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The Predictive Power of the Term Structure during Recent Monetary Regimes

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ABSTRACT

I use weekly Treasury-bill rates with maturities of one to twenty-six weeks to examine the information in forward rates during the 1970s and 1980s. Forward rates contain better information about future changes in spot rates than the information captured by autoregressive and vector-autoregressive models. Forward rates also have considerable predictive power, which increased after October 1979 and remained strong after October 1982. The results show no necessary connection between interest rate predictability and the degree to which the Fed adheres to interest rate targeting.

MOST OF THE RECENT empirical literature on the term structure of interest rates is concerned with the informative content of forward rates. According to the expectations hypothesis, forward rates are, up to a constant, unbiased predictors of future realized spot rates. The expectations hypothesis has been tested extensively and rejected. (See Jones and Roley [13] and their references.) However, rejection of the expectations hypothesis does not imply that there is no information in the term structure. Fama [6] examines one- to six-month Treasury bills from 1959 through 1982 and, although he rejects the expectations hypothesis, finds predictive power in forward rates that lasts about three to five months during the first half of his sample and one month during the second half of his sample. Others provide contrary evidence. Shiller, Campbell, and Schoenholtz [20] use quarterly post-World War II data and claim that forward rates predict the wrong direction in the subsequent change of long-term interest rates.[1]

Recently, Mankiw and Miron [16] extend the sample period back to year 1890. They use three- and six-month rates and find strong predictive ability during the period 1890 through 1914 before the establishment of the Federal Reserve, some predictive ability during the period 1915 through 1933, and, contrary to Fama's results, no predictive ability after 1933 or after 1959. They argue that the predictive ability of the term structure during the earlier periods was due to the

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[1] A regression of the realized change in the long rate on the lagged spread between the long rate and the short rate yields a negative coefficient. However, Mankiw [15] uses similar data and finds that forward rates predict the correct direction in the subsequent realization of short-term interest rates, but his regression results are not statistically significant.

absence of Fed intervention in the money market. In their view, the Fed's targeting of interest rates makes interest rate changes unpredictable.[2]

In this paper, I examine the predictive ability of the term structure across recent monetary regimes that are characterized by different degrees of interest rate targeting by the Federal Reserve. One of my aims is to scrutinize the Mankiw-Miron hypothesis. My sample runs from January 1972 through November 1985, which encompasses three separate official monetary "regimes": the subperiods January 1972 to October 1979, October 1979 to October 1982, and October 1982 to November 1985. Huizinga and Mishkin [12] have argued that these are three separate monetary regimes, indeed. A similar conclusion is reached in the money-announcements literature. (See Cornell [2], Hardouvelis [11], and Roley [19].) A rough characterization of the three regimes in chronological order would be interest rate targeting, lack of any interest rate targeting, and partial interest rate targeting.[3]

A second aim of this paper is to examine empirically how well the market is able to predict, a question that has not been addressed sufficiently in the literature. As Fama [6] emphasizes, the presence of time-varying risk premia can diminish and perhaps destroy the predictive ability of forward rates; yet forward rates may optimally incorporate publicly available information on future spot rates.

Finally, the lack of a consensus among researchers about the predictive power of the term structure during the recent past suggests re-examining the evidence using a complementary data set. I use weekly data on Treasury bills with maturities of one to twenty-six weeks. This rich array of maturities provides precision in detecting how far into the future the term structure is able to predict. The use of weekly rather than monthly or quarterly data may also provide more statistical power for the detection of predictive ability. Furthermore, a weekly or a biweekly holding period is closer to the holding-period horizon of major players in the money market (the dealers, the Fed, some portfolio managers) than the holding period of one month or one quarter that previous investigators have used.

[2] For example, suppose that there is a seasonal demand for credit that increases interest rates during a particular month of the year. Market participants would figure out such an empirical regularity, and their expectations of a temporary future increase in interest rates would be reflected in forward rates. However, if the Federal Reserve were to follow interest rate targets, it would not allow such a seasonal variation in interest rates. Thus, forward rates would lose their ability to predict future changes in interest rates.

[3] Until October 6, 1979, the Federal Reserve's intermediate targets were the growth rate of M1 and the level of short-term interest rates. Interest rates were not allowed to fluctuate freely even on a daily basis. From October 7, 1979, to October 5, 1982, the Fed officially adopted M1 as its only intermediate target and nonborrowed reserves as its inter-week and intra-week instrument for monetary control. This implied that interest rates could fluctuate freely. After October 6, 1982, the Fed pronounced that, besides M1, it would follow additional wider monetary aggregates, and it switched to borrowed reserves as its inter-week instrument for monetary control but kept nonborrowed reserves as its intra-week instrument. Interest rates cannot fluctuate as freely as in the October 1979 through October 1982 period, but they can fluctuate more freely than in the pre-October 1979 period, at least at the daily and weekly frequencies.

Section I derives the equations that will be estimated and examines the role that time-varying risk premia play in the ability of forward rates to predict future spot rates. Section II presents the empirical evidence, and Section III summarizes the principal conclusions.

I. Theoretical Framework

A. Definitions

I begin by assuming a two-week holding-period horizon, which conforms to the subsequent presentation of the empirical evidence. Let $P(n)_t$ denote the price at the beginning of week t of a Treasury bill that matures at the beginning of week $t + n$ and pays one dollar. Define $R(n)_t$, the n-week continuously compounded annualized interest rate observed at the beginning of week t, as

$$R(n)_t = -(365/7n)\ln\{P(n)_t\}, \quad n = 1, 2, \cdots, 26, \tag{1}$$

and $f(2, n)_t$, the two-week continuously compounded annualized forward rate n weeks ahead, as

$$f(2, n)_t = (365/14)\ln\{P(n)_t/P(n+2)_t\}, \quad n = 1, 2, \cdots, 24. \tag{2}$$

This forward rate can always be decomposed into $\mathrm{E}_t\theta(2)_{nt}$, an expected time-varying risk premium, and $\mathrm{E}_tR(2)_{t+n}$, the expected two-week rate of week $t + n$, as follows (see Fama [4, 5]):

$$f(2, n)_t = \mathrm{E}_t\theta(2)_{nt} + \mathrm{E}_tR(2)_{t+n}, \tag{3}$$

where E_t denotes the expectations operator conditional on information available at the beginning of week t. Equation (3) provides an implicit definition of the expected risk premium and is always true. The expectations hypothesis imposes the restriction that the expected risk premium is constant: $\mathrm{E}_t\theta(2)_{nt} = \theta(2)_n$, and thus,

$$f(2, n)_t = \theta(2)_n + \mathrm{E}_tR(2)_{t+n}. \tag{3'}$$

The expectations hypothesis is compatible with an arbitrage-pricing equilibrium and is less restrictive than the pure expectations hypothesis, which states that $\mathrm{E}_t\theta(2)_{nt} = 0$.[4]

B. Cumulative Predictive Power

In order to test for the presence of predictive power in forward rates, I will rewrite equation (3) in a more convenient form. For notational simplicity, let us suppress the number 2, the indicator of a two-week period, in $R(2)_t$, $f(2, n)_t$, and

[4] Cox, Ingersoll, and Ross [3] have criticized the different versions of the pure expectations hypothesis that exist in the literature as being incompatible with each other and with an arbitrage-pricing equilibrium. However, Campbell [1] has shown that their criticisms do not carry over to the more general expectations hypothesis that I consider in this paper.

$\theta(2)_{nt}$. Subtracting R_t from both sides of equation (3), we get

$$\mathrm{E}_t R_{t+n} - R_t = -\mathrm{E}_t \theta_{nt} + [f(n)_t - R_t]; \quad n = 1, 2, \cdots, 24. \tag{4a}$$

Let us now write the realized change in the two-week spot interest rate from the beginning of week t to the beginning of week $t + n$ as the sum of the expected change at t plus a forecast error, w_{t+n}:

$$R_{t+n} - R_t = \mathrm{E}_t[R_{t+n} - R_t] + w_{t+n}. \tag{5a}$$

Substituting (4a) into (5a), we get

$$R_{t+n} - R_t = -\mathrm{E}_t \theta_{nt} + [f(n)_t - R_t] + w_{t+n}. \tag{6a}$$

Equation (6a) does not impose any restrictions on the data. It is only a convenient representation of the realized change in the two-week interest rate. Note that, if expectations are formed rationally, the forecast error w_{t+n} is uncorrelated with variables in the information set at time t and, therefore, with the right-hand-side variables in equation (6a). Rational expectations will be part of our maintained hypothesis throughout the paper.

Our test for the presence of predictive power is based on the following regression equation:

$$R_{t+n} - R_t = a_n + b_n[f(n)_t - R_t] + u_{t+n}. \tag{7a}$$

We can think of equation (7a) as the empirical counterpart of equation (6a), in which the unobserved expected risk premium $\mathrm{E}_t\theta_t$ is excluded from the regression. From (3), $\mathrm{E}_t\theta_t$ is typically positively correlated with $f(n)_t - R_t$, and thus the estimated slope coefficient will be biased downward and will be less than unity. As in Fama [6] or Mankiw and Summers [17], we can compute its probability limit by utilizing equation (6a) and the assumption of rational expectations:

$$\operatorname{plim} \hat{b}_n = \frac{\sigma^2(\mathrm{E}_t[R_{t+n} - R_t]) + \rho_n \sigma(\mathrm{E}_t[R_{t+n} - R_t])\sigma(\mathrm{E}_t\theta_{nt})}{\sigma^2(\mathrm{E}_t[R_{t+n} - R_t]) + \sigma^2(\mathrm{E}_t\theta_{nt}) + 2\rho_n \sigma(\mathrm{E}_t[R_{t+n} - R_t])\sigma(\mathrm{E}_t\theta_{nt})}, \tag{8}$$

where ρ_n is the correlation coefficient between $\mathrm{E}_t\theta_{nt}$ and $\mathrm{E}_t(R_{t+n} - R_t)$, $\sigma^2(\cdot)$ denotes variance, and $\sigma(\cdot)$ denotes standard deviation. Note that the presence of a time-varying risk premium in equation (8) implies that $\sigma(\mathrm{E}_t\theta_{nt}) > 0$ and, thus, $\operatorname{plim} \hat{b}_n \neq 1$. However, under the expectations hypothesis, the expected risk premium is constant and, thus, $\sigma(\mathrm{E}_t\theta_{nt}) = 0$. Substituting this into equation (8) gives $\operatorname{plim} \hat{b}_n = 1$. This is a testable restriction. Its rejection is equivalent to rejection of the assumption of a constant expected risk premium. Previous authors have rejected the expectations hypothesis. Thus, the most interesting question is whether there is any predictive ability in the term structure, i.e., whether $b_n > 0$.[5]

[5] The expectations hypothesis does not restrict the actual risk premium, θ_{nt}, to be constant, and, therefore, it does not imply that $\sigma(\theta_{nt}) = 0$. Under the expectations hypothesis, the actual risk premium may vary over time, but it can only vary around a constant mean. Alternatively, recall the statistical property that the variance of a random variable equals the expected value of its conditional variance plus the variance of its conditional expectation; i.e., $\sigma^2(\theta_{nt}) = \mathrm{E}[\sigma_t^2(\theta_{nt})] + \sigma^2[\mathrm{E}_t\theta_{nt}]$. The expectations hypothesis states that $\sigma^2[\mathrm{E}_t\theta_{nt}] = 0$, which does *not* imply that $\sigma^2(\theta_{nt}) = 0$.

Equation (8) shows that plim $\hat{b}_n$ depends on two parameters: the correlation coefficient ρ_n and the ratio, say q_n, between the standard deviation of the expected change in the spot rate, $\sigma(\mathrm{E}_t[R_{t+n} - R_t])$, and the standard deviation of the expected risk premium, $\sigma(\mathrm{E}_t\theta_{nt})$. Figure 1 plots the relationship beween plim $\hat{b}_n$ and q_n for different values of ρ_n. Note that as $q_n \to \infty$, plim $\hat{b}_n \to 1$; i.e., when the variability in the expected future change in the spot rate overwhelms the variability in the expected risk premium, we approach the limiting case of a constant expected risk premium as described by the expectations hypothesis. Also note that it is possible to find negative $\hat{b}_n$s; i.e., it is possible that forward rates systematically predict the wrong direction of the subsequent change in spot rates even when they incorporate available information (perhaps) optimally. This can occur when $\rho_n < 0$ and $q_n > 1$ and provides a possible explanation for the counterintuitive result of Shiller et al. [20]. Finally, it is also possible to find the other extreme case of $\hat{b}_n$s larger than unity when $\rho_n < 0$ and $q_n > 1$.

C. Marginal Predictive Power

The dependent variable in equation (7a) is the sum of consecutive week-to-week changes in the interest rate that obscures individual week-to-week (marginal) changes. It is possible, for example, that the term structure is only able to predict for, say, five weeks ahead, but the contribution of $R_{t+5} - R_t$ in, say, $R_{t+13} - R_t$ is enough to give us a positive b_{13}. Thus, to be able to isolate how far into the future the term structure is truly able to predict, we have to examine differences between consecutive forward rates as predictors of future marginal changes in interest rates. Let us rewrite equation (3) as follows:

$$\mathrm{E}_t[R_{t+n} - R_{t+n-1}] = -\mathrm{E}_t(\theta_{nt} - \theta_{n-1,t}) + [f(n)_t - f(n-1)_t],\ n = 1, 2, \cdots, 24. \tag{4b}$$

and, utilitizing

$$R_{t+n} - R_{t+n-1} = \mathrm{E}_t[R_{t+n} - R_{t+n-1}] + e_{t+n}, \tag{5b}$$

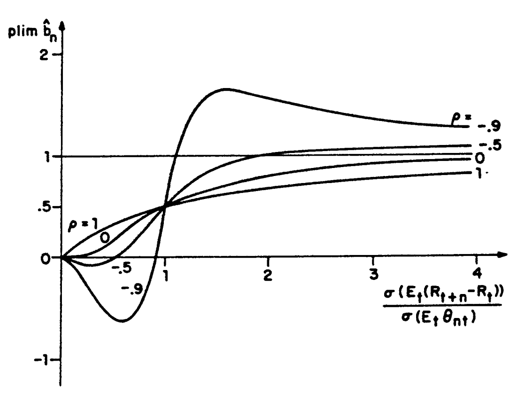

Figure 1. Equation (8)

where e_{t+n} is assumed to be a rational forecast error, we get

$$R_{t+n} - R_{t+n-1} = -\mathrm{E}_t(\theta_{nt} - \theta_{n-1,t}) + [f(n)_t - f(n-1)_t] + e_{t+n}. \tag{6b}$$

To find the marginal predictive ability of forward interest rates, we may, therefore, run the following regression equation:

$$R_{t+n} - R_{t+n-1} = c_n + d_n[f(n)_t - f(n-1)_t] + v_{t+n}. \tag{7b}$$

The expectations hypothesis imposes the restrictions that $d_n = 1$. Presence of marginal predictive ability implies that $d_n > 0$.

There is also an econometric reason for our interest in marginal predictive power. Observe that, in equation (7b), the independent variable, $f(n)_t - f(n-1)_t$, is observed at time t but the dependent variable, $R_{t+n} - R_{t+n-1}$, is observed at times $t + n - 1$ and $t + n$. The existing observation gap for $n > 1$ practically guarantees that, if measurement errors plague these variables, they will be uncorrelated. This implies that, if there is measurement error in $f(n)_t - f(n-1)_t$, it will bias the estimated coefficients $\hat{d}_n$ toward zero. Thus, positive and significant $\hat{d}_n$s reinforce conclusions on the presence of predictive power. This is not true for the estimated coefficients $\hat{b}_n$ of equation (7a). Measurement error can cause the bias to go in either direction. This is because the presence of measurement error in R_t implies a common, and thus correlated, error in the dependent and independent variables of equation (7a).[6]

D. How Well Does the Market Forecast Future Spot Rates?

Our previous analysis showed that time-varying risk premia obscure the ability of forward rates to capture the market's expectation of future spot rates. The size and significance of the estimated regression coefficients $\hat{b}_n$ and $\hat{d}_n$ provide information on how well forward rates predict, but they do not tell us how well the *market* predicts. To measure the market's ability to forecast future changes in interest rates, we need independent information on the parameters $\sigma(\mathrm{E}_t\theta_{nt})$ and ρ_n, which is difficult to find.

In this paper, I generate model-based forecasts of future spot rates and compare them with the forward rates of the term structure. Specifically, I compare the R^2s of regressions (7a) and (7b) with the R^2s from similar regressions in which the independent variables were generated from a model. If the R^2s of regressions (7a) and (7b) are larger than their model-based counterparts, then I conclude that forward rates incorporate better information about future spot rates than the information utilized by the models. This is because the R^2s of regression equations (7a) or (7b) represent *lower bound* statistics on the market's ability to forecast future spot rates. That is,

$$\text{plim } R^2 \le \{\sigma^2(\mathrm{E}_t R_{t+n} - R_t)/\sigma^2(R_{t+n} - R_t)\}. \tag{9}$$

[6] See, for example, the exposition in Maddala [14], page 302.

The proof is contained in the Appendix. A similar relationship holds for the R^2 of equation (7b).[7]

II. Empirical Evidence

A. Data

The original interest rate data are taken from the quotation sheets of the Federal Reserve Bank of New York and represent Thursday afternoon (3:30 P.M. Eastern time) bond-equivalent yields based on ask prices.[8] Bond-equivalent yields of T-bills with different maturities contain arbitrary compounding periods, and their use in the caculation of forward rates induces a maturity-dependent bias. (See Fielitz [7].) To avoid this bias, the bond-equivalent yields were transformed into continuously compounded annualized yields as follows. From the definition of the bond-equivalent yield,

$$i_{BEY} \equiv (365/d)(1/P - 1), \tag{10}$$

one can calculate the price of a T-bill, P, once d is known. d denotes the number of days between the settlement date and the maturity date. The New York Fed assumes a two-business-day settlement period in its calculation of the i_{BEY}s. This implies that $d = 3$ for the one-week T-bill, $d = 10$ for the two-week T-bill, etc. (Monday settlement to Thursday maturity). The price data can then be used to calculate forward rates according to equation (2). To calculate the spot interest rates $R(n)_t$, it is necessary to modify equation (1) to take into account a two-business-day settlement period (as in Stigum [21], page 30). Thus, spot interest rates were constructed as follows:

$$R(n)_t = -\{365/[3 + 7(n - 1)]\}\ln[P(n)_t]. \tag{1$'$}$$

Table I presents summary statistics for the change in the two-week rate, $R_{t+1} - R_t$, the typical dependent variable of our regressions. Notice that, until October 1979, the weekly volatility of $R_{t+1} - R_t$ was only thirty-one basis points, which is consistent with the strict interest rate targeting that the Fed followed on a daily and weekly basis. From October 1979 to October 1982, volatility increased to 140 basis points, which is also consistent with the Fed's claim that it allowed interest

[7] Fama [6] states this result without proof. Fama shows that, for $\rho_n \geq 0$, the quantity $R^2/\hat{b}_n$ represents an upper bound on the market's predictive ability. Also note that an alternative way to check whether forward rates have extra predictive power over the predictive power of benchmark models is to run Granger-Causality tests as follows: include lagged dependent variables in equations (7a) and (7b) and check the significance of $\hat{b}_n$ and $\hat{d}_n$. The results of these tests confirm the results of Tables II and III.

[8] Trading in T-bills with maturities of other than three or six months is relatively thin, especially for the very short-term bills. However, there exists a similar and very active market for repurchase agreements with maturities of one, two, and three weeks and one, two, three, and six months. Arbitrage between the two markets should ensure the quality of our T-bill data with short maturities. An examination of the data reveals that the one-week rate is distinctly more volatile than all other rates and that its bid-ask spreads are much larger. For this reason, I use a two-week holding period; yet the results for the one-week holding period are not substantially different.

Table I

Interest Rate Predictability[a]

	Univariate Autoregressions			Vector Autoregressions		
Period	1	2	3	1	2	3
No. of Observations	399	157	161	399	157	161
No. of Lags from Akaike Criterion	5	6	3	5	3	2
$\sigma[R_{t+1} - R_t]$	0.309	1.397	0.647	0.309	1.396	0.647
$\sigma[\mathrm{E}_t R_{t+1} - R_t]$	0.074	0.499	0.353	0.164	0.703	0.371
SEE	0.302	1.329	0.547	0.267	1.243	0.541
$\bar{R}^2$	0.046	0.093	0.285	0.253	0.208	0.302
F-Statistic	4.84*	3.66*	22.23*	9.98*	5.56*	12.53*
(Significance Level)	(0.0003)	(0.0020)	(0.0000)	(0.0000)	(0.0000)	(0.0000)

[a] Dependent variable: $R_{t+1} - R_t$. $R_{t+1} - R_t$ is the change in the two-week rate; $\sigma[R_{t+1} - R_t]$ is its standard deviation; $\sigma[\mathrm{E}_t R_{t+1} - R_t]$ is the standard deviation of the regression fit; SEE is the regression standard error; $\bar{R}^2$ is the coefficient of determination adjusted for degrees of freedom; the F-statistic refers to all independent variables except the constant. The vector autoregressions include the change in the one-week rate, $R(1)_{t+1} - R(1)_t$, the change in the two-week rate, $R_{t+1} - R_t$, and the change in the thirteen-week rate, $R(13)_{t+1} - R(13)_t$. The Akaike criterion picks the number of lags that minimize $-\ln$ (likelihood) + number of parameters. See Table II for the definition of each period.

rates to fluctuate freely. After October 1982, volatility decreased to sixty-five basis points, which shows the Fed's partial return to interest rate targeting.

Table I also offers some evidence on the connection between model-based predictability of future changes in interest rates and the degree to which the Federal Reserve adheres to interest rate targeting. Two types of models were fitted to each subperiod: a univariate autoregression (AR) and a vector autoregression (VAR) that includes the change in the one-, two-, and thirteen-week rates. The Akaike information criterion was used to select the appropriate number of lags. Predictability can be measured by the size of the adjusted coefficient of determination, $\bar{R}^2$. Notice that the $\bar{R}^2$ of the VAR model does not change very drastically across the three time periods. It decreases slightly after October 1979 and increases after October 1982. This is not what we would expect to observe under the Minkiw-Miron hypothesis, which predicts an increase in $\bar{R}^2$ after October 1979 and a decrease in $\bar{R}^2$ after October 1982.[9]

B. Econometric Issues

Before turning to the main empirical results, there are two technical issues that deserve attention: the method I used to generate model-based interest rate

[9] The Mankiw-Miron hypothesis is based on two claims. The first is that interest rate targeting makes future changes in interest rates unpredictable. The second is that this unpredictability is "rationally" reflected on a lack of forecasting power in forward rates. The evidence in Table I is about the first claim. The $\bar{R}^2$'s as well as the regression estimates of Subsection C may change across the subperiods for reasons unrelated to Fed policy. This is an important and obvious qualifier to all the comparisons that I make in this paper across the different subperiods. Also note that Mankiw and Miron concentrated on σ $(\mathrm{E}_t R_{t+1} - R_t)$, the numerator of the horizontal axis of Figure 1, assuming implicitly that $\sigma(\mathrm{E}_t\theta_{nt})$ and ρn remain constant. The changes in $\hat{\sigma}(\mathrm{E}_t R_{t+1} - R_t)$ of Table I are consistent with the Mankiw-Miron hypothesis, but the different SEEs suggest that $\rho(\mathrm{E}_t\theta_{nt})$ may have changed in a way that opposes the predictions of the Mankiw-Miron hypothesis.

forecasts, and the construction of consistent estimates of the standard errors of the regression slope coefficients of equations (7a) and (7b). The model-based forecasts were computed after estimating each model of Table I separately for each subperiod using the Kalman filter with random coefficients. The Kalman filter has the desirable property of utilizing information available only up to the time the forecast is made. With every new observation, it updates the coefficient estimates as well as their variance-covariance matrix. Furthermore, a Kalman filter with random coefficients allows the underlying true structure to be time varying. Re-initializing the Kalman filter at the beginning of each subperiod captures the fact that the regime changes that occurred were widely announced and may have, therefore, induced agents to reformulate their forecasting models somewhat abruptly, discarding very old time-series information.[10]

The random parameters of the AR and VAR models, α_j, were modeled as random walks:

$$\alpha_{j,t} = \alpha_{j,t-1} + \varepsilon_{j,t}, \quad j = 0, 1, \cdots, kL, \tag{11}$$

where L denotes the number of lags, $k = 1$ for the AR models, and $k = 3$ for the VAR models. Estimation of the random-coefficients model requires an assumption about the ratio of the variance of the ε_j s to the variance of the error term in the AR and VAR models. I chose this ratio to be 1/1000 in every subperiod.[11] Also, to initialize the Kalman filter, I used thirty observations from the beginning of the first subperiod and thirty observations prior to the beginning of the second and third subperiods. In the calculation of the R^2s of the first subperiod, I included forecasts from the first thirty observations. These forecasts are based on the parameter estimates of a single regression over the thirty-week period.

Let us turn now to the estimation of equations (7a) and (7b). The major econometric difficulty is by now familiar from the work of Hansen and Hodrick [10]. I use weekly observations, but the forecasting horizon in the regressions is n weeks, which generates a moving-average error term of order $n - 1$. The serial correlation of the error term does not affect the consistency of OLS coefficient estimates, $\hat{b}_n$ and $\hat{d}_n$, but it requires an adjustment in the variance-covariance matrix of these estimates.[12] Hansen [9] shows that a consistent estimate of the variance-covariance matrix is the following:

$$S = (X'X)^{-1}A(X'X)^{-1}; \; A = \textstyle\sum_{k=-n+1}^{n-1} \sum_{t=1}^{T} \hat{u}_t X_t' X_{t-k} \hat{u}_{t-k}. \tag{12}$$

X is the matrix of regressors; X_t is the vector of regressors at time t; $\hat{u}_t$ is the

[10] A random-coefficient model does have a better forecasting performance than a fixed-coefficient model. Also the R^2s do not change substantially when a single random-coefficient Kalman-filter model is fitted throughout the entire sample period, and the results are robust with respect to different lag lengths.

[11] The likelihood function is a nonlinear and highly complicated function of this ratio. In principle, this ratio can be chosen to maximize the likelihood function by searching over the parameter space. (See Garbade [8].) The value 1/1000 maximizes the likelihood over the whole sample period.

[12] Fama [6] reports uncorrected OLS standard errors that lead to incorrect statistical inferences. However, Fama does not emphasize the statistical significance of his $\hat{b}_n$s and $\hat{d}_n$s; he emphasizes the size of the regression R^2s.

residual at time t from the OLS regression; T is the sample size. For $n = 1$, the correction is for conditional heteroscedasticity as in White [22].

One problem with the above variance-covariance matrix is that it need not be positive semi-definite in finite samples. For this reason, I adopted Newey and West's [18] proposed modification on S, which is consistent and, by construction, positive semi-definite. The modification is to include, inside the summation terms of matrix A, the multiplicative factor $[1 - (k/n)]$. This factor downweights the higher order autocovariances of the residual $\hat{u}_t$ and achieves positive definiteness. Consistency is achieved by the fact that, as T increases, the downweighting is reduced.

C. The Results

Tables II and III contain the main results. Table II refers to cumulative predictive power, and Table III refers to marginal predictive power. The tables present the estimated slope coefficients $\hat{b}_n$ and $\hat{d}_n$ with their corrected standard errors, the regression R^2s of equations (7a) and (7b), R^2_{ar}, and R^2_{var}. R^2_{ar} and R^2_{var} come from regressions similar to (7a) and (7b) in which the AR and VAR forecasts take the place of forward rates. The minus signs in parentheses below R^2_{ar} and R^2_{var} denote a negative regression coefficient, i.e., that the model's predicted change is in the opposite direction from the realized change.

The AR and VAR models have very little predictive power. During the first two subperiods, predictive power lasts for only one week. During the third subperiod, there is predictive power that lasts for about six weeks. About half the time, the AR model predicts the wrong direction in the future change in interest rates. As expected, the VAR model predicts slightly better than the AR model, and predictions about cumulative changes in interest rates are slightly better than predictions about marginal changes. The overall evidence provides no support for the Mankiw-Miron hypothesis, which is what we also found in Table I.

In view of the models' poor forecasting performance, the term-structure results are impressive. Observe that, in each subperiod, the R^2s of regression equations (7a) and (7b) are always larger than the corresponding R^2_{ar}s and R^2_{var}s—by a substantial margin in subperiods 2 and 3. This is reassuring. Since the R^2 represents a lower bound on the market's predictive ability (recall inequality (9)), we conclude that the market utilizes much better information in its forecasting effort than what we have incorporated in our models.[13]

Let us examine now the predictive power of forward rates. Table II shows that, in the first subperiod, cumulative predictive power lasts for about seven weeks. During the second subperiod, predictive power increases substantially and lasts for the entire twenty-four-week horizon. In fact, coefficient estimates are so high that it is now hard to reject the expectations hypothesis. The expectations

[13] Observe that the AR and VAR models are reduced-form models and do not restrict agents' actual information set to lagged interest rates only. The indirect effects of many serially correlated third factors are also captured through their effects on lagged interest rates.

hypothesis is rejected in only three out of the twenty-four regressions (at $n = 3$, 4, 5). During the third subperiod, predictive power remains strong for fourteen weeks and then disappears. Overall, the change in the term structure's predictive power across the different monetary regimes is consistent with the Mankiw-Miron hypothesis. This contrasts with the conclusion we reached based on the results of the AR and VAR models. A consistent explanation for all the evidence is possible, however, when we recall that a change in the predictive power of forward rates does not necessarily originate from a change in the market's ability to predict. Figure 1 shows precisely how the size of the regression slope coefficient is affected by the parameters that describe the evolution of expected risk premia. Thus, based on the results from the models, I postulate that the market's predictive power did not necessarily change across the monetary regimes, but the parameters that describe the variability in expected risk premia changed and affected the predictive power of forward rates.

Table III confirms the presence of predictive power in forward rates. Indeed, recall that positive findings about marginal predictive power are very reliable since the estimated slope coefficients $\hat{d}_n$, if biased, can only be biased toward zero, i.e., against finding predictive power. The table also confirms an increase in predictive power after October 1979. However, it also shows some differences. Prior to October 1979, the predictive power of the term structure lasted for only one week. Thus, the results of Table II that show cumulative predictive power for about seven weeks are apparently driven by the ability to forecast only one week ahead. Beyond one week, the term structure is more likely to predict the wrong than the right direction in the subsequent marginal change in interest rates.[14] From October 1979 through October 1982, marginal predictive power increases but shows an irregular pattern. It disappears after six weeks and reappears from fourteen to twenty-one weeks ahead. Notice also that, contrary to the results of Table II, the expectations hypothesis is overwhelmingly rejected. The only exception occurs at $n = 1$, for which the regression equations in Tables II and III are the same. Two possible reasons for this discrepancy come to mind. The first is that the results of Table II are simply dominated by the $R_{t+1} - R_t$ component of $R_{t+n} - R_t$. The second is related to our discussion in Section I, Subsection C. A downward bias in $\hat{d}_n$ may have occurred together with a possible upward bias in $\hat{b}_n$. Finally, after October 1982, marginal predictive power remains strong for nine weeks. Then it falls to zero and shows up again from eighteen to twenty-one weeks ahead.

Table IV presents tests of structural change of the parameters b_n and d_n across the three subperiods. The tests take into account the heteroscedasticity across the subperiods and the MA $(n - 1)$ process of the error term.[15] The cumulative

[14] Figure 1 shows that it is possible to find a negative slope coefficient. Also note that the negative coefficient cannot be attributed to measurement-error bias. The existence of such bias cannot change the sign of an estimated coefficient, but it can decrease the absolute magnitude of the coefficient.

[15] The tests come from a weighted least-squares regression with subperiod dummies and weights equal to one over the standard error of estimate of each subperiod. The variance-covariance matrix of the coefficient estimates in the WLS regression is corrected to take into account the moving average in the error term and the conditional heteroscedasticity. WLS provides the same results as OLS except for column 1-(2+3) in Table IV.

Table II

The Cumulative Predictive Power of the Term Structure and of Time-Series Models[a]

	Period 1 (01/06/72–10/04/79) *NOBS* = 405				Period 2 (10/11/79–10/07/82) *NOBS* = 157				Period 3 (10/14/82–11/07/85) *NOBS* = 161 − *n*			
n	$\hat{b}_n$	R^2	R^2_{ar}	R^2_{var}	$\hat{b}_n$	R^2	R^2_{ar}	R^2_{var}	$\hat{b}_n$	R^2	R^2_{ar}	R^2_{var}
1	0.860*	0.192	0.027	0.141	1.057*	0.386	0.035	0.072	0.883*	0.652	0.168	0.254
	(0.199)				(0.108)				(0.063)			
2	0.605*	0.131	0.015	0.123	0.847*	0.426	0.024	0.112	0.762*	0.609	0.273	0.354
	(0.147)				(0.088)				(0.049)			
3	0.545*	0.101	0.003	0.088	0.656*	0.223	0.000	0.087	0.729*	0.354	0.091	0.218
	(0.111)				(0.089)		(−)		(0.088)			
4	0.409*	0.064	0.001	0.036	0.573*	0.172	0.014	0.046	0.394*	0.148	0.136	0.037
	(0.105)		(−)		(0.123)		(−)		(0.100)		(−)	
5	0.357*	0.043	0.001	0.060	0.620*	0.209	0.000	0.091	0.463*	0.218	0.048	0.101
	(0.143)		(−)		(0.103)				(0.092)			
6	0.356*	0.046	0.001	0.035	0.746*	0.210	0.004	0.078	0.511*	0.216	0.153	0.133
	(0.146)		(−)		(0.130)				(0.079)			
7	0.255*	0.034	0.000	0.027	0.650*	0.132	0.001	0.037	0.462*	0.126	0.003	0.044
	(0.129)		(−)		(0.172)		(−)		(0.122)			
8	0.147	0.015	0.000	0.018	0.648*	0.108	0.015	0.002	0.341*	0.083	0.132	0.004
	(0.107)		(−)		(0.202)		(−)		(0.107)		(−)	
9	0.145	0.012	0.000	0.016	0.865*	0.201	0.022	0.009	0.421*	0.129	0.003	0.008
	(0.117)		(−)		(0.247)		(−)		(0.098)			
10	0.095	0.005	0.000	0.018	0.939*	0.191	0.007	0.006	0.469*	0.135	0.046	0.023
	(0.109)				(0.176)		(−)		(0.098)			
11	0.038	0.001	0.000	0.022	0.720*	0.118	0.009	0.011	0.492*	0.110	0.035	0.006
	(0.096)				(0.165)		(−)		(0.132)		(−)	
12	0.350*	0.065	0.000	0.015	0.747*	0.133	0.016	0.000	0.274*	0.061	0.162	0.002
	(0.115)				(0.201)		(−)	(−)	(0.125)		(−)	(−)
13	0.283*	0.062	0.000	0.007	0.731*	0.108	0.022	0.000	0.373*	0.104	0.007	0.006
	(0.086)				(0.250)		(−)	(−)	(0.134)			(−)
14	0.159	0.009	0.000	0.005	0.796*	0.124	0.023	0.003	0.407*	0.119	0.024	0.001
	(0.162)				(0.246)		(−)	(−)	(0.098)			(−)

15	0.252	0.019	0.000	0.005	0.769*	0.142	0.020	0.000	0.222	0.031	0.054	0.005
	(0.169)				(0.232)		(–)		(.148)		(–)	(–)
16	0.101	0.004	0.001	0.003	0.793*	0.150	0.038	0.004	0.116	0.010	0.088	0.015
	(0.162)				(0.249)		(–)	(–)	(0.116)		(–)	(–)
17	0.034	0.001	0.001	0.001	0.769*	0.134	0.041	0.001	0.159	0.023	0.034	0.015
	(0.162)				(0.235)		(–)	(–)	(0.100)			(–)
18	0.131	0.008	0.001	0.001	0.670*	0.115	0.041	0.002	0.136	0.017	0.032	0.011
	(0.153)				(0.235)		(–)	(–)	(0.105)			(–)
19	0.214	0.020	0.001	0.001	0.655*	0.118	0.040	0.006	0.062	0.003	0.068	0.030
	(0.149)				(0.232)		(–)	(–)	(0.148)		(–)	(–)
20	0.267	0.029	0.000	0.001	0.677*	0.101	0.061	0.033	−0.053	0.002	0.056	0.058
	(0.171)				(0.247)		(–)	(–)	(0.103)		(–)	(–)
21	0.213	0.020	0.000	0.000	0.621*	0.084	0.059	0.016	−0.087	0.005	0.027	0.058
	(0.134)				(0.265)		(–)	(–)	(0.112)			(–)
22	0.109	0.005	0.000	0.000	0.749*	0.118	0.062	0.040	−0.035	0.001	0.010	0.053
	(0.157)		(–)		(0.300)		(–)	(–)	(0.141)			(–)
23	−.020	0.000	0.000	0.000	1.008*	0.188	0.076	0.009	−0.159	0.015	0.089	0.073
	(0.158)		(–)		(0.290)		(–)	(–)	(.214)		(–)	(–)
24	−0.069	0.004	0.000	0.000	0.958*	0.211	0.074	0.027	−0.155	0.015	0.021	0.113
	(0.111)		(–)		(0.253)		(–)	(–)	(0.194)		(–)	(–)

[a] $R_{t+n} - R_t = a_n + b_n\,[f(n)_t - R_t] + u_{t+n}$. R_t denotes the two-week interest rate observed at time t; $f(n)_t$ denotes the two-week forward rate n weeks ahead observed at time t. Method-of-moments standard errors are shown in parentheses. R^2 is the coefficient of determination. R^2_{ar} and R^2_{var} are the R^2s when the independent variable is a forecast based on a random-coefficient Kalman-filter estimation of the AR and VAR models of Table I. The minus sign in parentheses below R^2_{ar} and R^2_{var} denotes a negative regression coefficient. *NOBS* denotes number of observations.

* Significant at the five percent level.

Table III

The Marginal Predictive Power of the Term Structure and of Time-Series Models[a]

	Period 1 (01/06/72–10/04/79) $NOBS = 405$				Period 2 (10/11/79–10/07/82) $NOBS = 157$				Period3 (10/14/82-11/07/85) $NOBS = 161 - n$			
n	$\hat{d}_n$	R^2	R^2_{ar}	R^2_{var}	$\hat{d}_n$	R^2	R^2_{ar}	R^2_{var}	$\hat{d}_n$	R^2	R^2_{ar}	R^2_{var}
1	0.860*	0.192	0.027	0.141	1.057*	0.386	0.035	0.072	0.883*	0.652	0.168	0.254
	(0.199)				(0.108)				(0.063)			
2	0.132	0.006	0.019	0.038	0.653*	0.260	0.026	0.018	0.652*	0.488	0.123	0.161
	(0.116)				(0.092)				(0.059)			
3	0.118	0.008	0.014	0.023	0.522*	0.130	0.022	0.023	0.543*	0.276	0.080	0.081
	(0.105)				(0.122)				(0.096)			
4	−0.083	0.005	0.008	0.012	0.306*	0.109	0.016	0.001	0.298*	0.120	0.077	0.059
	(0.056)				(0.079)				(0.058)			
5	−0.097	0.007	0.000	0.003	0.411*	0.106	0.009	0.002	0.397*	0.138	0.029	0.034
	(0.054)		(−)		(0.089)			(−)	(0.088)			
6	−0.001	0.000	0.003	0.003	0.147*	0.018	0.006	0.013	0.465*	0.112	0.015	0.020
	(0.043)		(−)	(−)	(0.073)			(−)	(0.124)			
7	−0.078*	0.009	0.000	0.002	0.145	0.013	0.000	0.001	0.417*	0.122	0.021	0.027
	(0.040)			(−)	(0.074)			(−)	(0.102)			
8	−0.049	0.004	0.000	0.000	0.179*	0.020	0.004	0.002	0.390*	0.079	0.011	0.006
	(0.035)				(0.062)			(−)	(0.123)			
9	−0.034	0.002	0.002	0.002	0.150	0.013	0.000	0.010	0.494*	0.136	0.006	0.006
	(0.032)		(−)		(0.095)		(−)	(−)	(0.123)			
10	−0.046	0.005	0.001	0.002	0.149	0.011	0.001	0.003	0.146	0.023	0.004	0.007
	(0.042)		(−)	(−)	(0.083)		(−)	(−)	(0.083)		(−)	(−)
11	−0.056*	0.012	0.001	0.006	0.027	0.000	0.000	0.001	0.183	0.026	0.001	0.003
	(0.022)		(−)	(−)	(0.074)		(−)		(0.113)		(−)	(−)
12	0.024	0.003	0.002	0.007	0.144	0.014	.000	0.003	0.128	0.016	0.009	0.028
	(0.022)		(−)	(−)	(0.081)		(−)	(−)	(0.090)		(−)	(−)
13	−0.000	0.000	0.000	0.002	0.167	0.018	0.001	0.002	0.105	0.011	0.001	0.000
	(0.017)		(−)	(−)	(0.088)		(−)	(−)	(0.084)		(−)	
14	0.011	0.000	0.000	0.001	0.300*	0.061	0.009	0.000	0.162*	0.023	0.000	0.000
	(0.038)			(−)	(0.097)		(−)	(−)	(0.066)			

15	−0.054	0.005	0.001	0.000	0.298*	0.080	0.009	0.011	0.089	0.010	0.000	0.002
	(0.037)				(0.085)		(−)	(−)	(0.047)			(−)
16	−0.112*	0.025	0.001	0.000	0.364*	0.091	0.015	0.006	0.092	0.014	0.001	0.000
	(0.043)				(0.077)		(−)	(−)	(0.057)			(−)
17	−0.086	0.015	0.000	0.001	0.346*	0.105	0.000	0.000	0.068	0.007	0.008	0.000
	(0.043)				(0.063)				(0.046)			
18	−0.033	0.003	0.000	0.001	0.230*	0.046	0.004	0.001	0.109*	0.015	0.014	0.000
	(0.043)				(0.100)		(−)	(−)	(0.043)			
19	−0.019	0.001	0.000	0.000	0.223*	0.042	0.003	0.000	0.120*	0.020	0.013	0.000
	(0.043)			(−)	(0.061)		(−)		(0.061)			(−)
20	−0.027	0.002	0.000	0.000	0.178*	0.022	0.006	0.001	0.201*	0.036	0.001	0.001
	(0.036)		(−)		(0.066)		(−)		(0.057)			(−)
21	−0.026	0.002	0.000	0.001	0.214*	0.025	0.043	0.000	0.116*	0.013	0.002	0.000
	(0.033)		(−)	(−)	(0.066)		(−)	(−)	(0.041)			(−)
22	−0.073*	0.017	0.000	0.000	0.132	0.008	0.001	0.008	0.089	0.010	0.001	0.000
	(0.034)		(−)	(−)	(0.130)				(0.049)			(−)
23	−0.038	0.006	0.000	0.000	0.204	0.013	0.011	0.001	0.107	0.009	0.001	0.002
	(0.020)		(−)	(−)	(0.165)		(−)	(−)	(0.077)			(−)
24	−0.017	0.003	0.001	0.002	0.144	0.017	0.002	0.001	0.038	0.001	0.009	0.000
	(0.015)		(−)	(−)	(0.119)			(−)	(0.061)			(−)

[a] $R_{t+n} - R_{t+n-1} = c_n + d_n\,[f(n)_t - f(n-1)_t] + v_{t+n}$. See Table II, footnote a.

* Significant at the five percent level.

Table IV

Tests of Structural Change[a]

	Cumulative Predictive Power				Marginal Predictive Power			
	b_n				d_n			
n	1-2	1-3	2-3	1-(2+3)	1-2	1-3	2-3	1-(2+3)
1	0.8	0.0	2.0	0.1	0.8	0.0	2.0	0.1
2	2.0	1.0	0.7	1.7	12.4*	16.0*	0.0	17.0*
3	0.6	1.7	0.3	1.2	6.3*	8.9*	0.0	10.3*
4	1.0	0.0	1.3	0.3	16.3*	22.4*	0.0	26.4*
5	2.2	0.4	1.3	1.2	23.6*	22.9*	0.0	35.0*
6	4.0*	0.9	2.4	2.6	3.0	12.6*	4.8*	12.0*
7	3.4	1.4	0.8	2.9	7.0*	20.3*	4.6*	22.3*
8	4.8*	1.6	1.8	3.9*	10.3*	11.9*	2.4	16.0*
9	6.9*	3.3	2.8	6.5*	3.3	17.2*	4.9*	13.6*
10	16.6*	6.5*	5.4*	12.1*	4.5*	4.3*	0.0	6.5*
11	12.8*	7.7*	1.2	10.6*	1.2	4.3*	1.3	7.2*
12	4.5*	0.6	7.6*	1.5	2.0	1.3	0.0	2.7
13	2.9	0.3	1.6	2.7	3.5	1.5	0.3	3.9*
14	4.7*	1.7	2.1	4.4*	7.6*	4.0*	1.4	8.1*
15	3.2	0.0	3.9*	1.5	14.3*	5.8*	4.6*	12.6*
16	5.4*	0.0	6.0*	3.0	29.7*	8.2*	8.2*	13.1*
17	6.6*	0.4	5.7*	3.7	32.5*	6.1*	12.9*	9.6*
18	3.7	0.0	4.3*	1.7	5.9*	5.4*	1.2	6.8*
19	2.6	0.5	4.6*	1.0	10.6*	3.5*	1.4	7.8*
20	1.9	2.6	7.4*	0.1	7.6*	11.5*	0.1	15.8*
21	1.9	3.0	5.9*	0.1	10.6*	7.3*	1.6	13.1*
22	3.6	0.5	5.5*	1.1	2.3	7.5*	0.1	8.4*
23	9.7*	0.3	10.4*	2.3	2.2	3.3	0.3	5.2*
24	13.8*	0.1	12.2*	4.4*	1.9	0.7	0.6	2.4

[a] Method-of-moments statistic: χ^2 (1). 1-2 refers to period 1 versus period 2; 1-3 refers to period 1 versus period 3; 2-3 refers to period 2 versus period 3; 1-(2+3) refers to period 1 versus periods 2 plus 3. b_n and d_n are the slope coefficients of equations (7a) and (7b).

* Significant at the five percent level.

predictive-power results are mixed. However, the more reliable marginal predictive-power results show a clear pattern: a structural break in October 1979 but not in October 1982.

III. Conclusions

I found reliable evidence that forward rates have predictive power. Until October 1979, forward rates were only able to predict changes in interest rates that would occur one week later. However, when the Fed allowed interest rates to fluctuate relatively freely during the period October 1979 through October 1982, predictive power increased substantially lasting for at least six weeks into the future. (There is also predictive power from fourteen to twenty-one weeks ahead.) After October 1982, when the Fed returned to partial interest rate targeting, predictive power remained strong, lasting nine weeks into the future. Formal tests of structural change confirm a break in October 1979.

The predictive power of forward rates, however, is an ambiguous indicator of the *market's* ability to predict. This is because forward rates contain composite information on the market's expectations about both future spot rates *and* time-varying risk premia. Thus, I also investigated the question of how well the market is able to predict and found interesting evidence. I compared the R^2s from the forward rate regressions, which provide only a lower bound on the market's predictive ability, with the R^2s from similar regressions that used forecasts generated by autoregressive and vector-autoregressive benchmark models. The model-based R^2s indicate little predictability. They are small and increase slightly only after October 1982. Yet the R^2s of the forward rate regressions are always larger, and, throughout the post-October 1979 period, they are larger by a considerable margin. Apparently, markets utilize much better information when they try to forecast future interest rates than the information captured by my reduced-form AR and VAR models.

The Mankiw-Miron hypothesis predicts an increase in predictability after October 1979 and a decrease in predictability after October 1982. Yet the results of the AR and VAR models show no connection between predictability and the degree to which the Fed adheres to interest rate targeting. The term-structure results, on the other hand, conform more closely to the Mankiw-Miron hypothesis because they show an increase in predictive power after October 1979 (but do not show a decrease after October 1982). Of course, changes in the term structure's predictive power across the different monetary regimes are not necessarily due to changes in predictability since they can also originate from changes in parameters that characterize the evolution of expected risk premia. The effect of Fed policy on risk premia is an interesting topic for future research.

Appendix

In this Appendix, I prove inequality (9) of Section I. For notational simplicity, let $r \equiv R_{t+n} - R_t$, $f \equiv f(n)_t - R_t$, $x \equiv \mathrm{E}_t R_{t+n} - R_t$, $y \equiv \mathrm{E}_t \theta_t$, σ_i^2 denote the variance of variable i, and σ_{ij} denote the covariance between variables i and j.

Substituting the definition of plim R^2 into inequality (9), we rewrite it as

$$(\sigma_{fr}^2/\sigma_f^2\,\sigma_r^2) \leq (\sigma_x^2/\sigma_r^2) \leftrightarrow \sigma_{fr}^2 \leq \sigma_f^2\,\sigma_x^2. \tag{A1}$$

However, from equation (4a),

$$\sigma_{fr} = \sigma_x^2 + \sigma_{xy} \tag{A2}$$

and

$$\sigma_f^2 = \sigma_x^2 + \sigma_y^2 + 2\sigma_{xy}. \tag{A3}$$

Substituting (A2) and (A3) into inequality (A1), we may now rewrite the original inequality (9) as

$$\sigma_x^4 + \sigma_{xy}^2 + 2\sigma_x^2\sigma_{xy} \leq \sigma_x^4 + \sigma_x^2\sigma_y^2 + 2\sigma_x^2\sigma_{xy} \leftrightarrow \sigma_{xy}^2 \leq \sigma_x^2\sigma_y^2 \leftrightarrow (\sigma_{xy}/\sigma_x\sigma_y)^2 \leq 1, \tag{A4}$$

which is true because the item in parentheses is a correlation coefficient. Note that inequality (9) becomes an equality when the correlation between x and y, i.e., the parameter ρ of Figure 1, is one.

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