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THE CHANGING BEHAVIOR OF THE TERM STRUCTURE OF POST-WAR U.S. INTEREST RATES AND CHANGES IN THE FEDERAL RESERVE CHAIRMAN. IS THERE A LINK?[*]

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Abstract

Using U.S. interest rate data covering the period 1950:1-1992:7, this paper tests the rational expectations model of the term structure of interest rates. We show evidence that the rational expectations model of the term structure is supported by the data during the seventies and a period lasting from the mid-eighties to the end of the sample. However, during the fifties, sixties and a period that covers most of the Volcker's office term (from September 1979 to April 1986) the term structure model is rejected by the data. Moreover, we find evidence of regime changes in the short-term rate process and the term structure of interest rates. These regime switches roughly coincide with changes in the Federal Reserve chairman. The switches in monetary policy taking place when the chairmanship of the Federal Reserve changes therefore seem to play an important role in characterizing the term structure of interest rates.

Key words: term-structure, rational expectations, regime changes
JEL classification numbers: E43

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1 INTRODUCTION

In a recent paper Blinder (1997) argues that the term structure model is a key element for macroeconomic policy in order to bridge the gap between the nominal short-term interest rate set by monetary policy and the real long-term rates that presumably in‡uence aggregate demand. The expectations theory of the term structure of interest rates postulates that a nominal long-term interest rate is the present value of current and expected future nominal short-term interest rates plus a term premium. There is a great deal of literature showing evidence that the data reject the joint hypothesis of the expectations theory of term structure and rational expectations.[1] Recently, McCallum (1994) has argued that the empirical evidence found in previous studies can be reconciled with the expectations theory under rational expectations by assuming that monetary policy involves smoothing of a short-term interest rate, responses to the level of the spread between a long-term rate and a short-term rate, and an exogenous random term premium.[2]

By allowing for changes in the short-term rate process, this paper provides mild evidence supporting the rational expectations hypothesis of the term structure for long periods of time during U.S. post-war. More important, using the 1-month U.S. Treasury bill rate and the U.S. Treasury 20-year yields from 1950 to 1992, we …nd four di¤erent sub-periods in which the process characterizing the short-term interest rate and the term structure of interest rates have changed.[3] Interestingly, these switches in the term structure roughly coincide with changes in the chairmanship of the Federal Reserve. This …nding is consistent with the evidence found by Peek and Wilcox (1987) that signi…cant changes in monetary policy parameters took

[1] See, for instance, Shiller (1979), Campbell and Shiller (1987), Chow (1989) and Campbell (1995). Recent papers by Hardouvelis (1994), Gerlach and Smets (1997), and Domínguez and Novales (2000) have found empirical evidence in favor of the rational expectations hypothesis of the term structure using international data. However, the …rst two papers also found empirical evidence that the rational expectations hypothesis of the term structure does not …t well U.S. interest rate data.

[2] McCallum's model is a formalized extension of Mankiw and Miron's (1986) argument that the failure of the rational expectations hypothesis of the term structure is due to the interest rate smoothing characterizing the Fed's monetary policy after its founding in 1914. McCallum's model has been tested by Hsu and Kugler (1997). Using data at the short end of the maturity spectrum (one and three month Euro dollar rates), they …nd evidence supporting McCallum's model and the rational expectations hypothesis of the term structure for the most recent sub-sample considered (period 1987-1995).

[3] Mankiw and Miron (1986) suggested in their conclusions that a test of the rational expectations hypothesis of the term structure of interest rates under di¤erent monetary regimes using short-term and long-term rates would be useful for macroeconomic policy. This paper follows their suggestion.

place leading to changes in the reduced form for interest rates when the Federal Reserve chairman changed. Moreover, the evidence provided by Mankiw and Miron (1986) and Mankiw, Miron and Weil (1987) when examining the e¤ects of the establishment of the Federal Reserve in 1914 and the evidence reported in this paper during the U.S. post-war suggest a long-standing causal relation between institutional changes and the behavior of the term structure of interest rates.

We argue that McCallum´s argument of the recurrent failure of the empirical tests of the rational expectations of the term structure found in many studies can be viewed as a particular argument associated with a more general explanation that itself involves several aspects. First, one would expect in general a feedback relationship from the long-term rate to the short-term rate. This feedback can be rationalized, as it was by McCallum, as the result of a monetary policy in which the short-term rate responds to the level of the spread between the long-term rate and the short-term rate. However, there are other ways of explaining this feedback. For instance, in a context with asymmetric information, the long-term rate can summarize private information about the future behavior of interest rates and, therefore, a long-term rate can be used to forecast the future evolution of a short-term interest rate. Second, given the nature of the forces (monetary policy, aggregation of information) summarized by the feedback relationship, this relationship is likely to change over time. A possibility is that the short-term rate may react di¤erently to the spread depending on how tight monetary policy is. Another possibility is that the feedback relationship may change due to variations in the long-term rate volatility. The intuition is that the information content of the long-term rate to forecast the short-term rate may depend on the volatility of the long-term rate. One expects that the higher (lower) the volatility of a long-term rate is, the lower (higher) the informational content given to a long-term rate must be when the short-rate is forecast. As shown below in Figure 1, the volatility of the long-term rate seems to have drastically changed over the post-war period. To sum up, we argue that any short-term rate process assumed in empirical studies in order to test the rational expectations hypothesis of the term structure should be viewed as a reduced form that summarizes both behavioral relationships and economic policy rules. Therefore, the parameters characterizing this reduced form are likely to vary over time.

These considerations suggest that a 'fair' test of the rational expectations hypothesis of the term structure of interest rates should be carried out by taking into account the possibility of changes in the process characterizing the short-term interest rate. This strategy was also followed by Mankiw and Miron (1986) and Hamilton (1988), although this paper di¤ers in many aspects from their papers. First, Mankiw and Miron use OLS regression. Hamilton uses his Markov regime-switching maximum likelihood technique to estimate the model. Here, we use the method of simulated moments. Second, Mankiw and Miron study the term structure of interest rates at the short end of the maturity spectrum (3-month and 6-month rates) using quarterly data from 1890 to 1979. Hamilton uses quarterly yields on 3-month Treasury bills and 10-year Treasury bonds from 1962 to 1987. We use monthly yields data on different terms covering the post-war period. Third, our estimation results point to the presence of regime changes in 1970 and 1986 as well as the one detected in 1979 by Hamilton's study. One possible explanation for these differences (apart from the obvious ones such as the use of different data sets and different econometric techniques) is that Hamilton only allows for the presence of two states since the focus of his paper is to detect the major regime-switching in monetary policy occurring in October 1979. Thus, the presence of minor changes in regime such as those occurring in 1970 and 1986 may have passed unnoticed in Hamilton's analysis.

The rest of the paper is organized as follows. Section 2 introduces the present value model of interest rates under rational expectations which allows for the presence of feedback from the long-term rate to the short-term rate. Section 3 presents and discusses the empirical evidence. Moreover, the robustness of the estimation results is assessed. Finally, Section 4 shows the conclusions.

2 THE PRESENT VALUE MODEL OF INTEREST RATES

As shown by Shiller's (1979) seminal paper, the rational expectations theory of the term structure of interest rates postulates the following relation between a long-term rate and a short-term rate

$$R_t = (1 - \delta)\sum_{i=0}^{\infty}\delta^i E_t r_{t+i} + u_t; \qquad (1)$$

where R_t denotes a long-term rate at time t, r_t is a short-term rate at time t, E_t denotes the conditional expectation operator given the information set, I_t, available to the economic agents at the beginning of time t. I_t includes current and past values of all random variables included in the model. δ denotes the discount factor and u_t is a random error term. We assume that u_t follows an AR(1)

$$u_t = \lambda_0 + \lambda_1 u_{t-1} + z_t, \tag{2}$$

where λ_0 is a constant, $|\lambda_1| \le 1$ and z_t is an i.i.d. random error with mean zero and variance σ_z^2. u_t is often associated with a term premium that is usually assumed constant. However, we share McCallum's view (1994) that it seems implausible that there would not be some period-to-period variability in the error term u_t in (1) since a term such as this reflects changes regarding the need for financial flexibility, measurement error and other disturbing influences. The important point is that the inclusion of u_t in (1) keeps the essence of the expectations theory of the term structure, that is, the long-term rate differs from a weighted average sum of expected future short-term rates only randomly.

We further assume that the short-term interest rate r_t is characterized by the following process

$$r_t = \rho_0 + \rho_1 R_{t-1} + \rho_2 r_{t-1} + v_t, \tag{3}$$

where ρ_0 is a constant, ρ_1 and ρ_2 are both included in the interval $[-1, 1]$, and v_t is a random variable with mean zero.[4] v_t is included in I_t since r_t is also included. Equation (3) is a reduced form characterizing the short-term rate that allows for the presence of a positive feedback from the long-term-rate to the short-term rate. This positive feedback relationship can be rationalized in several ways. One possibility is that the feedback arises by aggregation of asymmetric information, thus, a long-term rate aggregates private information that can be used to predict the evolution of a short-term rate. Another possibility is that the feedback appears when monetary policy uses short-term interest rate as an instrument (as in McCallum (1994)).

Taking into account equation (1) to evaluate $E_t R_{t+1}$ and subtracting $\delta E_t R_{t+1}$ from (1) we obtain

$$R_t = (1 - \delta)\, r_t + \delta E_t R_{t+1} + u_t. \tag{4}$$

Equations (3) and (4) form a bivariate system of difference equations. Using the undetermined coefficient method (Muth (1961), McCallum (1983) among others) we begin by writing R_t as a linear function of a minimal set

[4] We assume that v_t is an i.i.d. random variable with mean zero and variance σ_v^2. In addition, we also estimate the model by allowing for v_t following an AR(1) process: $v_t = \tau_1 v_{t-1} + s_t$, where s_t is an i.i.d. random variable with mean zero and variance σ_s^2. As shown in Table 3 below, we cannot reject the null hypothesis that τ_1 is statistically equal to zero (that is, v_t is a white noise). These results suggest that considering more lagged values of R_t and r_t other than those appearing in (3) is not required.

of state variables u_t, r_t, plus a constant,

$$R_t = \pi_0 + \pi_1 r_t + \pi_2 u_t. \tag{5}$$

For appropriate real values of π_0 and π_1, the expectational variable $E_t R_{t+1}$ will then be given by

$$\begin{aligned} E_t R_{t+1} &= \pi_0 + \pi_1 E_t r_{t+1} = (\pi_0 + \pi_1 \rho_0 + \pi_1 \pi_0 \rho_1) + \\ &\quad \pi_1(\pi_1 \rho_1 + \rho_2) r_t + \pi_1 \pi_2 \rho_1 u_t. \end{aligned} \tag{6}$$

To evaluate the π's, we substitute (3), (5) and (6) into (4), which gives

$$\pi_0 + \pi_1 r_t + \pi_2 u_t = \delta(\pi_0 + \pi_1 \rho_0 + \pi_0 \pi_1 \rho_1) +$$

$$[\delta \pi_1 (\pi_1 \rho_1 + \rho_2) + (1 - \delta)] r_t + (1 + \delta \pi_1 \pi_2 \rho_1) u_t.$$

This equation implies identities in the constant term, r_t and u_t as follows:

$$\begin{aligned} \pi_0 &= \delta(\pi_0 + \pi_1 \rho_0 + \pi_0 \pi_1 \rho_1), \\ \pi_1 &= \delta \pi_1 (\pi_1 \rho_1 + \rho_2) + (1 - \delta), \\ \pi_2 &= 1 + \delta \pi_1 \pi_2 \rho_1. \end{aligned} \tag{7}$$

After some algebra, we can show that there are two solutions to the system of equations (7),

$$\pi^1 \equiv \left(\pi_0^1, \pi_1^1, \pi_2^1\right) = [\tau_1, \alpha_1, \varphi_1], \tag{8}$$

$$\pi^2 \equiv \left(\pi_0^2, \pi_1^2, \pi_2^2\right) = [\tau_2, \alpha_2, \varphi_2], \tag{9}$$

where

$$\begin{aligned} \alpha_1 &= \frac{1}{2\delta\rho_1}\left(1 - \delta\rho_2 + \sqrt{(1 - \delta\rho_2)^2 - 4\rho_1\delta(1 - \delta)}\right), \\ \alpha_2 &= \frac{1}{2\delta\rho_1}\left(1 - \delta\rho_2 - \sqrt{(1 - \delta\rho_2)^2 - 4\rho_1\delta(1 - \delta)}\right), \\ \tau_i &= \frac{\alpha_i \rho_0 \delta}{1 - \delta(1 + \alpha_i \rho_1)} \qquad i = 1, 2, \\ \varphi_i &= \frac{1}{1 - \alpha_i \rho_1 \delta}, \qquad i = 1, 2. \end{aligned}$$

The coefficients α_1 and α_2 are the roots of the second-order characteristic polynomial in π_1 given by the second equation in system (7). In addition to

RE solutions (8) and (9), called α_1-fundamental and α_2-fundamental, respectively; the present value model, equation (4), exhibits another RE equilibrium solution called the backward solution, given by

$$R_t = \delta^{-1} R_{t-1} + \left(1 - \delta^{-1}\right) r_{t-1} - \delta^{-1} u_{t-1} + \varepsilon_t; \qquad (10)$$

where ε_t is an arbitrary martingale difference with respect to I_{t-1} (that is, $\varepsilon_t = R_t - E_{t-1} R_t$ is the rational prediction error).[5]

Each rational expectations solution represents an alternative term structure that relates the long-term rate to the short-term rate in a particular manner. In order to simplify our exposition of the empirical results, we focus our attention on the fundamental solution yield by the minimal state variable criterion suggested by McCallum (1983). Provided that α_1 and α_2 are real numbers (that is, $(1 - \delta\rho_2)^2 - 4\rho_1\delta(1 - \delta) \geq 0$), it is straightforward to show that the α_2-fundamental solution (9) is the one pointed out by McCallum's criterion. Moreover, as shown in Section 3, our estimation results show that this fundamental solution fits the data better than the backward solution (10) for all the sub-samples studied.

Standard unit root and cointegration tests suggest that long- and short-term interest rates are individually I(1) process and the two rates are cointegrated.[6] In order to analyze the implications of cointegration we rewrite the α_1-fundamental and α_2-fundamental equilibrium solutions and the short-term interest rate process (3) in matrix form

$$\begin{pmatrix} \Delta R_t \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} \pi_1\rho_1 - 1 & \pi_1\rho_2 \\ \rho_1 & \rho_2 - 1 \end{pmatrix} \begin{pmatrix} R_{t-1} \\ r_{t-1} \end{pmatrix} + \begin{pmatrix} \pi_2 & \pi_1 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} u_t \\ v_t \end{pmatrix}; \qquad (11)$$

where $\Delta R_t = R_t - R_{t-1}$, and $\Delta r_t = r_t - r_{t-1}$. The long- and short-term interest rates are cointegrated if the error-correction companion matrix is singular.

[5] See Broze and Szafarz (1991, chapter 2) for a detailed discussion on the non-uniqueness issue.

[6] Augmented Dickey-Fuller tests and Phillips-Perron tests were carried out for the whole sample and alternative subsamples. Table A.1 in the appendix shows the Phillips-Perron Z_ρ statistics. For the whole sample the test results are qualitatively similar to those provided by Campbell and Shiller (1987). For the alternative subsamples the empirical evidence on cointegration is much weaker than that found for the whole sample. In spite of this weak evidence on cointegration when dealing with alternative subsamples, we estimate below the model imposing the cointegration restriction for each subsample. The reason is that cointegration is usually viewed as a long-run equilibrium relation. Thus, the fact that Z_ρ statistics show weaker evidence on cointegration when dealing with some subsamples than with the whole sample is interpreted as a lack of power of the cointegration test in small samples.

Formally, this condition imposes the following restriction $\pi_1\rho_1 + \rho_2 = 1$. By using the second equation of (7) we can easily show that the former cointegration restriction implies that $\rho_1 + \rho_2 = 1$ and that the cointegrating vector is $(1; -1)$.[7] The latter result simply says that the long-short spread (that is, $R_t - r_t$) is stationary when R_t and r_t are cointegrated. Moreover, the empirical evidence on cointegration suggests the existence of a close relationship between two of the parameters characterizing the short-term rate process given by the cointegration restriction $\rho_1 + \rho_2 = 1$. This evidence thus suggests than the feedback parameter ρ_1 seems to adjust in a particular manner when the other parameter describing the short-term rate process, ρ_2, changes.

3 EMPIRICAL EVIDENCE

We test the rational expectations hypothesis of the term structure on two monthly U.S. Treasury yield series (1-month U.S. Treasury bill rate and the U.S. Treasury 20-year yields) available from 1950 to 1992 from Salomon Brothers' Analytical Record of Yields and Yield Spreads (1992).[8]

Figure 1 displays the time series of the long-term and short-term interest rates. Figure 2 displays the time series of the long-short spread. Looking at the evolution of the three variables over the whole sample, we observe that each variable behaves remarkably differently over different periods of time. In particular, we observe a first period from 1950:1 to approximately 1970:6 in which the long-term rate is rather smooth. In the following period from 1970:6 to 1979:9 the long-term rate is quite volatile. The period from 1979:9 to 1986:4 presents huge fluctuations in the long-term rate. Finally, the period from 1986:4 to 1992:7 is characterized by a long-term rate that is more volatile than in the first sub-sample but much smoother than in the other two previous sub-samples. Similar conclusions can be drawn looking at the evolution of the short-term rate and the long-short spread. Notice that

[7] Notice that if R_t and r_t are cointegrated the α_1-fundamental and the α_2-fundamental solutions collapse in a single fundamental solution since the second-order characteristic polynomial in π_1 given by the second equation in (7) becomes a first-order polynomial that implies $\pi_1 = 1$.

[8] The 1-month Treasury bill rates are shown on a discount basis whereas the Treasury 20-year yields are shown on a bond yield basis. In order to get the appropriate bond yield associated with the 1-month Treasury bill rate we use the Conversion Table for issues Quoted on a Discount Basis, displayed in Salomon Brothers' Analytical Record of Yields and Yield Spreads. Thus, by adding the appropriate percentage shown in the Conversion Table to the discount yield, we obtain the 1-month Treasury bill rate on a bond yield basis.

the above split of the sample delivers four sub-samples each of which roughly coincides with the term of o¢ce of a Federal Reserve chairman.[9]

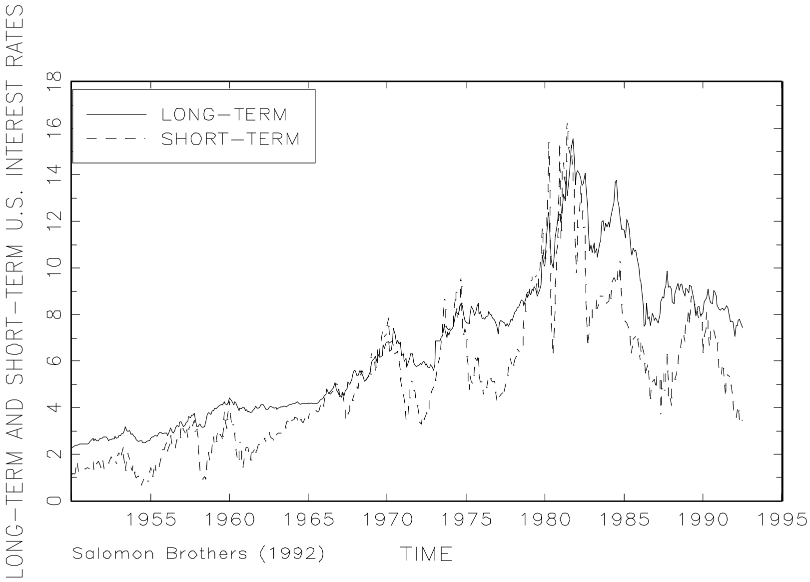

FIGURE 1: POSTWAR U.S. INTEREST RATES

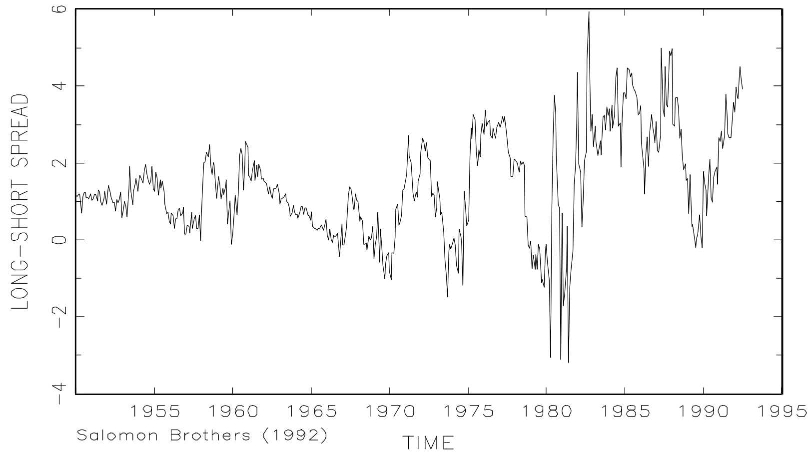

FIGURE 2: SPREAD OF POSTWAR U.S. INTESREST RATES

[9] There have been …ve Federal Reserve chairmen since 1951: Martin, Burns, Miller, Volcker and Greenspan (see Thornton (1996) for more details). The …rst subsample almost covers the term of o¢ce of Martin (1951:4-1970:1), the second subsample covers the terms of o¢ce of Burns (1970:2-1978:1) and Miller (1978:3-1979:8). The third subsample covers most of Volcker's o¢ce term (1979:8-1987:8). Finally, most of the fourth subsample is in Greenspan o¢ce term (1987:8-present).

3.1 Estimation Procedure

Given the evidence on cointegration between long- and short-term interest rates and the error-correction representation implied by the rational expectations hypothesis of the term structure, as shown in (11), we estimate the model in two steps. First, we estimate an error-correction model in order to summarize the relationship between the variables. Thus, an error-correction representation is the auxiliary model used to capture the empirical regularities displayed by the data. Second, we apply the simulated moments estimator (SME) suggested by Lee and Ingram (1991) and Duffie and Singleton (1993) to estimate the structural parameters of the term structure model.

The SME is a specific type of the generalized method of moments (GMM) estimator using time series data sets, which makes use of a set of statistics computed from the data set used and from a number of different simulated data sets generated by the model being estimated. Since a sufficient condition for the SME estimator to be consistent and asymptotically normal is that the time series used in the estimation should be covariance stationary, we work with first-differences of long-term and short-term interest rates and with the long-short spread. More specifically, the statistics used to carry out the SME are the coefficients from an error-correction two-variable system formed by first differences of R_t and r_t with eight-lags. To find the appropriate lag of the error-correction representation, the likelihood ratio test was used. The null hypothesis tested was s lags versus $s+1$ lags. The lowest number of lags s associated with the non-rejection of the null was chosen. For the whole sample we find $s = 8$ whereas for the sub-samples 1950:1-1970:6, 1970:6-1979:9, 1979:9-1986:4 and 1986:4-1992:7, we find s equal to 6, 3, 1 and 0, respectively.[10]

To implement the method, we construct a p x 1 vector with the coefficients of the error-correction representation obtained from real data, denoted by $H_T(\mu_0)$, where p in this application is 37,[11] T denotes the length of the time series data, and μ is a k x 1 vector whose components are the structural parameters of the model being estimated. The true parameter

[10] These differences in the lag lengths observed for different subsamples can be taken as rough evidence of regime switching in the short-term rate process and in the term structure. Moreover, as shown by Hosking (1981), a portmanteau test for checking whether residuals from the error-correction model are white noise may be viewed as a sequence of Lagrange ratio tests for zero restrictions on the coefficients of the error-correction representation as the one carried out in this paper.

[11] We have 32 coefficients from an eight-lag, two variable system, two coefficients for the error correction term since the cointegration vector is known, and three more coefficients from the non-redundant elements of the covariance matrix of the residuals. Notice that p becomes 29, 17, 9 and 5 for the first, second, third and fourth subsamples, respectively.

values are denoted by θ_0. In our model, the structural parameters are $\theta = (\rho_1; \lambda_1; \sigma_z; \sigma_v; \delta)$. Recall that the cointegration restriction imposes that $\rho_2 = 1 - \rho_1$ and $\pi_1 = 1$.

Given that the real data are by assumption a realization of a stochastic process, we decrease the randomness in the estimator by simulating the model n times. Time series of simulated data are recursively obtained using the matrix system (11) imposing the cointegration restriction $\rho_2 = 1 - \rho_1$. Since we estimate the model many times (we analyze two alternative rational expectations solutions and five data sets: the whole sample and four sub-samples), as a compromise, we make $n = 5$ in this application. For each simulation a $p \times 1$ vector of error-correction model coefficients, denoted by $H_{Ni}(\theta)$, is obtained from the simulated time series of R_t and r_t generated from the model being estimated, where $N = nT$ is the length of the simulated data. Averaging the n realizations of the simulated error-correction model coefficients, i.e., $H_N(\theta) = \frac{1}{n}\sum_{i=1}^{n} H_{Ni}(\theta)$, we obtain a measure of the expected value of the simulated coefficients of the error-correction representation, $E(H_{Ni}(\theta))$. To generate simulated values of ΔR_t, Δr_t and $(R_t - r_t)$ we need the starting values of the first-differences of long- and short-term rates, and the long-short spread. In the estimation, we have arbitrarily compute these starting values using the first two observed values of long-term and short-term rates in our sample. For the SME estimator to be consistent, the initial values must have been drawn from a stationary distribution. In practice, to avoid the influence of the starting values we follow Lee and Ingram's suggestion of generating a realization from the stochastic processes of ΔR_t and Δr_t of length $2N$, discard the first N-simulated observations, and use only the remaining N observations to carry out the estimation. After N observations have been simulated, the influence of the initial conditions must have disappeared.

The SME estimator of θ_0 is obtained from the minimization of a distance function of error-correction representation coefficients from real and simulated data. Formally,

$$\min_{\theta} J_T = [H_T(\theta_0) - H_N(\theta)]' W [H_T(\theta_0) - H_N(\theta)];$$

where the weighting matrix W^{-1} is the covariance matrix of $H_T(\theta_0)$:

Denoting the solution of the minimization problem by $\hat{\theta}$, i.e., the SME estimator, Lee and Ingram (1991) and Duffie and Singleton (1993) prove the following results:

$$\sqrt{T}(\hat{\theta} - \theta_0) \rightarrow N[0; (B'WB)^{-1}];$$

$$T J_T \rightarrow \chi^2(p - k);$$

where B is a full rank matrix given by $B = E(\frac{\partial H_N(\theta)}{\partial \theta})$. For small values of n the variance of the estimated parameter vector is $(1+\frac{1}{n})(B'WB)^{-1}$; and the statistic in the latter expression should be $(1+\frac{1}{n})TJ_T$.

The objective function J_T was minimized using the optimization package OPTMUM programmed in GAUSS language. The Broyden-Fletcher-Glodfard-Shanno algorithm was applied. To compute the covariance matrix we need to obtain B. Computation of B requires two steps. First, obtaining the numerical first derivatives of the coefficients of the error-correction model with respect to the estimates of the structural parameters θ for each of the n simulations. Second, averaging the n-numerical first derivatives to get B.

3.2 Empirical Results

Table 1 shows the estimation results for the whole sample (from January 1950 to July 1992). The estimated value of the feedback parameter ρ_1 is small but statistically significant. The estimated value of the parameter characterizing the term premium process λ_1 shows that we can reject the possibility that the term premium follows a random walk. Moreover, the value of the goodness-of-fit statistic $(1+\frac{1}{n})TJ_T = 168.755$, which is distributed as a $\chi^2(32)$ for the whole sample, shows that the cross-equation restrictions imposed by the rational expectations model of the term structure are rejected by the data at standard critical values.

As shown in Figure 1, the changing behavior of both long-term and short-term rates indicates the possibility of different regimes characterizing the term structure of interest rates during the post-war period. In order to investigate this hypothesis we split the sample into four sub-samples. As was discussed above, each sub-sample roughly coincides with the office term of a Federal Reserve chairman.

Table 2 shows the estimation results for the four sub-samples considered. The goodness-of-fit statistic $(1+\frac{1}{n})TJ_T$ is distributed as a $\chi^2(24)$, $\chi^2(12)$, and $\chi^2(4)$, for the first, second and third sub-samples, respectively. These statistics show that the data for the second sub-sample reject the cross-equation restrictions imposed by the rational expectations model of the term structure at the 5% critical value but not at the 1%. For the fourth subsample the number of lags of the error-correction representation that best fit the data is zero as suggested by the error-correction representation implied by the rational expectations hypothesis of term structure model.[12] Thus, the model fits well the data for this subsample. However, for the first and third

[12] For this subsample, the model is then exactly identified since the coefficients of the error-correction representation are five, which is equal to the number of structural parameters being estimated.

sub-samples the data clearly reject the model at any conventional critical value.

Table 1. SME estimation results for the whole sample

Period	1950:1-1992:7
$(1 + \frac{1}{n})TJ_T$	168:75524
½$_1$	0:08259 (0:01707)
¾$_z$	0:25349 (0:00179)
¾$_v$	0:66922 (0:00270)
±	0:98698 (0:00266)
¸$_1$	0:92980 (0:03771)

Table 2. SME estimates of the fundamental solution for the alternative sub-samples

Period	1950:1-1970:6	1970:6-1979:9	1979:9-1986:4	1986:4-1992:7
$(1 + \frac{1}{n})TJ_T$	123:65399	24:36343	34:17034	0:00000
½$_1$	0:12898 (0:01913)	0:06422 (0:03195)	0:15195 (0:05925)	0:15040 (0:03340)
¾$_z$	0:10896 (0:00095)	0:18217 (0:00438)	0:48675 (0:04890)	0:29222 (0:01994)
¾$_v$	0:30424 (0:00153)	0:56145 (0:00747)	1:33725 (0:02069)	0:67778 (0:00955)
±	0:99130 (0:00132)	0:98881 (0:00551)	0:96651 (0:01290)	0:98717 (0:00332)
¸$_1$	0:84062 (0:06078)	0:74247 (0:13204)	1:0 (0:25969)	0:83951 (0:24623)

The different estimation results of the parameters characterizing the short-term rates obtained for the alternative sub-samples shows the presence of regime changes in the short-term rate process. Thus, the feedback parameter ρ_1 is statistically different from zero at standard levels for all sub-samples and its estimated value is different depending on which sub-sample is being estimated. The estimated value of σ_v varies widely depending on which sub-sample is being estimated, and is positively related to the short-term rate volatility observed in each sub-sample.

Table 2 also shows that the estimated value of the discount factor parameter δ is reasonable (around 0.985) and quite similar for all the sub-samples considered. For the first and second sub-samples, the estimated value of the persistence parameter characterizing the term premium process λ_1 is statistically lower than one, showing a high degree of persistence. This high degree of persistence should be viewed as good news for the rational expectations model of the term structure because, as pointed out by Shiller (1979), the term premium is usually described as reflecting public attitudes toward and perceptions of risk and is usually modeled as slow moving or assumed constant.The estimated value of λ_1 is close to one for the third sub-sample, showing that the term premium follows a random walk. The presence of a unit root characterizing the term premium implies that the long-term rate movements are not well accounted for by the rational expectations model of the term structure. For the fourth subsample, the estimated value of λ_1 is not statistically different from one, but this result may be due to the relatively large standard deviation for this estimated value. We can alternatively test whether the term premium u_t is an I(1) process as follows. First, by imposing the cointegration restriction the second equation in (11) can be written as follows

$$\Delta r_t = \rho_1(R_{t-1} - r_{t-1}) + v_t.$$

Since Δr_t and $R_{t-1} - r_{t-1}$ are stationary, it follows that v_t is stationary. Moreover, by using the estimated value of ρ_1 in this subsample we can obtain that $\hat{v}_t = \Delta r_t - \hat{\rho}_1(R_{t-1} - r_{t-1})$. A value of the Phillips-Perron Z_ρ statistic of -87.93 shows strong evidence that $\hat{v}_t$ is not an I(1) process. Second, by imposing the cointegration restriction and after some small algebra, the first equation in (11) can be written as follows

$$\Delta R_t = (\rho_1 - 1)(R_{t-1} - r_{t-1}) + \frac{1}{1 - \rho_1\delta}u_t + v_t.$$

Since v_t is stationary we can then test whether u_t is an I(1) process by testing whether the variable $\Delta R_t + (1 - \hat{\rho}_1)(R_{t-1} - r_{t-1})$ is I(1). The Phillips-Perron Z_ρ statistic for this variable is -11.80. Thus, the null hypothesis that the term premium u_t is a unit root process is not rejected at the 5% critical value, but it is rejected at the 10% critical value.

Andrews and Fair (1988) suggested a Wald test of the null hypothesis that $\theta_1 = \theta_2$, where θ_i is the parameter vector θ that characterizes a particular sub-sample i of size T_i, for $i = 1; 2$. Let ζ_T be defined by

$$\zeta_T = T(\hat{\theta}_1 - \hat{\theta}_2)'[\pi^{-1}\hat{V}_1 + (1-\pi)^{-1}\hat{V}_2]^{-1}(\hat{\theta}_1 - \hat{\theta}_2);$$

where $T = T_1 + T_2$, $\pi = \frac{T_1}{T_1+T_2}$, $\hat{V}_i$ is the estimated covariance matrix of $\hat{\theta}_i$ for $i = 1; 2$. Andrews and Fair show that $\zeta_T \rightarrow \chi^2(k)$ under the null hypothesis that $\theta_1 = \theta_2$. The values of this statistic when testing parameter stability between first and second sub-samples, first and third sub-samples, first and fourth sub-samples, second and third sub-samples, second and fourth sub-samples, and third and fourth sub-samples are 2013.8, 3396.2, 2497.1, 1514.7, 290.5 and 822.1, respectively. These test results imply overwhelming rejection of the parameter stability hypothesis between alternative sub-samples, and provide additional evidence of regime changes in the term structure of interest rates during the post-war period.

3.3 Robustness of the estimation results

In the remainder of this section we assess the robustness of the estimation results displayed in Table 2 along several dimensions. First, we also estimate the model by allowing for serial correlation in the perturbation in the short-term rate process v_t. In particular, we assume that $v_t = \tau_1 v_{t-1} + s_t$, where s_t is an i.i.d. random variable with mean zero and variance σ_s^2. One would expect that omitting significant lagged values of R_t and r_t other than those appearing in (3) would result in a significant estimated value for τ_1. Table 3 shows that τ_1 is not statistically significant for any sub-sample studied indicating that including more lagged values of R_t and r_t other than those considered in (3) is not required. Moreover, the estimation results in Table 3 are quite similar to those displayed in Table 2. These results indicate that the introduction of serial correlation in v_t does not add any explanatory power to the term structure model for any of the sub-samples analyzed.

Table 3. SME estimates for the alternative sub-samples allowing for v_t to follow an AR(1) process

Period	1950:1-1970:6	1970:6-1979:9	1979:9-1986:4
$(1+\frac{1}{n})TJ_T$	122:31423	24:26108	34:07455
ρ_1	0:11922	0:06729	0:16513
	(0:02028)	(0:03573)	(0:07721)
σ_z	0:10906	0:18206	0:48530
	(0:00095)	(0:00440)	(0:04834)
σ_s	0:30422	0:56146	1:33725
	(0:00153)	(0:00747)	(0:02066)
δ	0:99195	0:98828	0:96371
	(0:00139)	(0:00616)	(0:01664)
λ_1	0:84192	0:74201	0:99999
	(0:06077)	(0:13205)	(0:25464)
τ_1	$-$0:07299	0:03320	0:04183
	(0:07480)	(0:12452)	(0:15250)

Second, the second sub-sample covers the office terms of Burns and Miller. We estimate the term structure model for the period (1970:7-1978:1) that covers most of Burns' office term in order to assess whether a term structure change occurred in the beginning of Miller's office term. The estimation results for this sub-sample are shown in Table 4. Comparing the estimation results in Table 4 with those displayed in Table 2, we see that they are quite similar, indicating that the term structure behavior during Burns' office term was similar to the term structure behavior observed taking into account the two office terms together. It would be also interesting to estimate the model using only data from Miller's office term to compare the estimation results with those obtained with the Burns sub-sample. However, the few observations available (only eighteen) prevent us from doing so.

Finally, as stated in Section 2, the term structure model exhibits another rational expectations solution, called the backward solution (10). We also estimate the model under the backward solution imposing the cointegration restriction. Table 5 shows the estimation results for this solution.[13] Comparing the goodness-of-fit statistics $(1+\frac{1}{n})TJ_T$ obtained with this solution for the different sub-samples with the goodness-of-fit statistics obtained with the fundamental solution (11) displayed in Table 2, we observe that the fundamental solution fits the data much better than the backward solution for all

[13] The results shown in Table 5 are qualitatively unchanged by varying the number of lags of R_t and r_t considered in the short-term rate process (3).

sub-samples studied. These results are important because some researchers (for instance, Chow (1989)) have rejected the rational expectations hypothesis of the term structure by analyzing only the backward solution. Our estimation results point out that tests of the term structure of interest rates based on the backward solution should be viewed at least with caution.

Table 4. SME estimation results for Burns' sub-sample

Period	1970:6-1978:1
$(1 + \frac{1}{n})TJ_T$	24:60644
½$_1$	0:07941 (0:03187)
¾$_z$	0:18742 (0:00502)
¾$_v$	0:57082 (0:00918)
±	0:98444 (0:00616)
¸$_1$	0:74946 (0:14409)

Moreover, the estimation of the term structure model under the backward solution (10) is also of interest because this solution characterizes the only rational expectations equilibrium, being immune to the Lucas Critique since changes in the (policy) parameters describing the short-term rate process involve changes in the reduced form parameters characterizing the fundamental solution (11). However, the backward solution (10) is invariant to those (policy) changes. Immunity to the Lucas Critique was suggested by Farmer (1991) as a selection criterion in a context of multiple equilibria.[14] Our estimation results point out that the data do not support the hypothesis that the term structure of interest rates is characterized by an equilibrium that is immune to the Lucas Critique (that is, solution (10)). However, the

[14] The idea underlying Farmer's paper is that in a model in which there are multiple rational expectations equilibria agents may …nd it useful to coordinate in a unique rational expectations equilibrium. This equilibrium can be supported by a self-ful…lling forecast rule having the property of being independent of the parameters characterizing the probability distribution of the fundamentals of the economy. In particular, the forecast rule might be independent of the parameters governing the process of economic policy.

data support equilibrium solution (11), characterized by reduced form parameters that vary with changes in the parameters describing the short-term rate process.

Table 5. SME estimates of the backward solution for the alternative sub-samples

Period	1950:1-1970:6	1970:6-1979:9	1979:9-1986:4	1986:4-1992:7
$(1 + \frac{1}{n})TJ_T$	999:27975	1801:84340	614:84633	44:97073
½1	0:14275 (0:02796)	0:09189 (0:03258)	0:24694 (0:06176)	0:16762 (0:07451)
¾z	0:11135 (0:00149)	0:15243 (0:00395)	0:47133 (0:01400)	0:30170 (0:31282)
¾v	0:28671 (0:00217)	0:36864 (0:00904)	1:05940 (0:01951)	0:66265 (0:00920)
±	1:0 (0:01104)	0:99999 (0:01278)	1:0 (0:02736)	1:0 (0:23730)
¸1	¡ 0:03580 (0:11113)	¡ 0:04487 (0:10157)	0:02494 (0:11718)	¡ 0:13446 (5:94267)

4 CONCLUSIONS

This paper shows evidence that the rational expectations model of the term structure is supported by U.S. data over two long periods. One period is the seventies and the other period lasts from the mid-eighties to at least the end of the sample. However, for the …fties, sixties and Volcker's o¢ce term (from September 1979 to April 1986) the term structure model is rejected by the data.

Moreover, we …nd evidence of regime changes in the short-term rate process and the term structure of interest rates. These regime changes roughly coincide with changes in the Federal Reserve chairman. Therefore, the switches in monetary policy taking place when the chairmanship of the Federal Reserve changes seem to play a major role in the appearance and disappearance of the term structure of interest rates implied by the rational expectations hypothesis. The evidence provided by Mankiw and Miron (1986) when examining the e¤ects of the establishment of the Federal Reserve in 1914 and the evidence reported in this paper during the U.S. post-war suggest a causal relation between institutional changes and the behavior of the term structure of interest rates.

Many papers (for instance, Antoncic (1986) and Huizinga and Mishkin (1986)) have studied the prominent change in Fed operating procedure in October 1979 that coincided with the beginning of Volcker's office term. However, the literature has not paid too much attention to the analysis of how changes in the Federal Reserve chairmanship influence monetary policy. An exception is Thornton (1996), which studies the discount rate policies of all Federal Reserve chairmen during U.S. post-war. An interesting line for future research is to investigate what aspects of monetary policy have been introduced after changes in the Federal Reserve chairman that have induced switches in the term structure of interest rates. We believe that this knowledge of past monetary policies would be helpful for a better understanding of the changes in regime observed in the term structure of interest rates and, perhaps more importantly, the macroeconomic implications of those policies. Moreover, a monetary policy involving a stable (not necessarily smooth) short-term interest rate process seems desirable, since it would help to consolidate a stable term structure of interest rates that would allow monetary authorities to assess with more confidence the effects of monetary policy on the aggregate variables of the economy.

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APPENDIX

Table A.1 Phillips-Perron $Z_{½}$ tests

Period	Variable	With Trend	Without Trend
	R_t	¡ 6:27	¡ 3:22
Sample	r_t	¡ 21:38	¡ 11:76
1950:1-1992:7	¢ R_t	¡ 417:02	¡ 417:05
	¢ r_t	¡ 512:45	¡ 512:69
	R_t ¡ r_t	¡ 62:19	¡ 50:82
	R_t	¡ 6:27	¡ 3:22
Subsample	r_t	¡ 21:38	¡ 11:76
1950:1-1970:6	¢ R_t	¡ 230:53	¡ 231:01
	¢ r_t	¡ 249:04	¡ 249:54
	R_t ¡ r_t	¡ 37:72	¡ 26:26
	R_t	¡ 8:40	¡ 5:99
Subsample	r_t	¡ 24:73	¡ 21:54
1970:6-1979:9	¢ R_t	¡ 105:99	¡ 105:87
	¢ r_t	¡ 107:09	¡ 107:33
	R_t ¡ r_t	¡ 7:99	¡ 8:35
	R_t	¡ 5:63	¡ 5:39
Subsample	r_t	¡ 17:85	¡ 17:69
1979:9-1986:4	¢ R_t	¡ 57:62	¡ 55:99
	¢ r_t	¡ 75:00	¡ 75:03
	R_t ¡ r_t	¡ 27:73	¡ 15:15
	R_t	¡ 6:89	¡ 4:90
Subsample	r_t	¡ 19:32	¡ 15:50
1986:4-1992:7	¢ R_t	¡ 72:59	¡ 72:94
	¢ r_t	¡ 81:22	¡ 82:81
	R_t ¡ r_t	¡ 9:05	¡ 9:00

Note: The Phillips-Perron $Z_{½}$ statistics are corrected for fourth-order serial correlation. The results are qualitatively similar to those obtained when considering Phillips-Perron $Z_{t_{½}}$ and augmented Dickey-Fuller tests, or when considering alternative orders of the serial correlation correction in computing Phillips-Perron statistics. For a sample size of 500 observations, the critical values for the Phillips-Perron $Z_{½}$ test are: with trend: 10%, -18.1; 5%, -21.5; 1%, -28.9; without trend: 10%, -11.2; 5%, -14.0; 1%, -20.5. A table displaying the critical values for the Phillips-Perron $Z_{½}$ test is reported in Fuller (1976, p. 371).