The information on short-maturity forward interest rates and spot rate mean reversion

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[1] I am grateful to Eugene Fama for his remarks on a previous version.

1. INTRODUCTION

In this paper, we use the regression technique in Fama (1984a, 1986) and Fama and Bliss (1987) to examine the forecast power of forward rates and the behavior of expected term premiums on 1 to 6 months to maturity Treasury Bills, and find evidence of some (weak) ability of the forward-spot spread to predict the changes in the spot rate out to 6 months.

The hypothesis advanced in Fama and Bliss (1987) to explain the finding of some power in the long maturity forward rates to forecast far-term changes in interest rates is the presence of a slow mean-reverting tendency in interest rates. Here we apply to the interest rate data considered some formal tests for unit roots, to check for the stationarity or lack thereof in the series.

The analysis, based on the information extracted from the price quotes for 6-month Treasury Bills with 1 to 6 months left to maturity, shows that the expected returns are time-varying, that the observed variability observed in the forward-spot spreads is due mostly to variation in the expected term premiums and not to forecasts of changes in the yields. We are able to document some forecast power of the forward-spot spreads on the changes in the spot rates, although it is not very strong and is concentrated mostly in the shorter maturities. We present evidence indicating that the term structure of expected returns is not upward sloping with maturity, and instead there are some periods of inversion and alterations in the term structure. Finally, a battery of tests applied to the spot rates seem to indicate that there is no mean reversion in the spot rate for these maturities.

2. THEORIES OF THE TERM STRUCTURE

The term structure of interest rates is concerned with the relationships among characteristics (whether the yields, the returns, the interest rates or the prices) on securities of the same type that differ only in their term to maturity. Most of the empirical work has concentrated on money-market securities (like Treasury bills, bankers' acceptances, commercial paper) that have no coupons and produce a return by selling at a discount price over the face value prior to maturity, with some studies focusing on longer term default-free securities like Treasury bonds.

In the absence of uncertainty, forward rates must be equal to future spot rates. But when there is uncertainty about the values of future variables, the analysis is complicated and risk considerations must be introduced.

Term structure models generally advance hypothesis about how the risks of holding-period returns should be measured and what are the relations between these measures of risk and expected returns.

(i) The **liquidity preference** model of Hicks (1946), Kessel (1965) and others assumes risk averse market participants, and states that the risks relevant to investors are the variances of the returns. Expected returns must be larger for longer maturity bonds to induce investors to hold riskier securities, since the variances of holding-period returns (for short holding periods) are larger for longer maturity bonds.

(2) The **expectations hypothesis** places emphasis on the expected values of future spot rates or holding-period returns. In the simplest version, the "pure expectations" hypothesis of Lutz (1940), Meiselman (1962) and others, the pricing of securities is dominated by risk neutral investors, with the implication that expected rates of return for a given holding period are the same for bonds of all maturities.

(3) The **preferred habitat** model of Culbertson (1957) and Modigliani and Sutch (1966) uses market segmentation arguments. While agreeing with the liquidity preference theory in that risk is measured by the variance of the returns, it states that differences in the expected holding-period return of bonds with different maturities respond to differences in the demands and supplies of bonds in different preferred habitats, so that bonds of different maturities might actually be traded in distinctly separate markets. This implies that although expected returns may differ across maturities, they need not be a monotonous function of maturity. This theory, thus, imposes no restriction on the behavior of expected returns.

(4) The **models based on portfolio theory** have in common the perception of the risks of individual securities as measured not by their variances, but by their contributions to the risks of individual portfolios. These include all the well-known modern theories of financial decision: the capital asset pricing model (CAPM) of Sharpe (1964) and Lintner

(1965), the intertemporal capital asset pricing model (ICAPM) developed by Merton (1973) and Long (1974), the consumption-based asset pricing theory of Breeden (1979), and the arbitrage pricing theory (APT) of Ross (1976)[2]. An interesting study in this strand of literature focused on the term structure of interest rates is Cox, Ingersoll and Ross (1985b).

[2] For a concise exposition of modern asset valuation theories, see Constantinides (1989). For a more in-depth view, see Ingersoll (1987) or Huang and Litzenberger (1988). Singleton (1989) provides a survey on the general equilibrium modeling of the term structure.

3. FORWARD RATES, SPOT RATES AND EXPECTED RETURNS

The notation here will follow the conventions in Fama and Bliss (1987) work on the term structure of interest rates[3]. Let us denote:

$P(x:t)$ the price at time t of a bill that promises $1.00 at maturity with x months to maturity.

$r(x:t)$ the yield to maturity on an x month bill. Notice that $r(1:t)$ will be the continuously compounded spot interest rate. It's linked to the price by:[4]

$$P(x:t)=\exp\{-r(x:t)\} \quad [1]$$

$h(x,x-1:t)$ the one-month continuously compounded holding period return. It's linked to the price by:

$$h(x,x-1:t+1)=\ln[P(x-1:t+1)/P(x:t)] \quad [2]$$

and is a random variable for $x \geq 2$ since $P(x-1:t+1)$ is uncertain.

$f(x,x-1:t)$ It is the forward rate (continuously compounded) at time t on a bill with x months to maturity. It is linked to the price by:

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[3] A brief exposition of the notational conventions can be found in Fama (1985).

[4] Notice that the relationship between this continuously compounded rate with the simple rate, $R(x:t)$, that is compounded just once during the period is:

$$r(1:t) = \ln[1+R(1:t)] < R(1:t)$$

$$f(x,x-1:t)=\ln[P(x-1:t)/P(x:t)] \quad [3]$$

All rates will be expressed in continuously compounded rates, following the convention in the financial community and the practice in recent literature.

Notice that the time structure of bill prices at time t must be downward sloping, i.e., the prices of longer maturity bills cannot be greater than the prices of shorter maturity bills. Otherwise, one could, for instance buy one-month bills at t and hold the $1 received at t+1 until t+2, for a cheaper price than buying a two-month bill at t that also gives $1 at t+2.

An important relationship links the time t price of an x years to maturity discount bond (paying $1 at maturity) with the expected future values of the one year returns on the security:

$$P(x:t)=\exp\{-E_t h(x,x-1:t+1)-E_t h(x-1,x-2:t+2)-\ldots-E_t r(1:t+x-1)\} \quad [4]$$

This equation as such is no more than a tautology implied by the definition of the variables. To give it content, we will take this expectations to be the rational forecast of equilibrium expected returns.

Algebraic manipulation of the above equations can be shown[5] to yield a relationship between the forward-spot spread, the expected change in the spot rate and the expected premium of

[5] Substitute [4] into [3], and subtract r(1:t) on both sides.

the return of the one year return of an x year bill over the spot rate, that is the basis for the strategy of empirical research:

$$[f(x,x-1:t)-r(1:t)]=[E_t r(1:t+x-1)-r(1:t)] + +[E_t h(x,1:t+x-1)-r(x-1:t)] \qquad [5]$$

FORWARD-SPOT = EXPECTED SPOT RATE + EXPECTED TERM
SPREAD CHANGE PREMIUM

Fama's methodology consists of analyzing the information in the forward rate about the future spot rate by running the regression:

$$[r(1:t+x-1)-r(1:t)]=a_1+b_1*[f(x,x-1:t)-r(1:t)] + u_1(t+x-1) \qquad [6]$$

i.e., regressing the realized spot rate change on the forward-spot spread. If $\mathbf{b}_1$ in [6] is found to be nonzero, then the forward-spot spread at t has power to forecast the change in the spot x-1 period ahead, and the R^2 adjusted by degrees of freedom will indicate the portion of the variance of the change in the spot rate (x-1) months ahead explained by the corresponding forward spot spread[6]. Results for this regression are presented in table 2 and discussed later.

Notice that if we run the realized premium on the forward-spot spread, since [5] holds also for realized values, the regressions would be complementary: intercepts would sum up to zero, slopes up to one, residuals would cancel out period by period.

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[6] Shiller, Campbell and Schoenholtz (1983), p. 191, interpret the constant in their regression as representing a constant risk premium. This is not of direct applicability in this case.

Strictly speaking, this interpretation of the regression [6] holds exactly true only when the covariance between the two unobserved conditional expectations in [5] is zero; but the general direction of the statements continues to be valid even if this is not the case[7].

It can also be shown[8] that an important relationship links the forward-spot spread with the expected premium of the one-month return on the x-month to maturity bill over the one month spot rate:

$$[f(x,x-1:t)-r(1:t)]=[E_t\ h(x,x-1:t+1)-r(1:t)] + +[E_t r(x-1:t+1)-r(x-1:t)] \quad [7]$$

FORWARD-SPOT = EXPECTED PREMIUM + EXPECTED CHANGE
SPREAD (ONE MONTH RETURN) IN THE YIELD

Equation [7] indicates that the forward-spot spread contains both the expected premium at time t of the one month return on a x-month to maturity bill, and the expected yield change from t to t+1 on a (x-1) months to maturity bills.

Fama's methodology involves running a regression of the term premium on the forward-spot spread to learn about how much information the forward rates contain about the future expected

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[7] See the Appendix or Fama (1984a), pp. 523-26 for details.

[8] Add up the last x-1 expected returns in [4], plug the resulting equation into [3] and subtract r(1:t) from both sides.

$$[h(x,x-1:t+1)-r(1:t)] = a_2+b_2*[f(x,x-1:t)-r(1:t)] + u_2(t+1) \quad [8]$$

By an argument similar to the one above, since [7] holds for realized values as well, then a regression of the one period change in the x-1 yield [r(x-1:t+1)-r(x-1:t)] on the forward-spot spread at t is complementary to [8]: intercepts add up to zero, slopes add up to one, and residuals are the same with different sign. That is, the slope in [8] splits the variation in the forward-spot spread between the expected term premium and the expected yield change terms appearing in [7].

This means that if $\mathbf{b_2}$ in [8] is found to be different from one, the slope coefficient in the regression complementary to [8] is different from zero, and the forward-spot spread would have forecast power for the change in the yield.

Finding that $\mathbf{b_2}$ is not significantly different from 1 would be consistent with the hypothesis that all variation in the forward-spot spread reflects a time-varying expected term premium. Conversely, if $\mathbf{b_2}$ is not reliably different from zero, then the coefficient on the complementary regression would not be significantly different from one and the forward-spot spread could forecast accurately the changes in the yield. Evidence that $\mathbf{b_2}$ is -at the statistically significant levels- between zero and one would mean that the forward-spot spread contains both a time-varying expected term premium and power to forecast changes in the yield. Results for regression [8] appear on table 3 and are discussed in a later section.

Again, this interpretation of [8] only holds exactly true when the covariance between the two unobserved terms in the right hand side of [7] is zero.

We can also gather some knowledge on the term structure (whether flat, upward sloping, etc.) of the forward rates by looking at the time averages of one month forward rates across bills with different number of months to maturity. These are presented in table 1, along with some other information on the correlation properties of the variables considered in regressions [6] and [8].

A richer information than the one offered by the time averages is revealed by the paths of the forward-spot spread, since the structure of forward rates varies through time and looking only at the averages could be misleading. Plots against time are presented as figures 2 to 6, and figures 7 and 8, and discussed below.

Finally, the parallels should be noted between this line of work on the term structure of the interest rates and the investigation of efficiency of the forward foreign exchange markets. Levich (1985), Boothe and Longwoth (1986) and Hodrick (1987) provide surveys of the empirical literature on the efficiency of forward foreign exchange markets. The evidence in this strand of literature points overwhelmingly towards the rejection of the proposition that the forward exchange rate is an unbiased predictor of the future spot rate. This rejection is viewed by some as evidence against the efficiency of the market, while others interpret it as evidence of time-varying risk premiums. The bulk of ongoing research on this area is on developing and testing manageable econometric models models of the risk premium.

4. MEAN REVERSION IN INTEREST RATES

In Fama and Bliss (1987) it is found that for long-maturity bills the forward-spot spread seems to have some forecasting power for the change in the spot rate for long horizons. This is attributed to a mean-reverting tendency in the spot interest rate. It is shown there that if the spot-rate is mean-reverting, then the expected change in the spot rate due to the mean reversion explains about one half of the variance of the actual change in the spot for long forecast horizons[9]. Some further tests are provided there, using the forecasts of an AR1 model.

Here I intend to provide some formal tests for the stationarity hypothesis (i.e., mean-reversion) as opposed to the unit root case, following some work by Dickey & Fuller (1981), Evans & Savin (1984) and Nankervis & Savin (1985).

Dickey & Fuller (1981) set the model:

$$x(t) = \delta + \alpha\ (t-1-(n/2)) + \beta\ x(t-1) + u(t) \qquad \textbf{[9]}$$

where x(1) is fixed and u(t) is white noise (n is sample size), to test the null hypothesis

$$H_0\text{: } (\delta,\ \alpha,\ \beta) = (\delta\ 0\ 1).$$

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[9] To see this, consider a stationary process z(t). Since cov[z(t+t),z(t)] approaches 0 as t grows large, var[z(t+t)-z(t)] approaches 2*var(z) as t grows large. If z(t) is an AR1 with parameter f, it can be shown that $var[E_t\ z(t+t)-z(t)]=var(z)*(f^t-1)$, that approaches var(z) as t grows large.

The likelihood ratio of the restricted and unrestricted models is a monotone transformation of the Dickey-Fuller statistic ϕ_3. It is equivalent to the common F test for linear hypothesis[10]. The empirical distribution of the ϕ_3 statistic is given in the Table VI of Dickey and Fuller (1981)[11]. Results for our data are presented later.

A problem with the Dickey-Fuller test is the low power of the test, that makes it difficult to discriminate against departures of the hypothesis. The empirical power of a 2-sided test size .05 for samples of size 100 is given in Table VII of Dickey and Fuller (1981)[12].

As an alternative, Evans and Savin (1984) present a non-similar test[13] based on the OLS estimator of the autoregression:

$$x(t) = \delta + \beta \; x(t-1) + u(t) \qquad [10]$$

where the empirical distribution depends on the so-called nuisance parameter

$$\psi = [\delta + (\beta-1) \; x(0)] \; / \; \sigma(u) \qquad [11]$$

The empirical distribution is presented in their Table V, and the empirical powers of the size .05 test in their Table VI[14]. Intuitively, this approach decreases the power at the explosive (i.e., β bigger than 1) alternatives, and increases

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[10] See Dickey and Fuller (1981), p. 1059.

[11] Dickey and Fuller (1981), p. 1063.

[12] Dickey and Fuller (1981), p. 1067.

[13] See Evans and Savin (1981, 1984) for details of the derivation.

[14] See Evans and Savin (1984), pp. 1260.

the power at stable (stationary) alternatives.

Results for the data at hand are discussed below.

In a similar vein, Nankervis & Savin (1985) present the empirical quantiles for the t-statistic in the autoregression [10] and the empirical powers of size .05 for values β=1.0, β=.9 for the null hypothesis[15].

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[15] See their Tables 3a and 3b, pp. 154-55.

5. DESCRIPTION OF THE DATA

The spot rates, holding period returns and forward rates are calculated from the data on prices of Treasury Bills, using equations [1], [2] and [3]. The price data are monthly observations and come from the U.S. Government Bond File of the Center for Research in Security Prices (CRSP) of the University of Chicago[16].

They refer to the averages of the end-of-month quotes of bid and asked prices of one-month to six-month Treasury bills. Six months to maturity Treasury bills are issued weekly. The last trading day of each month, the bill with maturity closest to six months is chosen as our six-months to maturity bill. At the end of the following month, this same bill is chosen as the five-month bill, and so on. So at the end of each month we get the data for prices of bills with one to six months to maturity.

It is evident that the bills selected in this fashion will seldom have exactly one to six months to maturity. To correct for this fact, for each end of the month quote date the exact number of days to maturity are used to express the monthly continuously compounded spot rates, forward rates and holding period returns on a per day basis. The per day values of the variables for each month are then multiplied by 30.4, to get a uniform monthly basis.

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[16] The CRSP database is available on a subscription basis.

The sample covers the period from March 1959 (when Treasury bills with six months to maturity begin to be consistently available) to December 1986.

6. EMPIRICAL RESULTS

Some of the descriptive time series characteristics of the variables considered are captured in table 1.

The high autocorrelation in the one month spot rate suggest a possible non-stationarity in the mean and provide further rationale for the particular implementation of [6] as a regression of the spot rate changes on the forward-spot spread, rather than of the spot rate on the forward rate. Changes in the spot rate show almost no autocorrelation.

Looking at the average forward rates in table 1 we notice that the forward rates have different means across maturities and seem, on average, to increase with longer maturities, indicating an upward sloping term structure for the forward rates. It will be noticed, too, that the average value of the forward rates exceeds the average value of the spot by at least .6 % a year; and that the standard deviation of the forward rates is higher than that of the spot. These figures should be taken with caution, though: if the spot and forwards rates do follow a random walk, these sample measures would not be meaningful.

Realized term premiums averages increase with maturity, as do the forward rates and the forward-spot spreads. This should not be constructed, however, as evidence in favor of the liquidity preference hypothesis, since the average returns need not capture accurately the behavior of expected returns[17]. In fact, Fama (1984b) performing test on both average returns (Hotelling T^2 test, Bonferroni multiple comparisons) and regression tests (adjusting realized returns from unanticipated surprises, and estimating expected returns from ex-ante forward rates) presents evidence that expected returns do not increase monotonically with maturity, and tend to peak at maturities of about 8 or 9 months. Further evidence that the expected returns do not monotonically increase with maturity is discussed bellow.

The term premiums display positive first order autocorrelation across maturities considered, suggesting that expected premiums may evolve through time in an autocorrelated way.

Forward-spot spreads for all maturities also show some positive autocorrelation. Given the very low autocorrelation displayed by changes in the spot, that the forward-spot spread variability is lower than that of the term premium, and the relationship implied by equation [5], there seems to be an indication here that it is the variation in expected term premiums what shows up in the properties of the forward-spot spread series. The regression results seem to strengthen this view.

Insofar as the forward-spot spread represents expected term premiums, the varying values of the averages for different maturities is

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[17] This is particularly true of longer-maturity bonds, where the high variability of returns makes inferences more imprecise. This point is raised by Fama (1984b), Fama and Bliss (1987) and McCulloch (1975) among others.

evidence against the pure expectations hypothesis, that predicts the same expected holding-period returns across maturities. Ample evidence against the constancy across maturities of the expected returns is documented also in Hamburger and Platt (1975), Fama (1984a,b, 1986), Fama and Bliss (1987), Shiller, Campbell and Schoenholtz (1983) among others.

The average values increase with maturity, seems to indicate an upward sloping term structure for the expected term premiums in maturities up to six months, that is consistent with the prediction of the liquidity preference hypothesis and with the evidence in Fama (1984a); however, as noted above, existing evidence for longer maturities do not support a monotonically increasing term structure. Furthermore, a closer look at the forward-spot spreads shows that this appearance can be deceiving, as it is argued below.

Plots of the forward-spot spreads for different maturities are presented in figures 2 to 6. Inasmuch as forward-spot spreads are expected term premiums in one-month returns, these figures show the time path of the expected term premiums.

The results of regression [6] appear in table 2 and the results of regression [8], in table 3, as it was mentioned before.

Both these regressions share a estimation problem. The OLS coefficient estimators are consistent, since all is needed is orthogonality of

the disturbance and the regressors[18], and it can be readily verified that
$E\{[f(x,x-1:t)-r(1:t)]*u_1(t+x-1)\}$ =
$= E\{[f(x,x-1:t)-r(1:t)]*u_2(t+1)\} = 0$.

However, the OLS standard errors for the coefficient estimates are not appropriate, since the conditions for their justification, namely, that $E(u_1\ u_1') = \sigma_1^2\ \mathbf{I}$, and $E(u_2\ u_2') = \sigma_2^2\ \mathbf{I}$ do not hold: we have to correct for possible heteroskedasticity and for the autocorrelation (of the moving average type) induced by the sampling of the data at more frequent intervals than the maturity period, that is, the overlap of monthly observations of multimonth rates. The GLS techniques (the standard procedure in other contexts) require the strict econometric exogeneity of the regressors In the notation of the previous note, that $E(u_{t+k} | x_t, x_{t+1}, x_{t-1}, x_{t+2}, x_{t-2}, \ldots) = 0$. If this condition is not satisfied, GLS may render an inconsistent estimator of b. See Hansen and Hodrick (1980), pp. 832-33, or Cumby, Huizinga and Obstfeld (1983), pp. 338-39. That condition cannot be expected to hold either in the present context, and therefore the usual autocorrelation correction is inapplicable here. A correction for the computation of a strongly consistent asymptotic covariance matrix in the presence of stochastic regressors and heteroskedastic errors was proposed first by White (1980); and later, the GMM technique developed by Hansen (1982) allowed to take care of the problem of correlated and hetaroskedastic residuals with

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[18] In general, if we consider a regression of the type $y_{t+k} = x_t\ b + u_{t+k}$, a set of conditions that guarantees consistency of the OLS estimator of b is: (i) y_t and x_t are stationary and ergodic; (ii) $E(x_t'\ x_t)$ is nonsingular; and (iii) $E(x_t'\ u_{t+k})=0$.

greater generality and in a wider range of contexts[19]. In the case under consideration, this procedure yields a consistent estimator of the variance-covariance matrix of the OLS coefficient estimates; and $(\boldsymbol{\beta}_{OLS}-\boldsymbol{\beta})$ (where $\boldsymbol{\beta}_{OLS}$ is the OLS estimator of the coefficients, $(\mathbf{a}_1\ \mathbf{b}_1)$ for [6] and $(\mathbf{a}_2\ \mathbf{b}_2)$ for [8]) converges in distribution to a normally distributed variable with zero mean and covariance the one being computed by the procedure[20]. Moreover, Hansen and Hodrick (1980) show that the procedure is more efficient than the alternative of sampling less frequently to avoid inconsistency of the OLS coefficient covariance estimators due to induced serial correlation.

As for the results, let us consider first equation [8], the term premium regression.

Recall that [8] attempts to measure variation through time in the levels of expected premiums in multi-month bill returns over one-month bill returns. The estimates of $\mathbf{b}_2$ are all positive and lie well beyond the $2*\sigma$ level away from zero; and in all but the first case unity lies within the $2*\sigma$

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[19] For an early application, see Hansen & Hodrick (1980). Hansen (1982) is the best reference for technical details, but a more accessible introduction to estimation problems in rational expectation environments can be found in Cumby, Huizinga and Obstfled (1983).

[20] Justification of the procedure requires the ergodicity and stationarity of the processes followed by the regressand and the regressors; and that the covariance of the one-step ahead forecast errors of the linear projections of the regressand and the regressors on the information set formed by all past values of regressand and regressors, is a matrix of constants that do not depend on the elements of that information set. See Hansen and Hodrick (1980), pp. 833-34 for a statement of the assumptions, and Hansen (1982) for a proof.

band from the estimated coefficient. This means that the sample provides **evidence of the time variance of the expected term premiums.** This result of time-varying expected premiums to maturity is also obtained by Fama (1984a, 1986), Startz (1982) and Shiller, Campbell and Schoenholtz (1983) for T-bills, and Fama and Bliss (1987) for bonds.

Intercepts are positive, although statistically negligible, and slope coefficients are also positive. Since forward-spot spreads are almost always positive as indicated by figures 2 to 6, this indicates a regression forecast of positive term premiums as a rule.

Note also how the regressions seem to absorb the autocorrelations present in the term premiums, by comparing the corresponding rows of table 1 with the residual autocorrelation numbers in table 3.

As mentioned before, except for the first -and only marginally in the second case- the estimates are within a 2 standard deviation boundary from 1. This is to be interpreted (according to the previous comments on complementary regressions) in the sense that the forward-spot spread has little or no forecast power for the one-month change in the yield in the longer horizons; it appears to have some, however, for the shorter horizons. For instance, if we take the .6 coefficient in the first regression at face value, it would imply a .4 coefficient in the regression of the one-month change in the spot interest rate [r(1:t+1)- r(1:t)] on the spread between the implicit forward rate computed from the prices of bills maturing in two months and one month and the spot interest rate [f(2,1:t)-r(1:t)]; similarly, we would have a .32 coefficient in the regression of [r(1:t+2)- r(1:t)] on [f(3,2:t)-r(1:t)].

The values of these estimates resemble (although are somehow higher) the ones obtained in Fama (1984a) with a similar data set for a different time period; the ones obtained in Fama (1986) for a variety of money-market financial instruments (Treasury Bills and several private-issuer money-market securities: bankers' acceptances, negotiable Certificates of Deposit and commercial paper); and (although values obtained here seem somehow lower) those in Fama and Bliss (1987) for 1- to 5-year Treasury Bonds. The evidence in the two first mentioned studies also points to the fact that, for the shorter maturities of up to three months, the forward-spot spreads have some power to forecast the (short-term) changes in the yield. But the bulk of the evidence -and most significatively for longer maturities- indicates that **the variation in the forward-spot spreads is capturing mostly variation in the expected term premiums, and not forecasts of changes in the yields.**

If forward-spot spreads represent then expected term premiums in one-month returns (at least for longer maturities), the time paths of the forward-spot spreads capture the evolution of the expected term premiums through time. Figures 2 to 6 present this evolution. Notice the variability of the spreads, implying that a single central tendency or dispersion sample measure like those appearing in table 1 is not going to provide a good enough description. It is interesting to notice that the spreads (that we interpret as to some extent representing the behavior of expected premiums) are almost always positive for all maturities considered, whereas the longer-term data of Fama and Bliss (1987) show alternate runs of positive and negative values.

Let us turn now to the regressions of the change in the spot on the forward-spot spread (equation [6]). Recall that this regression measures the forecast power of the forward-spot spread (computed from bills with different months to maturity) with respect to the (several months) change in the one-month yield.

The generally positive slope coefficients are accompanied by negative (though statistically insignificant) intercept coefficients. Since forward-spot spreads are generally positive, these negative intercepts allow the regression to mimic the behavior of multi-month changes in the spot, that split more evenly between positive and negative values.

Residual autocorrelations are high. This is due to the overlapping of the monthly observations in the [r(1:t+x-1)- r(1:t)] series.

We find that for most cases all the estimates of $\mathbf{b}_1$ are positive, and except for the last regression, stay on or outside the 2 standard deviation boundary. This is **evidence of some (albeit slight) forecast power of the forward rates on the future spot rates.** Again, this is more marked for shorter horizon, one-month changes, where the forward-spot spread has definitely some forecast power on the change in the interest rate (although this should be qualified: it only explains 13% of its variance at best). Notice by the way how this was already implied by the results in the previous regression, since for this horizon, [6] and the complementary of [8] are the same (a one-month change in the spot interest rate, not a multiperiod yield change as the other complementary regressions to [8] are), and we expected a coefficient of about .4. Again, this result is remarkable, since it is evidence of some -however small (explanatory power ranges from a meek 1.6% to just 13.3%)- forecasting power out to 4 months, with the coefficients for the forward-spot spread variable going from .43 to .21. It has the significance of qualifying our rejection of the expectation hypothesis that in various versions asserts that market efficiency arguments suggest that securities are priced so that implied forward rates equal expected future rates, and the implied forward rates should give some prediction of the future spot rates.

Previous findings in Hamburger and Platt (1975), Fama (1976), and Shiller, Campbell and Schoenholtz (1983) indicated that forward rates, unadjusted for variation in premiums, were poor forecasts of future spot rates[21]. But the results obtained here reverse that position and are in tune with the results in Fama (1984a) for Treasury Bills, in Fama (1986) for commercial paper and certificates of deposit, and in Fama and Bliss (1987) for Treasury Bonds, indicating a weak but nevertheless existing prediction ability of the implied forward rates on the future spot rates for

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[21] Fama (1986) also obtains poor results for T-Bills and banking acceptances.

the one-month changes, prediction ability that debilitates as we stretch maturities up to one or two years and that again strengthens considerably when longer horizons of up to 4 years are taken into account[22].

As we have seen, slope coefficient estimates close to one (specially for longer maturities) in the term premium regressions [8] are an indication that the forward-spot spreads characterize the behavior of expected term premiums on one-month returns. These expected term premiums are generally interpreted as rewards towards risk, and we might have expected that the term structure of (implied) forward rates that captures the term structure of expected premiums be monotonically increasing with maturity, as a superficial look at the average values in table 1 shows.

Surprisingly enough, this is not the case. When the forward-spot spreads for different maturities are plotted together, the expected pattern of an upward sloping term structure does not emerge. Rather, one finds inversions and humps in the term structure, hinting that **the term structure of expected returns is not monotonically increasing with maturity.**

This will be documented here with two examples, establishing in each case the comparison of two different maturity forward-spot spreads for a subsample of the data (figures 7 and 8). This is done to simplify the presentation and avoid clutter in the graphs, but the same points can be made

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[22] Fama and Bliss (1987) report that for Treasury Bonds the forward-spot spread explains 48% of the variance of the 4-year change in the yield.

about the other maturities and subsamples not presented here.

Figure 7 presents the time paths of [f(6,5:t) - r(1:t)] (dotted line) and [f(2,1:t) - r(1:t)] (continuous line) from January 1978 to December 1986; and Figure 8 shows the time paths of [f(6,5:t) - r(1:t)] (dotted line) and [f(4,3:t) - r(1:t)] (continuous line) from January 1970 to December 1977.

Although in general the forward-spot spreads (taken as representing expected term premiums) for longer maturities are higher, one certainly finds periods when this is not the case. Moreover, **the periods of inversions and humps in the term structure seem to be associated with recessionary phases in the business cycle.** This phenomenon was first documented by Fama (1986) and later by Fama and Bliss (1987).

This not only shows that expected returns do not always increase with maturity, and therefore the main prediction of the liquidity theory is not supported by the data. Also, the evidence that the ordering of expected returns across maturities changes with the business cycle implies an stochastic variation in the investment opportunity set that is not accounted for in the Sharpe-Lintner CAPM model.

A richer model within the class of equilibrium models based on financial decision theories is the intertemporal capital asset pricing model. An example of this modelling strategy is Cox, Ingersoll and Ross (1985b), that -building on the general equilibrium models of Lucas (1978) and

Cox, Ingersoll and Ross (1985a)- present a model of multiperiod consumption-investment decisions and asset pricing in continuous time. Empirical implementation of this type of models with a focus on term structure issues presents a challenge for future research in the area. An attempt to tackle this problem can be found in Pearson and Sun (1989). For an example of this line of econometric modelling strategy in the field of international finance, see Cumby (1988).

As for the test for stationarity, notice first that r(1:t) is highly autocorrelated (see table 1), suggesting either a unit root or a high stationary parameter in an AR1. The autocorrelations decay for longer lags. This pattern is the same exhibited by the bond data in Fama and Bliss (1987), leading them to advance the hypothesis of mean reversion.

They argue that the forecast power of the forward-spot spread is largely due to mean-reversion in the spot rates, and try to document this hypothesis by using the forecasts of an AR1 model fitted to the spot rate. In their view, the changes in the spot rate as forecasted by the AR1 should be able, under mean-reversion, of explaining the actual changes in the spot; and they argue that this shows in the statistically significant coefficients for the forecasted spot rate change and in a proportion of the explained variance that grows with the forecast horizon. Then they regress spot rate changes on both the forecasted changes and the forward-spot spreads, to ascertain if all the forecasting power of the forward-spot spread previously obtained was due only to mean reversion.

A similar exercise has been performed here. An AR1 model was estimated with the following results (standard deviation in parenthesis):

$$r(1{:}t) = \underset{(.001)}{.002} + \underset{(.015)}{.959}\, r(1{:}t-1) + e(t) \qquad [12]$$

ssr=.0214

first autocorrelations of residuals: -.069 -.005 -.058 -.094 -.011 .028

and the results of the regressions that use the forecast of r(1:t+x-1) obtained from [12] appear in table 4.

Although significant coefficients are obtained for [(x,x-1:t)-r(1:t)] in the first type of regression, and the explained variance goes from .5% for the one-month change to 50.9% for the five-month change, a comparison of tables 2 and 4 demonstrates that the inclusion of the AR1 forecast do not debilitate the significance of the forward-spot spreads in the spot rate regressions. If anything, the t statistics are now bigger.

This is not the pattern shown in the Fama and Bliss (1987) results. Furthermore, we obtain significant values in the first type of regression shown in table 4, as Fama and Bliss, but the results discussed earlier did not seem to be pointing in the direction of a mean-reversion explanation. Recall that the forward-spot spread showed more forecasting power on the spot rate changes for the shorter horizons, not the longer ones like in Fama and Bliss (1987).

This indicates that something is wrong with the implementation of the mean reversion hypothesis, and that results of the type of the ones appearing in the first part of table 4 do not provide enough justification for it. A more formal test of the hypothesis is called for.

To this effect, tests of unit roots were performed on the spot rate. The results of running a regression of the Dickey-Fuller type [9] are as follows:

$$r(1{:}t) = \underset{(.002)}{.005} + \underset{(.67E-05)}{.000014} * trend(t) + \underset{(.022)}{.925}\, r(1{:}t-1) + e(t)$$

ssr=.0212

The Dickey-Fuller ϕ_3 statistic has here the value:

$$\phi_3 = 5.74$$

and the values given by the tables of the empirical distribution are:

prob(ϕ_3 less than 6.34) = .95 for a 250 sample

prob(ϕ_3 less than 6.30) = .95 for a 500 sample

under the null of a unit root. Therefore, for our sample size (333), we accept the hypothesis of nonstationarity at the 95% level. The corresponding levels for the 90% level are 5.39 for a 250 sample and 5.36 for a 500 sample. Therefore, the hypothesis would be rejected at the 10% significance level.

For the test of Evans and Savin, first note that under the null of a unit root, the expression for the nuisance parameter (see [11] above) reduces to $\psi = \delta/\sigma(u)$.

For our particular configuration of parameters, we accept the unit root hypothesis at 95 level for the 3 type of tests proposed by Evans and Savin. For the equal-tailed test at $\gamma = 0$, the lower and upper critical values they are .80733 and 1.010753[23] for their bigger sample size of 100 under the null of a unit root. The corresponding values for the maximally unbiased test are .855816 and 1.010854; and for the one-tailed test the critical value is .86468. The outcome of a test of a stationary coefficient of .9 as the null hypothesis, is rejection for both the equal-tailed test (critical values are .7215 and .944523) and the maximally unbiased test (critical values of .7355 and .95).

Using the results stated above and the tables provided by Nankervis and Savin we reject the β=.9 hypothesis and accept the unit root hypothesis at the 95% level.

So, we have consistent evidence that **there is no mean reversion in the spot rate** as defined before, i.e., the one-month continuously compounded interest rate derived from prices of Treasury Bills issued with 6 months to maturity.

This seems to be consistent with the results of Nelson and Plosser (1982) for a long sample of bond yields.

[23] The sizes of the test vary if less decimal places for the critical values are used. See Evans and Savin (1984) for details.

8. CONCLUSIONS

We have studied the behavior of the term structure of interest rates for a particular short-maturity instrument, the 6-month U.S. Treasury Bills. The information contained in the bill prices of up to six-months to maturity has been examined and analyzed, using regression techniques pioneered by Fama (1984a, 1986) and Fama and Bliss (1987).

We have found evidence of time-varying expected term premiums. We argued that the detected variation in the forward-spot spreads captures mostly variation in the expected term premiums and not forecasts of changes in the yields. We also found evidence that the forward rates have -specially for the shorter horizons- some power to forecast the future spot rates. We showed how the term structure of expected returns is not monotonically increasing with maturity and rather shows periods of inversions and humps, that seem to be related to the business cycle. Finally, our tests of unit roots showed that there is no evidence of mean reversion in interest rates, contrary to the results in Fama and Bliss (1987) for longer maturities.

TABLE 1: Means, Standard Deviations and Autocorrelations

SPOT & FORWARD RATES

VARIABLE	MEAN	S.D.	AUTOCORRELATIONS 1	2	3	4	5	6	12	24	36	48	60
r(1:t)	5.87	2.91	.96	.92	.89	.87	.85	.83	.74	.54	.39	.27	.20
r(1:t)-r(1:t-1)			-.07	-.01	-.06	-.09	-.01	.03	.01				
f(2,1:t)	6.43	3.15	.97	.94	.92	.91	.87	.86	.78	.55	.39	.28	.20
f(3,2:t)	6.63	3.17	.96	.93	.91	.88	.87	.85	.77	.55	.40	.28	.19
f(4,3:t)	6.61	3.13	.96	.93	.91	.88	.87	.85	.78	.57	.42	.29	.20
f(5,4:t)	6.95	3.16	.94	.92	.90	.87	.86	.85	.78	.58	.44	.30	.20
f(6,5:t)	6.89	3.18	.93	.90	.87	.84	.83	.81	.76	.57	.43	.28	.20

TERM PREMIUMS

VARIABLE	MEAN	S.D.	AUTOCORRELATIONS 1	2	3	4	5	6	12
h(2,1:t+1) - r(1:t)	.07	.10	.41	.15	.16	.11	.15	.12	.04
h(3,2:t+1) - r(1:t)	.09	.17	.24	.04	.04	.02	.06	-.08	-.06
h(4,3:t+1) - r(1:t)	.09	.25	.18	.03	-.04	-.03	.04	-.10	-.13
h(5,4:t+1) - r(1:t)	.13	.33	.22	.04	-.03	-.02	.04	-.07	-.12
h(6,5:t+1) - r(1:t)	.13	.40	.17	.00	-.05	-.07	.03	-.08	-.08

FORWARD-SPOT SPREADS

VARIABLE	MEAN	S.D.	AUTOCORRELATIONS 1	2	3	4	5	6	12
f(2,1:t) - r(1:t)	.067	.084	.22	.30	.29	.22	.15	.41	.32
f(3,2:t) - r(1:t)	.091	.085	.27	.34	.20	.24	.11	.22	.35
f(4,3:t) - r(1:t)	.088	.084	.30	.34	.09	.19	.11	.21	.21
f(5,4:t) - r(1:t)	.129	.113	.37	.34	.31	.07	.11	.15	.18
f(6,5:t) - r(1:t)	.136	.118	.35	.30	.18	.11	.06	.02	.23

NOTATION: r(1:t) is the spot rate observed at time t
f(x,x-1:t) is the forward rate observed at t, for the commitment to buy a 1 month to maturity bill at time t+x-1
h(x,x-1:t) is the holding period return (t to t+1) on a bill x months to maturity
All rates are continuously compounded, adjusted for the different duration of the months throughout the year, and are extracted from means of bid and ask prices for 6 month T-Bills.

SAMPLE: Monthly data, 59-3 to 86-12 (one year less for holding period returns)

TABLE 2: Spot Rate Change Regressions

REGRESSION EQUATION: $[r(1:t+x-1)-r(1:t)]=a_1+b_1[f(x,x-1:t)-r(1:t)]+u_1(t+x-1)$

DEPENDENT VARIABLE	a_1	$s(a_1)$	b_1	$s(b_1)$	RBAR**2	RESID. AUTOCORREL. 1	2	3	4	5	6
r(1:t+1)-r(1:t)	-.002	.006	.43	.09	.133	.11	.01	-.01	-.07	.04	-.01
r(1:t+2)-r(1:t)	-.002	.009	.28	.14	.031	.46	-.06	-.13	-.15	-.07	-.09
r(1:t+3)-r(1:t)	-.002	.001	.33	.16	.028	.63	.21	-.17	-.19	-.17	-.09
r(1:t+4)-r(1:t)	-.002	.001	.21	.09	.016	.68	.36	.05	-.27	-.18	-.09
r(1:t+5)-r(1:t)	-.001	.001	.14	.09	.005	.73	.46	.16	-.07	-.24	-.14

NOTATION: r(1:t) is the spot rate observed at time t
f(x,x-1:t) is the forward rate observed at t, for the commitment to buy a 1 month to maturity bill at time t+x-1
Both refer to 6-month Treasury Bills, and are continuously compounded (with an adjustment for the different duration of the month). They refer to the mean of bid and ask prices.

SAMPLE: 59-3 to 86-12, monthly data.

METHOD: Regressor coefficients are estimated consistently by OLS. OLS estimates of the covariance matrix of the regression coefficients are inconsistent, though; the method applied here to obtain consistent estimates of this matrix was suggested by White(80) and Hansen(82), and it has been computed using the RATS software package.

TABLE 3: Term Premium Regressions

REGRESSION EQUATION: $h(x,x-1:t+1)-r(1:t)=a_2+b_2[f(x,x-1:t)-r(1:t)]+u_2(t+1)$

DEPENDENT VARIABLE	a2	s(a2)	b2	s(b2)	RBAR**2	RESID.AUTOCORREL. 1	2	3	4	5	6
h(2,1:t+1)-r(1:t)	.002	.006	.60	.10	.226	.14	-.00	-.07	-.05	.03	-.03
h(3,2:t+1)-r(1:t)	.002	.001	.68	.16	.113	.14	-.10	-.01	-.08	.00	-.13
h(4,3:t+1)-r(1:t)	.0002	.002	.95	.25	.100	.10	-.05	-.07	-.09	.01	-.16
h(5,4:t+1)-r(1:t)	.001	.002	.90	.20	.093	.13	-.05	-.10	-.09	.03	-.14
h(6,5:t+1)-r(1:t)	.001	.003	.89	.26	.065	.14	-.07	-.07	-.13	.02	-.11

NOTATION: r(1:t) is the spot rate observed at time t
f(x,x-1:t) is the forward rate observed at t, for the commitment to buy a 1 month to maturity bill at time t+x-1
h(x,x-1:t+1) is the holding period return (from t to t+1)
Both refer to 6-month Treasury Bills, and are continuously compounded (with an adjustment for the different duration of the month). They refer to the mean of bid and ask prices.

SAMPLE: 59-3 to 86-12, monthly data (one less year for the holding period return).

METHOD: Regressor coefficients are estimated consistently by OLS. OLS estimates of the covariance matrix of the regression coefficients are inconsistent, though; the method applied here to obtain consistent estimates of this matrix was suggested by White(80) and Hansen(82), and it has been computed using the RATS software package.

TABLE 4: AR1 forecast regressions

REGRESSION EQUATION: $[r(1:t+x-1)-r(1:t)]=\alpha_1+\beta_1[\hat{r}(x,x-1:t)-r(1:t)]+u_1(t+x-1)$

DEPENDENT VARIABLE	α_1	$s(\alpha_1)$	β_1	$s(\beta_1)$	RBAR**2	RESID.AUTOCORREL. 1	2	3	4	5	6
r(1:t+1)-r(1:t)	.6E-4	.4E-3	-.09	.12	.005	-.00	-.04	-.09	-.12	-.04	-.00
r(1:t+2)-r(1:t)	.5E-4	.5E-3	.45	.12	.188	.16	-.32	-.07	-.14	-.01	-.05
r(1:t+3)-r(1:t)	.6E-4	.5E-3	.61	.13	.339	.14	.07	-.46	-.06	-.13	-.01
r(1:t+4)-r(1:t)	.9E-4	.6E-3	.68	.11	.424	.11	.06	.03	-.57	-.05	.00
r(1:t+5)-r(1:t)	.6E-4	.6E-3	.74	.09	.509	.09	.14	-.11	-.09	-.46	-.01

REGRESSION EQUATION: $[r(1:t+x-1)-r(1:t)]=\alpha_1+\beta_1[\hat{r}(1:t+x-1)-r(1:t)] +$
$+\beta_2[f(x,x-1:t)-r(1:t)] + u_1(t+x-1)$

DEP. VARIABLE	α_1	β_1	β_2	RBAR**2	RESID.AUTOCORREL. 1	2	3	4	5	6
r(1:t+1)-r(1:t)	-.002	.030	.441	.131	.09	.01	-.00	-.06	.05	-.00
	(.5E-3)	(.13)	(.08)							
r(1:t+2)-r(1:t)	-.002	.460	.298	.223	.18	-.26	-.04	-.11	.04	-.06
	(.8E-3)	(.13)	(.12)							
r(1:t+3)-r(1:t)	-.003	.632	.436	.391	.21	.05	-.37	-.02	-.07	-.02
	(.8E-3)	(.13)	(.11)							
r(1:t+4)-r(1:t)	-.003	.691	.267	.453	.15	.08	.04	-.52	-.02	.02
	(.9E-3)	(.11)	(.08)							
r(1:t+5)-r(1:t)	-.002	.747	.160	.518	.11	.14	-.08	-.10	-.44	.01
	(.8E-3)	(.08)	(.07)							

(standard errors in parenthesis)

NOTATION: r(1:t) is the spot rate observed at time t
$\hat{r}$ (1:t+x-1) is the forecast from the AR1 model
f(x,x-1:t) is the forward rate observed at t, for the commitment to buy a 1 month to maturity bill at time t+x-1
Both refer to 6-month Treasury Bills, and are continuously compounded (with an adjustment for the different duration of the month). They refer to the mean of bid and ask prices.

SAMPLE: 59-3 to 86-12, monthly data.

METHOD: Regressor coefficients are estimated consistently by OLS. OLS estimates of the covariance matrix of the regression coefficients are inconsistent, though; the method applied here to obtain consistent estimates of this matrix was suggested by White(80) and Hansen(82), and it has been computed using the RATS software package.

Figure 1

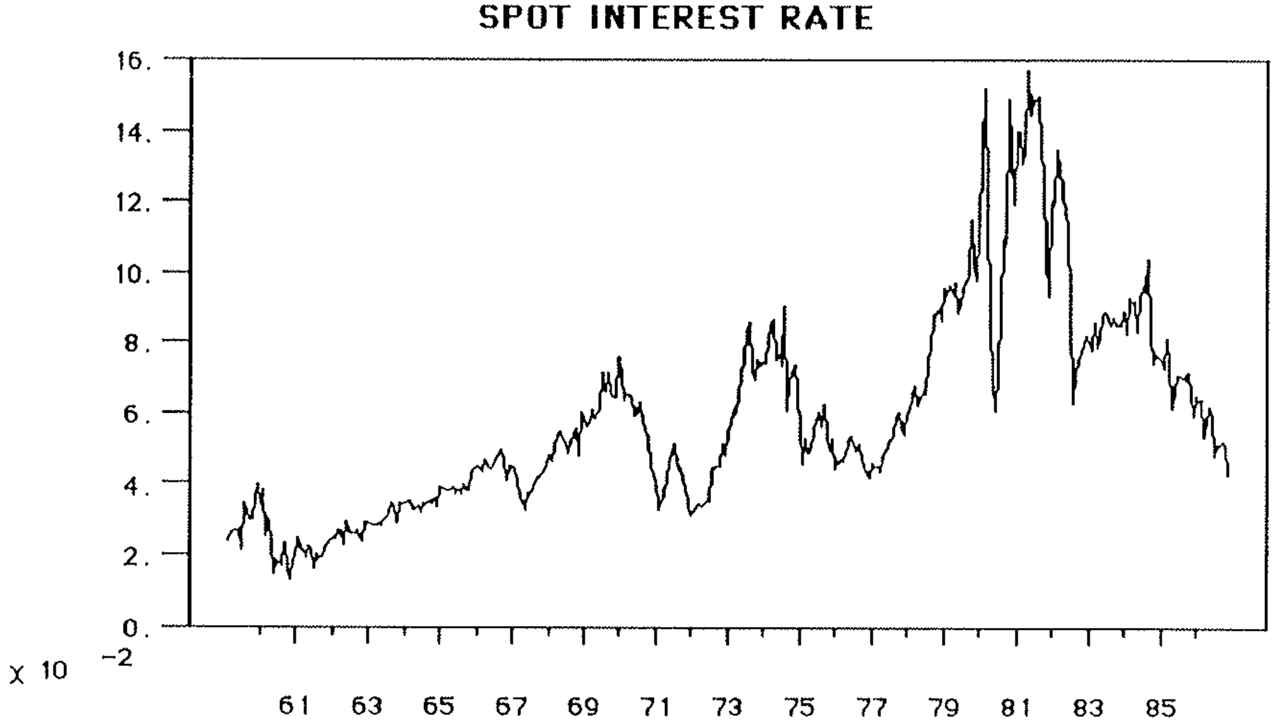


Figure 2

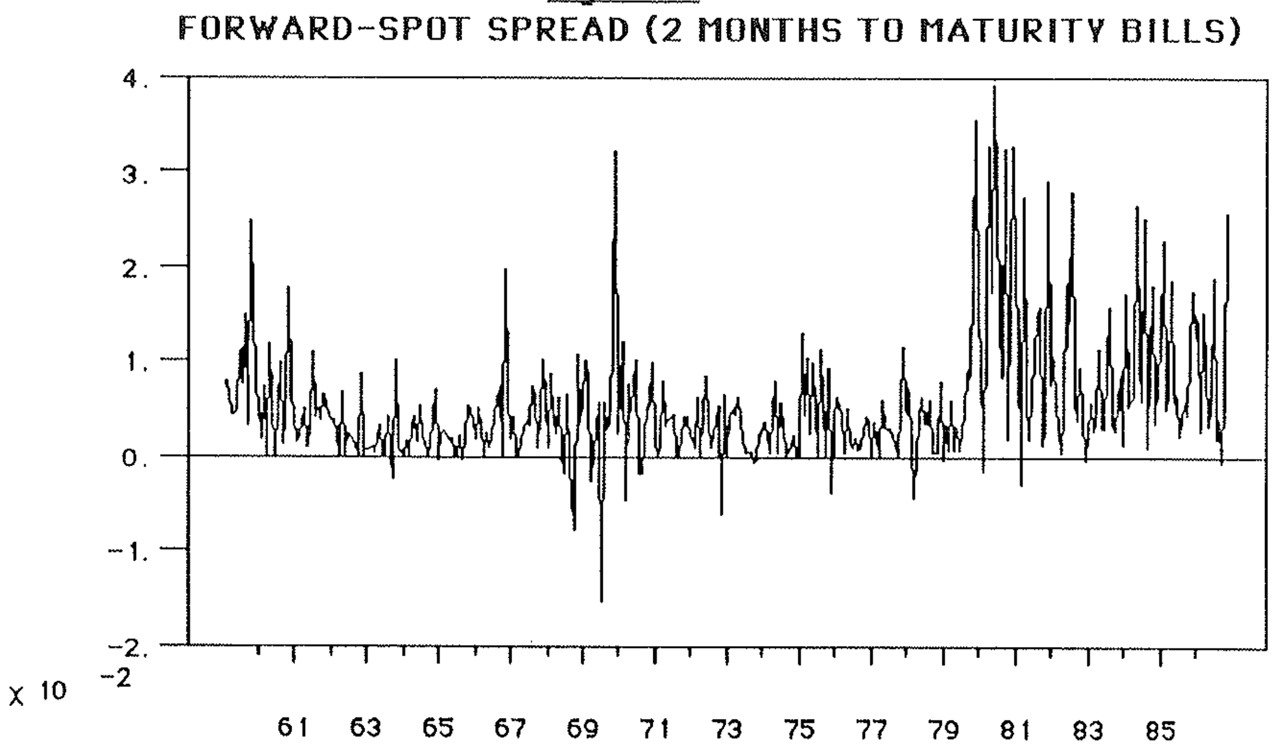


Figure 3

FORWARD-SPOT SPREAD (3 MONTH TO MATURITY BILLS)

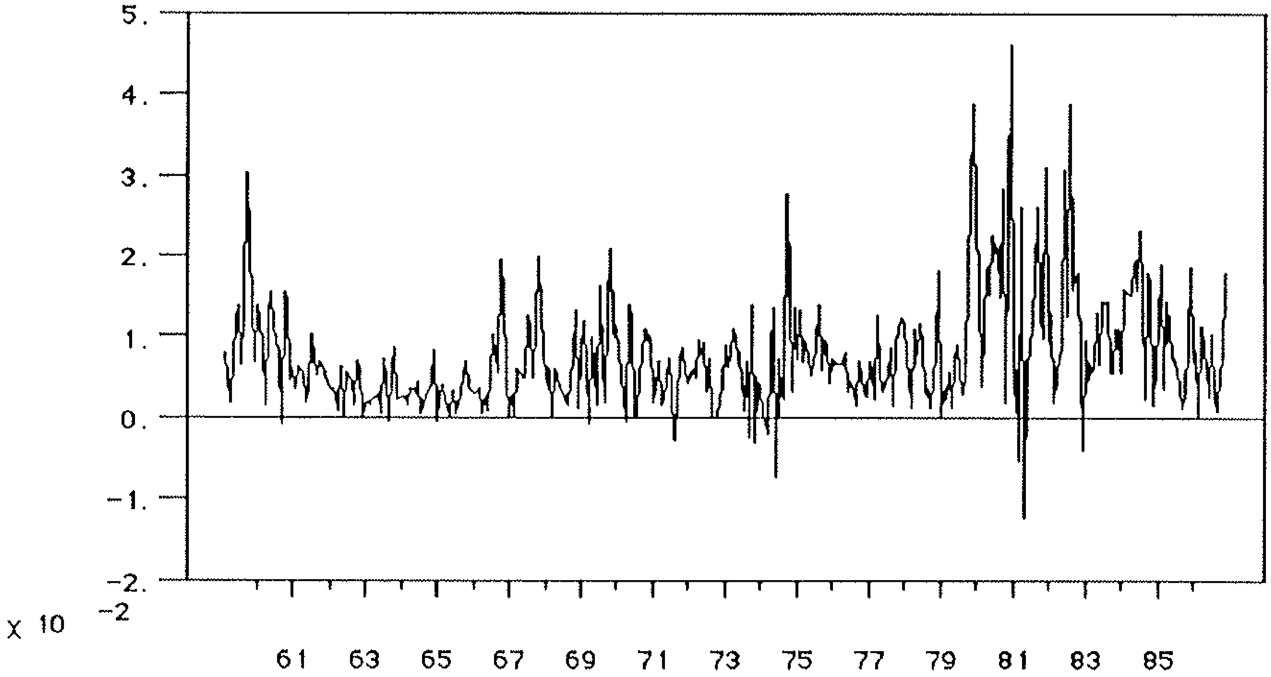


Figure 4

FORWARD-SPOT SPREAD (4 MONTH TO MATURITY BILLS)

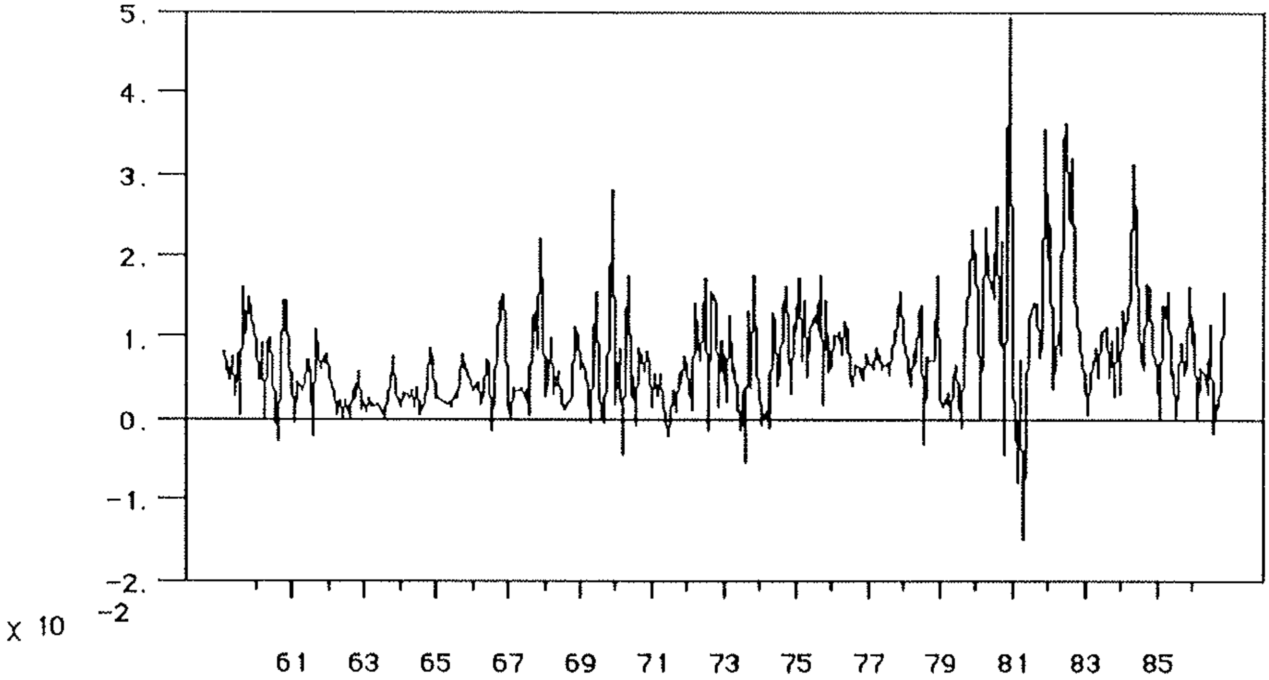


Figure 5

FORWARD-SPOT SPREAD (5 MONTH TO MATURITY BILLS)

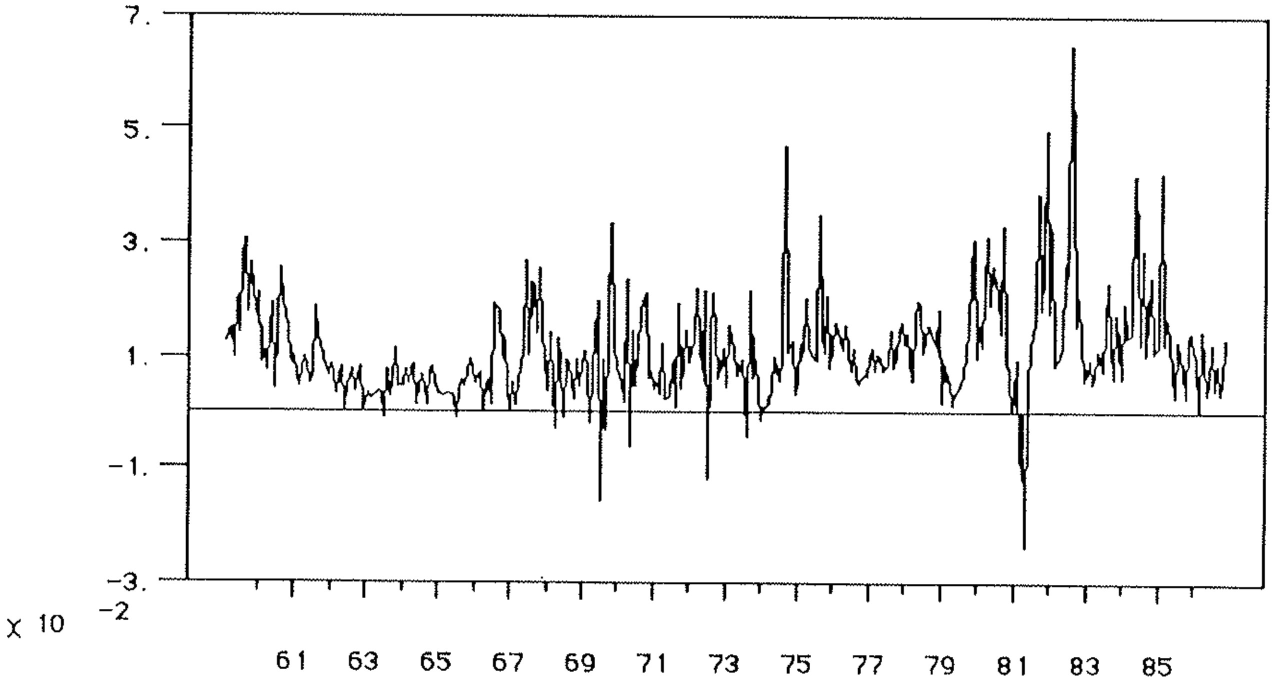


Figure 6

FORWARD-SPOT SPREAD (6 MONTH TO MATURITY BILLS)

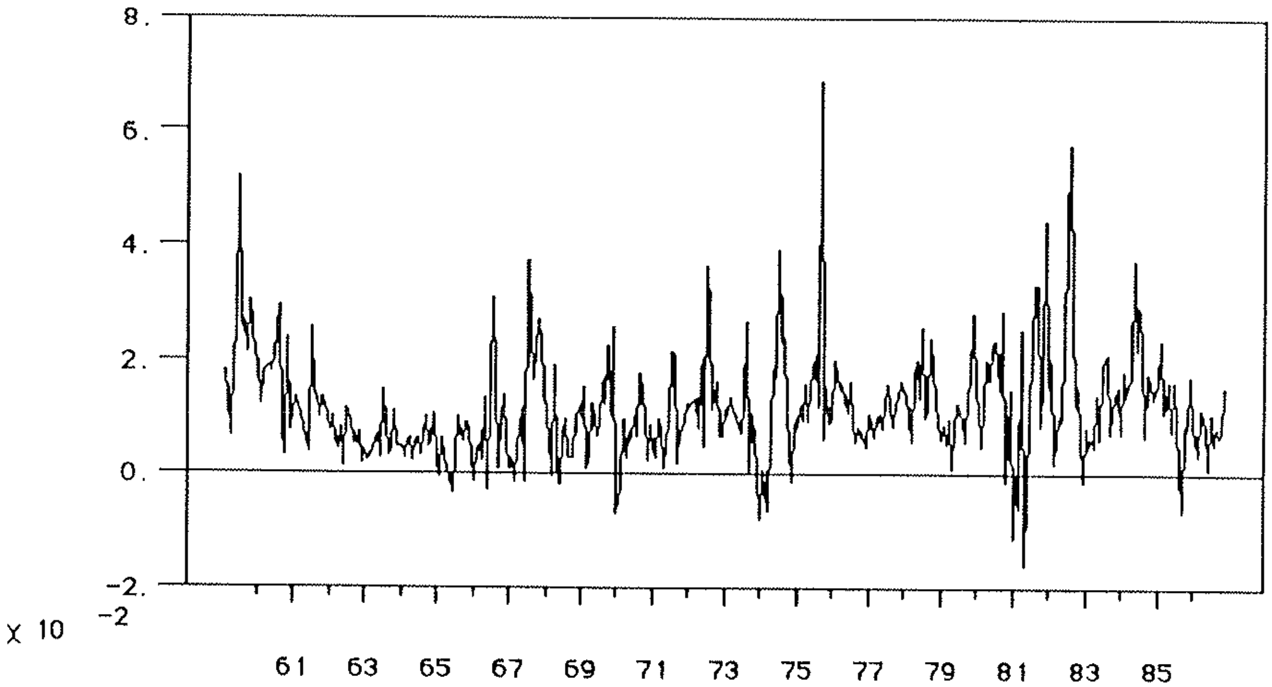
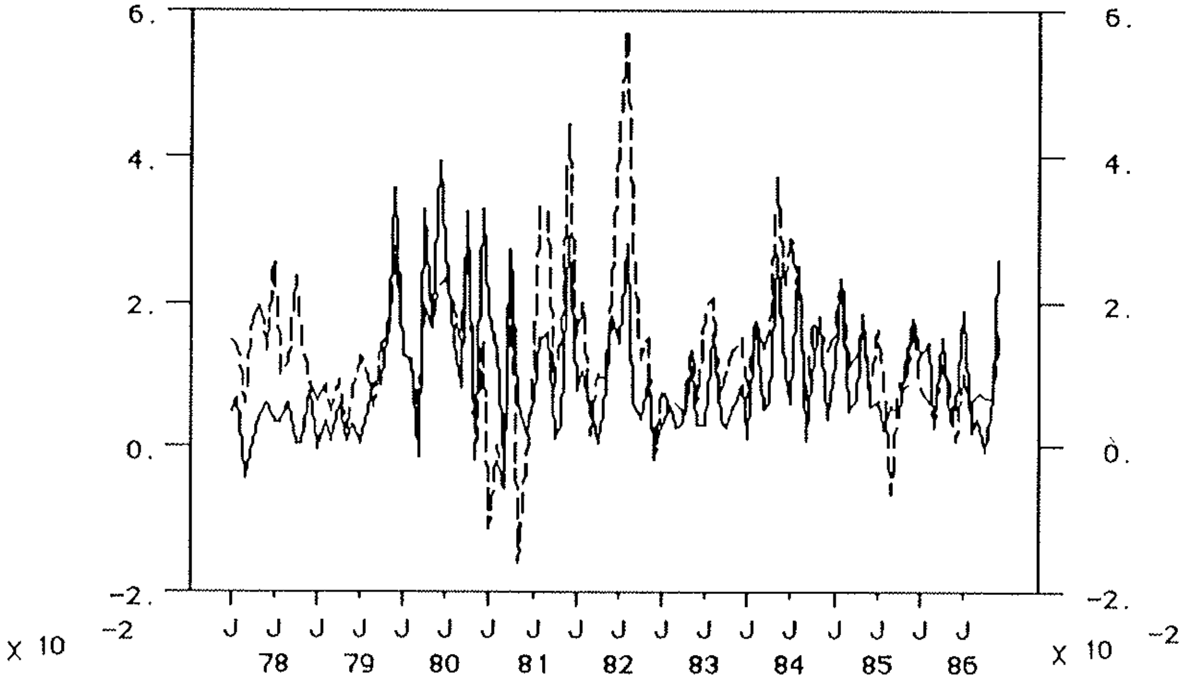
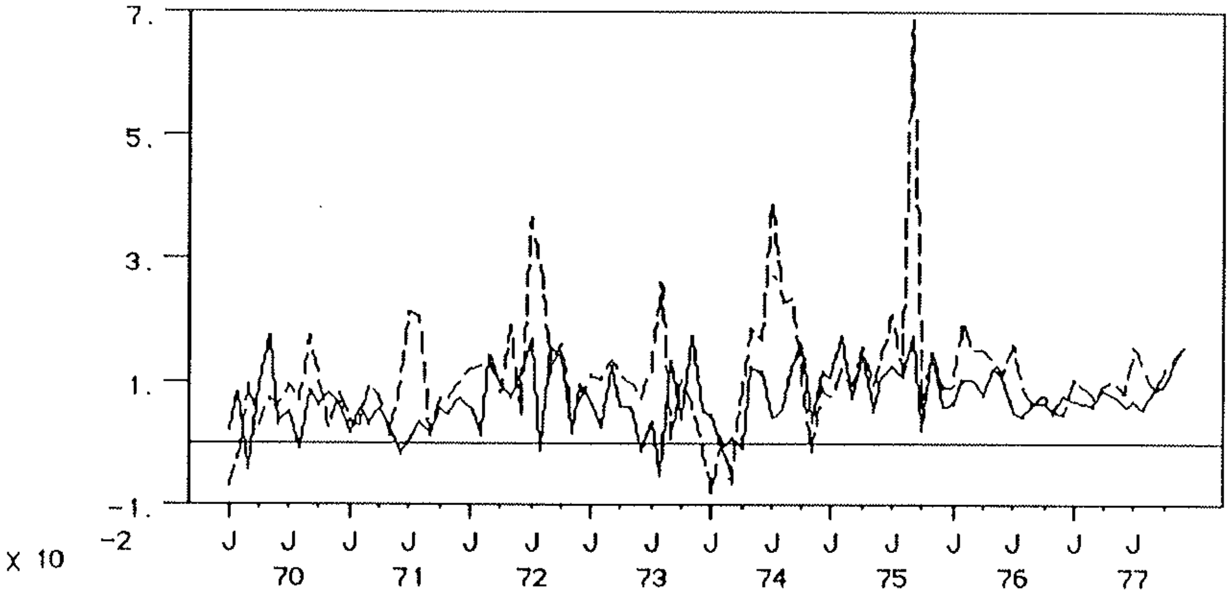


Figure 7
Forward-spot spread term structure



[f(2,1:t)-r(1:t)] ________ [f(6,5:t)-r(1:t)] _ _ _ _ _

Figure 8
Forward-spot spread term structure



f(4,3:t)-r(1:t) ______ f(6,5:t)-r(1,t)_ _ _ _

APPENDIX

Start from [5]:

$$[f(x,x-1:t)-r(1:t)=E_t[r(1:t+x-1)-r(1:t)]+ + E_t[h(x,x-1:t+x-1)-r(x-1:t)]$$

That must hold also for realized values:

$$[f(x,x-1:t)-r(1:t)] = [r(1:t+x-1)-r(1:t)] +[h(x,x-1:t+x-1)-r(x-1:t)]$$

To save notational burden, let us denote these equations compactly as:

$$(f_t-r_t) = E_t(r_{t+x-1}-r_t)+E_t(h_{t+x-1}-r_t) \qquad [A1]$$

$$(f_t-r_t) = (r_{t+x-1}-r_t)+(h_{t+x-1}-r_t) \qquad [A2]$$

Now consider the forecasting equation [6] and its complementary. In compact notation:

$$(r_{t+x-1}-r_t) = a_1+b_1(f_t-r_t)+u_{t+x-1} \qquad [A3]$$

$$(h_{t+x-1}-r_t) = \alpha_1 +\beta_1(f_t-r_t)+v_{t+x-1} \qquad [A4]$$

That [A3] and [A4] are complementary follows directly from [A2]. To see this, notice that:

$$b_1 = \frac{Cov\ [(f_t-r_t),\ (r_{t+x-1}-r_t)]}{Var(f_t-r_t)} =$$

$$= \frac{Var(r_{t+x-1}-r_t) + Cov[h_{t+x-1}-r_t),\ (r_{t+x-1}-r_t)]}{Var(f_t-r_t)}$$

and similary

$$\beta_1 = \frac{Var(h_{t+x-1}-r_t) + Cov[(h_{t+x-1}-r_t),\ (r_{t+x-1}-r_t)]}{Var(f_t-r_t)}$$

so that $b_1+\beta_1 =1$

The interpretation of the slope coefficients given in the text follows from [A1].

$$b_1 = \frac{\mathrm{Cov}\{[E_t(r_{t+x-1}-r_t) + E_t(h_{t+x-1}-r_t)],[r_{t+x-1}-r_t]\}}{\mathrm{Var}(f_t-r_t)} =$$

$$= \frac{\mathrm{Var}\,[E_t(r_{t+x-1}-r_t)] + \mathrm{Cov}[E_t(h_{t+x-1}-r_t),\ E_t(r_{t+x-1}-r_t)]}{\mathrm{Var}(f_t-r_t)}$$

since the error $(r_{t+x-1}-r_t)-E_t[r_{t+x-1}-r_t)$ must be orthogonal to anything in the time t information set.

Similarly

$$\beta_1 = \frac{\mathrm{Var}[E_t(h_{t+x-1}- r_t)] + \mathrm{Cov}[E_t(h_{t+x-1}-r_t),E_t(r_{t+x-1}-r_t)]}{\mathrm{Var}(f_t-r_t)}$$

so the interpretation given in the text is only exactly valid when $\mathrm{Cov}[E_t(h_{t+x-1}-r_t),E_t(r_{t+x-1}-r_t)]=0$. If this term is important relative to the variance terms, it rests significance to the interpretation of the slope coefficients as spliting the variation in the contemporaneously observed (f_t-r_t) into variation of the unobserved expected spot rate change and variation of the unobserved expected termpremium. This would be the case specially if the covariance is negative.

However, the evidence that both b_1 and β_1 are positive and less than one, implies (given the restriction $\beta_1+b_1=1$) that the variances of the expected spot rate change and expected term premium are reliably positive, and the interpretation given remains generally valid (although subject to those caveats).

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