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WHAT DOES THE TERM STRUCTURE TELL US ABOUT FUTURE INFLATION?

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Abstract

This paper examines empirically what the term structure of interest rates tells us about future inflation. The evidence indicates that the information in the term structure about the future path of inflation is quite different at the shortest end of the term structure (maturities six months or less) than it is for maturities of nine to twelve months. For maturities of six months or less, in all the sample periods examined -- February 1964 to December 1986, 1964 to October 1979, November 1979 to October 1982, November 1982 to December 1986 -- the term structure provides almost no information about the future path of inflation. On the other hand at this end of the term structure, the results do indicate that the term structure of nominal interest rates contain a great deal of information about the term structure of _real_ interest rates. This finding is quite important because it suggests that researchers can examine observable data on the shortest end of the nominal term structure to provide them with information about the behavior of the _real_ term structure.

For maturities of nine and twelve months, the term structure does appear to contain information about the future path of inflation in the full sample period and in the sub-periods before October 1982. At these longer maturities, however, there does not appear to be much information in the nominal term structure about the term structure of real interest rates.

The evidence in this paper suggests that some caution should be exercised in using the term structure of interest rates as a guide for assessing inflationary pressures in the economy, as is currently under consideration by the Federal Reserve. Although there is apparently significant information in the term structure about the future path of inflation for maturities greater than six months, there is no information about the future path of inflation that can be obtained from the shorter end of the term structure.

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I. Introduction

Much recent research has focused on the information in the term structure of interest rates. Specifically, this research explores to what extent the relationship between interest rates for different maturities helps to predict the future movement of short-term interest rates. Although Shiller, Campbell and Shoenholtz (1982) and Mankiw and Summers (1984) have questioned the value of the term structure in predicting future short-term interest rates, recent evidence in Fama (1984), Fama and Bliss (1987), Campbell and Shiller (1986), Hardouvelis (1986) and Mishkin (1988) indicates that the term structure is capable of forecasting future interest rates.

The information in the term structure can also be examined from a somewhat different perspective. Fama's (1975) classic study on interest rates as predictors of inflation suggests that movements in interest rates primarily reflect fluctuations in expected inflation (the so-called Fisher effect) rather than changes in real interest rates. Subsequent work supports this conclusion for the postwar United States except for the period of the change in Federal Reserve operating procedures from October 1979 until October 1982.[1] This evidence along with the evidence that the term structure contains information about future interest rate movements suggests that the term structure might also contain information about the future path of inflation.

This paper examines empirically what the term structure of interest rates tells us about future inflation. This is an important topic

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[1]For example, Nelson and Schwert (1977), Mishkin (1981), Fama and Gibbons (1982) and Huizinga and Mishkin (1986).

because inflation is a major concern of policymakers. Indeed, the Federal Reserve is currently considering using the slope of the term structure as a policy guide for assessing inflationary pressures in the economy. The evidence in this paper bears directly on whether such a Fed strategy makes sense. This paper will also provide evidence on the prevalent view that a downward sloping yield curve reflects expectations of a falling rate of inflation, while a steeply upward sloping yield curve indicates expectations of a rising rate of inflation. The behavior of the term structure has thus been a central element in debates about whether Federal Reserve anti-inflation policy has been credible or not. [2]

In addition, this paper provides evidence on whether movements in the term structure of real interest rates (which is not directly observable) are revealed by movements in the term structure of nominal interest rates (which is observable). This is also an important topic because it will tell researchers whether using stylized facts about the observable nominal term structure provides them with information about the behavior of the <u>real</u> term structure, which has an important role in understanding asset pricing and in theories of the business cycle.

II. The Basic Methodology

The empirical analysis in this paper primarily focuses on estimation of two forecasting equations. The first, which will be referred to as the "inflation forecasting equation," is a regression of the level of the m-period inflation rate (π_t^m) on the m-period interest rate (i_t^m):

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[2]Blanchard (1984).

$$\pi_t^m = \alpha_m + \beta_m i_t^m + \eta_t^m \tag{1}$$

Tests of the statistical significance of the β_m coefficient and whether it differs from one provide information about whether the overall level of interest rates helps predict the overall level of future inflation. This forecasting equation, which was first studied by Fama (1975), is estimated in this paper in order to allow a comparison of the results here with those in previous research. Estimates of this equation, however, do not indicate whether the term structure contains information about the future path of inflation because the term structure could have no marginal explanatory power for predicting the path of future inflation and yet β_m estimates for many maturities could be statistically significant.

The main focus of the paper is, therefore, on estimates of a forecasting equation that does tell us whether the term structure helps to predict future inflation. This equation, which will be referred to as the "inflation change equation," is a regression of the change in the future m-period inflation rate from the n-period inflation rate $(\pi_t^m - \pi_t^n)$ on the "slope" of the term structure $(i_t^m - i_t^n)$.

$$\pi_t^m - \pi_t^n = \alpha_{m,n} + \beta_{m,n}[i_t^m - i_t^n] + \eta_t^{m,n} \tag{2}$$

Tests of the statistical significance of the $\beta_{m,n}$ coefficient and whether it differs from one reveal how much information there is in the slope of the term structure about future changes in inflation.

The regression framework outlined above is quite simple, but we need to carefully examine the link between nominal nominal interest rates, real interest rates and expected inflation in order to understand how to clearly interpret the results. According to the Fisher (1930) equation, expected inflation over <u>m</u> periods is equal to the m-period nominal interest rate minus the m-period real interest rate:

$$(3) \qquad E_t\pi_t^m = i_t^m - rr_t^m$$

where,

E_t = expectations at time t,

π_t^m = the inflation rate from time t to t+m,

i_t^m = the m-period nominal interest rate at time t.

rr_t^m = the m-period (ex-ante) real interest rate at time t -- i.e., the ex-ante real return on an m-period bond from t to t+m.

The realized inflation rate over the next <u>m</u> periods can be written as the expected inflation rate plus the forecast error of inflation:

$$(4) \qquad \pi_t^m = E_t\pi_t^m + \varepsilon_t^m$$

where,

ε_t^m = the forecast error of inflation = $\pi_t^m - E_t\pi_t^m$

Substituting in for $E_t\pi_t^m$ from equation (3), we obtain,

$$(5) \qquad \pi_t^m = i_t^m - rr_t^m + \varepsilon_t^m$$

The equation above can be rewritten in the form of the inflation forecasting equation (1) as follows:

$$\pi_t^m = \alpha_m + \beta_m i_t^m + \eta_t^m$$

where,

α_m = minus the mean real rate = $-\overline{rr}^m$

$\beta_m = 1$

$\eta_t^m = \varepsilon_t^m - u_t^m$,

$u_t^m = rr_t^m - \overline{rr}^m$.

If as in Fama (1975), we assume that expectations are rational and the real interest rate is constant, then ordinary least squares (OLS) estimates of the forecasting equation above produce a consistent estimate of $\beta_m = 1$ and the nominal interest rate adjusted by a constant is the optimal predictor of the inflation rate. We see this by first recognizing that rational expectations implies that the forecast error of inflation, ε_t^m, is orthogonal to the right-hand-side regressors because under rational expectations $E_t\varepsilon_t^m = 0$ -- i.e., the forecast error of inflation must be unforecastable conditional on all available information at time t which includes i_t^m. Constancy of the real rate then makes OLS estimates consistent because u_t^m disappears, leaving an error term for the forecasting equation of ε_t^m, which is orthogonal to the right-hand-side regressors under rational expectations and is also a minimum variance error.

If the real rate is not constant, then the nominal interest rate can still contain information about the future inflation rate but it is no longer an optimal predictor because u_t^m is no longer equal to zero. Furthermore, if the nominal interest rate is correlated with the real rate, then u_t^m and hence η_t^m is correlated with i_t^m and the OLS estimate of β_m will have a probability limit different from one. In addition, correlation of η_t^m and i_t^m implies that the OLS estimate of α_m also will not be a consistent estimate of minus the mean real rate $(-\overline{rr}^m)$.

The discussion above leads to the following conclusions about tests on the β_m coefficient. If the null hypothesis $\beta_m = 0$ is rejected statistically, then nominal interest rates contain significant information about the level of future inflation. If the null hypothesis $\beta_m = 1$ is rejected, then the empirical evidence indicates that the real rate of interest is not constant.

We can look at tests of the β_m coefficient in a slightly different way by subtracting i_t^m from both sides of the forecasting equation (4) and then multiplying both sides by -1 to obtain the following ex-post real rate regression equation:

$$eprr_t^m = -\alpha_m + [1 - \beta_m]i_t^m - \eta_t^m \tag{6}$$

where,

$eprr_t^m = i_t^m - \pi_t^m$, the ex-post real interest rate on an m-period bond at time t (i.e., the realized real return from t to t+m.)

Now we see that $\beta_m = 1$ implies that the coefficient $[1 - \beta_m]$ on the nominal interest rate i_t^m equals zero in the ex-post real rate regression. Since the ex-ante real interest rate is just the conditional expectation at time t of the ex-post real interest rate, a rejection of $\beta_m = 1$ (or equivalently, $1 - \beta_m = 0$) in the inflation forecasting equation indicates that nominal interest rates contain significant information about ex-ante real interest rates.[3] Using similar reasoning, a rejection of $\beta_m = 0$ (or equivalently, $1 - \beta_m = 1$) means that the null hypothesis that ex-ante real interest rates move one-for-one with nominal interest rates can be rejected.

A natural way to examine the information in the term structure about future changes in the inflation rate is to subtract equation (5) for the n-period inflation rate from equation (5) for the m-period inflation rate to obtain,

$$\pi_t^m - \pi_t^n = i_t^m - i_t^n - rr_t^m + rr_t^n + \varepsilon_t^m - \varepsilon_t^n \tag{7}$$

[3] For a more detailed demonstration of this point, see Mishkin (1981).

This equation can be rewritten in the form of the inflation change forecasting equation (2) as follows,

$$\pi_t^m - \pi_t^n = \alpha_{m,n} + \beta_{m,n}[i_t^m - i_t^n] + \eta_t^{m,n}$$

where,

$$\alpha_{m,n} = \overline{rr}^n - \overline{rr}^m,$$

$$\beta_{m,n} = 1,$$

$$\eta_t^{m,n} = \varepsilon_t^m - \varepsilon_t^n - u_t^m + u_t^n.$$

Since the above forecasting equation has been derived in a similar fashion to the m-period inflation forecasting equation, the interpretation of the $\beta_{m,n}$ coefficient follows straightforwardly along the lines outlined for β_m. If $\beta_{m,n}$ is significantly different from zero, then the "slope" of the term structure -- i.e, $i_t^m - i_t^n$ -- contains significant information about the change in the future m-period inflation rate from the n-period inflation rate. A significant rejection of $\beta_{m,n} = 0$ (or equivalently, $1 - \beta_{m,n} = 1$) also indicates a rejection of the null hypothesis that the slope of the term structure of real interest rates ($rr_t^m - rr_t^n$) moves one-for-one with the slope of the nominal term structure ($i_t^m - i_t^n$). If $\beta_{m,n}$ is significantly different from one, then we can reject the null hypothesis that the slope of the real term structure ($rr_t^m - rr_t^n$) remains constant over time. A significant rejection of $\beta_{m,n} = 1$ (or equivalently, $1 - \beta_{m,n} = 0$) also indicates that changes in the slope of the nominal term structure contains significant information about changes in the slope of term structure of real interest rates.

III. Some Additional Econometric Issues

Before going on to a discussion of the data and the empirical results, several additional econometric issues that have important consequences for hypothesis testing need to be discussed. One important econometric consideration is that the error term η_t^m will be serially correlated when $\underline{m} > 1$ -- i.e., when the number of periods for the interest rate and the inflation rate are greater than the observation interval, as occurs in the following empirical analysis. In this overlapping data case, the forecast error ε_t^m is not realized until t+m. Thus ε_t^m is likely to be correlated with $\varepsilon_{t+1}^m, \varepsilon_{t+2}^m, \ldots, \varepsilon_{t+m-1}^m$, and ε_t^m will follow an MA(m-1) process. Because of the resulting serial correlation, the standard errors of the OLS parameter estimates will be incorrect. Correct standard errors are generated using the method outlined by Hansen and Hodrick (1980), with a modification due to White (1980) that allows for conditional heteroscedasticity and a modification suggested by Newey and West (1985) that insures the variance-covariance matrix is positive-definite by downweighting the off-diagonal elements.[4]

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[4]The formula for the variance-covariance matrix of the parameter estimates, V, is

$$V = (X'X)^{-1}X'\Omega X(X'X)^{-1}$$

where,

X = the matrix of explanatory variables, which is T×k (T = the number of observations, k = the number of explanatory variables),
Ω = the variance-covariance matrix of the residuals $E(\eta\eta')$,

and the (i,j)th element of the estimated Ω is defined as

Because contemporaneous errors in forecasting inflation for different horizons may be highly correlated, seemingly unrelated regression (SUR) estimates of a system of equations with different horizons may produce substantial gains in efficiency.[5] The SUR standard error estimates will again be incorrect because of the serial correlation of the error terms. The Hansen-Hodrick, Newey-West estimate of the variance-covariance matrix allowing for conditional heteroscedasticity can be generalized to apply to a seemingly unrelated regression system of *g* equations as follows.[6]

$$\omega(i,j) = [1 - p/(q+1)]\hat{\eta}_i^m \hat{\eta}_j^m \qquad \text{for } p \leq q$$

$$= 0 \qquad \text{otherwise}$$

where $p = |i - j|$ and q = the order of the MA process, $m - 1$.

Note that in constructing the corrected standard errors, u_t^m is assumed to have a MA process of order less than m-1. If u_t^m follows a higher order MA process then the composite error term η_t^m will also follow a higher MA process. In order to allow for more gradual down-weighting of the off-diagonal elements of the the Ω matrix and the possibility that u_t^m has a MA process of order greater than m-1, I also constructed Newey-West estimates of the variance-covariance matrix where q is replaced by 2q in the equation above. The estimated standard errors of the coefficients typically were quite close to those reported in the tables and none of the conclusions were affected. The results thus do not appear to be sensitive to allowance for a higher order MA process for the error terms or to use of a more gradual down-weighting scheme.

[5]The same conditions for the consistency of OLS estimates of $\beta_{m,n} = 1$ -- the constancy of the real rate differentials, $rr_t^m - rr_t^n$ -- also apply for consistency of the SUR estimates. Constancy of the real rate differentials implies that the error terms $\eta_t^{m,n}$ just equals $\varepsilon_t^m - \varepsilon_t^n$. Since under rational expectations these forecast errors are uncorrelated with all information available at time t, which includes i_t^m for all m, constancy of the real rate differentials implies the condition for consistency of the SUR estimates, that all of the explanatory variables in the equations are orthogonal to all the error terms.

[6]The variance-covariance matrix, V, is derived as follows. The SUR estimation method assumes that the variance-covariance matrix of the residuals is $\Sigma \otimes I_T$. Using the Choleski decomposition $\Sigma^{-1} = P'P$, we get the GLS (i.e., the SUR) estimates by premultiplying the system by $P \otimes I_T$ and then proceed with OLS estimation. Allowing for conditional

$$V = (X'(\Sigma^{-1}\otimes I_T)X)^{-1}X'(\Sigma^{-1}\otimes I_T)\Omega(\Sigma^{-1}\otimes I_T)X(X'(\Sigma^{-1}\otimes I_T)X)^{-1} \tag{8}$$

where,

V = variance-covariance matrix of estimated coefficients

$$X = \begin{bmatrix} X_1 & 0 & . & . & . & 0 \\ 0 & X_2 & 0 & . & . & 0 \\ . & & & & & \\ . & & & & & \\ 0 & . & . & . & 0 & X_g \end{bmatrix}$$

Σ = variance-covariance matrix of the contemporaneous residuals from the g equations,

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heteroscedasticity, the Hansen-Hodrick variance-covariance matrix of the parameter estimates of the transformed system is then,

$$V = (\tilde{X}'\tilde{X})^{-1}\tilde{X}'\widetilde{\eta\eta}'\tilde{X}(\tilde{X}'\tilde{X})^{-1}$$

where,

$$\tilde{X} = (P\otimes I_T)X,$$

$$\tilde{\eta} = (P\otimes I_T)\eta.$$

Writing the variance-covariance matrix out results in

$$V = (X'(P'\otimes_T)(P\otimes I_T)X)^{-1}X'(P'\otimes I_T)(P\otimes I_T)\eta\eta'$$
$$(P'\otimes I_T)(P\otimes I_T)X(X'(P'\otimes I_T)(P\otimes I_T)X)^{-1}$$

Making use of the fact that $P'P = \Sigma^{-1}$ and $\eta\eta' = \Omega$ we have

$$V = (X'(\Sigma^{-1}\otimes I_T)X)^{-1}X'(\Sigma^{-1}\otimes I_T)\Omega(\Sigma^{-1}\otimes I_T)X(X'(\Sigma^{-1}\otimes I_T)X)^{-1}$$

as in the text.

I_T = T×T identity matrix,

Ω = the Newey-West modification of the $\eta\eta'$ matrix,[7]

Now that we have completed our discussion of the econometric details, we can go on to discuss the data used in the empirical analysis.

IV. The Data

The empirical analysis makes use of monthly data on inflation rates and one to twelve-month U.S. Treasury bills for the period February 1964 to December 1986. The sample starts with February 1964 because this is the first date that data on all the Treasury bills became available (twelve-month Treasury bills were not issued until late 1963). Because six month Treasury bills were first issued in late 1958, tests using one to six-month Treasury bills can be conducted with sample periods starting

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[7]The (i,j)th element of the $\Omega_{k,\ell}$ block of Ω (where k and ℓ reference equations) is defined as follows,

$$\begin{aligned} \omega_{k,1}(i,j) &= [1 - p/(q + 1)]\hat{\eta}_i^k\hat{\eta}_j^\ell \qquad && \text{for } p \leq q \\ &= 0 && \text{otherwise} \end{aligned}$$

where p = |i - j| and q = the highest order of the MA process for the error term of any equation in the system. The order of the MA process is assumed to be the same for all the equations in the system because otherwise the Newey-West method used here does not insure that the variance-covariance matrix is positive definite.

Note that the presence of the Ω matrix in the formula for the variance-covariance matrix in (8) takes account of serial correlation of the error terms both within an equation and across equations as well as for conditional heteroscedasticity in these covariances. Thus even though SUR estimation only takes account of contemporaneous correlation of error terms across equations, the variance-covariance matrix above is corrected for serial correlation both within and across equations as well as for conditional heteroscedasticity.

in January 1959. The results are very similar to those for sample periods starting with February 1964 and are reported in the appendix. End of month T-bill data were obtained from the Center for Research in Security Prices (CRSP) at the University of Chicago. The one-month bill was defined to have a maturity of 30.4 days, the three-month bill 91.25 days, on up to the twelve month bill with a maturity of 365 days. For each defined maturity the interest rate was interpolated from the two bills that were closest to the defined maturity. In effect, this means that the slope of the term structure is assumed to be constant between these two bills.[8] The interest rates are expressed on a continuously compounded basis at an annual rate in percent as are the inflation rates. The inflation data is calculated from a CPI series which appropriately treats housing costs on a rental-equivalence basis throughout the sample period. For more details on this series see Huizinga and Mishkin (1984, 1986).

The timing of the variables is as follows. A January interest rate observation uses the end of December bill rate data. A January observation for a one-month inflation rate is calculated from the December and January CPI data; a three-month inflation rate from the December and March CPI data; and so on.[9]

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[8]Fama (1984) instead chooses a bill that has a maturity closest to six months and then keeps on taking the interest rate from this same bill every month as its maturity shortens in order to get interest rates on one to six-month bills. In effect, Fama is assuming that the slope of the term structure is flat around the chosen bill. The differences between these two procedures is very slight and makes no appreciable difference to the results.

[9]The appropriate dating for the CPI is a particular month is not clear since price quotations on the component items of the index are collected at different times during the month. As a result, there is

V. The Empirical Results

Table 1 contains the estimates of the inflation forecasting equations for horizons of one, three, six, nine and twelve months. Panel A contains the results for the full sample period, February 1964 to December 1986, while Panels B, C and D contain the results for three sub-periods, February 1964 to October 1979, November 1979 to October 1982, and November 1982 to December 1986. The sample has been split into these three sub-periods because results in Clarida and Friedman (1984), Huizinga and Mishkin (1986) and Roley (1986) indicate that the relationship of nominal interest rates and inflation shifted with the monetary regime changes of October 1979 and October 1982.

For the full sample period, β_m is significantly different from zero for all time horizons, indicating that one to twelve month Treasury bill rates do contain a significant amount of information about future inflation. This finding is especially strong for the pre-October 1979 sample period (Panel B) where the t-statistics on the β_m coefficients range from 9.76 to 11.85. However, after October 1979, the one to twelve month nominal interest rates contain much less information about future inflation. In the October 1979 to October 1982 period of the Fed's nonborrowed reserves target operating procedure, none of the β_m coefficients are

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some misalignment of the inflation data and the interest rate data which is collected at the end of the month. In order to see if this misalignment could have an appreciable affect on the results, I also estimated the regressions in this paper lagging the interest rate data one period (i.e., for the January observation I used the end of November bill rate). The results with the lagged interest rate data are very similar to those found in the text and none of the conclusions of the paper changes.

Table 1

Estimates of Inflation Equations

$$\pi_t^m = \alpha_m + \beta_m i_t^m + \eta_t^m$$

m (months)	α_m	β_m	R^2	SE	t-test of $\beta_m = 0$	t-test of $1-\beta_m = 0$
Panel A: February 1964 - December 1986 Sample Period						
1	1.2232 (0.4482)	0.5966 (0.0714)	0.207	3.200	8.36**	5.65**
3	1.4486 (0.5659)	0.5296 (0.0845)	0.248	2.669	6.27**	5.57**
6	1.7363 (0.7573)	0.4730 (0.1129)	0.237	2.465	4.19**	4.67**
9	2.1852 (0.9062)	0.4075 (0.1322)	0.189	2.428	3.08**	4.48**
12	2.5011 (1.0302)	0.3647 (0.1485)	0.156	2.407	2.46**	4.28**
Panel B: February 1964 - October 1979 Sample Period						
1	-2.2721 (0.6330)	1.3746 (0.1216)	0.439	2.590	11.30**	-3.08**
3	-2.2135 (0.6887)	1.2941 (0.1187)	0.549	1.976	10.90**	-2.48**
6	-2.6634 (0.6739)	1.3236 (0.1117)	0.649	1.654	11.85**	-2.90**
9	-2.6410 (0.7421)	1.3070 (0.1266)	0.657	1.595	10.32**	-2.42**
12	-2.6099 (0.7906)	1.3009 (0.1332)	0.648	1.589	9.76**	-2.26**

Panel C: November 1979 - October 1982 Sample Period

1	7.1035 (1.8326)	0.0890 (0.1552)	0.005	3.498	0.57	5.87**
3	5.0256 (3.4120)	0.2353 (0.2526)	0.036	2.937	0.93	3.03**
6	7.0521 (4.1291)	0.0356 (0.2887)	0.001	2.674	0.12	3.34**
9	10.7631 (3.3672)	-0.2785 (0.2129)	0.055	2.382	-1.31	6.00**
12	10.6754 (2.7065)	-0.2918 (0.1567)	0.064	2.239	-1.86	8.25**

Panel D: November 1982 - December 1986 Sample Period

1	-1.7349 (1.9260)	0.6341 (0.2362)	0.112	2.474	2.68**	1.55
3	-0.1532 (1.6798)	0.4054 (0.1910)	0.099	1.806	2.12*	3.11**
6	1.2817 (1.7622)	0.2351 (0.1867)	0.077	1.301	1.26	4.10**
9	1.8158 (1.7917)	0.1706 (0.1803)	0.061	1.109	0.95	4.60**
12	2.4821 (1.5415)	0.0927 (0.1518)	0.024	1.017	0.61	5.98**

Notes for all tables

Standard errors of coefficients in parentheses.
SE = standard error of the regression.
* = significant at the 5% level.
** = significant at the 1% level.

significantly different from zero and for about half of the time horizons they are even negative. Although there is a positive relationship between inflation and nominal interest rates at all time horizons in the post-October 1982 period, β_m is only significant at two time horizons, one month and three months.

The t-test of $1 - \beta_m = 0$ for the full sample period indicates that the constancy of the real interest rate can be rejected and that nominal interest rates do contain information about real interest rates. However, $1 - \beta_m$ is negative at all time horizons in the pre-October 1979 sample period (Panel B), indicating that real interest rates are negatively correlated with nominal interest rates. In the October 1979 to October 1982 period, $1 - \beta_m$ flips sign and become strongly positive with t-statistics of $1 - \beta_m = 0$ ranging from 3.03 to 8.25. Indeed, as the t-test of $\beta_m = 0$ indicates, we cannot reject the hypothesis that $1 - \beta_m$ equals one, i.e., that the nominal interest rate moves one-for-one with the real interest rate. After October 1982, $1 - \beta_m$ remains positive, and is significant in four of the five time horizons, but relative to results for the October 1979 to October 1982 period it is smaller for all the time horizons.

As our discussion of the β_m coefficient estimates indicates, there seems to be large changes of the β_m coefficients over the three sub-periods examined here. Table 2 presents formal tests (Wald tests) of parameter instability. The tests provide very strong evidence that β_m is not the same in the three sub-periods. The marginal significance levels for the tests of shifts in β_m indicate that the probability of obtaining that high a value of the test statistic under the null hypothesis that β_m are equal in the three sub-periods is always less than 1 in

Table 2

Tests of Parameter Shifts in the Inflation Equations

m (months)	Test of shifts in β_m $\chi^2(2)$ =	Marginal significance level	Tests of shift in α_m and β_m $\chi^2(4)$ =	Marginal significance level
1	43.49**	3.59×10^{-10}	118.20**	1.30×10^{-24}
3	24.71**	4.30×10^{-6}	104.25**	1.22×10^{-21}
6	36.03**	1.50×10^{-8}	124.15**	6.93×10^{-26}
9	52.08**	4.91×10^{-12}	101.80**	4.07×10^{-21}
12	67.22**	2.53×10^{-15}	94.37**	1.55×10^{-19}

Tests of null hypothesis that the parameters are equal in the three periods, February 1964 - October 1979, November 1979 - October 1982, and November 1982 - December 1985.

100,000 and is in one case lower than 1 in 10^{14} (for m = 12). Tests for shifts in both α_m and β_m produce even stronger rejections of parameter stability with marginal significance levels of the test statistics hovering around 10^{-20}.

The results in Table 1 and Table 2 are consistent with earlier findings in the literature which have examined the relationship between future inflation and nominal interest rates for a more limited range of time horizons (one to three months). To summarize the results here: nominal interest rates with maturities of one to twelve months do contain a great deal of information about future inflation, especially in the pre-October 1979 sample period; the constancy of the real rate is strongly rejected; nominal interest rates are negatively correlated with real interest rates in the pre-October 1979 period and are positively correlated with real interest rates thereafter; only during the October 1979 to October 1982 period of the Fed's nonborrowed reserves target operating procedure do nominal interest rates move one-for-one with real rates; and the relationship between nominal interest rates and future inflation does undergo significant shifts with the changes in monetary policy regimes in October 1979 and October 1982.

Although we have found that nominal interest rates contain significant information about the level of future inflation at all the maturities we have looked at here, this does not mean that the slope of the term structure contains information about the future path of inflation. The significant β_m coefficients at all maturities may not reflect any marginal explanatory power for the term structure, but might rather just reflect the fact that the overall level of nominal interest rates is correlated with the overall level of future inflation. To see if the term structure helps to forecast future inflation, we turn to the inflation change equations which regress the change in the future m-period inflation rate from the n-period inflation rate ($\pi_t^m - \pi_t^n$) on the "slope" of the term structure ($i_t^m - i_t^n$). These results are found in Table 3.

As we can see in Table 3, at the shortest end of the term structure (maturities six months or less) the results for the inflation change equations is strikingly different than that found for the inflation level equations in Table 1. For all sample periods and time horizons of six months or less in Table 3, the R^2s of the inflation change equations are near zero and none of the $\beta_{m,n}$ coefficients is statistically significant. Further tests, found in the appendix, which look at all combinations of m and n, for m and n less than or equal to six months, also strongly confirm these findings. Apparently, the term structure for maturities of six months or less contains almost no information about the path of future inflation.[10]

The results for the longer maturities, nine and twelve months, tell a different story, however. For the full sample period in Panel A, the difference between the twelve month and six month rate and between the twelve month and nine month rate do have substantial information about the path of future inflation. The $\beta_{12,6}$ and $\beta_{12,9}$ coefficients are quite

[10] The word "information" is being used in this paper quite narrowly. Information in the term structure about the path of future inflation refers only to the ability of the slope, $i_t^m - i_t^n$ to predict the change in the inflation rate, $\pi_t^m - \pi_t^n$. None of the evidence in this paper rules out predictive power for forecasting equations that make use of more complicated interactions of interest rates at maturities of six months or less or that use additional economic variables in combination with the slope of the term structure. This paper focuses on the predictive power of the slope term, $i_t^m - i_t^n$, because it is the most natural piece of information in the term structure to examine.

Table 3

Estimates of Inflation Change Equations

$$\pi_t^m - \pi_t^n = \alpha_{m,n} + \beta_{m,n} [i_t^m - i_t^n] + \eta_t^{m,n}$$

m,n (months)	$\alpha_{m,n}$	$\beta_{m,n}$	R^2	SE	t-test of $\beta_{m,n} = 0$	t-test of $1-\beta_{m,n} = 0$
	Panel A: February 1964 - December 1986 Sample Period					
3, 1	0.1686 (0.1420)	-0.3182 (0.2595)	0.005	1.883	-1.23	5.08**
6, 3	-0.0433 (0.1236)	0.2330 (0.4150)	0.003	1.197	0.56	1.85
9, 6	-0.0411 (0.0712)	0.4902 (0.2901)	0.020	0.726	1.69	1.76
12, 6	-0.1351 (0.1108)	0.9493 (0.3327)	0.073	0.920	2.85**	0.15
12, 9	-0.0498 (0.0411)	1.1799 (0.3150)	0.087	0.501	3.75**	-0.57
	Panel B: February 1964 - October 1979 Sample Period					
3, 1	0.1421 (0.1851)	-0.3127 (0.4498)	0.002	1.754	-0.70	2.92**
6, 3	0.0379 (0.1427)	0.1813 (0.5499)	0.001	1.088	0.33	1.49
9, 6	0.0826 (0.0647)	0.0014 (0.2695)	0.000	0.692	0.01	3.71**
12, 6	0.0249 (0.1251)	0.7108 (0.4873)	0.033	0.895	1.46	0.59
12, 9	-0.0199 (0.0397)	1.5336 (0.3527)	0.145	0.461	4.35**	-1.51

Panel C: November 1979 - October 1982 Sample Period

3, 1	0.0086	-0.1774	0.003	2.397	-0.48	3.16**
	(0.5638)	(0.3727)				
6, 3	-0.6397	0.5696	0.036	1.442	0.91	0.69
	(0.1767)	(0.6269)				
9, 6	-0.3646	1.3277	0.239	0.765	3.94**	-0.97
	(0.1304)	(0.3373)				
12, 6	-0.5736	1.4179	0.335	0.875	7.17**	-2.11*
	(0.1452)	(0.1978)				
12, 9	-0.2292	0.6740	0.037	0.599	1.31	0.63
	(0.0691)	(0.5152)				

Panel D: November 1982 - December 1986 Sample Period

3, 1	0.6458	-0.9540	0.027	2.009	-1.05	2.14*
	(0.5678)	(0.9125)				
6, 3	0.2666	-0.5976	0.007	1.346	-0.61	1.64
	(0.3702)	(0.9741)				
9, 6	0.0086	0.1060	0.000	0.739	0.16	1.38
	(0.2315)	(0.6492)				
12, 6	0.2676	-0.7101	0.021	0.894	-0.71	1.71
	(0.4582)	(1.0017)				
12, 9	0.1379	-1.1027	0.033	0.509	-0.77	1.48
	(0.2003)	(1.4242)				

close to one and have t-statistics of 2.85 and 3.75, making them significantly different from zero at the one percent level. In looking at the sub-periods, we see the $\beta_{12,9}$ coefficient is significantly different from zero in the February 1964 - October 1979 sample period, as are the $\beta_{9,6}$ and $\beta_{12,6}$ coefficients in the November 1979 - October 1982 sample period. Indeed, the R^2's are particularly high for the regressions with m,n = 9,6 and 12,6 in the November 1979 - October 1982 period, indicating a large amount of forecast power for changes in inflation.[11]

The t-tests of $1 - \beta_{m,n} = 0$ in Table 3 indicate that the term structure of nominal interest rates contains information about the term structure of real interest rates for the shortest maturities. For $\underline{m}$ = 3, in all the sample periods the $[1 - \beta_{m,n}]$ terms are always positive, statistically significant from zero, and never statistically significant

[11]Another way to look at the information in the term structure follows along the lines of Fama (1984) who looks at whether forward rates help predict future changes in one-month interest rates. This involves regressions of the change in the one-period inflation rate, $\pi^1_{t+m} - \pi^1_t$, on the forward-spot differential at time t, $f_{t+m} - i^1_t$ (f_{t+m} = the forward rate at time t for the one-period interest rate at time t+m). For the full sample period, the coefficient on the forward-spot differential is never significantly different from zero, and it is significant for only a few time horizons in other sample periods. The evidence thus does not reveal a great deal of information in forward rates about future changes in one-month inflation rates.

Because $\pi^1_{t+m} - \pi^1_t$ is the cumulative change in the one-month inflation rate from t to t+m, the regressions described above do not tell us whether forward rates have predictive ability for successive changes in one-month inflation rates. Regressions of $\pi^1_{t+m} - \pi^1_{t+m-1}$ on $f_{t+m} - f_{t+m-1}$ produce similar results to those described above; only in 3 out of 44 cases is the coefficient on the forward rate differential significantly different from zero. Forward rates, therefore, do not seem to contain much information about successive future changes in one-month inflation rates either. A somewhat similar test involves regressing $\pi^m_t - \pi^{m-1}_t$ on $i^m_t - i^{m-1}_t$ and the results are also very similar; only 5 out of the 44 estimated $\beta_{m,m-1}$ coefficients are statistically significant at the five percent level. Furthermore, for both of these tests, significant coefficients are as likely to be negative as positive.

from one. Indeed, in three of the four sample periods the $[1 - \beta_{3,1}]$ term is statistically significant from zero at the one percent level, and for the full sample period, its t-statistic is very large, exceeding five. The evidence thus suggests that most fluctuations in the slope of the term structure at the very short end reflect changes in the slope of the term structure of real interest rates on a one-for-one basis and do not reflect changes in expectations about future changes in inflation.

The t-tests of $1 - \beta_{m,n}$ for the intermediate maturities, six and nine months, provide some evidence for the the nominal term structure containing information about the real term structure, but the evidence is not always strong. For $\underline{m}$ = 6 and 9, $[1 - \beta_{m,n}]$ is positive in seven out of eight cases in Table 3, but is statistically significant only once (when $\underline{m}$ = 9 in Panel B, the February 1964 - October 1979 sample period). However, additional results in the appendix lend stronger support for the view that the term structure of nominal interest rates provides information about the term structure of real interest rates for maturities up to six months.

At the longest end of the term structure studied here, however, the story is quite different. For $\underline{m}$ equal to twelve months, $[1 - \beta_{m,n}]$ is only significantly different from zero in one case, and in this case it is negative and is only barely significant at the five percent level. The evidence for maturities around twelve months thus suggests that the term structure of nominal interest rates does not provide information about the term structure of real interest rates, but is rather more likely to reflect changes in expectations about future changes in inflation.

The nature of the shifts in the $\beta_{m,n}$ coefficients in the three sub-periods is completely different than is true for the inflation level equations. Where Table 2 shows very strong evidence that the β_m coefficients for $m \leq 6$ are not stable in the inflation level equations, Table 4 indicates that we cannot reject the equality of the $\beta_{m,n}$ in the three sample periods for $m \leq 6$; indeed, the marginal significance levels are quite high with values above one-half. For m,n = 9,6 and 12,6, Table 4 indicates that $\beta_{m,n}$ do shift from one sub-period to another, but as the comparison of the Table 3 versus Table 1 results indicates, these shifts are in exactly the opposite direction for the inflation change equations than they are for the inflation level equations. In the inflation level equations, the β_m coefficients decline in November 1979 - October 1982 period and become statistically insignificant, thus indicating that the level of interest rates is unable to forecast the future level of inflation. In stark contrast, the $\beta_{9,6}$ and $\beta_{12,6}$ coefficients increase in the November 1979 - October 1982 period and become statistically significant, thus indicating a greater ability to forecast the future path of inflation. Tests for shifts in both $\alpha_{m,n}$ and $\beta_{m,n}$ reveal evidence of parameter stability when $m > 3$, but the marginal significance levels are far higher than in Table 2.

As was discussed in the section on some additional econometric issues, if equation residuals across different time horizons are correlated, more efficient estimates can be obtained by exploiting this information with seemingly unrelated regression (SUR) estimation. Table 5, which contains the SUR estimates of the same inflation change equations found in Table 3, indicates that SUR estimation sometimes leads to large increases in efficiency -- nearly half of the standard errors of the $\beta_{m,n}$ coefficients decline more than 20% in going from Table 3 to

Table 4

Tests of Parameter Shifts in the Inflation Change Equations

m,n (months)	Test of shifts in $\beta_{m,n}$ $\chi^2(2)$ =	Marginal significance level	Tests of shifts in $\alpha_{m,n}$ and $\beta_{m,n}$ $\chi^2(4)$ =	Marginal significance level
3, 1	0.62	0.7324	0.89	0.9263
6, 3	0.96	0.6175	10.94*	0.0272
9, 6	10.70**	0.0047	20.54**	0.0004
12, 6	6.17*	0.0457	17.58**	0.0015
12, 9	4.46	0.1075	11.23**	0.0241

Tests of null hypothesis that the parameters are equal in the three periods, February 1964 - October 1979, November 1979 - October 1982, and November 1982 - December 1985.

Table 5

Seemingly Unrelated Regression (SUR) Estimates of Inflation Change Equations

$$\pi_t^m - \pi_t^n = \alpha_{m,n} + \beta_{m,n} [i_t^m - i_t^n] + \eta_t^{m,n}$$

m,n (months)	$\alpha_{m,n}$	$\beta_{m,n}$	t-test of $\beta_{m,n} = 0$	t-test of $1-\beta_{m,n} = 0$
		Panel A: February 1964 - December 1986 Sample Period		
3, 1	0.1717 (0.1303)	-0.3250 (0.2717)	-1.20	4.88**
6, 3	-0.0499 (0.1186)	0.2565 (0.4182)	0.61	1.78
9, 6	-0.0743 (0.0555)	0.7552 (0.2231)	3.39**	1.10
12, 6	-0.1022 (0.0944)	0.7694 (0.2059)	3.74**	1.12
12, 9	-0.0268 (0.0450)	0.7806 (0.2768)	2.82**	0.79
		Panel B: February 1964 - October 1979 Sample Period		
3, 1	0.1539 (0.1211)	-0.3492 (0.3495)	-1.00	3.86**
6, 3	0.0236 (0.1327)	0.2325 (0.5889)	0.39	1.30
9, 6	0.0026 (0.0643)	0.6565 (0.3815)	1.72	0.90
12, 6	0.0231 (0.1048)	0.7209 (0.2673)	2.70**	1.04
12, 9	0.0239 (0.0437)	0.7968 (0.4117)	1.94	0.49

Panel C: November 1979 - October 1982 Sample Period

3, 1	0.2554 (0.3745)	-0.4318 (0.3387)	-1.27	4.23*
6, 3	-0.7206 (0.1434)	0.8235 (0.4506)	1.83	0.39
9, 6	-0.3603 (0.1205)	1.1949 (0.1903)	6.28**	-1.02
12, 6	-0.5733 (0.1452)	1.1677 (0.1907)	6.12**	-0.88
12, 9	-0.2165 (0.0670)	1.0814 (0.3256)	3.32**	-0.25

Panel D: November 1982 - December 1986 Sample Period

3, 1	0.5609 (0.6342)	-0.8067 (1.0911)	-0.74	1.66
6, 3	0.2600 (0.2726)	-0.5715 (0.6782)	-0.84	2.32*
9, 6	0.1108 (0.2613)	-0.3958 (0.9463)	-0.42	1.47
12, 6	0.1842 (0.4229)	-0.4485 (0.9885)	-0.45	1.47
12, 9	0.0670 (0.1735)	-0.4858 (1.0407)	-0.47	1.43

Table 5.[12] Even greater increases in efficiency can be achieved by choosing a system of equations in which the error terms are highly correlated. For example, in the system of eleven equations in which $\underline{m}$ = 2, 3, ... 12 and $\underline{n}$ = 1, SUR estimation results in standard errors which typically decline by 50% and in a few cases decline by as much as 80%.

The increased efficiency of the SUR estimates, however, does not change the conclusions we reached before; indeed, in the cases where there are large efficiency gains it only tends to strengthen them. In the full sample of Panel A, we find that, in addition to $\beta_{12,6}$ and $\beta_{12,9}$, $\beta_{9,6}$ now is significantly different from zero at the one percent level. $\beta_{12,6}$ also becomes significant in the February 1964 to October 1979 sample period and in the November 1979 to October 1982 sample period, we now also find that all three of these coefficients are statistically significant at the one percent level. SUR estimation thus strengthens the conclusion that the slope of the term structure around maturities of nine and twelve months provides substantial information about the path of future inflation.

VI. Conclusions.

The empirical evidence in this paper indicates that the information in the term structure about the future path of inflation is quite

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[12]Note, however, that in several cases, the estimated standard errors are higher in Table 5 than they are in Table 3. Even though the SUR estimates are asymptotically more efficient than OLS estimates, in small samples estimated SUR standard errors can turn out to be larger than OLS standard errors, especially when the order of the MA process for each equation is taken to be the highest order of the MA process for the error term of any equation in the system. (This is done to assure positive-definiteness of the variance-covariance matrix. See footnote 7.)

different at the shortest end of the term structure (maturities six months or less) than it is for maturities nine months or greater. For maturities of six months or less, in all the sample periods examined -- February 1964 to December 1986, 1964 to October 1979, November 1979 to October 1982, November 1982 to December 1986 -- the term structure provides almost no information about the future path of inflation. On the other hand at this end of the term structure, the results do indicate that the term structure of nominal interest rates contain a great deal of information about the term structure of real interest rates. This finding is quite important because it suggests that researchers can examine observable data on the shortest end of the nominal term structure to provide them with information about the behavior of the real term structure.[13]

For maturities of nine and twelve months, the term structure does appear to contain information about the future path of inflation. The

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[13]The inability of the shortest end of the term structure to provide information about the future path of inflation is related to research that suggests that term premiums undergo substantial fluctuations over time (Jones and Roley (1983), Mankiw and Summers (1984) Shiller, Campbell and Schoenholtz (1983), and Startz (1982)). If these premiums are an important enough source of fluctuations in the slope of the term structure, then they can obscure any information that the term structure might contain about expected changes in inflation and hence future changes in inflation. Furthermore since fluctuations in these term premiums are reflected in both nominal and real interest rates, it is logical that the slopes of the nominal and real term structures move together. The fact that the term structure of nominal interest rates for very short maturities contains a great deal of information about the term structure of real interest rates is then just a consequence of substantial fluctuations in term premiums which dominate fluctuations in expected changes in inflation. One way that this could occur is if changes in the inflation rate are not very predictable for short horizons (as would be the case if the inflation rate were close to a random walk over short horizons). Then fluctuations of expected changes in inflation would be small, while fluctuations in the slope of the term structure of real interest rates might be high because term premiums vary substantially over time.

$\beta_{m,n}$ coefficients for these maturities are often highly significant in the full sample period and in the sub-periods before October 1982.[14] At these longer maturities, however, there does not appear to be much information in the nominal term structure about the term structure of real interest rates.

The results in this paper on the information in the term structure strongly contrast with those which find that, for much of the postwar period, nominal interest rate movements reflect changes in expected inflation rather than real interest rates. Furthermore, while the degree to which the overall level of nominal interest rates reflects expectations of future inflation decreases dramatically with the change in the monetary policy regime in October 1979 and October 1982, the information in the term structure about future inflation remains unchanged for the shortest maturities, but increases for maturities of nine and twelve months in the October 1979 to October 1982 period. The nature of the relationship between the term structure and future inflation is thus strikingly different from that found between the overall level of interest rates and future inflation.

The evidence in this paper suggests that some caution should be exercised in using the term structure of interest rates as a guide for assessing inflationary pressures in the economy, as is currently under consideration by the Federal Reserve. Although there is apparently significant information in the term structure about the future path of

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[14]These results are consistent with those in Fama (1988) who finds that the interest rate spread between five year bonds and one year bonds helps forecast future one-year changes in inflation for horizons of two to five years.

inflation for maturities greater than six months, there is no information about the future path of inflation that can be obtained from the shorter end of the term structure.

Appendix

Table A1
Additional Estimates of Inflation Change Equations

$$\pi_t^m - \pi_t^n = \alpha_{m,n} + \beta_{m,n} [i_t^m - i_t^n] + \eta_t^{m,n}$$

m (months)	α_m	β_m	R^2	SE	t-test of $\beta_m = 0$	t-test of $1-\beta_m = 0$
Panel A: January 1959 - December 1986 Sample Period						
2, 1	0.0214 (0.0822)	-0.0468 (0.2312)	0.000	1.525	-0.20	4.53**
3, 1	0.1236 (0.1247)	-0.2433 (0.2406)	0.003	1.856	-1.01	5.17**
4, 1	0.1422 (0.1515)	-0.2287 (0.2427)	0.002	2.076	-0.94	5.06**
5, 1	0.2063 (0.1684)	-0.2773 (0.2166)	0.004	2.174	-1.28	5.90**
6, 1	0.2563 (0.1846)	-0.3026 (0.2085)	0.005	2.237	-1.45	6.25**
3, 2	0.0367 (0.0608)	-0.1682 (0.2620)	0.001	0.929	-0.64	4.46**
4, 2	0.0395 (0.0889)	-0.0943 (0.2762)	0.000	1.254	-0.34	3.96**
5, 2	0.0645 (0.1239)	-0.1126 (0.3016)	0.001	1.453	-0.37	3.69**
6, 2	0.0976 (0.1518)	-0.1492 (0.3000)	0.001	1.562	-0.50	3.83**
4, 3	-0.0064 (0.0353)	0.2011 (0.2929)	0.001	0.693	0.69	2.73**
5, 3	-0.0042 (0.0711)	0.1085 (0.3304)	0.000	0.969	0.33	2.70**
6, 3	-0.0303 (0.1121)	0.1888 (0.3827)	0.002	1.146	0.49	2.12*
5, 4	-0.0056 (0.0341)	0.1182 (0.1926)	0.001	0.544	0.61	4.58**
6, 4	0.0171 (0.0710)	-0.0145 (0.2741)	0.000	0.781	-0.05	3.70**
6, 5	-0.0076 (0.0274)	0.1516 (0.2145)	0.002	0.448	0.71	3.95**

Panel B: January 1959 - October 1979 Sample Period

m (months)	α_m	β_m	R^2	SE	t-test of $\beta_m = 0$	t-test of $1-\beta_m = 0$
2, 1	0.0016 (0.1025)	0.0437 (0.4121)	0.000	1.545	0.11	2.32*
3, 1	0.0521 (0.1546)	-0.0733 (0.3780)	0.000	1.747	-0.19	2.84**
4, 1	-0.1531 (0.1873)	0.5089 (0.3809)	0.005	1.933	1.34	1.29
5, 1	-0.0474 (0.2082)	0.2421 (0.3549)	0.001	2.038	0.68	2.14*
6, 1	0.0178 (0.2086)	0.1360 (0.3213)	0.001	2.110	0.42	2.69**
3, 2	-0.0018 (0.0722)	0.1360 (0.3874)	0.000	0.874	0.35	2.23*
4, 2	-0.0766 (0.0995)	0.5501 (0.3746)	0.007	1.100	1.47	1.20
5, 2	-0.0256 (0.1282)	0.2850 (0.3802)	0.002	1.288	0.75	1.88
6, 2	0.0616 (0.1487)	0.0695 (0.3610)	0.000	1.404	0.19	2.58**
4, 3	0.0021 (0.0400)	0.3027 (0.4028)	0.002	0.652	0.75	1.73
5, 3	0.0426 (0.0710)	0.0287 (0.3828)	0.000	0.869	0.07	2.54*
6, 3	0.0428 (0.1242)	0.1086 (0.4753)	0.001	1.042	0.23	1.88
5, 4	0.0199 (0.0356)	0.0448 (0.2199)	0.000	0.518	0.20	4.34**
6, 4	0.1085 (0.0692)	-0.2786 (0.2910)	0.006	0.711	-0.96	4.39**
6, 5	0.0211 (0.0277)	0.0425 (0.2216)	0.000	0.430	0.19	4.32**

Panel C: November 1979 - October 1982 Sample Period

m (months)	α_m	β_m	R^2	SE	t-test of $\beta_m = 0$	t-test of $1-\beta_m = 0$
2, 1	0.0329 (0.3621)	-0.1436 (0.3830)	0.003	1.585	-0.37	2.99**
3, 1	0.0086 (0.5638)	-0.1774 (0.3727)	0.003	2.397	-0.48	3.16**
4, 1	0.1149 (0.6199)	-0.3788 (0.3539)	0.013	2.734	-1.07	3.90**
5, 1	-0.1063 (0.5763)	-0.2805 (0.2863)	0.008	2.815	-0.98	4.47**
6, 1	-0.3064 (0.5270)	-0.2447 (0.2573)	0.007	2.810	-0.95	4.84**
3, 2	-0.0531 (0.2283)	-0.1570 (0.3607)	0.002	1.187	-0.44	3.21**
4, 2	-0.0893 (0.2798)	-0.3413 (0.3046)	0.008	1.799	-1.12	4.40**
5, 2	-0.3927 (0.2622)	0.0059 (0.4059)	0.000	2.045	0.01	2.45*
6, 2	-0.5788 (0.2423)	0.0306 (0.4063)	0.000	2.089	0.08	2.39*
4, 3	-0.1222 (0.1170)	-0.0683 (0.4713)	0.000	0.874	-0.14	2.27*
5, 3	-0.4146 (0.1775)	0.4858 (0.6155)	0.021	1.299	0.79	0.84
6, 3	-0.6397 (0.1767)	0.5696 (0.6269)	0.036	1.442	0.91	0.69
5, 4	-0.2203 (0.0789)	0.4017 (0.2817)	0.024	0.639	1.43	2.12*
6, 4	-0.4523 (0.1225)	0.5836 (0.4515)	0.048	0.944	1.29	0.92
6, 5	-0.1880 (0.0714)	0.3052 (0.5104)	0.008	0.498	0.60	1.36

Panel D: November 1982 - December 1986

m (months)	α_m	β_m	R^2	SE	t-test of $\beta_m = 0$	t-test of $1-\beta_m = 0$
2, 1	0.0197 (0.3064)	0.0854 (0.6791)	0.000	1.431	0.13	1.35
3, 1	0.6458 (0.5678)	-0.9540 (0.9125)	0.027	2.009	-1.05	2.14*
4, 1	0.9349 (0.7593)	-1.2462 (1.0638)	0.041	2.236	-1.17	2.11
5, 1	0.9392 (0.7432)	-0.9963 (0.7704)	0.035	2.337	-1.29	2.59**
6, 1	0.9694 (0.8053)	-0.9139 (0.7086)	0.032	2.401	-1.29	2.70**
3, 2	0.2267 (0.2225)	-1.0718 (0.8695)	0.026	1.011	-1.23	2.38*
4, 2	0.2377 (0.3689)	-0.6693 (0.9812)	0.008	1.508	-0.68	1.70
5, 2	0.4725 (0.4634)	-0.9685 (0.8167)	0.020	1.714	-1.19	2.41*
6, 2	0.5175 (0.5559)	-0.8467 (0.8737)	0.019	1.829	-0.97	2.11*
4, 3	-0.0242 (0.1126)	0.8361 (0.6221)	0.014	0.759	1.34	0.26
5, 3	0.1493 (0.2160)	-0.3739 (0.6152)	0.002	1.154	-0.61	2.23*
6, 3	0.2666 (0.3702)	-0.5976 (0.9741)	0.007	1.346	-0.61	1.64
5, 4	0.0500 (0.1293)	-0.0538 (0.8213)	0.000	0.596	-0.07	1.28
6, 4	0.1074 (0.3104)	-0.1421 (1.3090)	0.001	0.921	-0.11	0.87
6, 5	-0.0153 (0.1179)	0.8724 (1.2750)	0.021	0.487	0.68	0.10

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