The apparent non-monotonicities in the data are consistent with sampling error.

The low significance values suggest that existing stylized facts may need to be reevaluated. One way to do this reevaluation is to partition the information into finer elements, as in our description of case (ii) above. Here, each term premium is associated with its corresponding forward rate, so that we condition on states in which only this particular forward rate is declining. Table 1B provides tests of the restrictions in case (ii). Using the maturity-specific instruments, the annualized difference in expected returns on the four nearby different holding periods are all negative, that is, -0.345%, -0.659% and -0.761%, respectively.

The appropriate multivariate one-sided test statistics for the difference in average premiums are 8.388 (for 1/0 instruments) and 8.426 (for informative ones), which represent P-values of 0.011 and 0.012, respectively. Similar to Fama (1986) (Table 2), and in contrast to our initial tests above, there is strong evidence that term premiums time-vary and that they are not monotonically increasing in maturity in some states of the world. These states are related to periods in which the term structure of forward rates is non-monotonic over the equivalent holding period to the term premium for the corresponding maturity.

The contrast between the results in case (i) and case (ii) emphasizes the relation between conditioning information and the power of tests. Here, negative term premiums are associated with particular non-monotonic forward-rate curves. Thus, while any non-monotonicity of the forward-rate curve implies that $\hat{\theta}_2$ is 0.170%, a non-monotonicity in the forward rate of corresponding maturity implies that $\hat{\theta}_2$ is -0.659%. This is consistent with the theory outlined in Section 2.1, which stresses the relation between a j-period bond's return and the conditional covariance between short-term and j-period stochastic discount factors. This helps explain why the one-sided statistics differ in cases (i) and (ii), and shows how these apparently contradictory results can in fact be consistent with Fama's (1986) story of negative term premiums.

4 Empirical Results

One of the motivations for this paper is to explore how conditioning information affects our interpretation of the LPH. Some previous work has analyzed expected bond returns over both short and long maturity U.S. Treasury bills and notes (see Fama (1984b), Fama and Bliss (1987), Fama and French (1989) and Klemkosky and Pilotte (1992)); however, the majority of work has focused on shorter term maturities. To the extent that both theory and empirical work comment on real rates of interest (either through equilibrium models of business cycles or asset pricing), longer-term maturities are especially important. In this section, we document and test properties of short-horizon holding period returns across the entire maturity spectrum.

4.1 Data Description

From the Fama bond files, we collected data on one-month T-bills and monthly holding period returns on 2-120+ month T-notes and T-bonds. We then formed six equally-weighted portfolios of 2-6 month bills, 7-12 month bills, 12-36 month notes, 36-60 month notes, 60-120 month notes and 120+ notes. The maturities and monthly horizon cover those looked at by Klemkosky and Pilotte (1992).[9] We investigate the sample period January 1972 to December 1994.[10]

Given these returns on bonds of different maturities, we want to test whether the *ex ante* term premiums at different maturities (i.e., holding period return minus one-month T-bill rate) are increasing in maturity. To do this, we need to choose a set of instruments. To coincide with the theory in Section 2.1, and thus enhance the power of the inequality testing methodology, we wish to focus on states of the economy which are the least supportive theoretically of the LPH.

In particular, data on 1-6 month bills and 1-5 year spot rates are available from the Fama bond files over the sample period. We define a non-monotonic yield curve as one in which the yield on one of the six maturities used (i.e., 1-6, 12, 24, 36, 48 and

[9]We choose *a priori* to form portfolios of these bonds over the different maturities to avoid a loss of power of the inequality testing methodology. For example, if we break the portfolio of 36-60 month notes into three bonds (as in Klemkosky and Pilotte) of 36, 48 and 60 months and the returns on these bonds are highly cross-correlated, then it is likely that the benefit of the added information does not offset the increase in the degrees of freedom of the weighted chi-squared test.

[10]Some authors have suggested that, due to changes in monetary policy over the 1979 to 1983 period, the interest rate process may have shifted during this time. Therefore, we also performed the empirical analysis with this period omitted. The results do not change in any substantive way.

60 months) falls relative to the yield on the shorter nearby maturity bond (where the shortest yield, 1-6, is the average yield on 1,2,...,6 month treasuries). That is, there is a violation of the following condition, $y_{1-6} \leq y_{12} \leq y_{24} \leq y_{36} \leq y_{48} \leq y_{60}$, where y_i is the yield on a bond with i-months to maturity.

In addition to conditioning on the shape of the yield curve, we provide more information by conditioning on its magnitude. That is, if the yield curve is non-monotonic in maturity (as defined above), we let I_t^* equal the maximum of the difference between any of the above six maturities (so long as they are negative). The idea here is that extreme declines in the yield curve are associated with large drops in forward rates, and thus corresponding falls in term premiums (Fama (1986)).

One of the problems with conditioning on a non-monotonicity over the entire yield curve is that potentially important information is thrown away. For example, it maybe that term premiums on shorter maturity instruments are declining only for downward sloping term structures at the short-end of the yield curve. To better understand this possibility, we condition on periods in which the short-end of the yield curve (defined by y_{1-6}) lies above the long-end (defined by y_{60}).

4.2 Term Premiums and Interest Rates

Table 2 provides summary statistics on the instruments. In particular, we provide an estimate of the probability of each state, and the state's serial dependence properties (its transition probabilities and corresponding autocorrelation).

Over the sample period, a large fraction of the periods are captured by non-monotonic states. For example, the unconditional probability of the yield curve being non-monotonic is 0.362, while over 50% of these (i.e., 0.185) involve a negative spread between the short-end and long-end of the yield curve. Note, though, that once a state occurs, such as monotonicity or non-monotonicity, the probability of remaining in the state from month to month is very high, e.g., 0.908 for non-monotonicity. Thus, if term premiums time-vary depending on the particular state, these premiums will be autocorrelated. Hence, from a statistical viewpoint, it may be important to adjust the distribution of conditional bond returns for serial dependence.

As an introduction to the interaction between these instruments and the bond returns of different maturities, Figure 1A provides a bar graph of the average term premiums on the different bond portfolios under three possible scenarios: unconditional, upward sloping between the short- and long-end of the spot curve, and downward sloping between the short- and long-ends.

Over the entire sample period, the unconditional monthly expected returns on long-term bonds exceed the one month rate. Moreover, these premiums increase with maturity, that is, from 0.904% (annualized) to 2.187% across the bond portfolios of increasing maturity. These results coincide with the term structure of term premiums documented elsewhere, albeit over slightly different time periods.

The conditional average term premiums, however, tell a different story. In periods of upward sloping yield curves, the premiums are also monotonically increasing, but at much higher rates (from 0.907% to 4.344%). In contrast, in periods of downward sloping yield curves, these premiums are negative for all the bond portfolios except the 2-6 month maturity. Moreover, in these periods, the term structure of average premiums is downward sloping, the exact opposite of the implication of the LPH. While these premiums are not measured very precisely (e.g., see Table 3), and thus must be treated cautiously, the magnitudes of the premiums are economically meaningful. For example, the premium for holding the longest maturity bond portfolio over the short rate is -7.331% on an annualized basis! This means that, in a rational expectations setting, long-term bonds must be a substantial hedge for short-term nominal consumption.

Figure 1B reports the volatility of the term premiums on the different bond portfolios under the same three scenarios. While the term structure of term premiums is increasing (decreasing) with upward (downward) sloping term structures, the term structure of volatilities is always increasing, irrespective of the term structure slope. For example, in periods of downward sloping term structures, the volatilities of bond premiums increase sharply in maturity (from 0.957% to 12.148%), while their ex ante returns drop dramatically (from 0.893% to -7.331%). Standard errors aside, these results suggest that an argument based on a mean/volatility tradeoff cannot explain the conditional distribution of bond returns. In Section 5, we address this issue in more detail.

4.3 Tests of the LPH

In this section, we extend the analysis of Section 4.2 to include formal tests of the LPH. These tests will aid researchers in their evaluation of the significance of the average term premium cross-sectionally, as well as point to particularly important states for determining the term structure of term premiums.

Table 3A provides the difference in average term premiums across the maturities, conditional on non-monotonic yield curves. Also given are the corresponding standard errors of these differences, adjusted for serial dependence and heteroskedasticity, and the one-sided Wald statistics for each instrument and joint across the instruments. Although most of the individual differences in premiums for the non-monotonic instrument are negative, the one-sided test conditional on non-monotonic yield curves has a Wald statistic of only 0.105. This represents a P-value of 0.513. This statistical result again illustrates the need to be cautious when interpreting individual estimates (i.e., a given maturity) in a joint setting (i.e., across maturities).

With respect to non-monotonic yield curves, recall that a declining term structure at the short-end of the yield curve may imply nothing about term premiums on long-term bonds. Thus, following Section 4.1, we break up non-monotonic periods into states in which the short-end of the yield curve lies above its long-end. As described in Section 4.2, conditional on these states the estimated term premiums are decreasing almost everywhere. Table 3B provides the difference in average term premiums of the bond portfolios, corresponding standard errors and the one-sided Wald statistic for this case. The differences in premiums are always negative, with a 1.63% difference between nearby bond portfolios and an average standard error of 1.34. The Wald statistic is 3.081 which represents a P-value of 0.092. Given the conservative nature of inequality constraint testing procedures, it is reasonable to consider this as evidence against the LPH.

However, for those researchers with strong prior beliefs about the LPH, the statistical evidence here is somewhat weaker than that implied by the current literature. It is convenient to criticize these tests as having low power, but this criticism is not well-founded. Inequality restrictions are generally weaker than the more standard equality restrictions; that is, by their nature, inequality constraints provide a higher hurdle. Since tests of the LPH involve inequality restrictions with highly correlated variables, the real question is which inequality constraints-based test has the most power. It is well known, for example, that Bonferroni-type procedures are generally weaker than the statistical techniques advocated here.

As we discussed in Section 2.3, a related power issue is that the 1/0 instruments ignore information about the magnitude of the non-monotonicity of the yield curve.

Tables 3A and 3B provide extensions of the results to include instruments conditioned on the magnitude of the event (see the description in Section 2.3). From Table 3A, the joint one-sided test for the non-monotonic case provides stronger (though perhaps not enough) evidence against the null — the W-statistic equals 2.473, with a P-value 0.146. However, the differences in premiums increase substantially. This implies, for example, that (standard errors aside) the premiums are not only related to the periods in which the term structure is downward sloping, but also vary depending on the degree of the decline. Table 3B shows a similar pattern in terms of the magnitude of the premium differences; however, the increase in the estimation error associated with these weighted premiums actually leads to a reduction in the P-value (from 0.092 to 0.133). While it is difficult to compare P-values across different test statistics, these results do point to the difficulty in measuring the means of short-horizon long-term bond returns.

Nevertheless, while there are few observations on downward sloping term structures, the shape in the term structure of term premiums is still illuminating. Consistent with Fama (1986), Fama and Bliss (1987), and the theory in Section 2.1, it suggests that term premiums vary depending on the current term structure of interest rates. However, because the statistical significance of these results is not overwhelming, the final conclusion about the LPH depends on prior beliefs about the relation between term premiums and conditioning information. Asset pricing theory suggests that the term structure of term premiums may be closely related to the term structure of interest rates, and not the underlying volatility of these premiums. The evidence in this paper supports this theory.

5 Concluding Remarks

The results presented in Sections 3 and 4 provide the first formal (albeit weak) evidence that short-horizon (i.e., monthly) long-term bond returns may be less risky than short-term bond returns under certain circumstances. These circumstances are related to the shape of the yield curve. Interestingly, while expected bond returns are decreasing in maturity, the volatilities of bond returns are still increasing. Of course, while this means that bond volatility is not priced in equilibrium, it does not explain why long-term premiums are so small. That is, what can occur over the next month that makes holding long-term bonds so valuable?

From an economic perspective, these term premium results can be linked to the relation between the term structure and stages of the business cycle. Specifically, it is well documented that at troughs (i.e., the beginning of expansions) the term structure tends to be upward sloping, while at peaks (i.e., the end of expansions) the term structure is flat or inverted (Harvey (1988)). Since the term structure relates directly to the business cycle, the results in this paper suggest that expected returns on bonds may increase with maturity only during particular stages of the business cycle. For example, when the term structure is upward sloping, future covariances between stochastic discount factors tend to be negative. Thus, shocks to the economy next month which cause there to be low discount factors, imply higher discount factors in the future. Since high discount factors mean that longer-term bond returns are low when payoffs are more valuable, the monthly term premium on long-term bonds is high. In contrast, when the term structure is humped or inverted, covariances tend to be positive. Thus, low discount factors today imply even lower discount factors in the future, and long-term bonds provide a good hedge. Thus, expected returns on bonds are decreasing in some maturities. Overall, the empirical results suggest a world in which the autocovariances of the stochastic discount factor, and hence the term structures of term premiums and yields, vary over the course of the business cycle in a systematic way.

The above analysis calls for a structural model of bond pricing that can generate these types of results.[11] As a first pass at developing a structural model of the conditional distribution of the term structure of bond premiums, Figure 2 provides nonparametric kernel estimates of expected excess returns on the six bond portfolios against the yield spread between the long-end and short-end of the yield curve.[12] The figure clearly shows that *ex ante* bond returns on long-term bonds lie below those of

[11]Multifactor models, such as Brennan and Schwartz (1982), Chen and Scott (1989), Longstaff and Schwartz (1989), and Duffie and Kan (1992), all provide frameworks in which the term structure of bond premiums can change sign. For example, Longstaff and Schwartz (1989) make assumptions about the instantaneous short rate and volatility which leads to term premiums being a function of these two variables. However, the implications of these models for maturity effects on premiums and the LPH are not well developed.

[12]Using a multivariate normal kernel, we estimate the distribution of the term premiums conditional on the yield spread. The figure plots the mean of the conditional distribution of these premiums over a continuous range of spreads for which there is enough data. The nonparametric estimation of the mean is fairly intuitive. For a given spread, the mean is essentially a weighted average of the observed *ex post premiums*, where the weights depend on how *close* the given spread is to the observed spreads in the data.

short-term bonds for downward sloping term structures, while the exact opposite is true for upward sloping term structures. Moreover, casual observation suggests the relation may not be linear but instead globally concave in the slope of the term structure. On the face of it, our nonparametric estimates of negative term premiums provide sharp restrictions on theoretical models of the term structure.

Interestingly, the negative term premiums documented here coincide with states associated with negative *ex ante* equity risk premiums (see, for example, Boudoukh, Richardson and Smith (1993) and Boudoukh, Richardson and Whitelaw (1996)). This link suggests that it is not the variability in stocks' cash flows (i.e., dividends) which produces a hedge against short-term consumption, but more the fact that equities can be thought of as a long-term securities, albeit with different cash flow structures than long-term bonds. Negative premiums in general may be due to the hedging benefits of holding long-maturity instruments in certain periods, such as recessions, than anything else per se. These ideas warrant further investigation in the future.

Appendix

The statistical procedure for estimating econometric models in the presence of inequality constraints is described in detail in Wolak (1989). (For a detailed description in the context of conditional asset pricing models, see Boudoukh, Richardson and Smith (1993)). For expositional purposes we review the main steps as they apply to our problem:

- *Step 1*

  Estimate the sample (probability weighted) mean of the difference in term premiums, conditional on non-monotonic term structures:

  $$\hat{\theta} = \frac{1}{T}\sum_{t=1}^{T}[(r_{t,t+1}(\tau) - r_{t,t+1}(\tau-1)) \times z_t] \qquad \forall \tau \geq 3.$$

- *Step 2*

  Estimate the same mean, but now under the LPH restriction that it must be non-negative.[13] Denote this restricted estimator $\hat{\theta}^R$.

- *Step 3*

  A natural test statistic of the restriction, $\theta \geq 0$, is to compare the vector of unrestricted conditional means, $\hat{\theta}$, to the vector of restricted conditional means, $\hat{\theta}^R$. One way to do this is to apply a multivariate one-sided *Wald* statistic, i.e.

  $$W \equiv T(\hat{\theta}^R - \hat{\theta})'\hat{\Omega}^{-1}(\hat{\theta}^R - \hat{\theta}),$$

  where $\hat{\Omega}^{-1}$ is the sample variance-covariance matrix of the conditional means on the bonds.[14] The statistic can then be evaluated at some appropriate level of significance using its asymptotic distribution,

  $$\sum_{k=0}^{N} Pr[\chi_k^2 \geq c]\; w\left(N, N-k, \frac{\hat{\Omega}}{T}\right), \tag{10}$$

[13]This constrained estimation can be performed using standard statistical packages, such as IMSL subroutine DBCONF or DNCONF.

[14]The sample variance-covariance matrix can be constructed to take account of serial correlation in the data, as we have here with month-to-month correlation in term structure shapes. We employ the procedure suggested by Andrews (1991) with quadratic spectral kernel and bandwidth determined by fitting an AR(1) model for each element of $\hat{\theta}$.

where $c \in R^{+}$ is the critical value for a given size, N is the number of restrictions, and the weight $w\left(N, N-k, \frac{\hat{\Omega}}{T}\right)$ is the probability that $\hat{\theta}^{R}$ has exactly $N-k$ positive elements.[15]

[15]The weights can be calculated using a Monte Carlo simulation procedure outlined in Wolak (1989) (statistical routines such as the IMSL subroutine DRNMVN can generate the Monte Carlo data).The critical value itself can then be calculated via IMSL subroutine DUVMIF.

Table 1

A Test of the Liquidity Preference Hypothesis: 1971-1984 (Fama (1986))

Tables 1A and 1B provide tests of whether conditional term premiums are monotonically increasing in maturity, motivated by results in Fama's (1986) Tables 1 and 2. The data are collected from the Fama files for one to eleven month bills (percent annualized) over the period 1971-1984. Each table provides the average difference in term premium measures and corresponding standard errors, conditional on either non-monotonic forward rates (i.e., case (i)), or corresponding declining forward rates over a particular maturity (i.e., case (ii)).

Table 1A

Case (i): Non-Monotonicity of Forward Rates

Diff. in Premiums	Zero/One $\hat{\theta}$	Zero/One (s.e.)	Informative $\hat{\theta}$	Informative (s.e.)
$H3 - H1$	0.568	(0.205)	0.530	(0.360)
$H6 - H3$	0.170	(0.277)	-0.181	(0.356)
$H11 - H6$	-0.614	(0.595)	-1.558	(1.345)
W (pval)	1.066	(0.382)	2.170	(0.301)

Table 1B

Case (ii): Declining Forward Premium (per Maturity)

Diff. in Premiums	Zero/One $\hat{\theta}$	Zero/One (s.e.)	Informative $\hat{\theta}$	Informative (s.e.)
$H3 - H1$	-0.345	(0.206)	-0.675	(0.518)
$H6 - H3$	-0.659	(0.290)	-1.260	(0.602)
$H11 - H6$	-0.761	(0.693)	-2.895	(2.102)
W (pval)	8.388	(0.011)	8.426	(0.012)

Table 2
Statistical Properties of the Instruments

Table 2 provides summary statistics for different states of the world over the sample period — 1/72-12/94. These instruments are defined as follows: $z_{1t} = 1$ when the yield curve is monotonic, $z_{2t} = 1$ when the yield curve is non-monotonic, and $z_{3t} = 1$ when the yield curve is downward sloping. The column acr(z_t) is the autocorrelation of the instrument, q is the probability that $z_{it+1} = 1$ given that $z_{it} = 1$, and p is the probability that $z_{it+1} = 0$ given that $z_{it} = 0$.

Instrument	$Pr(Z_t = 1)$	acr(Z_t)	q	p
Monotonicity	0.634	0.750	0.842	0.908
Non-Monotonicity	0.366	0.750	0.908	0.842
Downward Sloping	0.185	0.783	0.960	0.824

Table 3
Test of the Liquidity Preference Hypothesis Across All Maturities: 1/72-11/94

Tables 3A and 3B use data from the Fama bond files. One month excess holding period returns (term premiums) are calculated for six equally weighted portfolios of 2-6 month bills, 7-12 month bills, 12-36 month notes, 36-60 month notes, 60-120 month notes and 120+ notes. Table 3A provides the difference in the average term premiums, conditional on non-monotonicity, while Table 3B provides this difference, conditional on whether the short-end of the yield curve lies above the long-end. Non-Monotonicity is defined as a state in which there is a violation of the following condition, $y_{1-6} \leq y_{12} \leq y_{24} \leq y_{36} \leq y_{48} \leq y_{60}$, where y_i is the yield on a bond with i-months to maturity. In Table 3A, the corresponding informative instrument is defined as the maximum of the difference between any of the above six maturities (so long as they are negative); in Table 3B, the informative instrument is just the magnitude of the spread between the short- and long-rate. The table also provides one-sided multivariate Wald test statistics of the hypothesis that the difference in these term premiums are positive. The tests are performed for each instrument and across the maturities jointly. The statistic's P-value is calculated using a Monte Carlo simulation.

Table 3A

Instrument and Maturity	Zero/One $\hat{\theta}$	Zero/One (s.e.)	Informative $\hat{\theta}$	Informative (s.e.)
Non-Monotonicity				
$P_{7-12} - P_{2-6}$	-0.100	(0.649)	-0.094	(0.738)
$P_{12-36} - P_{7-12}$	0.187	(1.020)	0.197	(1.467)
$P_{36-60} - P_{12-36}$	-0.115	(0.924)	-1.228	(1.246)
$P_{60-120} - P_{36-60}$	-0.312	(0.883)	-2.237	(1.507)
$P_{120+} - P_{60-120}$	-0.315	(1.342)	-2.622	(1.777)
W(pval)	0.105	(0.513)	2.473	(0.146)

Table 3B

Instrument and Maturity	Zero/One $\hat{\theta}$	Zero/One (s.e.)	Informative $\hat{\theta}$	Informative (s.e.)
Downward Sloping				
$P_{7-12} - P_{2-6}$	-0.972	(0.627)	-0.248	(0.916)
$P_{12-36} - P_{7-12}$	-1.307	(0.949)	-0.189	(1.340)
$P_{36-60} - P_{12-36}$	-1.626	(1.110)	-2.117	(1.651)
$P_{60-120} - P_{36-60}$	-1.545	(1.086)	-2.744	(1.776)
$P_{120+} - P_{60-120}$	-2.774	(1.647)	-3.201	(2.166)
W(pval)	3.081	(0.092)	2.555	(0.133)

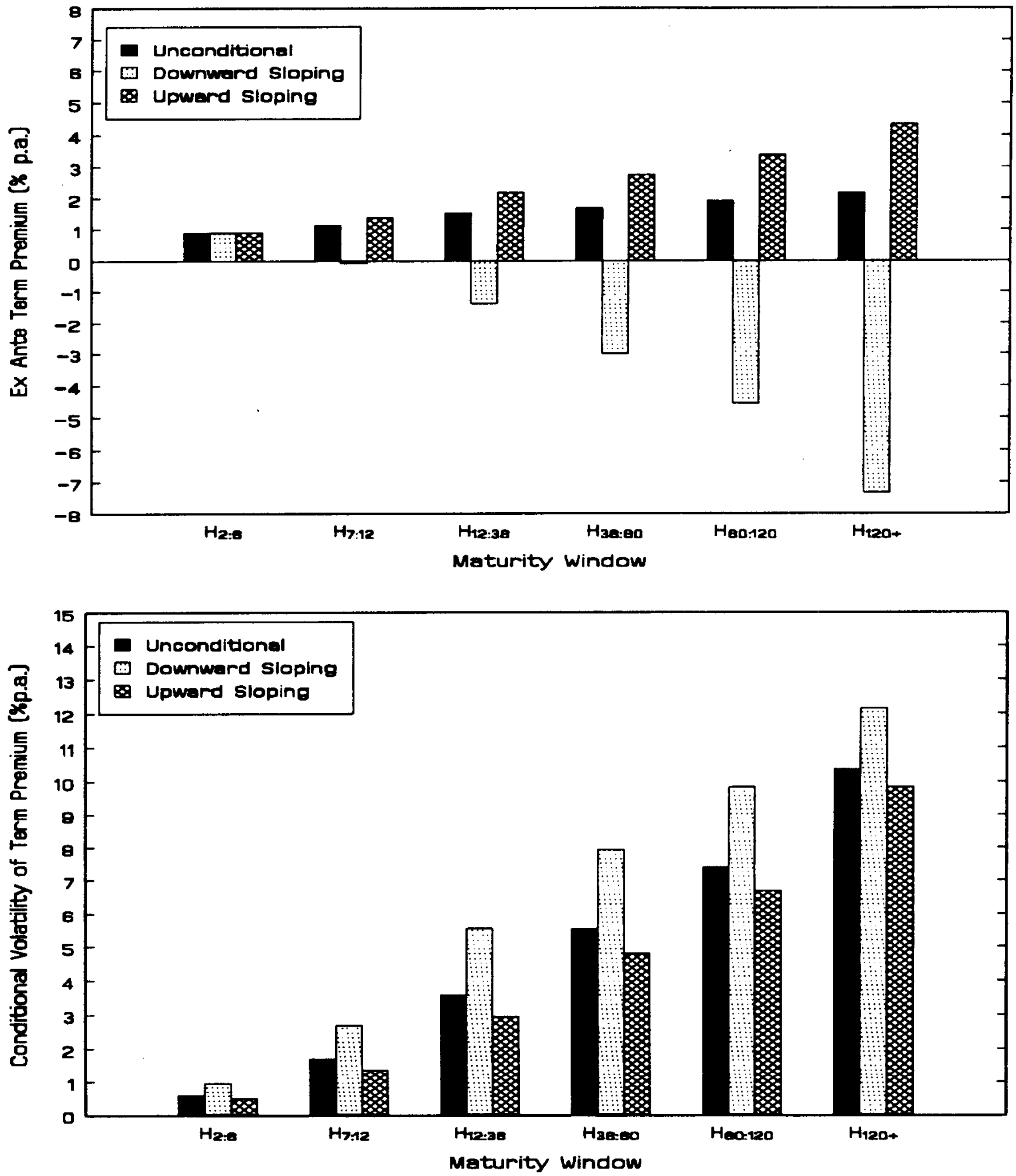

Figure 1: The Expectation and Volatility of the Term Premiums
Figure 1A (top) provides the average term premiums, conditional on one of three possible states: (i) unconditional, (ii) downward sloping yield curves, and (iii) upward sloping yield curves. The term premiums are calculated (in excess of the one-month rate) for six equally weighted portfolios of 2-6 month bills, 7-12 month bills, 12-36 month notes, 36-60 month notes, 60-120 month notes and 120+ notes over the sample period January 1972 to November 1994. Figure 1B (bottom) provides the volatility of the six term premiums conditional on the same information.

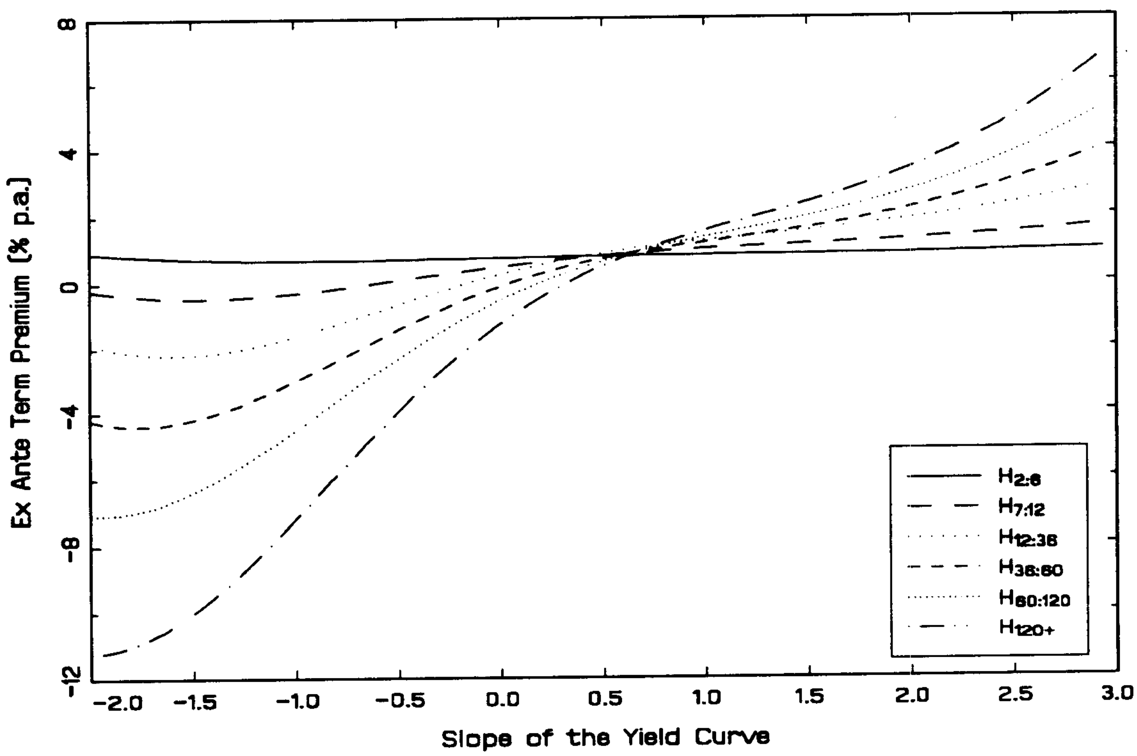

Figure 2: Kernel Estimation of Ex Ante Term Premium

Figure 2 provides a nonparametric estimate of the ex ante term premium against the spread (i.e., slope) between the long- and short-end of the yield curve. The term premiums, denoted H, are calculated (in excess of the one-month rate) for six equally weighted portfolios of 2-6 month bills, 7-12 month bills, 12-36 month notes, 36-60 month notes, 60-120 month notes and 120+ notes over the sample period January 1972 to November 1994.

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