Testing for Segmentation in the Term Structure: Operation Twist Revisited

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While the expectations hypothesis has been investigated extensively in a score of studies, the segmentation hypothesis remains largely unexplored. Using the recent inference procedures in cointegration analysis, this paper investigates the segmentation hypothesis for the entire term structure by restricting the cointegrating relation between yields and volumes. Our results indicate that a change in the relative availability of one security impacts its yield spread and provides the Treasury with an opportunity to twist the yield curve. This suggests that investors are bound by asset and liability constraints that restrict them to a specific maturity.

Background

What determines the shape of the term structure? Several theories have been advanced to explain the observed shape of the yield curve: the expectations theory with its various forms, the liquidity preference theory, the preferred habitat theory, and the market segmentation theory.

The expectations theory posits that the yield curve is influenced primarily by the market's expectations about future short-term rates. In the traditional unbiased expectations hypothesis, risk-neutral investors treat bonds with different maturities as perfect substitutes, making the long rate an unbiased average of current and expected shorter rates.[1]

The liquidity preference theory contends that the shape of the term structure is affected by the presence of premiums primarily at the long end due to investors' preference for lending short and borrowing long. The market's expectations regarding

[1] More realistic versions of the expectations hypothesis allow for risk-averse investors and time-varying premia. These adjustments lead to imperfect substitutability between securities of different maturity.

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future interest rates notwithstanding, the slope of the yield curve is assumed to be generally positive due to a larger premium required for holding longer-term securities.

In the market segmentation hypothesis, agents have preferred habitats and do not treat Treasury securities of various maturities as substitutes. Rather, investors have needs for specific maturities and cannot alter their portfolios to take advantage of a higher yield in one segment. This leaves the equilibrium yield in each market segment determined by the relative supply and demand of securities.

The origins of the term premium theory can be traced to Hicks (1946) who posits that the premium is an increasing function of a bond's maturity. Modigliani and Sutch (1966) reject his argument on the grounds that it is inconceivable for individuals to always prefer to lend short-term. They propose the preferred habitat theory in which the term premia fluctuate with the supply and demand of loanable funds at each maturity. In this setting, individuals abandon their preferred habitats only when enticed by higher expected returns. In 1961, during the Kennedy Administration, the loanable funds theory was put to a major test. *Operation Twist* was a plan established by the Treasury to twist the yield curve and reduce long-term rates by shortening the average maturity of its outstanding debt. The results were disastrous, and the policy quickly was abandoned when the yield curve reacted in the opposite direction from the one intended. Recently, however, there has been a resurgence of interest in the segmentation hypothesis and theoretical belief in the relation between the slope of the term structure and the flow of funds.[2] For example, if more private savings are invested in long-term bond funds, the demand for long-term debt should rise, their yields would fall, and the term structure would flatten. Conversely, if an increasing proportion of the public debt were financed by issuing long-term bonds, their supply and yields would rise, and the term structure would become steeper.

This paper focuses on the market segmentation of Treasury securities over the entire term structure. The theory replaces the assumption of perfect substitutability between securities of different maturities and recognizes that investors have preferred habitats dictated by saving and investment decisions. The shape of the yield curve is influenced primarily by asset/liability management constraints of borrowers and creditors restricted to a specific maturity sector. These restrictions are assumed to hinder their abilities to shift from one maturity to another in order to take advantage of yield differentials due to differences between expected and forward rates.

Our results find ample evidence of causal relationships between Treasury supplies and yields over the term structure. Without directly testing the expectations theory,[3] we find more than one long-term relationship between yields on securities for various terms, a result that some authors argue does not mesh with the expectations hypothesis.

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[2]See the Office of Management and Budget (1993) for a proposal to shorten the average maturity of government debt outstanding.

[3]See Crockett (1998) and Froot (1989) for a summary of scores of recent and old failures of the expectations hypothesis.

The major focus of the paper, however, is examining the relation between a Treasury security yield and relative volume. We identify three common stochastic trends between yields and volumes and present a model for market segmentation. The model is tested, and the results confirm the segmentation hypothesis.

We investigate the cointegrating relation between eight independent interest rates by constructing a vector autoregressive model that we estimate by maximum likelihood. We present an alternative theory for the term structure and construct a model for market segmentation. We test the segmentation hypothesis by restricting the cointegrating relation between yields and volumes using likelihood ratio techniques derived in Johansen and Juselius (1990, 1992). We illustrate the yield response of a security to a supply shock of its own and that of other securities.

The Empirical Literature on the Term Structure

Studies on the theories of the term structure are extensive. Earlier empirical studies include Engel, Lilien, and Robins (1987) who provide an Arch-M model that allows for the conditional variance to affect the excess holding yield on long-term bonds. The novel part of their paper is the introduction of a time-varying premium on securities of different maturities. The assumption of nonconstant heteroskedasticity in the disturbances was prompted by the findings of Shiller (1979), Shiller, Campbell, and Shoenholtz (1983), Mankiw and Summers (1984), and Campbell and Shiller (1987) who suggest that the expectations hypothesis fails to explain the size of the variance of long-term rates.

Bing-Song Lee (1989) elaborates on the model of Engel, Lilien, and Robins (1987) by allowing for time-varying risk premia and conditional heteroskedasticity. The results suggest that allowing for conditional heteroskedasticity improves the results significantly more than the introduction of only time-varying premia.

Bradley and Lumpkin (1992) show that rates for Treasury securities are cointegrated with unit roots. This suggests that interest rates are bound by a long-term equilibrium[4] and not likely to deviate permanently and significantly.

Simon (1991) finds evidence of segmentation in the cash management bill market. Simon uses an ANOVA model to relate maturity, size, and length of when-issued period to the interest rate differential between cash management bills and adjacent maturity bills. His results indicate that unexpected increases in Treasury bill supply have a significant effect on interest rates, with segmentation increasing at the short end. Further evidence of segmentation at the short end of the term structure is provided in Simon (1994) where the yields and supplies of Treasury securities are examined. Using regression analysis, his study considers pairs for a maximum maturity of 26 weeks. While no rigorous causality tests are undertaken, the findings suggest that the yield

[4]Bradley and Lumpkin (1992, p. 455) argue that the cointegration relation they find is not necessarily unique.

spreads between 13-week bills and 12-week bills are correlated with their differential supplies.

Taylor (1992) tests a variety of models for the U.K. term structure. His findings reject the rational expectations hypothesis. An effort to allow for time-varying premia using a GARCH model also fails to explain the risk premium model. His results, however, seem to support the segmentation theory. Taylor suggests that the U.K. government efforts to reduce the relative amount of debt outstanding explain the excess holding return.

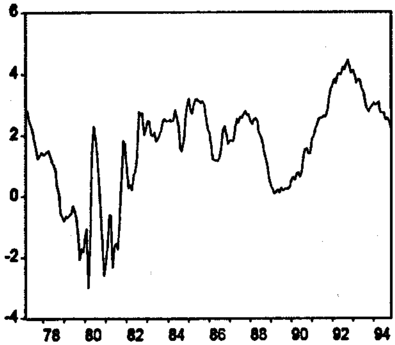
Hall *et al.* (1992) find that 11 US Treasury bill rates form a cointegrated system. Two key features in their study are noteworthy. First, they find that when the Fed was targeting interest rates, the cointegration results support the expectations hypothesis. Therefore, they argue that it is reasonable to treat the interest rate spread as a basis for the cointegrating vector. The cointegrating relation breaks down, however when the Fed switches targets and adopts the growth in bank reserves. Engsted and Tangaard (1994) report similar failures of the cointegrating relations.

Current research can be grouped into two broad categories. First are the studies that overlook the supply effects and focus exclusively on the relation between yields. These studies lack an alternative explanation to the theory of the term structure during the time periods when they note a failure of the expectations hypothesis. The models in the second category account for the correlation between yields and volumes but focus on one individual segment of the term structure and ignore the broader interaction with other portions of the yield curve.

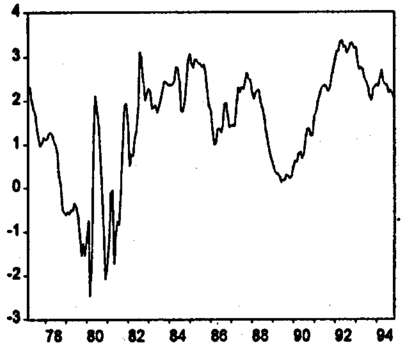
This paper complements the existing literature on the term structure by investigating the segmentation hypothesis. We test for causality between security volumes and their corresponding yields and yield spreads, and we build a vector autoregression model that we estimate by maximum likelihood. We identify three common stochastic trends among eight interest rates ranging in maturity from three months to 30 years. Our investigation of the relationship between the volume of securities of different maturities and their respective yields reveals that an increase in the relative volume of long-term securities impacts the spread between short- and long-term yields and provides an opportunity for the Treasury to manipulate the yield curve.

Our data consist of constant maturity yields of eight Treasury securities (three- and six-month, one-, three-, five-, seven-, ten-, and 30-years) from March 1977 through December 1994. The eight interest rates are matched with the supply of securities in five segments of the yield curve: $3m \leq . < 1y$, $1y \leq . < 5y$, $5y \leq . < 10y$, $10y \leq . < 20y$, $20y \leq . \leq 30y$. The interest data are obtained from *Salomon Brothers' Analytical Record of Yields and Yields Spreads* and represent the nominal yield to maturity. The observed yield on each security is derived from its price on the last trading day of the month. The nominal yields are derived by taking the average bid and ask quotes on the last trading day of the month. The debt volume data are taken from several *Quarterly Reports*.

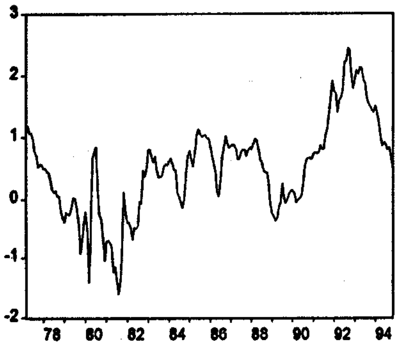
Figures 1 through 6—Yield Spreads of Selected Interest Rates



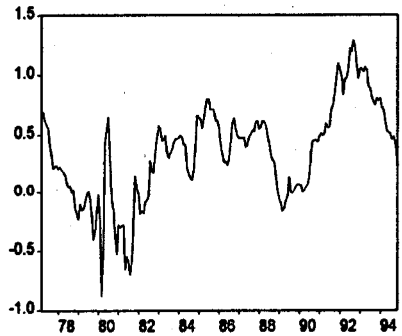
— Y30-Y1



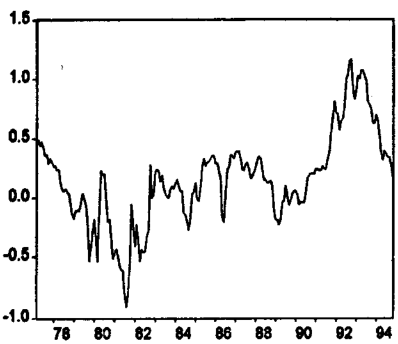
— Y10-Y1 Spread



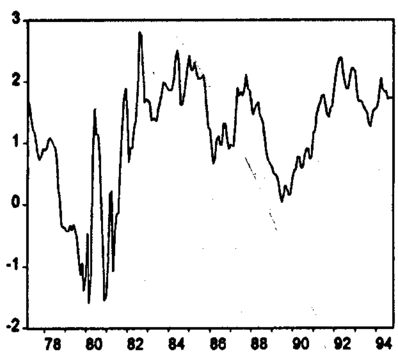
— Y30-Y5 Spread



— Y10-Y5 Spread



— Y30-Y10 Spread



— Y5-Y1 Spread

(1977 through 1996) published by the Treasury Department on the *Maturity Composition of Outstanding Debt.* Our data are monthly; the sample contains 237 observations. We deliberately avoid the period prior to March 1977 because this marked the first time that the Treasury issued the benchmark 30-year bond.

Causal Relations between Treasury Yields and Volumes

In this section, we conduct causality tests between yield spreads and supply changes of Treasury securities. To our knowledge, this is the first time that the entire term structure has been subjected to a similar causal investigation. Figure 1 provides plots of the yield spreads during the time period of the study. The causality tests are conducted for all supply and yield series with respect to the benchmark 30-year bond. Causality tests are run on changes in those variables to ensure stationarity[5]. Formally, let $\Delta Y \equiv Y_{30} - Y_i$ and $\Delta S \equiv S_j - S_{30}$ represent the spread and the volume differential of securities, respectively, where i = 3m, 6m, 1y, 3y, 5y, 7y, 10y, and j represents a grouping variable for the volume of securities with maturity between: 3m-1y, 1y-5y, 5y-10y, and 10y-20y. The Granger causality tests boils down to the question of whether ΔQ is linearly informative about a future ΔY. This would require that the event ΔQ precedes the event ΔY or that the current and past observations of ΔQ help in the forecast of ΔY. The forecast is based on the minimization of the mean squared error of future changes in yields based on current and past changes: $\mathrm{MSE}\{E(\Delta Y_{t+1} \mid \Delta Y_t, \Delta Y_{t-1},...)\}$. If this is not significantly larger than $\mathrm{MSE}\{E(\Delta Y_{t+1} \mid \Delta Y_t, \Delta Y_{t-1},...; \Delta Q_t, \Delta Q_{t-1},...)\}$, we conclude that ΔQ is not linearly informative about ΔY.

To conduct the test, each series is represented as a VAR of lag length p:

$$\Delta Y_t = \beta_0 + \beta_1 \Delta Y_{t-1} + \beta_2 \Delta Y_{t-2} + ... + \beta_p \Delta Y_{t-p} + \mu_1 \Delta S_{t-1} + ... + \mu_p \Delta S_{t-p} + u_t$$

and estimated by least squares. Then an F test is conducted for the null hypothesis:

$$H_0\colon \mu_1 = \mu_2 = ... = \mu_p = 0$$

The null hypothesis is ΔY does not Granger cause ΔQ. Our prediction is that the volume differentials affect the corresponding spreads, while the reverse is not true.

The results of the Granger causality F tests are reported in Table 1. There is a clear directional cause from volume changes to changes in yields, mainly at the short-end of the yield curve. For the other directional cause, from yields to volumes (Y_{30}-Y_i =/=> Q_j-Q_{30}), the null hypothesis of no causation cannot be rejected for virtually all the series (i.e., yields do not cause quantities). We find that differences in supply between the 30y bond and a shorter-term security (1y, 5y) cause changes in yield spreads throughout the yield curve. The results reject: Q_j-Q_{30} =/=> Y_{30}-Y_i for j = 1y, 5y

[5] The choice of the 30-year bond as a benchmark is arbitrary. Another choice would be the other extreme of the yield curve, the 3m Treasury bill. We run the tests using the 3m T-bill and find no significant change in the results.

Table 1—Granger Causality Tests, March 1977 to December 1996, n = 211, lags = 2

H_0: X does not Granger Cause (=/=>) Y

Variable	Variable	F-stat	Variable	Variable	F-stat
Q1-Q30	=/=> Y30-Y1	6.41**	Q1-Q30	=/=> Y30-Y10	3.48*
Q1-Q30	=/=> Y30-Y6m	6.04**	Q1-Q30	=/=> Y30-Y3	5.21**
Q1-Q30	=/=>Y30-Y3m	5.37**	Q1-Q30	=/=> Y30-Y7	3.94*
Q10-Q30	=/=> Y30-Y1	1.19	Q10-Q30	=/=> Y30-Y10	2.42
Q10-Q30	=/=> Y30-Y3m	0.15	Q10-Q30	=/=> Y30-Y3	2.36
Q10-Q30	=/=> Y30-Y6m	0.50	Q10-Q30	=/=> Y30-Y7	4.69**
Q20-Q30	=/=> Y30-Y1	2.24	Q20-Q30	=/=> Y30-Y10	2.88*
Q20-Q30	=/=> Y30-Y3m	1.89	Q20-Q30	=/=> Y30-Y3	2.74
Q20-Q30	=/=> Y30-Y6m	1.83	Q20-Q30	=/=> Y30-Y7	3.30*
Q5-Q30	=/=> Y30-Y1	3.08*	Q5-Q30	=/=> Y30-Y10	2.99*
Q5-Q30	=/=> Y30-Y3m	2.97*	Q5-Q30	=/=> Y30-Y3	3.42*
Q5-Q30	=/=> Y30-Y6m	2.96*	Q5-Q30	=/=> Y30-Y7	3.11*
Y30-Y1	=/=> Q1-Q30	0.04	Y30-Y10	=/=> Q1-Q30	0.05
Y30-Y1	=/=> Q10-Q30	0.97	Y30-Y10	=/=> Q10-Q30	2.50
Y30-Y1	=/=> Q20-Q30	1.97	Y30-Y10	=/=> Q20-Q30	0.96
Y30-Y1	=/=> Q5-Q30	0.21	Y30-Y10	=/=> Q5-Q30	0.08
Y30-Y3m	=/=> Q10-Q30	0.15	Y30-Y3	=/=> Q1-Q30	0.12
Y30-Y3m	=/=> Q20-Q30	1.27	Y30-Y3	=/=> Q10-Q30	1.55
Y30-Y3m	=/=> Q5-Q30	0.37	Y30-Y3	=/=> Q20-Q30	1.76
Y30-Y3m	=/=> Q1-Q30	0.35	Y30-Y3	=/=> Q5-Q30	0.18
Y30-Y6m	=/=> Q1-Q30	0.20	Y30-Y7	=/=> Q1-Q30	0.20
Y30-Y6m	=/=> Q10-Q30	0.37	Y30-Y7	=/=> Q10-Q30	4.05*
Y30-Y6m	=/=> Q20-Q30	1.53	Y30-Y7	=/=> Q20-Q30	2.49
Y30-Y6m	=/=> Q5-Q30	0.32	Y30-Y7	=/=> Q5-Q30	0.09

* Significant at the 5 percent level

** Significant at the 1 percent level

and i = 3m, 6m, 1y, 3y, 7y, 10y, a total of 15 rejections of the hypothesis that supply changes do not Granger cause changes in spreads. The impact of supply changes between long- and medium-term securities (Q_{10}-Q_{30}, Q_{20}-Q_{30}) on yield spreads, however, is mixed. The supply effects may be more pronounced when they involve securities with dissimilar maturities that preclude their treatment as substitutes. Because the tests are sensitive to the lag order, we test several lag lengths. There is no significant change in the results.

Two important caveats are worth noting. First, while the preceding causality tests demonstrate that the volume data can be a useful tool for forecasting interest rates, they do not indicate a functional form or a magnitude for that relation. Moreover, any potential statistical correlation should have a theoretical basis before it is established empirically. In a following section we suggest an alternative framework to the expectations hypothesis and test it empirically to demonstrate its validity.

Second, to suggest that the preceding correlations between volume changes and yield differentials are causal, one must establish that the change in the volume of securities is exogenous and not responding to an independent. One possibility of this occurrence is when an increase in the budget deficit produces a simultaneous increase in securities and interest rates. A more circumspect description is to argue that the volume changes help forecast changes in yield spreads.

Testing for Cointegration and Unit Roots

We begin by examining the stochastic nature of yields and supplies on securities of various maturities. We show that each set of series represents a unit root process that we assume can be written in first difference form as:

$$\Delta y_t = \beta_0 + \beta_1 t + (\alpha - 1) y_{t-1} + u_t$$

where:

β_0 = A drift parameter; and

$\beta_0 t$ = Time trend.

The test for unit root[6] is essentially one that tests $\alpha = 1$.

The test statistics for all variables (reported in Table 2) clearly reject the null hypothesis of unit root in the differences but not in the levels. This leads us to believe that each variable is an I(1) process that becomes stationary after differencing it once. To see whether our results are sensitive to serial correlation in the disturbances, we run the tests with AR(2), AR(3), and AR(4) errors[7] without a change in our results. In the end, we base the model selection on Schwartz's (1978) criterion.[8]

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[6] If the errors are assumed to follow an AR(1) process, $u_t = \rho u_{t-1} + \varepsilon_t$, the model can be rewritten as:

$$\Delta y_t = Z_t \beta \ + (\alpha - 1)(\rho - 1) y_{t-1} + \alpha \rho \Delta y_{t-1} + \varepsilon_t$$

where Z_t combines the drift term and the line trend. The test of $(\alpha\text{-}1)(\rho\text{-}1) = 0$ is the augmented Dickey Fuller ($\hat{\tau}$).

[7] MacKinnon (1990) provides a more complete set of critical values for $\hat{\tau}$. A nonparametric analog to $\hat{\tau}$ that allows for a wide range of serial correlation and heterogeneity patterns is suggested by Phillips (1987a and 1987b) and Phillips and Perron (1988). The critical values for the statistic are presented in Phillips and Ouliaris (1990). The finite sample properties of the statistic, however, are not fully investigated.

[8] Schwartz's (1978) criterion is defined as:

$$\text{BIC} = \log |\hat{\Sigma}| + (p^2 q) T^{-1} \log T$$

where:

p = The dimension of the VAR;

q = The order of the VAR model being tested for q = m against q = m-1;

m = The maximum order considered; and

$\hat{\Sigma}$ = The estimated covariance of the errors.

In effect, BIC includes a penalty adjustment for the number of estimated parameters. Note that Schwartz's (BIC) is similar to Akaike's information criterion (AIC) with the distinction that the former can be shown to be strongly consistent.

Table 2—Unit Root Tests for Interest Rates and Volumes of Treasury Securities*

ADF of the Augmented Dickey Fuller t-Statistic: 1 percent MacKinnon critical values for rejection of hypothesis of a unit root for 208 observations = -4.0053 (5 percent critical value=-3.4325), D(.) is the first time difference

Variable	ADF t-Statistic	Variable	ADF t-Statistic
Y3m	-3.058888	D(Y3m)	-6.362421
Y6m	-3.023224	D(Y6m)	-6.155429
Y1	-2.947030	D(Y1)	-5.787000
Y3	-2.697422	D(Y3)	-5.574008
Y5	-2.591100	D(Y5)	-5.363195
Y7	-2.515532	D(Y7)	-5.261418
Y10	-2.498701	D(Y10)	-5.191406
Y30	-2.469824	D(Y30)	-5.303718
Q1	-1.397448	D(Q1)	-6.158390
Q5	2.297918	D(Q5)	-5.031249
Q10	2.6244115	D(Q10)	-6.254076
Q20	0.672145	D(Q20)	-8.721677
Q30	-2.665527	D(Q30)	-6.169314

*All variables become stationary after taking the first time difference

Given the evidence of the preceding section suggesting the unit root nature of the series, we now examine their cointegrating relationship. Let Y_t represent a $k \times 1$ vector of nonstationary/I(1) time series. A $k \times r$ matrix of cointegrating vectors γ is said to exist if the linear combination $\gamma' Y_t$ is stationary or I(0). Each cointegrating vector suggests a long-term relationship of the series Y(t), namely an equilibrium.

We consider a multivariate model of the form:

(4.1) $Y_t = \mu + A_1 Y_{t-1} + A_2 Y_{t-2} + \ldots + A_q Y_{t-q} + \varepsilon_t, \qquad t = 1,2,\ldots,T$

where:

Y_t = An eight-dimensional vector of yields $Y_t \equiv [y_{3m}, y_{6m}, y_1, y_3, y_5, y_7, y_{10}, y_{30}]'$; and

ε_t = Independent k-dimensional Gaussian error with a fixed covariance matrix.

Johansen (1988) and Johansen and Juselius (1990) develop the procedures to maximize the likelihood function for Y_t.

The λ-trace test is the LR statistic $-2\log[Q(H(r)|H_0)]$ derived under the null hypothesis that there are r cointegrating vectors, H(r), against the alternative of no cointegrating vectors, H_0. The cointegration results among the eight interest rates, simultaneously rather than in pairs using Johansen's full information maximum likelihood, are reported in Table 3.

The lambda-trace and lambda-max statistics indicate the existence of five cointegrating vectors and three common stochastic trends.[9] Generally, with n yields to maturity, it is possible to form n - 1 linearly independent yield spreads. The yields are I(1) cointegrated processes, and the spreads between them represent cointegrating relations. If the expectations hypothesis is valid, this leaves *one* common nonstationary

[9] We reject the hypothesis that there exist at most four but not five cointegrating vectors. We therefore conclude that there exist five cointegrating vectors. The number of common trends is the number of interest variables less the number of cointegrating vectors.

Table 3—Tests of Cointegration of Interest Rates

3m, 6m, 1y, 3y, 5y, 7y, 10y, and 30y, March 1977 to December 1996, lags = 12

				90 Percent Critical Value	
Eigenvalue	L-max	Trace	H0:r	L-max90	Trace90
0.2609	63.78*	227.22*	0	32.709	155.748
0.2525	61.41*	163.44*	1	29.121	123.039
0.1548	35.49*	102.03*	2	25.237	93.918
0.1110	24.82*	66.54	3	21.473	68.681
0.0915	20.24*	41.72	4	17.832	47.208
0.0471	10.18	21.48	5	14.036	29.376
0.0417	8.98	11.30	6	11.499	15.340
0.0109	2.32	2.32	7	3.841	3.841

Likelihood ratio test (-2 ln Q) indicates five (four) cointegrating equations at 5 percent (*) or 1 percent (**)

I(1) factor that represents an element exogenous to the system of yields, as for example, the growth in monetary aggregates that derives from Fed monetary policy. The evidence in support of a unique common factor is found in several studies (Engle and Granger, 1987; Stock and Watson, 1988).

A noted exception is Hall *et al.* (1992) who examine 12 different interest rates and find there exists at least two nonstationary common factors.[10] Our findings of one additional common factor may be due to the wider scope of interest rates we investigate, particularly because the data of Hall *et al.* (1992) are limited to the short-end of the yield curve. To summarize, our results seem to lend little support to the expectations hypothesis and more to the findings of Hall *et al.* (1992) and, more recently, Zhang (1993), as opposed to Engle and Granger (1987) or Stock and Watson (1988). The final determination of the number of cointegrating vectors should not hinge on formal testing, however, but should include graphs depicting the cointegrating relations[11] and should provide an economic interpretation of the results.

A Segmentation Model

Next, we turn our investigation to the segmentation hypothesis. We propose a simple test of the segmentation theory to determine whether the yield differential between two non-identical securities is explained by their relative availability in each market segment. Under the market segmentation hypothesis, the relative proportion of one security should be statistically and economically significant in explaining the yield spread. We propose the following theoretical model:

(5.1) $y_{30t} - y_{1t} = f(S_{30t} / S_{1t}) + \varepsilon_t$

where the left side is the slope in a segment of the term structure and the right side is the relative volume of debt instruments i and j at time t. The model can be rewritten as:

(5.2) $Y_{it} = \delta_0 Y_{jt} + \delta_1 \log Q_{it} + \delta_2 \log Q_{jt}$

10 Hall *et al.* (1992) argue that the additional nonstationary common factor is attributed to a change in the Fed monetary policy.

11 For further elaboration on this point, see Johansen and Juselius (1992, p. 221).

where we expect $\delta_0 > 0, \delta_1 < 0$. The first inequality allows for various forms of spreads ($Y_{it} - \delta_0 Y_{jt}$), while the second inequality stems from the belief that increases in the supply of one security raises its yield. Furthermore, if segmentation holds, we expect $\delta_2 = -\delta_3$. With the proper normalization, expression (5.2) is equivalent to: $Y_{it} = \beta_0 Y_{jt} + \beta_1 \log(Q_{it}/Q_{jt})$. The model in (5.2) cannot be estimated using standard regression techniques, however, because the variables are nonstationary. To confirm the structure of (5.2), we rewrite the segmentation model in vector error-correction form where, instead of Y_t in (4.1), we now define a new variable $Z_t \equiv [Y_i, Y_j, Q_i, Q_j]'$ consisting of a four-dimensional vector of Treasury yields and volumes for i, j = 1y, 5y, 10y, and 30y, $i \neq j$. In total we have six segmentation relations. Again, the maximum likelihood estimation procedures developed by Johansen (1988) and Johansen and Juselius (1990) are applied. For testing the restrictions on the cointegrating vectors, we follow the methods of Johansen (1991) and Johansen and Juselius (1992). Our objectives are twofold. First, we will demonstrate that the individual elements of Z_t are tied together by a long-term (mean reversion) relation defined in a cointegrating vector. Second, the results will show that the cointegrating vector is represented by the segmentation relation of the two pairs of yields and volumes defined in Z_t.

To test the linear restrictions on the cointegrating vectors, we define an $r \times s$ matrix

$$H^* = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & -1 \end{bmatrix}$$

such that $H^{*\prime} Z$ affects only two of the four cointegrating coefficients. The H^* matrix represents the hypothesis that increases in the relative volume of one security raises its yield spread and is essentially equivalent to the test of $\delta_2 = -\delta_3$ in model (5.2).

To form the likelihood ratio test for the segmentation model, we follow the same methodology[12] as earlier.

Table 4 shows the existence of two cointegrating vectors for each of the six segmentations relations. In the same table, we provide the likelihood ratio for the test hypothesis of r cointegrating vectors against the alternative of no cointegrating vector. All six models have at least two cointegrating relations at 5 percent significance (critical value of 5.99 for two degrees of freedom). This establishes the first objective of this section. Table 5 suggests evidence of segmentation for several parts of the term structure. We report whether the segmentation relation is contained in the cointegrating space. Overall, these likelihood ratio tests confirm the expected sign of the parameters

[12] We compare the eigenvalues obtained from the restricted model ($\tilde{\lambda}_i$) with those obtained from the unrestricted model (λ_i) for r = 2 cointegrating vectors to obtain an expression of the form

$$-2\log[Q(H^*|H(r))] = T\sum_{i=1}^{r}\ln\left[\frac{(1-\hat{\lambda})}{(1-\tilde{\lambda})}\right].$$

Table 4—Tests for Unrestricted Cointegrating Coefficients Π, β', and α

Panel A: Z = (Y30 Y1 Q30 Q1)

Eigenvalue	L-max	L-Trace	
0.1451	33.08*	66.84*	
0.1004	22.33*	33.76*	
0.0520	11.26	11.44	
0.0008	0.18	0.18	
Ho: r = 2 accepted			
Π Y30	Y1	Q30	Q1
-0.104	0.031	-0.125	0.207
0.021	-0.088	-0.129	-0.193
-0.001	0.001	-0.007	0.008
0.003	0.001	-0.000	-0.008
β' Y30	Y1	Q30	Q1
1.375	-0.634	0.424	-3.452
-0.660	1.007	0.376	2.302
1.378	-0.834	2.253	-3.310
-1.001	0.692	-6.943	15.071
α Y30	Y1	Q30	Q1
-0.048	-0.036	-0.046	-0.002
0.003	-0.124	-0.050	-0.004
0.001	-0.000	-0.002	0.000
0.005	0.002	-0.002	-0.000

Panel B: Z = (Y10 Y1 Q10 Q1)

Eigenvalue	L-max	L-Trace	
0.1852	43.21*	79.11*	
0.1075	24.00*	35.89*	
0.0464	10.02	11.90	
0.0089	1.88	1.88	
Ho: r = 2 accepted			
Π Y10	Y1	Q10	Q1
-0.098	0.027	-0.023	-0.103
0.010	-0.066	-0.025	-0.271
0.014	-0.004	-0.008	0.002
-0.004	0.005	-0.016	0.029
β' Y10	Y1	Q10	Q1
1.140	-0.276	-0.778	1.115
1.784	-1.572	2.347	-4.942
-0.041	-0.331	-0.546	-2.147
0.065	-0.180	-6.003	8.959
α Y10	Y1	Q10	Q1
-0.053	-0.020	0.057	-0.002
-0.031	0.027	0.091	0.011
0.014	-0.001	0.003	-0.001
0.002	-0.003	-0.000	0.001

Table 4 (cont.)—Tests for Unrestricted Cointegrating Coefficients Π, β', and α

Panel C: Z = (Y5 Y1 Q5 Q1)

Eigenvalue	L-max	L-Trace	
0.1187	26.67*	63.16*	
0.0952	21.10*	36.49*	
0.0540	11.71*	15.39*	
0.0173	3.68	3.68	
Ho: r = 3 accepted			
Π Y5	Y1	Q5	Q1
-0.171	0.087	-0.914	1.036
-0.054	-0.032	-0.330	0.103
-0.003	0.002	-0.089	0.118
-0.001	0.003	0.007	-0.005
β' Y5	Y1	Q5	Q1
-1.553	1.591	-14.950	21.459
2.476	-1.614	11.730	−14.867
0.433	-0.896	-17.389	21.231
0.688	-0.629	-0.397	-2.065
α Y5	Y1	Q5	Q1
-0.024	-0.092	0.011	0.021
-0.075	-0.085	0.026	0.041
0.003	0.000	0.003	0.001
0.003	0.001	-0.002	0.002

Panel D: Z = (Y30 Y5 Q30 Q5)

Eigenvalue	L-max	L-Trace	
0.1255	28.31*	62.87*	
0.0854	18.84*	34.57*	
0.0551	11.96*	15.73*	
0.0177	3.77	3.77	
Ho: r = 3 accepted			
Π Y30	Y5	Q30	Q5
-0.124	0.052	0.170	-0.345
-0.010	-0.057	0.089	-0.292
-0.003	0.003	-0.002	-0.004
0.008	-0.005	0.016	-0.026
β' Y30	Y5	Q30	Q5
-2.078	1.166	-0.015	0.088
3.168	-2.810	-3.407	3.931
-0.265	-0.002	0.857	-3.727
-2.308	2.168	-8.311	13.379
α Y30	Y5	Q30	Q5
0.072	-0.002	0.033	-0.017
0.078	0.036	0.042	-0.021
-0.002	-0.001	0.004	0.001
-0.003	-0.001	0.000	-0.002

Table 4 (cont).—Tests for Unrestricted Cointegrating Coefficients Π, β', and α

Panel E: Z = (Y10 Y5 Q10 Q5)

Eigenvalue	L-max	L-Trace	
0.1863	43.51*	69.01*	
0.0711	15.55*	25.50	
0.0419	9.03	9.94	
0.0043	0.91	0.91	
Ho: r = 2 accepted			
Π Y10	Y5	Q10	Q5
-0.160	0.094	0.127	-0.275
-0.030	-0.024	0.102	-0.265
0.024	-0.012	0.003	-0.006
0.001	-0.000	-0.009	0.013
β' Y10	Y5	Q10	Q5
-3.117	2.131	0.513	-0.773
4.824	-4.353	-2.856	4.103
-5.037	4.972	-0.504	2.570
0.688	-0.559	6.738	-7.313
α Y10	Y5	Q10	Q5
0.047	-0.044	-0.041	-0.006
0.040	-0.036	-0.054	-0.007
-0.014	-0.005	-0.000	-0.000
-0.000	0.001	0.001	-0.001

Panel F: Z = (Y30 Y10 Q30 Q10)

Eigenvalue	L-max	L-Trace	
0.2041	48.17	93.16	
0.1389	31.55	44.99	
0.0480	10.39	13.44	
0.0144	3.05	3.05	
Ho: r = 2 accepted			
Π Y30	Y10	Q30	Q10
-0.174	0.114	0.204	-0.297
-0.073	0.018	0.101	-0.197
-0.008	0.007	-0.046	0.049
0.016	-0.002	0.007	-0.007
β' Y30	Y10	Q30	Q10
1.196	-0.294	-1.746	2.103
-2.146	2.218	8.709	-9.379
-6.403	5.585	-1.647	1.306
3.194	-3.087	4.561	-7.006
α Y30	Y10	Q30	Q10
-0.044	0.007	0.030	0.026
-0.044	-0.010	0.023	0.032
0.002	-0.004	0.003	-0.000
0.016	0.003	0.000	0.002

of the segmentation hypothesis. We find that the market segmentation restriction (MSR) is supported for the pairs {30y, 1y}, {30y, 5y}, {5y, 10y}, {10y, 1y}, and {10y, 5y}. Finally, our results fail to show sufficient evidence of segmentation for the remaining pair {5y, 1y}.

To recapitulate, our findings have a valid economic interpretation in the existing theory of the term structure. Investors are often bound by asset and liability constraints that restrict them to a specific sector. These constraints hinder their ability to shift from one maturity segment to another in order to exploit yield differentials due to differences in relative supply. Therefore, our findings run against the assumption of perfect substitutability between securities of various terms and suggest that agents regard them as complements. A noted exception is the relative volume of the 5y to 1y notes that appear insignificant to their yield spread. We suspect that, given their shorter maturity, market participants treat these two securities as imperfect substitutes and arbitrage away any yield differential due to differences in supply. Overall, our results confirm the findings on causality by providing a theoretical support and a functional form from which the magnitude of the causality, not merely its direction, can be gauged.

Table 6 reports the results of the tests of weak exogeneity. If the tests of the hypothesis $H_0 : \alpha = 0$ is not rejected, we conclude that the independent variable is weakly exogenous to the dependent variable. This suggests that short-term movements in the dependent variables are insensitive to the long-term equilibrating relation binding yield spreads and their relative volumes. The weak exogeneity restriction is rejected for all pairs, a sign that the cointegrating equation helps explain the short-term dynamics of the dependent variable.

The rejection of the hypothesis of weak exogeneity can be interpreted as a confirmation of the causality tests of Table 1 that identify one way causality emanating from Treasury volumes to yield spreads.

Our results also indicate the existence of two cointegrating vectors for each of the six equations represented in model (5.2). These cointegrating relations represent two stabilizing forces that suggest that for each pair of securities the relation between their relative volumes and yield spreads is stationary and does not diverge over the 18 year time period we analyze. The segmentation results also suggest the existence of two common trends. These two nonstationary factors could stem from a change in the Treasury's fiscal policy dictated by exogenous elements such as attempts to reduce the budget deficit or micromanage the yield curve. Any factor has the potential to individually produce nonstationarity in liquidity premia and a breakdown in the cointegrating relationships.

The preceding results should be viewed as an invitation to examine the broader question of whether the Treasury should borrow short-term or long-term, a topic that filled center stage during the early period of the Clinton administration. Supported by new findings,[13] the Office of Management and Budget (1993) advised the Treasury to shorten the average maturity of its outstanding debt at a time when the yield curve was particularly steep. The policy was believed to save approximately \$10 billion in four years. Campbell (1995) also discussed the proposal and argued that the Treasury could reduce the average cost of debt by borrowing at the short end of the yield curve when it

[13] For an earlier discussion on whether an optimal debt policy exists, see Agell *et al.* (1992).

Table 5—Likelihood Ratio Tests for Market Segmentation Restriction (MSR) and Weak Exogeneity Restriction (WER)

Panel A: Z=(Y30 Y1 Q30 Q1)	
MSR Marginally Accepted	LR Test: $\chi^2(2) = 5.75$
WER Rejected	LR Test: $\chi^2(6) = 20.06$
Panel B: Z=(Y10 Y1 Q10 Q1)	
MSR Marginally Accepted	LR Test: $\chi^2(2) = 5.38$
WER Rejected	LR Test: $\chi^2(6) = 40.20$
Panel C: Z=(Y5 Y1 Q5 Q1)	
MSR Rejected	LR Test: $\chi^2(3) = 17.44$
WER Rejected	LR Test: $\chi^2(9) = 33.47$
Panel D: Z=(Y30 Y5 Q30 Q5)	
MSR Accepted	LR Test: $\chi^2(3) = 2.13$
WER Rejected	LR Test: $\chi^2(9) = 24.41$
Panel E: Z=(Y10 Y5 Q10 Q5)	
MSR Accepted	LR Test: $\chi^2(2) = 2.12$
WER Rejected	LR Test: $\chi^2(6) = 37.60$
Panel E: Z=(Y30 Y10 Q30 Q10)	
MSR Accepted	LR Test: $\chi^2(2) = 2.98$
WER Rejected	LR Test: $\chi^2(6) = 51.30$

is steep and at the long end when it is flat. The potential savings of this policy quickly will disappear, however, when a massive shift to one end of the yield curve will depress prices and raise yields thereby negating the policy's raison d'être. But regardless of the potential savings, this shift will alter the relative level of yields and therefore the yield spreads between securities.

Opponents of the new proposal contend that, to the extent that long-term rates are generally more volatile than short-term ones, such a policy would increase the variation of the Treasury's average cost of debt. Because a stable average cost is desirable, they recommend long-term borrowing. Another key factor is the impact of inflation on debt management. Because inflation effects are asymmetric on long- and short-term securities, borrowing long-term has the added advantage of a faster amortization of the real value of outstanding debt and real interest costs.[14] The fact that the Treasury has the potential to manipulate the yield curve is significant, regardless of whether significant interest savings can be exploited.

Summary

This paper investigates the joint and stochastic nature of supplies and yields of Treasury securities for all segments of the term structure. Using monthly data between 1977 and 1994, we show that the yield and supply series are nonstationary. Standard causality tests indicate a one directional impact from changes in supplies and to spreads in yields. An examination of the relationship between eight interest rates, independent of their volumes, suggests the existence of three common trends and five cointegrating

[14] For a discussion on the relation between debt burden and debt management policy, see Missale and Blanchard (1994).

relations. These findings would be difficult to reconcile with the narrow form of the expectations hypothesis, which posits that a unique trend between rates exists.

Our focus however is the segmentation hypothesis. We show that changes in the proportion of long- vs. short-term securities impacts their spread. This suggests that agents have preferred habitats dictated by asset/liability constraints that restrict them from arbitraging away spreads that owe to a modification in the Treasury's debt management policy. The segmentation seem more pronounced in the short and medium sectors of the yield curve, an asymmetry we attribute to our belief that the 10 y note and the 30y bond are often viewed as close substitutes rather than complements. Graphs of the yield responses to supply shocks show that the pattern varies with the maturity of Treasury securities.

Our results also shed some light on the Treasury's ability to twist the yield curve, a focus of recent debates on optimal debt management policy. A natural consequence of our findings is the logical question of whether the Treasury should borrow short- or long-term. The answer to this question is beyond the scope of this paper. But while we do not advocate any specific theory on this contentious issue, our results indicate the Treasury's ability to twist is determined to some extent by the degree of substitutability or complementarity that financial markets attribute to Treasury securities. Specifically, our results indicate that the likelihood of a twist is strong when it involves securities that are further apart on the maturity spectrum and diminishes when investors are prone to treat these securities as substitutes.

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