The Term Structure of Commercial Paper Rates

Chris Downing and Stephen Oliner*

Federal Reserve Board

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Abstract

This paper tests the expectations hypothesis in the market for commercial paper. Our main dataset, which is new to the literature, consists of daily indexes constructed from the actual market yields for nearly all commercial paper issued by U.S. corporations between January 1998 and August 2003. We show that the term premia built into commercial paper yields rise dramatically at year-end, causing the expectations hypothesis to be rejected. However, once we control for these predictable year-end effects, we find the reverse—that commercial paper yields largely conform with the expectations hypothesis.

Key words: Term structure; Expectations hypothesis; Commercial paper
Classification: E43

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1 Introduction

Interest in the expectations hypothesis of the term structure of interest rates dates back at least a century, largely because it is rooted in one of the most basic concerns of capital market participants: Is one better off committing funds for a long or short period? Tests of the expectations hypothesis also command attention because they shed light on the formation of expectations, the operation of financial markets, and the transmission of monetary policy through the markets.

Accordingly, an enormous amount of research has been devoted to testing the expectations hypothesis.[1] In general, although not uniformly, the expectations hypothesis has been rejected in various U.S. credit markets.[2] Recently, however, Longstaff (2000) found strong empirical support for the expectations hypothesis in the market for domestic U.S. repurchase agreements of less than 90 days maturity. Longstaff's results raise the question of whether the expectations hypothesis holds in other very short-term U.S. credit markets.

This paper tests the expectations hypothesis in the market for commercial paper. Commercial paper (CP) is an unsecured debt instrument of up to 270 days' maturity issued by investment-grade corporations. Firms use CP to finance inventories, to provide bridge financing in connection with various transactions, and for day-to-day cash management. The commercial paper market is large, with approximately $1.3 trillion outstanding at the end of 2003. In the fourth quarter of 2003, daily placements of commercial paper averaged about $100 billion, principally in maturities of 90 days or less, and each day an average of about

[1] For surveys of this literature, see Melino (1988) and Shiller (1990); Campbell, Lo and MacKinlay (1996) provide a textbook treatment.

[2] The empirical evidence from foreign markets has been somewhat more supportive of the expectations hypothesis. See Hardouvelis (1994), Gerlach and Smets (1997), Dahlquist and Jonsson (1995), and Bekaert, Hodrick and Marshall (2001). Gerlach and Smets (1997) suggest that the better performance of the expectations hypothesis in some foreign markets owes to the greater predictability of short rates in those markets.

500 firms issued new paper. Despite this active new-issue market, secondary trading is thin, consisting primarily of trades between the major dealers and investors who wish to liquidate their holdings.[3]

In addition to having limited liquidity, commercial paper is subject to default. Hence, the difference in expected returns between a long-term position in commercial paper and a strategy of rolling over shorter-term paper will reflect not only the usual premia that compensate investors for interest-rate risk, but also premia to compensate for default risk.

Testing the expectations hypothesis in the commercial paper market is of interest for several reasons. First, this test offers additional evidence on the expectations hypothesis for very short-term interest rates, where only limited research has been conducted.[4] Second, as mentioned above, commercial paper is a defaultable instrument traded in a relatively thin secondary market. Accordingly, tests of the expectations hypothesis in this market offer insights into the interaction of default and liquidity risk with the expectations hypothesis. Third, commercial paper rates are strongly influenced by monetary policy. Hence, our results have implications for the transmission of monetary policy to private interest rates.[5]

We conduct our main tests of the expectations hypothesis using a comprehensive database on commercial paper yields published by the Federal Reserve. This database, which is new to the literature, consists of daily indexes constructed from the actual market yields on nearly all commercial paper issued by U.S. firms, beginning in January 1998. Using these transactions-based data, we show that term premia for commercial paper jump up at year-end, causing the expectations hypothesis to be rejected. The year-end jump probably reflects a combination

[3]For further discussion of the commercial paper market, see Stigum (1990).

[4]Besides Longstaff (2000), the only studies we are aware of that have characterized the overnight to 90-day portion of the yield curve are Simon (1990), Roberds, Runkle and Whiteman (1996), Balduzzi, Bertola and Foresi (1997), Lange, Sack and Whitesell (2003), and Swanson (2004).

[5]For recent studies that highlight the role of the expectations hypothesis in the transmission of monetary policy, see Rudebusch (1995), Balduzzi et al. (1997), Kozicki and Tinsley (2001), and Roush (2001).

of factors. One contributing factor appears to be "window dressing" on the part of some large institutional investors. Just prior to releasing their year-end financial statements, these investors evidently have an incentive to temporarily substitute Treasury bills and other safe instruments for their holdings of commercial paper—especially lower-grade paper—to present a strong balance sheet to investors.[6] Another likely factor is the desire of commercial paper issuers to lock-in longer-term funding over year-end when conditions in overnight markets tend to be volatile. Their willingness to pay a premium for this insurance boosts the yield on longer-term commercial paper at year-end.

We then test whether the expectations hypothesis holds after controlling for these predictable year-end effects. The key result of the paper is that, with these controls in place, we find strong support for the expectations hypothesis in the commercial paper market. Evidently, the term premia in commercial paper yields are fairly stable other than at year-end.[7]

For completeness, we also conduct the same tests with daily yield indexes constructed from dealer quotes spanning March 1989 to August 1997. Although these indexes served as the Federal Reserve's official series on commercial paper yields until 1997, they are less accurate than the indexes based on actual transactions, and the Fed stopped publishing the dealer quotes when the transaction data became available. In contrast to the transactions data, the yields based on dealer quotes reject the expectations hypothesis, even after controlling for year-end effects.

One possible explanation for the different results is that the dealer-quote data are of

[6]This argument requires that demand for commercial paper be inelastic, so that other investors do not fill the void at prevailing yields. See Musto (1997) and Musto (1999) for additional discussions of the window dressing hypothesis.

[7]The days following the terrorist attacks of September 11, 2001 are a notable exception. The attacks severely disrupted the operation of the commercial paper market, an unanticipated event that produced sharp movements in realized term premia. We treat these observations as outliers to prevent them from affecting our conclusions.

inferior quality—perhaps containing stale quotes, for example—and do not accurately reflect the dynamics of commercial paper rates. An alternative explanation relates to the conduct of monetary policy. As documented by Lange et al. (2003) and Swanson (2004), the Federal Open Market Committee (FOMC) has taken many steps since the late 1980s to make its views on the economy and its policy actions more transparent. Among these steps, the FOMC began in February 1994 to announce policy changes on the day of its meetings and to explain the reasons for the change. Likely reflecting the greater transparency of monetary policy, Lange et al. found that the federal funds rate became more predictable starting in 1994 and that the expectations hypothesis performed far better after that point than in earlier years.[8] We obtain similar results when we split the dealer-quote data for CP rates at February 1994: the earlier data soundly reject the expectations hypothesis, but we find more support for the hypothesis from 1994 on. This finding suggests that the later period covered by the transactions data (1998-2003) may help explain the differing results generated by these data and the full sample of dealer quotes. If we had a longer history for the transactions data, we could run a more definitive test of this explanation. However, in the absence of such a test, we cannot rule out that the inferior quality of the dealer-quote data also plays a role.

The only previous tests of the expectations hypothesis with commercial paper yields appear to have been conducted by Fama (1986) and Cook and Hahn (1990), neither of which dealt with year-end effects. Both studies rejected the expectations hypothesis using dealer-quote indexes, consistent with our results based on the pre-1994 dealer-quote data.

This paper is organized as follows. In section 2, we review the theory of the expectations

[8]Swanson (2004) finds some deterioration in the predictability of the federal funds rate since 2001. However, for the period since 1994 as a whole, he concludes that increased predictability remains a robust feature of the data and that changes in FOMC transparency have likely helped to make the funds rate more predictable.

hypothesis and our specification of time-varying term premia. Section 3 discusses our data and econometric results, and section 4 concludes with some final thoughts on the implications of our results and some directions for future work.

2 Theory

2.1 The Expectations Hypothesis

Denote by $r(m,t)$ the interest rate at time t on a spot loan to be repaid at time $t+m$ (a discount bond of maturity m). Similarly, let $f(m,t+i,t)$ denote the interest rate at time t on a forward loan that starts at time $t+i$ and ends at time $t+i+m$. We measure both $r(m,t)$ and $f(m,t+i,t)$ on a continuously compounded and annualized basis. Using $m=1$ to denote an overnight maturity, we can write the following relationships between forward overnight rates and future spot overnight rates:

$$f(1,t+0,t) - r(1,t) = 0 \quad (1)$$

$$f(1,t+1,t) - r(1,t+1) = \nu_{t+1} \quad (2)$$

$$f(1,t+2,t) - r(1,t+2) = \nu_{t+2} \quad (3)$$

$$\vdots$$

$$f(1,t+m-1,t) - r(1,t+m-1) = \nu_{t+m-1} \quad (4)$$

where the $\nu_{\cdot}$ are random errors. Equation 1 states that, by definition, the forward overnight rate zero days ahead equals the current spot overnight rate. Equations 2-4 state that the forward overnight rate set at time t for time $t+i$ ($i = 1, 2, \ldots, m-1$) will differ from the spot overnight rate realized at time $t+i$ by the random amount ν_{t+i}.

To develop a test of the expectations hypothesis, we first add the elements on each side of equations 1-4, producing:

$$\sum_{i=t}^{t+m-1} (f(1,i,t) - r(1,i)) = \sum_{i=t+1}^{t+m-1} \nu_i. \tag{5}$$

Noting that the m-period continuously-compounded spot rate $r(m,t)$ equals the average of the forward overnight rates over its term ($r(m,t) = \frac{1}{m}\sum_{i=t}^{t+m-1} f(1,i,t)$), we have:

$$r(m,t)m = \sum_{i=t}^{t+m-1} r(1,i) + \sum_{i=t+1}^{t+m-1} \nu_i. \tag{6}$$

Next, multiply equation 6 by $\frac{1}{m}$, subtract $r(1,t)$ from both sides, and rearrange terms, producing:

$$\frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) - r(1,t) = (r(m,t) - r(1,t)) - \frac{1}{m}\sum_{i=t+1}^{t+m-1} \nu_i. \tag{7}$$

Now introduce $\beta_m^* = 1$, and write $\sum_{i=t+1}^{t+m-1} \nu_i$ as the sum of a predictable component, $\alpha_{m,t}^*$, and a mean-zero error term, $\epsilon_{m,t}^*$, which produces:

$$\frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) - r(1,t) = -\frac{1}{m}\alpha_{m,t}^* + \beta_m^*\left(r(m,t) - r(1,t)\right) - \frac{1}{m}\epsilon_{m,t}^*. \tag{8}$$

Equation 8 can be rewritten in the following equivalent but more convenient form that we use for our regression tests:

$$r(m,t) - \frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_{m,t} + \beta_m(r(m,t) - r(1,t)) + \epsilon_{m,t}, \tag{9}$$

where $\alpha_{m,t} \equiv \frac{1}{m}\alpha^*_{m,t}$, $\beta_m \equiv 1 - \beta^*_m$, and $\epsilon_{m,t} \equiv \frac{1}{m}\epsilon^*_{m,t}$. Under rational expectations, $\epsilon_{m,t}$ is uncorrelated with information available at time t, including the yields appearing on the right-hand side of equation 9. As a result, the $\alpha_{m,t}$ and β_m coefficients can be estimated consistently with ordinary least squares.

The expectations hypothesis rules out any time variation in term premia.[9] In terms of the coefficients in equation 9, we characterize the expectations hypothesis with the restrictions $\alpha_{m,t} = \alpha_m$, a time-invariant constant for each maturity m, and $\beta_m = 0$ for all m. The so-called *pure expectations hypothesis* further requires that $\alpha_m = 0$. In other words, under the expectations hypothesis, the difference between an m-period yield and the average of the overnight yields over the same period equals a constant, possibly equal to zero, plus expectational error.

2.2 Year-End Premia in the Commercial Paper Market

Predictable year-end effects are a key component of term premia at all maturities in the commercial paper market. We illustrate these effects for the 30-day maturity in figure 1; the other maturities exhibit qualitatively similar year-end effects and so are omitted for the sake of brevity. Panel A of the figure shows 30-day term premia for CP issued by the highest quality firms—those with short-term credit ratings of A1/P1 and long-term bond ratings of AA or better.[10] The term premia are computed as the difference between the 30-day CP rate and the average overnight CP rate over the same 30-day period. The panel

[9]Many different versions of the expectations hypothesis have appeared in the literature. At longer maturities, some forms of the expectations hypothesis are inconsistent with each other. However, for maturities as short as we consider here, the various forms of the expectations hypothesis are virtually the same (see Longstaff (2000) for further discussion).

[10]Standard and Poor's and Moody's both issue short-term ratings on a three-point scale (1, 2, and 3), with 1 being the highest and 3 being the lowest. Standard and Poor's prefixes their short-term ratings with an 'A', while Moody's uses a 'P'. Thus, an A1/P1 rating means that the firm received a '1' rating by both rating agencies.

displays both the dealer-quote data (1989-1997) and the transaction-based data (1998-2003) that we discuss at length in section 3.1 below.[11] Panel B shows 30-day term premia for CP issued by A2/P2-rated firms with long-term bond ratings of BBB or A—issuers with credit quality substantially below those in the upper panel; this series begins in 1998 because the dealer-quote data for this rating class are very limited. As can be seen, the 30-day term premium tends to jump as year-end approaches, with the largest increases for A2/P2-rated issuers. This additional premium then disappears as soon as year-end passes. The size of the spike varies from year to year, and we allow for this variation in our empirical tests of the expectations hypothesis. In recent years, the largest jump occurred in advance of the century date change at the end of 1999.

Year-end effects in other money markets are more muted, on balance, than those in the commercial paper market. For example, as shown in figure 2, term premia do tend to rise at year-end in the markets for repurchase agreements (panel A) and federal funds (panel B). However, for repurchase agreements ("repo"), the year-end increases are relatively small and are often difficult to distinguish from the variation throughout the rest of the year. For federal funds, the year-end increases in the term premium are roughly on par with those for AA-rated commercial paper but are much smaller than the spikes observed for A2/P2 paper. These differences likely help to explain why Longstaff (2000) could not reject the expectations hypothesis in the repo market even without controls for year-end effects and why Lange et al. (2003) found that the expectations hypothesis has performed reasonably well in the federal funds market since the early 1990s, again without year-end controls.

Year-end effects are evident as well in the quantity of commercial paper outstanding.

[11] The break in the series reflects the Fed's switch from dealer quotes to the transactions data. Although 30-day rates based on the transactions data are available back to the beginning of 1997, the series for overnight rates starts a year later, which prevents us from using the 1997 data.

Figure 3 displays the level of outstanding commercial paper for "tier-1" and "tier-2" issuers.[12] The lower panel indicates that CP outstanding for tier-2 issuers declines systematically at year-end and that the drop is usually reversed in January. This pattern is much less evident for tier-1 issuers, further indicating that year-end effects are concentrated among lower-quality firms.

Another notable feature of the CP market is that the volume of maturing paper declines markedly at year-end. Figure 4 illustrates this point by showing the average maturity structure of outstanding commercial paper at two points in December–the first Wednesday of the month (panel A) and the third Wednesday (panel B). The bars depict the average amount of paper, as of the indicated date, that is set to mature in the maturity ranges shown on the horizontal axis. The dotted vertical lines indicate year-end, and the solid lines indicate the average maturity structure for all other Wednesdays.[13] The distributions of outstanding paper reveal a pronounced drop in the amount of paper maturing around year-end. As year-end approaches, the maturity structure tends to shift toward longer-dated paper, with the amount of CP maturing after year-end being substantially greater than would be predicted by the average maturity structure of paper at other times of the year. This paucity of maturing paper, combined with the limited amount of secondary market trading, points to a fall-off in market liquidity at year-end.

One explanation for these year-end patterns focuses on "window dressing" by some in-

[12]Rule 2a-7 of the Investment Company Act of 1940 limits the credit risk that money market mutual funds may bear by restricting their investments to "eligible" securities. An eligible security must carry one of the two highest ratings ("1" or "2") for short-term obligations from at least two of the nationally recognized statistical ratings agencies (which currently consist of Standard and Poor's, Moody's, and Fitch IBCA). A tier-1 security is an eligible security rated "1" by at least two of the rating agencies; a tier-2 security is an eligible security that is not a tier-1 security. We employ this quality split in order to display lengthy time series, as the Federal Reserve data provide only a limited history for CP outstanding based on the A1/P1 versus A2/P2 split.

[13]We use Wednesdays for the figure because the Federal Reserve releases its data on CP outstanding at a weekly-Wednesday frequency.

vestors in the commercial paper market (Musto (1997, 1999)). According to the Federal Reserve's Flows of Funds accounts, approximately one-half of all outstanding commercial paper is held by money market mutual funds, insurance companies, and pension funds. Most of these entities report on their portfolio holdings at year-end, and they show some propensity to window dress their balance sheets by temporarily substituting higher-quality investments for their usual holdings. Given this behavior, lower-quality issuers likely would have to offer a yield premium to induce investors to hold their paper over year-end. Rather than paying such a premium, some issuers might find it less costly to turn to other forms of finance at year-end, which would help explain the drop in outstanding paper.[14]

A second explanation focuses on uncertainty in short-term financing markets at year-end. Overnight interest rates tend to be highly volatile at that time of year because of the large—and variable—increases in demand for cash by financial institutions, nonfinancial businesses, and individuals. Figure 5 depicts one measure of this heightened rate volatility. As shown, the deviation of the federal funds rate from its target tends to be substantially larger around year-end than during the rest of the year. To the extent that this volatility is transmitted to other very short-term instruments like overnight commercial paper, firms might be willing to insure against this interest-rate risk by issuing longer-maturity paper in lieu of rolling over paper every day. The spike in term premia around year-end shown in figure 1 could partly reflect the price that issuers are willing to pay for this insurance.

[14] Window dressing by money fund managers might also help explain the year-end pricing effects observed in the Treasury bill market. If enough CP investors move their funds into Treasury bills for a short time around year-end, this flow could cause the temporary increase in Treasury bill prices documented by Duffee (1996).

2.3 Time Variation in Term Premia

The foregoing discussion suggests that term premia in the commercial paper market might rise at year-end as compensation for increases in liquidity risk, investors' aversion to holding lower-quality assets at that time, and heightened interest-rate risk. It is important to highlight that these year-end factors are *predictable*—and thus get embedded in the structure of commercial paper yields. This stands in contrast to unpredictable events, such as sudden defaults by large issuers (owing to fraud, for example) or the terrorist attacks of September 11, 2001, which have important but unanticipated effects on the market. In our rational expectations framework, it is only anticipated events that can systematically shift term premia.

As we showed in figure 1, the amplitude of the year-end jump in term premia varies markedly both over time and with credit quality. Our analysis of year-end effects suggests that a realistic specification of term premia ought to consist of three components: (i) the standard time-invariant component that is present at all times; (ii) a component that rises as year-end approaches; and (iii) a component that can capture the effects of heightened interest-rate volatility and reduced market liquidity right around year-end. To allow for these sources of variation, we use the following specification for term premia:

$$\alpha_{m,t} = \alpha_{m,0} + \sum_{yr}(\alpha_{m,yr,1} + \alpha_{m,yr,2}\tau_m + \alpha_{m,yr,3}\tau_m^2 + \alpha_{m,yr,4}\tau_m^3)DX_{m,yr,t} + \alpha_{m,yr,5}D_{yr,t}, \quad (10)$$

where $yr = 1998, \ldots, 2002$ for the transactions-based data and $yr = 1989, \ldots, 1996$ for the dealer-quote data.

In this specification, the coefficient $\alpha_{m,0}$ represents the usual time-invariant component of the term premium—component (i) above. Component (ii) is modeled using a cubic specification in time that is in effect only near year-end. Specifically, the variable $DX_{m,yr,t}$ equals one when the observation date t is before, and the maturity date $t+m$ is after, December 26 of the year indexed by yr. We allow $DX_{m,yr,t}$ to switch on for maturities slightly before year-end because figure 4 suggests that year-end effects begin to be seen about a week before the turn of the year. When this dummy variable equals one, the variable τ_m counts the number of days from December 26 of the given year to the maturity date $t+m$. The cubic term in τ allows the specification to capture a wide range of time patterns for the year-end term premium. Finally, component (iii) is modeled with the dummy variable $D_{yr,t}$, which equals one when the observation date t is between December 22 of the year indexed by yr and January 10 of the following year. Hence the coefficient $\alpha_{m,5}$ captures any level shifts in term premia right around year-end, when, as we discussed earlier, liquidity in the market appears low and short-term interest rates are relatively volatile. Allowing the coefficients to vary across years introduces additional flexibility. In the next section, we test traditional specifications of term premia, as well as our new specification that incorporates year-end premia.

3 Data and Estimation Results

3.1 Data

Large investment-grade corporations in the United States typically maintain commercial paper programs, often of significant size. The bulk of commercial paper issuers reside in the top size quintile of publicly traded corporations—firms with total assets, at book value, of more than $1.4 billion in 2003. For this top quintile of firms, commercial paper accounts, on average, for 30 percent of their current liabilities, making it an important source of short-term credit.[15]

We employ two sources of daily data on commercial paper yields.[16] Our primary source consists of the commercial paper discount yields currently published by the Federal Reserve Board for AA-rated and A2/P2-rated domestic nonfinancial companies. We focus on maturities of 90 days or less because the market for longer-maturity paper is quite thin. The Federal Reserve Board constructs these yield indexes from transaction-level data supplied by the Depository Trust Company, which handles clearing and settlement for more than 95 percent of all commercial paper trades in the United States. Federal Reserve staff fit a smooth curve to the transaction-level data. The end result is a daily series of constant-maturity, zero-coupon yields.[17] Our yield data cover each business day from January 2, 1998 through August 1, 2003, for a total of 1,365 daily observations.

The upper two panels of table 1 display summary statistics for these transactions-based yields. For AA-rated issuers (top panel), yields rise very little with increases in maturity, resulting in a mean spread between 90-day and overnight paper of only 3.09 basis points. In contrast, for A2/P2-rated issuers (middle panel), the mean 90-day spread is 19.09 basis points, reflecting the greater default and liquidity risks in this part of the market. The yields for both AA and A2/P2 paper are highly autocorrelated at all maturities, consistent with the behavior of other short-term interest rates.

We also make use of a second database, consisting of dealer quotes on yields for com-

[15]These figures are calculated using the Federal Reserve's commercial paper database and Compustat.

[16]The yield indices used to construct the term premia are from the Federal Reserve's daily commercial paper release. Details can be found at:
`http://www.federalreserve.gov/Releases/cp/about.htm`.

[17]This procedure does not have to deal with the effect of coupons on observed yields because the underlying data pertain to newly-issued discount instruments. Thus, the smoothing procedure directly estimates the discount function itself, as opposed to extracting the discount function implied by the prices of coupon securities.

mercial paper issued by a generic AA-rated domestic corporation, collected by the Federal Reserve Bank of New York between February 27, 1989 and August 29, 1997.[18] Over this period, the New York Fed surveyed the major dealers each day and constructed unweighted averages of the rates reported for various maturities. The lower panel of table 1 displays summary statistics for these data. As shown, the longer-term yields from the dealer survey exhibit higher spreads to overnight rates than the transactions-based AA yields, reaching 14.59 basis points at the 90-day maturity. One possible explanation for the higher spreads is that, while the dealers were instructed to report yields for AA-rated corporations, they may have in fact provided yields for firms of lesser credit quality.[19]

3.2 A First Look at Term Premia

To characterize term premia in the commercial paper market, table 2 reports estimates of the following regression:

$$r(m,t) - \frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_m + \epsilon_{m,t}, \tag{11}$$

where α_m represents the average term premium at maturity horizon m, and $\epsilon_{m,t}$ is a mean-zero error term. Recalling equation 9, this specification imposes the restrictions from the expectations hypothesis—namely, that $\alpha_{m,t}$ is time-invariant and that β_m equals zero—and then estimates the average term premium conditional on these restrictions.

As shown in the upper two panels of the table, the transactions data clearly indicate that term premia are non-zero, with the exception of the shortest maturity (7 days) for the

[18] The New York Fed surveyed dealers going back to the 1960s, but only for a limited range of maturities. In order to make our results comparable across the two sources of data, we restrict our analysis of the dealer-quote data to the period over which the dealer survey included the widest range of maturities.

[19] See Cook and Lawler (1983) for further discussion of the New York Fed's dealer survey.

highest-quality issuers. On average, AA-rated issuers pay 3.50 basis points more to place 30-day paper than to roll overnight paper for 30 days, while A2/P2-rated issuers pay 12.76 basis points more; over a 90-day horizon, these premia rise to 11.93 basis points for AA issuers and 27.85 basis points for A2/P2 issuers.[20] Based on these results, we would clearly reject the pure expectations hypothesis in the commercial paper market.

The dealer-quote data support a similar conclusion. The estimates of term premia in the lower panel are highly significant and larger in size than those based on the transactions data for AA firms, the part of the market to which the quotes should apply. In fact, for every maturity, the estimated term premium is closer to the premium for A2/P2-rated firms than to that for AA-rated firms.

3.3 Tests of the Expectations Hypothesis

3.3.1 Maturity-Specific Term Premia

We first consider the evidence for the standard version of the expectations hypothesis that allows term premia to vary by maturity but not over time:

$$r(m,t) - \frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_m + \beta_m \left(r(m,t) - r(1,t)\right) + \epsilon_{m,t}. \tag{12}$$

In contrast to equation 11, this specification tests whether $\beta_m = 0$ rather than imposing that restriction *a priori*. Table 3 displays the estimates of α_m and β_m, along with their t-statistics. The t-statistics are corrected for the overlap in the errors. However, as is well known in the term structure literature, interest rates are highly persistent, a feature that

[20] These results contrast with those in Longstaff (2000), who found a 90-day term premium of just 3 basis points using data on repurchase agreements. We would expect term premia to be greater in the commercial paper market both because repo contracts are almost free of default risk and because these contracts are more liquid than commercial paper.

can lead to distortions in the distributions of conventional test statistics (Bekaert, Hodrick and Marshall (1997)). Hence for each maturity the final two columns of the table display the critical values for a 95 percent confidence interval that maintains the correct test size in the presence of a regressor with a (possibly large) autoregressive root (Cavanagh, Elliott and Stock (1995)).[21]

Based on the transactions data for AA-rated issuers, we find some support for the expectations hypothesis. The point estimates of β_m at the 60- and 90-day maturities are relatively close to zero (0.13 and 0.04, respectively); comparing the t-statistics to the lower and upper critical values, these coefficient estimates are also statistically insignificant. The point estimate of β_m at the 7-day maturity is somewhat larger at 0.29, but it too is insignificant. At the same time, the estimate of β_m is significantly different from zero at the 15- and 30-day maturities. In contrast to these mixed results based on the data for AA-rated issuers, we decisively reject the expectations hypothesis with the data for A2/P2-rated issuers and with the dealer-quote data. For both sets of data, the estimate of β_m is well above zero and highly significant at every maturity.

The estimates of α_m are all significantly different from zero, with the exception of the 7-day maturity for AA-rated issuers. Thus, as in table 2, we obtain strong evidence against the pure expectations hypothesis. The estimates of α_m using the dealer-quote data are again more similar to those for A2/P2-rated issuers than for AA-rated issuers, underscoring our concerns about the accuracy of the dealer-quote data.

[21] Bekaert et al. (1997) develops a Monte Carlo approach for correcting the size of the test statistics in finite samples. This method is implemented by Longstaff (2000) for very short-term repo data. We have implemented this procedure on all of the data here, but found that the data generating process (4-factor VAR-GARCH) tended to produce too many explosive paths for the Monte Carlo results to be reliable.

3.3.2 Controlling for Time Variation in Term Premia

We now consider the evidence for the expectations hypothesis after we control for the rise in term premia around year-end and for the idiosyncratic effects of the September 11 terrorist attacks. In this case, our specification is given by:

$$
\begin{aligned}
& r(m,t) - \tfrac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_{m,0} + \beta_m(r(m,t) - r(1,t)) + \\
& \sum_{yr}(\alpha_{m,yr,1} + \alpha_{m,yr,2}\tau_m + \alpha_{m,yr,3}\tau_m^2 + \alpha_{m,yr,4}\tau_m^3)DX_{m,yr,t} + \alpha_{m,yr,5}D_{yr,t} + \\
& \alpha_{m,6}SX_{m,t} + \alpha_{m,7}S_t + \epsilon_{m,t}.
\end{aligned}
\tag{13}
$$

For the transaction data, $yr = 1998, \ldots, 2002$, while for the dealer-quote data, $yr = 1989, \ldots, 1996$. The variables $DX_{m,yr,t}$, $D_{yr,t}$, and τ_m are as defined in section 2.3, and $SX_{m,t}$ and S_t are variables designed to control for the effects of the September 11 attacks on the CP market.

These attacks severely disrupted the computer networks that money-market participants rely on to carry out their transactions. With some banks unable to transfer the funds necessary to redeem maturing commercial paper, a sizable portion of the paper that came due in the days following the attacks could not be honored.[22] Moreover, a large number of issuers could not roll over maturing paper. Many of these issuers instructed their banks to draw down liquidity backup lines in order to make payments on maturing paper, producing substantial draws at the discount window at the Fed.[23]

These market disruptions produced significant pricing effects in the commercial paper

[22] The situation was analagous to an individual attempting to cash a check, and upon finding the necessary funds unavailable, returning the next day to try again. In the jargon of the money markets, the maturity presentments were "failed" and presented again the next day.

[23] For a full description of the Federal Reserve's response to the stresses in the U.S. financial system, see the Monetary Report to Congress, February 2002, available at:
`http://www.federalreserve.gov/boarddocs/hh/2002/February/FullReport.htm`.

market. Yields on the Tuesday of the attacks showed little change from Monday, but they likely reflected pricing on deals done in the morning before the attacks.[24] On Wednesday, yields on overnight paper jumped about 30 basis points for AA-rated issuers, and about 60 basis points for A2/P2 issuers.[25] Overnight yields rose further on Thursday and Friday, as operational risks in the clearance and settlement systems remained significant, raising the possibility of defaults. Then, on the Monday following the attacks, yields dropped precipitously, owing in part to a 50 basis point cut in the target federal funds rate just prior to the reopening of the equity markets, as well as to the sizable liquidity injections by the Fed in the days following the attacks. Rates continued to gyrate for the remainder of the month, but by the beginning of October, the market had largely stabilized.

To account for these effects, we introduce the dummy variable $SX_{m,t}$ which takes the value one when the observation date t is before September 11, 2001 and the maturity date $t+m$ is after September 11. When $SX_{m,t}$ equals one, the yields at date t were unaffected by September 11, but all of the overnight yields realized after September 11—which we use to construct term premia—were highly elevated. We also introduce the variable S_t, which equals one when the observation date t is between September 11 and September 18, 2001, a period during which all maturities—including overnight paper—bore the imprint of September 11. We treat the observations that are dummied out in this way as outliers resulting from a singular event that ought not influence our conclusions regarding the expectations hypothesis.[26]

[24]The commercial paper market is a "morning market" in the sense that nearly all of the deals are completed before noon, in order to facilitate same-day settlement.

[25]The Fed's calculation of yields on these days was hampered by the sharp drop in market liquidity, and the figures cited here should be regarded as merely indicative.

[26]While these September 11 dummy variables are similar in their construction to the year-end dummy variables we introduced earlier, their interpretation is very different. Whereas the year-end dummies control for the influence of *predictable* events on term premia, the September 11 dummy variables are being used to remove the influence of an *unpredictable one-time* event during the period covered by our transactions data.

Table 4 presents the estimates of the $\alpha_{m,0}$ and β_m coefficients in equation 13; the tables in appendix A.2 display the estimates of all of the year-end and September 11 coefficients. Based on the transactions data, we find strong support for the expectations hypothesis once we control for year-end effects and September 11.[27] For AA-rated commercial paper, the point estimates of β_m are now much closer to zero than they were absent these controls (except for 90-day paper, for which β_m was already close to zero). Moreover, comparing the t-statistics to the robust critical values in the last two columns of the table, we see that all of the slope coefficient estimates are statistically insignificant. For A2/P2-rated commercial paper, we find that the point estimates of β_m are also in the neighborhood of zero. Only the estimate at the 30-day maturity is statistically significant.

The results for the dealer-quote data, shown in the bottom panel of the table, provide much less support for the expectations hypothesis. The point estimates of β_m shrink in size after controlling for year-end effects, but they remain statistically significant except at the 7-day maturity. In addition, these coefficient estimates are generally larger than those obtained with the transactions data.

As noted earlier, year-end changes in term premia are a dominant characteristic of commercial paper yields. Table 5 quantifies this fact by comparing the adjusted-R^2 statistics from the standard regression (equation 12) with those from the specification that accounts for year-end effects and September 11 effects (equation 13). In every case, the R^2 rises dramatically with the inclusion of these controls. Indeed, for the 90-day AA-rated transaction data, the R^2 rises from essentially zero to nearly 60 percent, and most of the other table entries show a jump of at least 30 percentage points.[28]

[27] If we omit the September 11 dummy variables, our qualitative results are unchanged, but some coefficients are less precisely estimated.

[28] In every regression, the September 11 dummy variables account for only a small part of the increase in R^2. For the AA-rated data, the September 11 dummy variables contribute between three and 14 percentage

3.3.3 Year-End and September 11 Effects

Recall that equation 13 includes a cubic polynomial function to capture the movements in term premia as year-end approaches. Figures 6 and 7 plot this estimated function for 30-day commercial paper yields.[29] As shown in figure 6, using the transactions-based data, we find that term premia for both AA- and A2/P2-rated issuers vary considerably from year to year. However, more often than not, the term premia display a concave pattern, initially rising and then declining as the issue date approaches year-end. This pattern was especially pronounced in 1999, just ahead of the century date change. In that year, term premia initially shot up on concerns that computer bugs could hinder the functioning of financial markets but plummeted shortly before the turn of the year, at least in part because the Federal Reserve committed to provide substantial liquidity in the event of a market disruption. In contrast to 1999, year-end premia were muted in some other years, notably for AA-rated issuers in 2001 and 2002. The year-to-year variation shown in figure 6 is highly significant: Using a standard F-test, the null hypothesis that the year-end coefficients are equal across the years is rejected with more than 95 percent confidence for both rating classes.

Figure 7 displays the analogous estimates for the dealer-quote data. The pattern of year-end term premia again varies across years. In 1990—a recession year—term premia kept rising as year-end approached, reaching an extremely high level. However, in 1992, 1993, and 1994, term premia peaked well before year-end, and in 1995, they were consistently small. As with the transactions data, an F-test indicates that these differences are highly

points to the increase in R^2, while for the A2/P2 data, these controls contribute between one and six percentage points.

[29]These plots are based on the estimates of $\alpha_{30,yr,1}, \ldots, \alpha_{30,yr,4}$ shown in tables A.1-A.3 in appendix A.2. The bulk of the estimated coefficients are statistically significant; thus, we can reject the hypothesis that the plots in each panel equal a horizontal line drawn at zero. To conserve space, we focus here on the 30-day maturity; the qualitative features of the estimates for the other maturities are similar.

significant.

The estimates of the September 11 coefficients are displayed in table 6. The coefficient $\alpha_{m,6}$ shows the estimated change in term premia in the period just before September 11, when m-day paper was priced without knowledge of the terrorist attacks but at least some of the subsequently realized overnight yields entering the calculated term premia pertain to days after September 11. As can be seen, the estimate of this coefficient is negative for 7-day paper and then turns positive for longer maturities. This sign change reflects the dynamics in overnight rates discussed above—rates rose in the days following September 11 but then fell sharply a week later when the Federal Reserve unexpectedly cut the federal funds rate by 50 basis points. The coefficient for 7-day paper picks up only the short-lived rise in overnight rates, which drives down the calculated term premium, while the coefficients for longer maturities reflect the unanticipated decline in overnight rates after the Fed's rate cut. The other September 11 coefficient, $\alpha_{m,7}$, shows the effect on term premia for m-day paper issued during the week after the terrorist attacks. All of the estimates of this coefficient are positive and highly significant, indicating the presence of a sizable term premium for paper placed between September 11 and September 18.

3.3.4 Reconciling the Dealer-Quote and Transactions-Based Results

Contrary to the results we obtained with the transactions data, the dealer-quote data generally reject the expectations hypothesis even after controlling for year-end effects. One factor behind the differing results could be the lower quality of the dealer-quote data. The ideal way to test this explanation would be to estimate equation 13 with the dealer-quote and transactions data over exactly the same period. Because we know that the transactions data accurately reflect the prevailing yields in the commercial paper market, any difference in estimation results would have to arise from problems with the dealer quotes. However, the two datasets have no time periods in common, which prevents us from running this test. Accordingly, we cannot rule out that differences in data quality help explain the divergent results.

Nevertheless, we can test an alternative explanation that relates to the conduct of monetary policy. As noted in the introduction, Lange et al. (2003) document that the FOMC has taken many steps over the past fifteen years to make monetary policy more transparent, which has enabled market participants to better anticipate policy actions. Indeed, Lange et al. found that the federal funds rate became much more predictable starting in February 1994, when the FOMC began to announce policy changes and the rationale for these actions shortly after the conclusion of its meetings. Consistent with the greater predictability of the funds rate, Lange et al. showed that the expectations hypothesis performed considerably better in the federal funds market after February 1994 than before.[30]

These developments in the federal funds market have direct implications for our tests of the expectations hypothesis in the commercial paper market. As shown in figure 8, both overnight CP rates (panel A) and longer-term yields (panel B) closely track the FOMC's target for the federal funds rate. Thus, we might expect the performance of the expectations hypothesis in the commercial paper market to have improved over the course of the 1990s, as it did for federal funds. If that were the case, this pattern could help explain why the more recent transactions-based data provide far more support for the expectations hypothesis than do the earlier dealer-quote data.

We examine this argument by splitting our dealer-quote data into two subsamples. The

[30]The link between the predictability of interest rate changes and the expectations hypothesis is analyzed in depth by Mankiw and Miron (1986); for related work, see Balduzzi et al. (1997), Gerlach and Smets (1997), and Rudebusch (1995).

first subsample covers the period from the beginning of the sample to February 4, 1994, the date of the FOMC meeting that corresponds to the Lange et al. breakpoint; the second subsample includes the rest of the dealer-quote data. Tables 7 and 8 present the coefficient estimates from re-running both the standard regression and the specification with year-end controls on each subsample. For the earlier subsample, the top panel of table 7 shows that the standard regression decisively rejects the expectations hypothesis, as the estimate of β_m is well above zero and highly significant at every maturity. Adding the year-end controls does not materially change this conclusion; as shown in the lower panel, the estimates of β_m become smaller but they remain significant at all maturities except seven days.

Turning to table 8, we find somewhat more support for the expectations hypothesis in the later subsample once we control for year-end effects. As shown in the lower panel of the table, the estimates of β_m at the 7-, 15-, and 30-day maturities are close to zero and statistically insignificant. The estimates for 60- and 90-day maturities are significant, but they are much smaller in size than their counterparts in table 7. Overall, these results are consistent with the view that the greater transparency of monetary policy accounts, at least in part, for the better performance of the expectations hypothesis in recent years.

4 Conclusion

In this paper, we rigorously tested the expectations hypothesis in the market for commercial paper. Our tests relied on daily yield indexes constructed by the Federal Reserve Board from the actual market yields for virtually all commercial paper issued by U.S. corporations. These transactions-based indexes provide an extremely accurate summary of market yields from 1998 onward. For completeness, we ran a parallel set of tests on the dealer-quote data that the Federal Reserve published before the transactions data became available; the dealer quotes are less accurate than the transactions-based indexes, but they span a longer period and have been used in previous research.

Both datasets reject the traditional specification of the expectations hypothesis that requires the term premium for a given maturity to be constant over time. This rejection is not surprising, as term premia typically rise in the commercial paper market at year-end, especially for lower-rated paper. Some plausible explanations for this year-end effect include "window dressing" by institutional investors, who hold a substantial amount of commercial paper, as well as a desire by issuers to insure against volatile interest rates and the attendant increase in rollover risk around year-end.

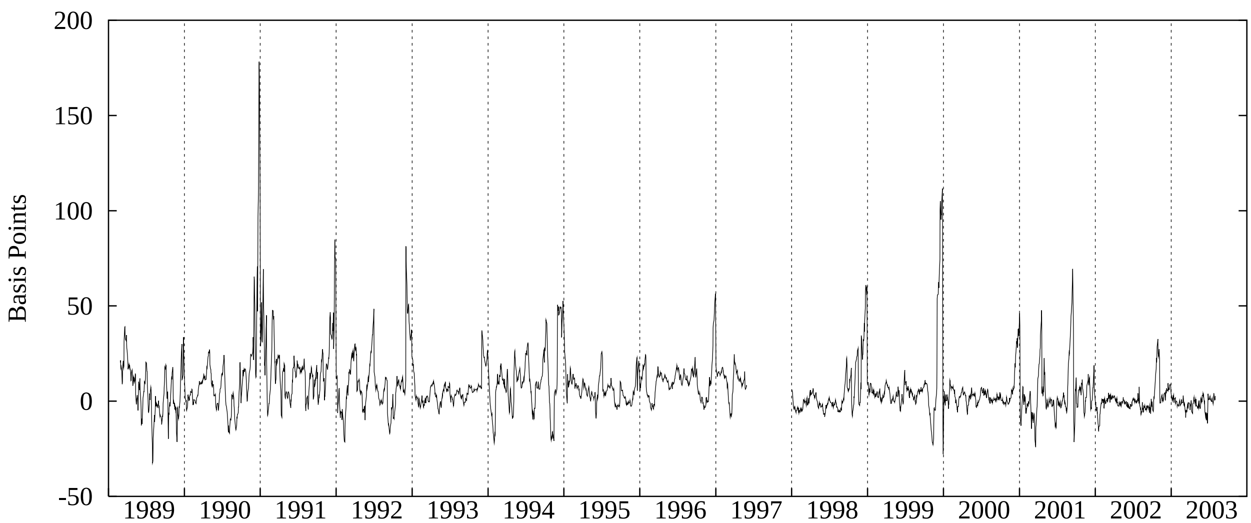
After we control for these year-end effects, we find strong support for the expectations hypothesis using the transactions data—the key result in the paper. The dealer-quote data, in contrast, largely reject the expectations hypothesis even after controlling for year-end effects. We argued that at least some of this difference likely reflects the increasing predictability of short-term interest rates over the 1990s, though we cannot rule out that the lower quality of the dealer-quote data also contribute to this result.

Finally, we should note that we have relied exclusively on composite yield indexes to carry out these tests. Our results with the transactions-based indexes indicate that the expectations hypothesis is valid for the "average firm" in the commercial paper market. However, the extent to which the hypothesis holds for individual firms remains an open question and would be an important topic for future research.

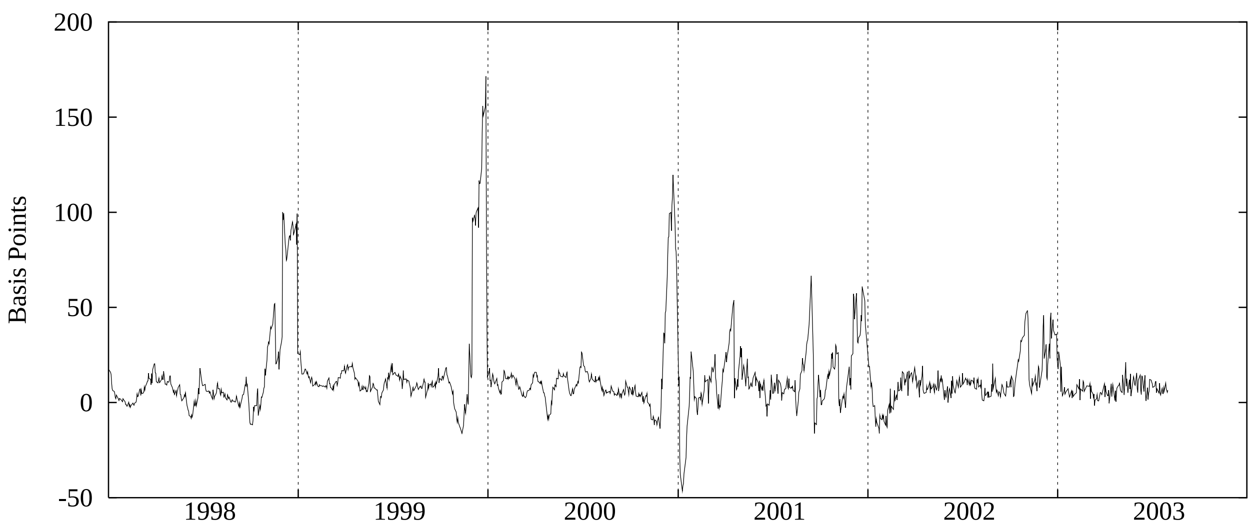
Figure 1: Term Premia for 30-Day Commercial Paper

The figure shows term premia for 30-day commercial paper issued by nonfinancial firms rated AA (panel A) and A2/P2 (panel B). Each daily observation is calculated as the difference between the 30-day rate and the average of the overnight rates subsequently realized over the term of the 30-day rate. The data in panel A cover the period from February 27, 1989 to August 1, 2003, with a break from August 30, 1997 to January 1, 1998. The data in panel B cover the period from January 2, 1998 to August 1, 2003. The vertical lines in both panels are drawn at year-end.

Panel A: AA-rated firms



Panel B: A2/P2-rated firms

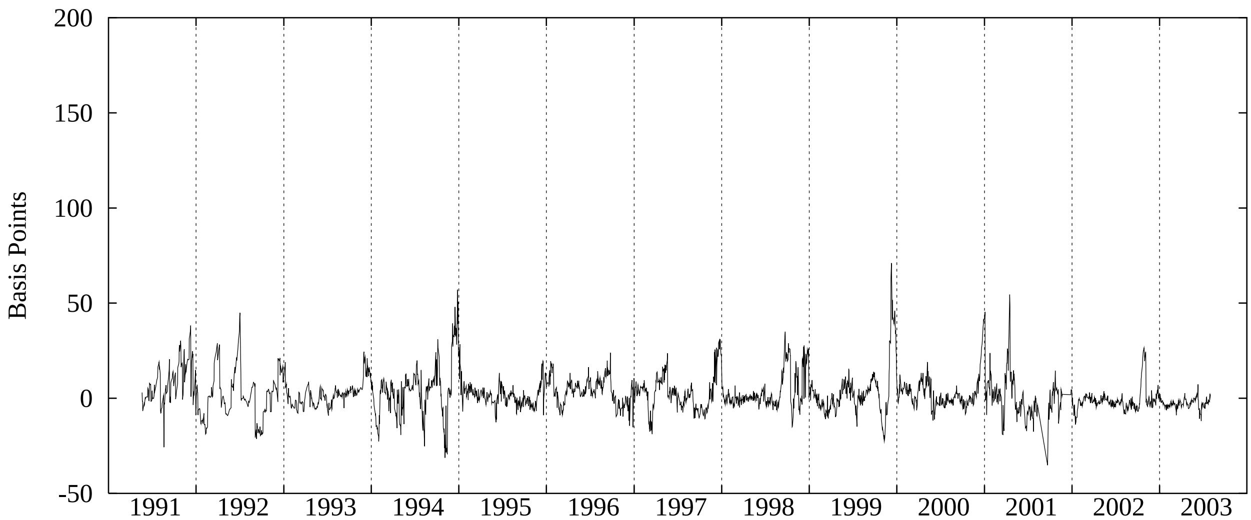


Source: Federal Reserve Board (based on data from the Federal Reserve Bank of New York market survey through 1997 and data from the Depository Trust Company after 1997).

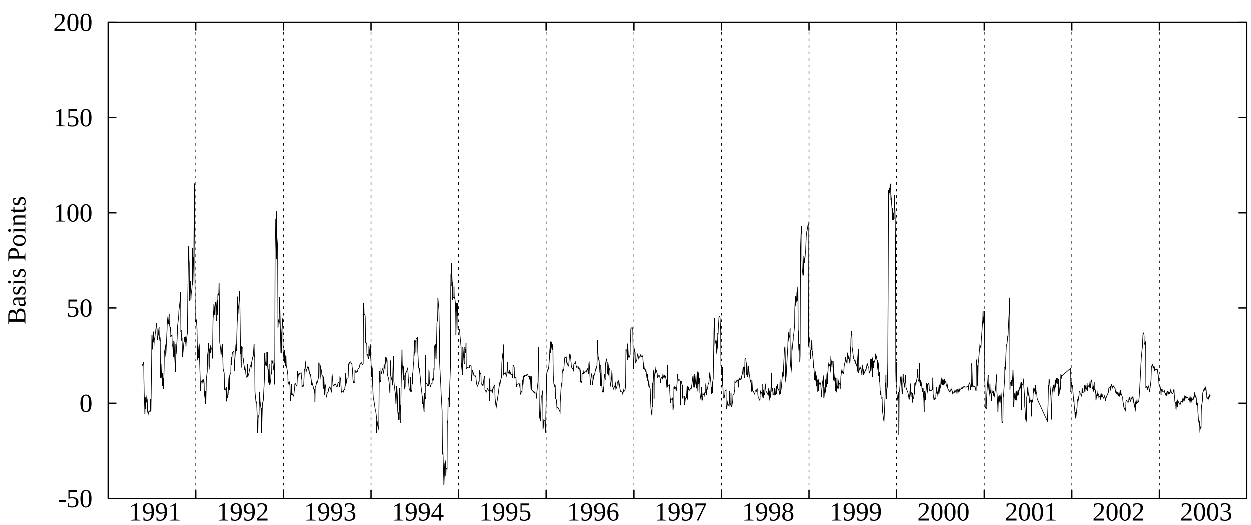
Figure 2: Term Premia for 30-Day Repurchase Agreements and Federal Funds

The figure shows term premia for 30-day repurchase agreements (panel A) and federal funds (panel B). Each daily observation is calculated as the difference between the 30-day rate and the average of the overnight rates subsequently realized over the term of the 30-day rate. The data are daily and cover the period May 21, 1991 through August 1, 2003. The vertical lines in both panels are drawn at year-end.

Panel A: Repurchase Agreements



Panel B: Federal Funds

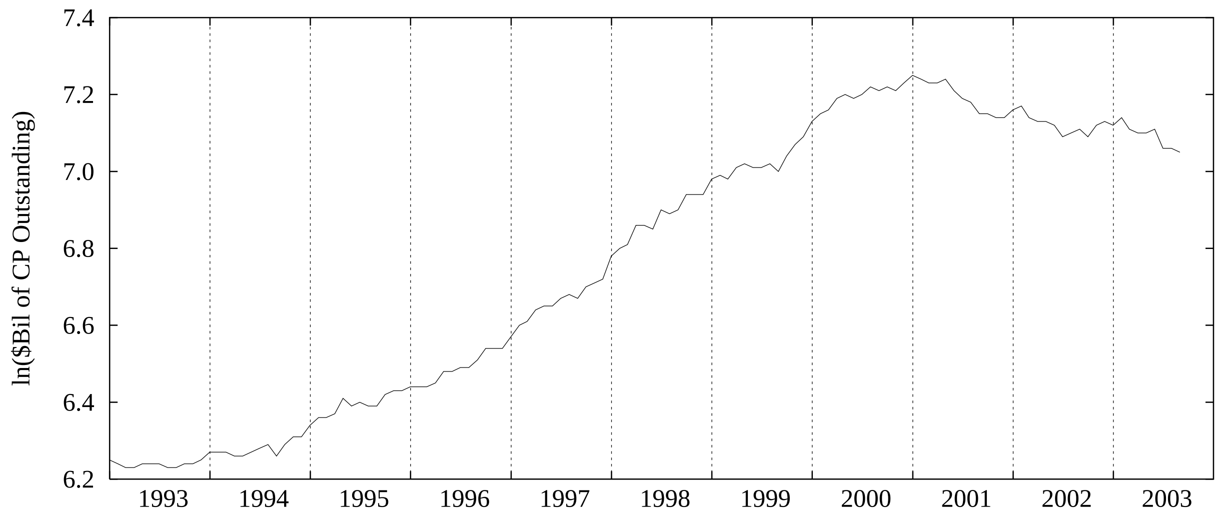


Sources: Repurchase agreement data are from Garban International; federal funds data are from the Federal Reserve Board.

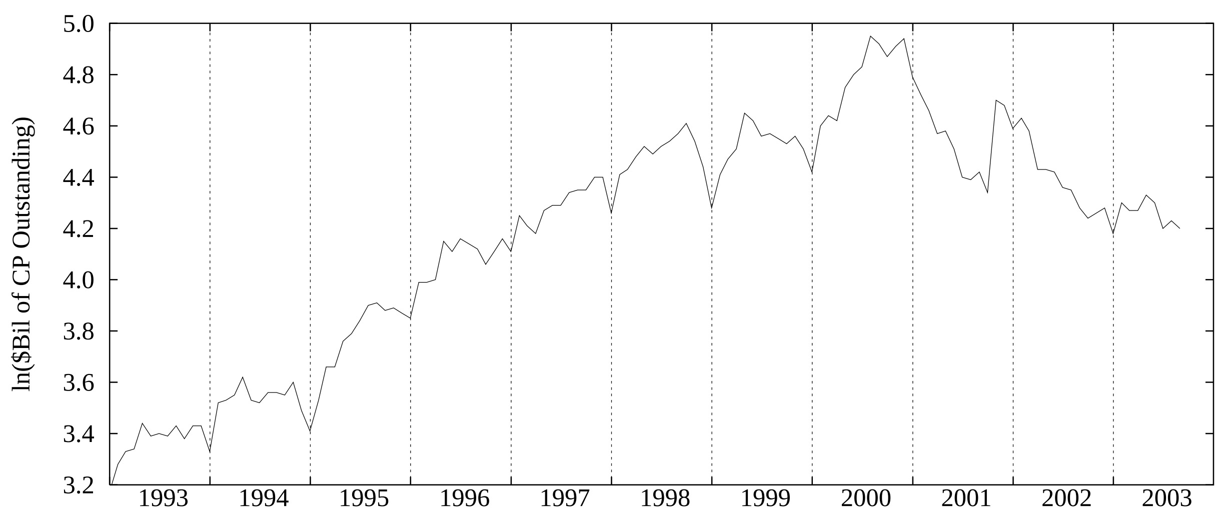
Figure 3: Tier-1 and Tier-2 Commercial Paper Outstanding

The panels display the log of commercial paper outstanding at month-end from January 1993 through August 2003 for all firms in the indicated short-term rating class. Both nonfinancial and financial firms are included in the samples. A tier-1 security is a money-market mutual fund eligible security rated "1" by at least two of the major rating agencies; a tier-2 security is an eligible security that is not a tier-1 security; all ineligible securities are excluded (refer to footnote 12 for further discussion). The vertical lines in both panels are drawn at year-end.

Panel A: Tier-1 Firms



Panel B: Tier-2 Firms

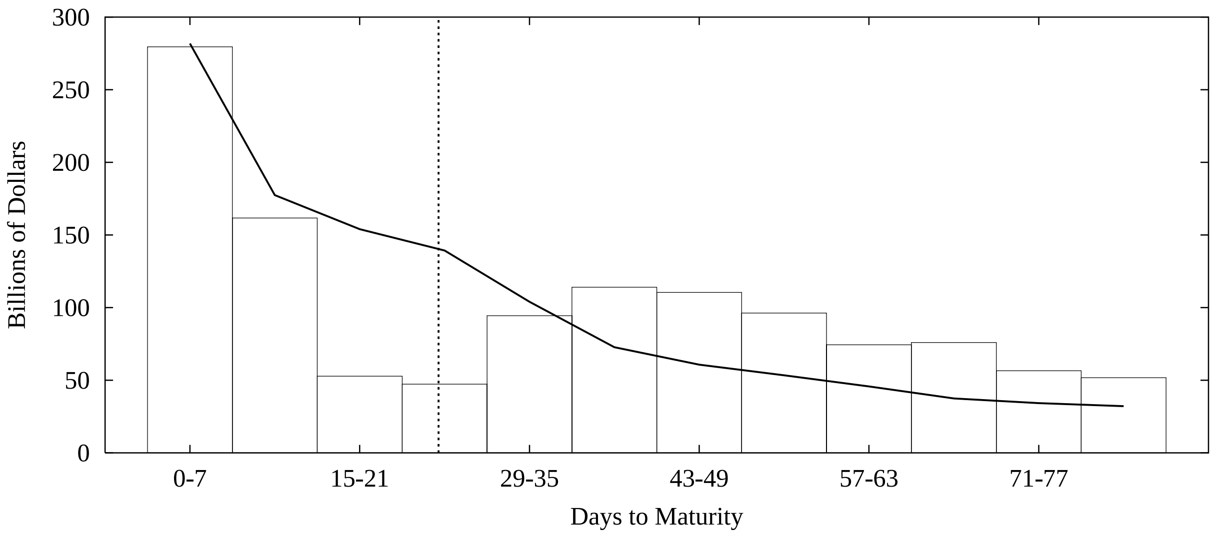


Source: Federal Reserve Board (based on data from the Federal Reserve Bank of New York market survey through 1997 and data from the Depository Trust Company after 1997).

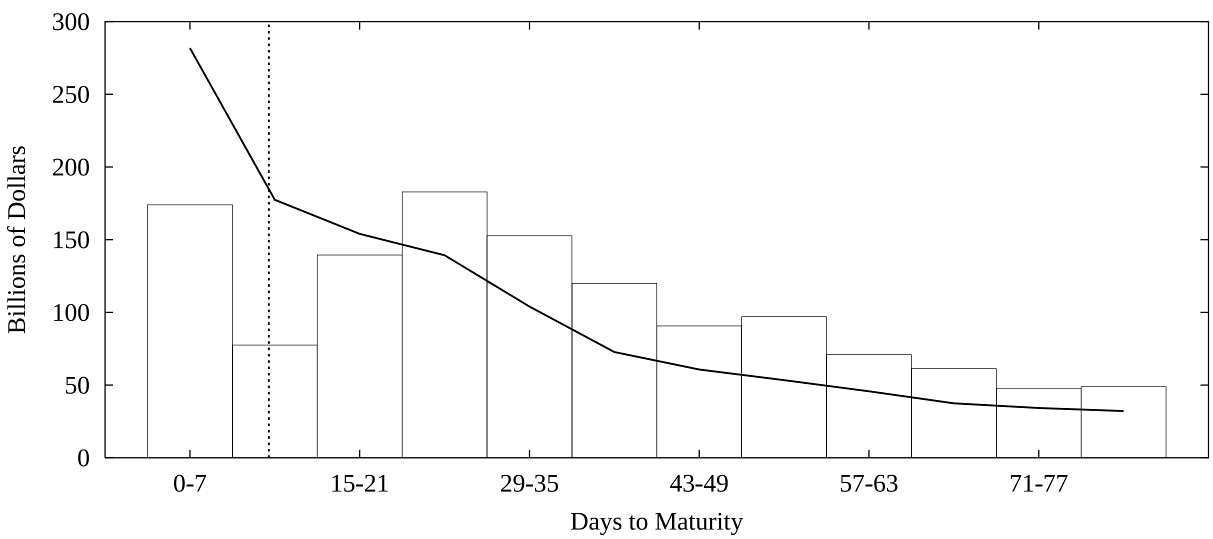
Figure 4: Year-End Maturity Structure of Commercial Paper

The figure shows the average amount of nonfinancial and financial commercial paper outstanding that is scheduled to mature in the indicated date ranges. Panel A shows the average amounts outstanding as of the first Wednesday of December for all of the years included in the transactions-data sample (1998-2002). Panel B shows the average amounts outstanding as of the third Wednesday in December for the same years. The dotted vertical line in each panel shows the approximate location of December 31. The solid line in each panel shows the average maturity structure over all other Wednesdays.

Panel A: First Wednesday of December



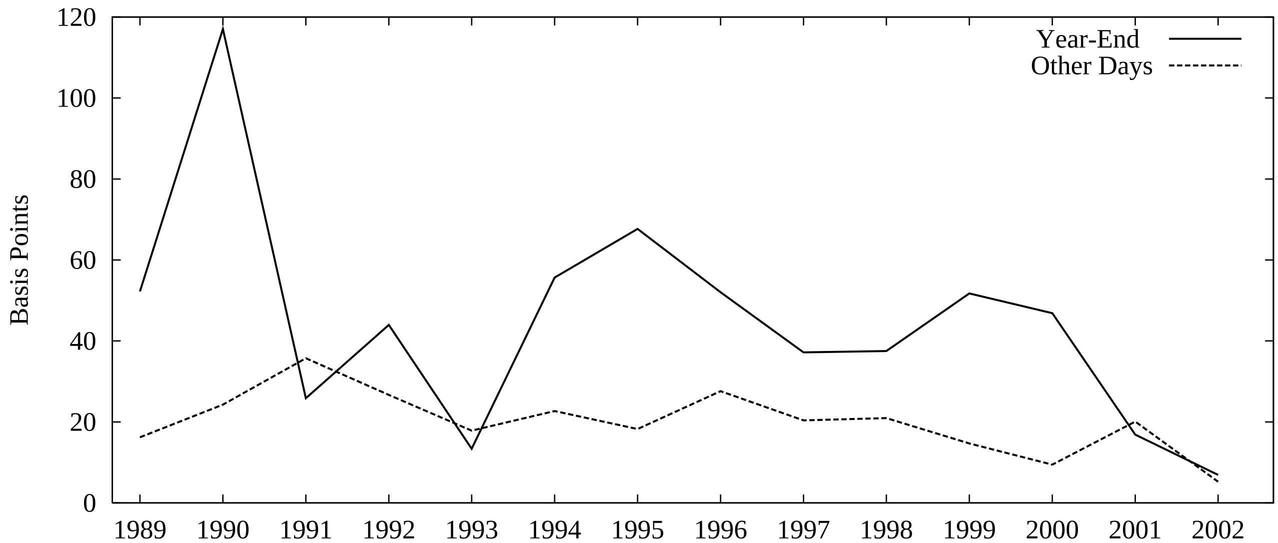
Panel B: Third Wednesday of December



Source: Federal Reserve Board, based on data from the Depository Trust Company.

Figure 5: Deviation of Federal Funds Rate from Target

The figure displays the standard deviations of the daily differences of the effective federal funds rate from the intended target rate. The solid line shows the standard deviations for year-end observations, defined as observations after December 25 of the indicated year or before January 5 of the following year. The dashed line displays the standard deviations for all other observations in the indicated year.



Source: Federal Reserve Board.

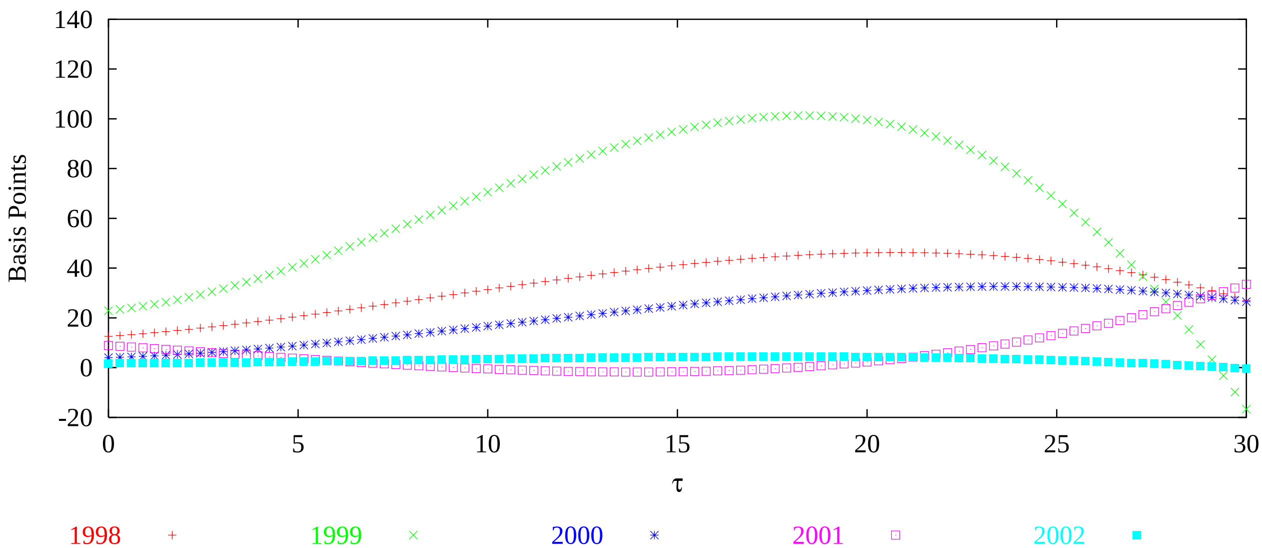
Figure 6: Year-End Term Premia for 30-Day Commercial Paper Rates, Transactions Data

The panels plot the cubic portion of the estimated year-end term premium functions for 30-day paper, given by

$$\hat{\alpha}_{m,yr,1} + \hat{\alpha}_{m,yr,2}\tau_m + \hat{\alpha}_{m,yr,3}\tau_m^2 + \hat{\alpha}_{m,yr,4}\tau_m^3$$

for $\tau = 0, 1, \ldots, 30$ and $m = 30$. When $\tau = 0$, the maturity date is December 26 and when $\tau = 30$ the maturity date is 30 days later (January 24 of the following year). The $\hat{\alpha}.$ denote estimated coefficients from the appropriate rows in tables A.1 and A.2.

Panel A: AA-Rated Firms



Panel B: A2/P2-Rated Firms

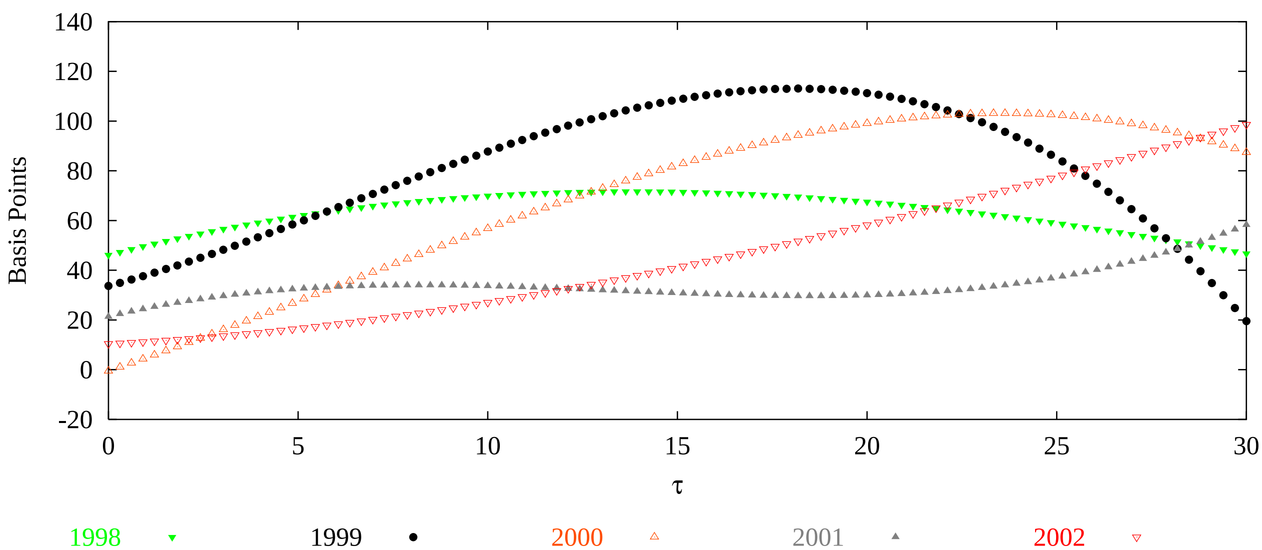


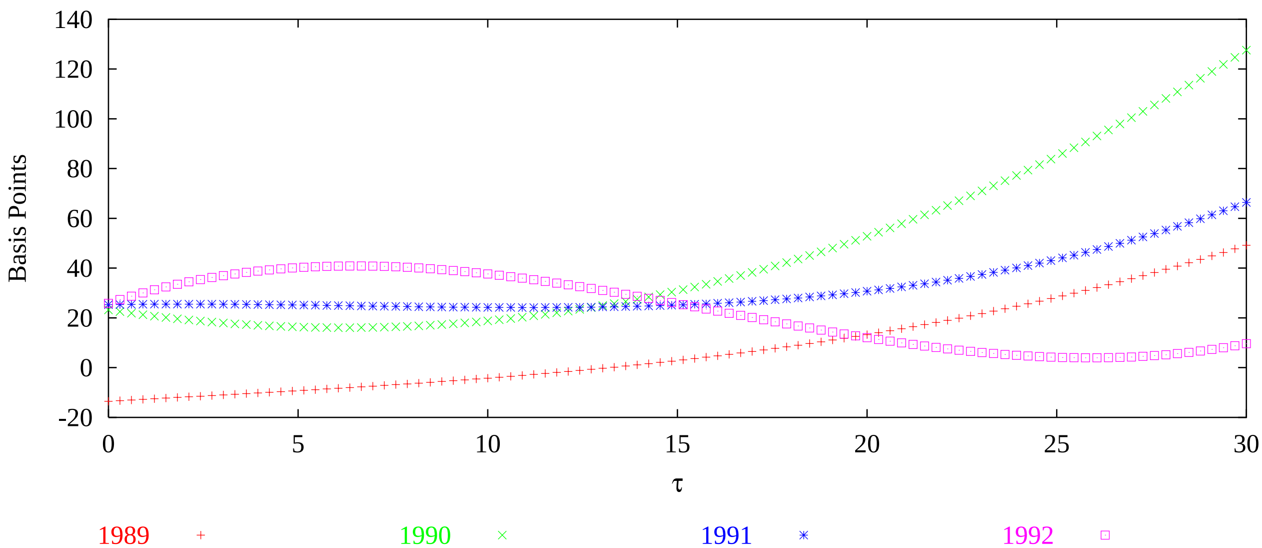
Figure 7: Year-End Term Premia for 30-Day Commercial Paper Rates, Dealer-Quote Data

The panels plot the cubic portion of the estimated year-end term premium functions for 30-day paper, given by

$$\hat{\alpha}_{m,yr,1} + \hat{\alpha}_{m,yr,2}\tau_m + \hat{\alpha}_{m,yr,3}\tau_m^2 + \hat{\alpha}_{m,yr,4}\tau_m^3$$

for $\tau = 0, 1, \ldots, 30$ and $m = 30$. When $\tau = 0$, the maturity date is December 26 and when $\tau = 30$ the maturity date is 30 days later (January 24 of the following year). The $\hat{\alpha}$. denote estimated coefficients from the appropriate rows in table A.3.

Panel A: 1989-1992



Panel B: 1993-1996

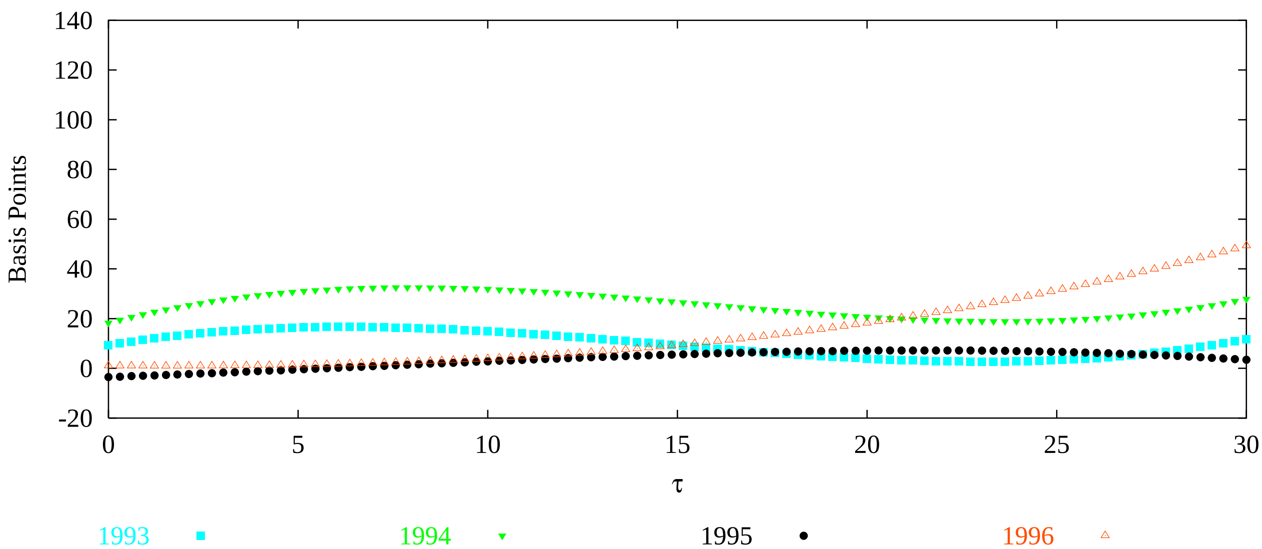
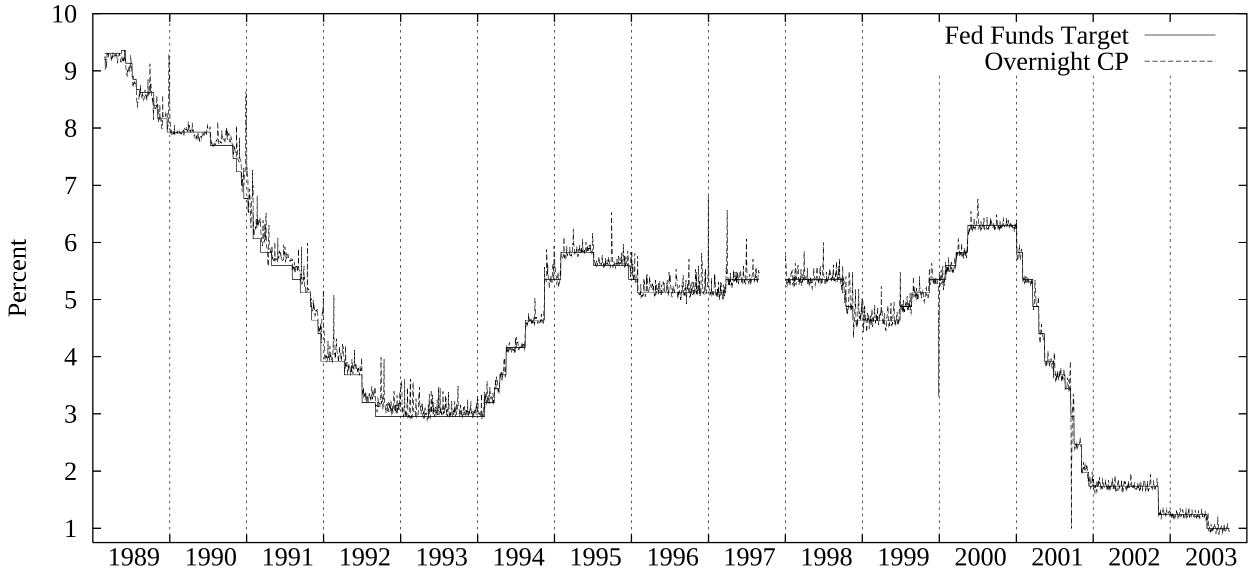


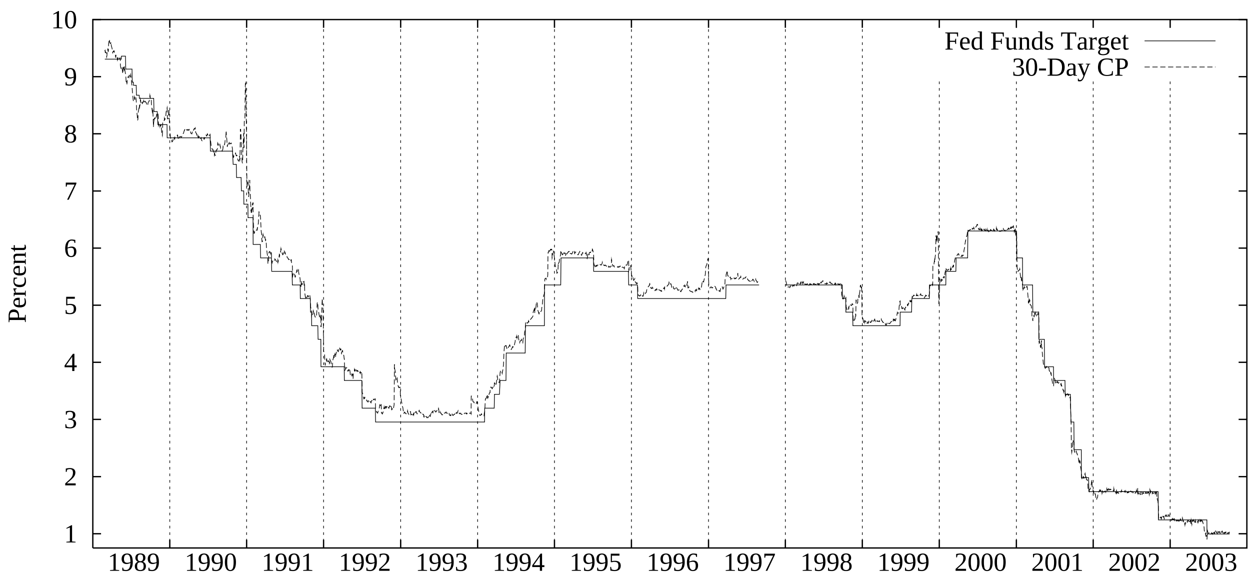
Figure 8: Commercial Paper Rates and the Target Rate for Federal Funds

The panels display rates on overnight and 30-day commercial paper issued by AA-rated nonfinancial firms, along with the target federal funds rate. Both panels show daily data covering the period from February 27, 1989 to August 1, 2003, with a break from August 30, 1997 to January 1, 1998.

Panel A: Overnight CP and Fed Funds Target Rates



Panel B: 30-Day CP and Fed Funds Target Rates



Source: Federal Reserve Board. The commercial paper rates are based on data from the Federal Reserve Bank of New York market survey through 1997 and data from the Depository Trust Company after 1997.

Table 1: Summary Statistics

This table displays univariate statistics for the commercial paper rates used in our empirical analysis. The upper two panels pertain to the transactions-based data from January 2, 1998 to August 1, 2003, for a total of 1,365 daily observations. The lower panel pertains to the dealer-quote data from February 27, 1989 to August 29, 1997, for a total of 2,046 daily observations. The means are expressed in percent; the spreads are expressed in basis points.

Transactions data, AA-rated

	Days to maturity					
	1	7	15	30	60	90
Mean yield	4.07	4.06	4.08	4.08	4.08	4.10
Std. dev.	1.80	1.80	1.82	1.83	1.84	1.86
Spread to overnight		-0.35	0.70	1.04	1.67	3.09
Std. dev.		10.60	15.41	18.09	21.34	24.92
Lag	Autocorrelations					
1	0.9959	0.9974	0.9979	0.9984	0.9985	0.9985
5	0.9849	0.9890	0.9903	0.9916	0.9924	0.9925
10	0.9791	0.9821	0.9822	0.9833	0.9846	0.9848

Transactions data, A2/P2-rated

	Days to maturity					
	1	7	15	30	60	90
Mean yield	4.28	4.31	4.34	4.38	4.43	4.47
Std. dev.	1.80	1.81	1.83	1.83	1.84	1.85
Spread to overnight		3.64	6.47	10.36	15.37	19.09
Std. dev.		13.78	20.14	23.96	25.62	28.49
Lag	Autocorrelations					
1	0.9956	0.9969	0.9975	0.9979	0.9977	0.9977
5	0.9843	0.9864	0.9868	0.9892	0.9898	0.9906
10	0.9779	0.9786	0.9756	0.9784	0.9802	0.9812

Dealer-quote data

	Days to maturity					
	1	7	15	30	60	90
Mean yield	5.43	5.47	5.49	5.51	5.55	5.58
Std. dev.	1.79	1.79	1.78	1.77	1.76	1.75
Spread to overnight		3.62	5.52	7.84	11.56	14.59
Std. dev.		14.17	16.93	17.74	20.75	24.18
Lag	Autocorrelations					
1	0.9962	0.9975	0.9978	0.9996	1.0014	1.0015
5	0.9907	0.9906	0.9926	0.9954	0.9963	0.9960
10	0.9943	0.9868	0.9911	0.9944	0.9939	0.9939

Table 2: Term Premia in the CP Market

This table reports the results from estimating equation 11. The term premium on the left-hand side of the equation is computed as the difference between the term CP rate at the indicated maturity and the average overnight CP rate for the horizon of the term CP rate:

$$r(m,t) - \frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_m + \epsilon_{m,t},$$

for $m = 7, 15, 30, 60, 90$ and $t = 1, 2, \ldots, T$. In the regressions using dealer-quote data, $T = 2,046 - m$. In the regressions using transactions data, $T = 1,365 - m$. The estimates of α_m are expressed in basis points. The t-statistics are adjusted for the overlap in the observations.

Transactions data, AA-rated

Days to maturity	α_m	t-stat.
7	-0.19	-0.46
15	1.57	1.96
30	3.50	2.88
60	7.32	4.14
90	11.93	4.51

Transactions data, A2/P2-rated

Days to maturity	α_m	t-stat.
7	3.77	6.14
15	7.27	5.89
30	12.76	6.30
60	20.94	6.95
90	27.85	7.13

Dealer-quote data

Days to maturity	α_m	t-stat.
7	4.33	8.48
15	6.58	7.38
30	9.80	7.83
60	15.45	7.67
90	20.35	7.25

Table 3: Tests of the Expectations Hypothesis

This table reports the results from estimating equation 12. The term premium on the left-hand side of the equation is computed as in table 2. The term spread on the right-hand side is computed as the difference between the term CP rate and the overnight CP rate. Equation 12, reproduced here for convenience, is given by:

$$r(m,t) - \frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_m + \beta_m \left(r(m,t) - r(1,t)\right) + \epsilon_{m,t}.$$

If the expectations hypothesis is true, then $\beta_m = 0$. If the pure expectations hypothesis is true, then $\alpha_m = \beta_m = 0$. The sample sizes and units are the same as in table 2. The columns labeled "Lower 95% crit." and "Upper 95% crit." display critical values of the t-statistic (corrected for the overlap in the observations) for a 95% test that $\beta_m = 0$, where the confidence intervals maintain the correct size of the test in the presence of a persistent regressor (Cavanagh et al. (1995)).

Transactions data, AA-rated

Days to maturity	α_m	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	-0.09	-0.23	0.29	1.56	-2.49	1.64
15	1.34	2.61	0.33	1.93	-2.40	1.64
30	3.20	3.41	0.29	2.20	-2.14	1.65
60	7.11	4.14	0.13	1.24	-1.76	1.82
90	11.79	4.39	0.04	0.35	-1.64	2.07

Transactions data, A2/P2-rated

Days to maturity	α_m	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	1.91	4.50	0.51	4.59	-2.38	1.64
15	3.40	4.52	0.60	4.61	-2.30	1.64
30	6.67	4.95	0.59	4.94	-2.02	1.65
60	13.52	4.16	0.48	3.79	-1.75	1.82
90	21.43	3.99	0.34	2.25	-1.64	1.99

Dealer-quote data

Days to maturity	α_m	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	2.23	4.78	0.58	3.95	-2.05	1.65
15	3.81	5.96	0.49	3.52	-2.27	1.64
30	6.77	8.42	0.38	4.93	-2.22	1.64
60	11.66	6.98	0.32	4.74	-1.88	1.69
90	16.32	5.79	0.27	4.30	-1.71	1.90

Table 4: Tests of the Expectations Hypothesis, Controlling for Time Variation in Term Premia

This table reports the results from estimating equation 13. The term premium and term spread are computed as in tables 2 and 3. Equation 13 is reproduced here for convenience:

$$r(m,t) - \frac{1}{m}\sum_{i=t}^{t+m-1} r(1,i) = \alpha_{m,0} + \beta_m(r(m,t) - r(1,t)) +$$
$$\sum_{yr}(\alpha_{m,yr,1} + \alpha_{m,yr,2}\tau_m + \alpha_{m,yr,3}\tau_m^2 + \alpha_{m,yr,4}\tau_m^3)DX_{m,yr,t} + \alpha_{m,yr,5}D_{yr,t} +$$
$$\alpha_{m,6}SX_{m,t} + \alpha_{m,7}S_t + \epsilon_{m,t},$$

The right-hand side includes dummy variables for shifts in term premia around year-end (the $DX.$ and $D.$) and for the effects of September 11, 2001 (the $SX.$ and $S.$); see the text for details and appendix A.2 for the estimates of the relevant coefficients. If the expectations hypothesis is true, then $\beta_m = 0$. The sample sizes and other details about the displayed results are the same as in the previous tables.

Transactions data, AA-rated

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	-0.90	-2.63	-0.14	-2.04	-2.38	1.64
15	0.14	0.33	-0.02	-0.35	-2.33	1.64
30	1.05	1.74	0.02	0.65	-1.91	1.64
60	3.26	3.63	-0.03	-0.73	-1.64	2.08
90	5.94	3.83	-0.08	-1.08	-1.64	2.35

Transactions data, A2/P2-rated

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	2.49	5.47	0.11	1.35	-2.42	1.64
15	4.47	6.42	0.11	1.47	-2.38	1.64
30	7.32	7.53	0.13	2.32	-1.84	1.76
60	13.11	5.69	0.01	0.19	-1.64	2.47
90	18.96	4.45	-0.10	-0.75	-1.64	2.81

Dealer-quote data

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	3.12	9.20	0.08	0.97	-2.44	1.64
15	4.38	8.95	0.11	2.74	-2.37	1.64
30	6.09	7.84	0.20	4.28	-2.16	1.64
60	10.13	6.70	0.20	3.89	-1.87	1.64
90	13.37	5.58	0.25	3.97	-1.73	1.94

Table 5: Goodness-of-Fit Measures

The table compares the adjusted-R^2 goodness-of-fit measure for the regressions in tables 3 and 4 above. The columns labeled "No Yr-end" show the goodness-of-fit measures for the regressions in table 3, while the columns labeled "Yr-end" show the goodness-of-fit measures for the corresponding regressions in table 4. The adjusted-R^2 figures are in percent.

	Transactions data					
	AA-rated		A2/P2-rated		Dealer-quote data	
Days to maturity	No Yr-end	Yr-end	No Yr-end	Yr-end	No Yr-end	Yr-end
7	7	45	26	48	34	63
15	16	60	43	72	31	65
30	16	66	45	78	20	54
60	4	61	28	73	15	52
90	0	60	14	67	11	49

Table 6: September 11 Coefficient Estimates

The table displays the estimates of the coefficients on the September 11 dummy variables in equation 13. The relevant portion of that equation is given by:

$$\ldots + \alpha_{m,6} SX_{m,t} + \alpha_{m,7} S_t + \ldots$$

where the dummy variable $SX_{m,t}$ takes the value one when the observation date t is before September 11, 2001, but the maturity date $t + m$ crosses September 11. The variable S_t equals one when the observation date t is between September 11 and September 18, 2001. The estimates of $\alpha_{m,6}$ and $\alpha_{m,7}$ are expressed in basis points.

AA-rated

Days to Maturity	$\alpha_{m,6}$	t-stat.	$\alpha_{m,7}$	t-stat.
7	-10.36	-6.82	86.21	18.28
15	12.89	1.87	47.95	10.07
30	17.89	2.72	51.97	22.73
60	25.18	3.55	48.29	33.48
90	28.83	2.74	51.95	20.17

A2/P2-rated

Days to Maturity	$\alpha_{m,6}$	t-stat.	$\alpha_{m,7}$	t-stat.
7	-16.81	-3.28	82.98	9.13
15	0.48	0.10	49.28	5.93
30	8.11	1.94	39.20	9.95
60	18.20	3.73	36.05	9.38
90	20.37	2.23	37.36	5.63

Table 7: Tests of the Expectations Hypothesis on Dealer-Quote Data; Early Period

The table reports results from estimating equation 13 with the dealer-quote data prior to February 4, 1994. The regression specification in the upper panel is identical to that in table 3, while the specification in the lower panel is identical to that in table 4. The sample sizes are $1,237 - m$.

Without Year-End Controls

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	2.36	6.36	0.69	5.61	-2.02	1.67
15	4.04	6.61	0.61	4.53	-2.06	1.64
30	7.43	6.51	0.50	5.39	-2.37	1.64
60	13.40	5.71	0.52	6.49	-2.15	1.64
90	19.18	5.11	0.48	6.23	-1.84	1.69

With Year-End Controls

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	2.57	7.08	0.15	1.35	-2.40	1.64
15	3.70	5.85	0.15	2.85	-2.49	1.64
30	6.06	5.73	0.31	5.06	-2.41	1.64
60	11.13	5.47	0.35	4.38	-2.28	1.64
90	15.51	5.03	0.43	4.50	-2.30	1.64

Table 8: Tests of the Expectations Hypothesis on Dealer-Quote Data; Late Period

The table reports results from estimating equation 13 with the dealer-quote data starting from February 4, 1994. The regression specification in the upper panel is identical to that in table 3, while the specification in the lower panel is identical to that in table 4. The sample sizes are $832 - m$.

Without Year-End Controls

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	4.11	7.36	0.12	1.27	-2.45	1.64
15	5.83	7.67	0.15	1.75	-2.39	1.64
30	7.05	5.80	0.21	2.18	-2.17	1.64
60	9.54	6.84	0.19	2.34	-2.11	1.64
90	10.39	6.85	0.23	3.64	-1.87	1.71

With Year-End Controls

Days to maturity	$\alpha_{m,0}$	t-stat.	β_m	t-stat.	Lower 95% crit.	Upper 95% crit.
7	4.73	9.15	-0.10	-1.66	-2.49	1.64
15	6.13	8.49	0.00	0.06	-2.32	1.64
30	7.40	7.27	0.06	1.02	-2.24	1.64
60	9.31	6.93	0.11	2.61	-2.10	1.64
90	9.35	4.38	0.21	4.17	-1.74	1.84

A.1 Robust Confidence Intervals

Cavanagh et al. (1995) develop formal results for testing the null hypothesis that $\gamma = \gamma_0$ for the recursive system:[31]

$$x_t = \mu_x + v_t, \quad (1 - \alpha L)b(L)v_t = \epsilon_{1t}, \tag{A.1}$$

$$y_t = \mu_y + \gamma x_{t-1} + \epsilon_{2t}, \tag{A.2}$$

where $b(L) = \sum_{i=0}^{k} b_i L^i$, $b_0 = 1$, and $\epsilon_t = (\epsilon_{1t}, \epsilon_{2t})'$ is a martingale difference sequence with $E(\epsilon_t \epsilon_t') = \sigma$ and $\text{corr}(\epsilon_{1t}, \epsilon_{2t}) = \delta$. In our implementation, y_t is the term premium and x_{t-1} is the term spread. The issue is that the term spread contains a large, perhaps unit, autoregressive root, modeled here as $\alpha = 1 + \frac{c}{T}$, where c is a fixed constant (the "local to unity" model).

Cavanagh et al. (1995) show that the t-statistic for testing $\gamma = \gamma_0$ has a limiting representation:

$$t_\gamma \to \delta \tau_{1c} + \sqrt{1 - \delta^2} z, \tag{A.3}$$

where $\tau_{1c} = \frac{\int J_c^\mu dB_1}{\sqrt{\int \left(J_c^\mu\right)^2}}$, J_c is the diffusion process defined by $dJ_c(s) = cJ_c(s)ds + dB_1(s)$ with $J_c(0) = 0$, B_1 is a standard Brownian motion, $J_c^\mu(s) = J_c(s) - \int_0^1 J_c(r)dr$, and z is a standard normal random variable that is independent of B_1 and J_c. The parameter δ is consistently estimated by the sample correlation between $\hat{\epsilon}_{1t}$ and $\hat{\epsilon}_{2t}$.

We construct a conservative test of $\gamma = \gamma_0$ using the extrema of the asymptotic local-to-unity critical values of t_γ. Formally, letting $d_{t_\gamma, c, \eta}$ denote the $100\eta\%$ quantile of the distribution of $\delta \tau_{1c} + \sqrt{1 - \delta^2} z$ for a given value of δ, we compute the following "sup-bound

[31] This appendix is a very brief distillation of a portion of Cavanagh et al. (1995). For more details and additional useful results, the reader is referred to the source.

confidence interval" by Monte Carlo simulation:

$$(\underline{d}_\eta, \bar{d}_\eta) = \left(\inf_c d_{t_\gamma,c,\eta}, \sup_c d_{t_\gamma,c,\eta} \right). \tag{A.4}$$

Intuitively, the procedure amounts to picking the most conservative confidence limits across a sequence of values of c.

The following Matlab code details our implementation of the Cavanagh et al. (1995) procedure for calculating the conservative sup-bound confidence intervals:[32]

```
function [tstat, crit] = supbound_dum(x, y, d, p);
%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%
% Computes sup-bound critical values in a regression of            %
% y(t) on d(t) and x(t-1,1); the routine will also add a constant. %
% Autocorrelation robust standard errors are used with truncation  %
% p.  d(t) is a set of calendar dummy variables with number of     %
% rows equal to the number of rows of y and x and any number of    %
% columns (subject to computational limits.)                       %
% It is assumed that the number of observations is large enough    %
% that the number of rows of d can serve as the number of Monte    %
% Carlo iterations in the calculation of the supbounds.            %
% NOTE: Before calling this routine, set the seed for Matlab's     %
% random number generator if you want results that can be          %
% replicated exactly.                                              %
%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%%
t=size(x,1);
q=y(1:t-1);
w=[ones(t-1,1) d(1:t-1,:) x(1:t-1)];
k=size(w,2);
bhat=inv(w'*w)*w'*q;
u=q-w*bhat;
z=kron(u,ones(1,k)).*w;
for j=1:p+1;
  lb(:,:,j)=((p+2-j)/(p+1))*z(j:t-1,:)'*z(1:t-j,:);
end; %NW
sigma=sum(lb,3)+(sum(lb,3)')-lb(:,:,1);
clear lb;
a=inv(w'*w)*sigma*inv(w'*w);
```

[32] This program incorporates minor modifications to a program kindly provided to the authors by Jonathan Wright. We modified Wright's program to allow for calendar dummy variables in the regression specification.

```
w=[ones(t-1,1) d(2:t,:) x(1:t-1)];
q=x(2:t,:);
v=q-(w*inv(w'*w)*w'*q);
z=[u v];
for j=1:p+1;
  lb(:,:,j)=((p+2-j)/(p+1))*z(j:t-1,:)'*z(1:t-j,:);
end; %NW
sigma=sum(lb,3)+(sum(lb,3)')-lb(:,:,1);
delta=sigma(1,2)/sqrt(sigma(1,1)*sigma(2,2));
tstat=bhat(k)/sqrt(a(k,k));
s=size(d,1);
dum=[ones(s,1) d];
dp=dum*inv(dum'*dum)*dum';
b1=zeros(s,1000);
sa=zeros(1000,1);
stat=zeros(1000,1);
sb1=zeros(3,31);
sb2=zeros(3,31);
u=randn(s,31000)./sqrt(s);
eps=randn(1000,31);
for k=1:31;
  alpha=1+((1-k)/1000);
  b1=filter(1,[1;-alpha],u(1:s,(k-1)*1000+1:k*1000));
  b1=b1-dp*b1;
  sa=sum(b1(1:s-1,1:1000).*u(2:s,(k-1)*1000+1:k*1000))./sqrt(mean((b1.^2)));
  stat=(delta.*sa')+(sqrt(1-(delta^2)).*eps(1:1000,k));
  stat=sort(stat);
  sb1(1:3,k)=stat([10;50;100]);
  sb2(1:3,k)=stat([990;950;900]);
end;
sb1(1:3,32)=norminv([0.01 0.05 0.10]');
sb2(1:3,32)=norminv([0.99 0.95 0.90]');
crit=[[0.01 0.05 0.10]' min(sb1')' max(sb2')'];
```

A.2 Dummy Variable Coefficient Estimates

The tables in this appendix display the estimated coefficients on the dummy variables in equation 13 that control for year-end and September 11 effects. Tables A.1-A.3 report the results from the transactions data for AA-rated firms, the transaction data for A2/P2-rated firms, and the dealer-quote data, respectively. Tables A.4 and A.5 show the estimates for the sample splits on the dealer-quote data. Below each coefficient estimate we report its t-statistic, where the underlying standard errors are adjusted for the overlap in the observations.

Table A.1: Dummy Variable Coefficient Estimates for Tests of the Expectations Hypothesis, Controlling for Time-Varying Term Premia, Transactions Data, AA-Rated Firms

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$	$\alpha_{m,6}$	$\alpha_{m,7}$
7	1998	0.64	21.29	4.65	-1.0915	4.37		
		0.09	13.78	7.18	-11.1876	1.02		
7	1999	82.03	33.10	-0.96	-0.9390	-4.88		
		10.14	5.52	-0.83	-3.1403	-2.96		
7	2000	-10.36	1.71	0.87	-0.1747	7.54		
		-1.09	2.05	2.45	-3.2217	0.85		
7	2001	-0.72	3.21	0.20	-0.1088	0.85	-10.36	86.21
		-0.82	17.03	2.27	-5.4875	1.21	-6.82	18.28
7	2002	1.38	1.81	0.02	-0.0444	2.54		
		1.66	6.99	0.30	-11.1876	4.65		
15	1998	21.06	8.61	-0.18	-0.0262	1.74		
		15.78	7.80	-0.97	-2.7824	1.52		
15	1999	50.92	20.97	-0.68	-0.0684	0.98		
		7.42	9.78	-0.76	-1.2795	0.96		
15	2000	6.13	0.49	-0.12	0.0188	4.97		
		4.14	2.30	-0.90	2.2146	0.86		
15	2001	-0.76	1.84	0.12	-0.0184	0.94	12.89	47.95
		-0.55	7.08	0.89	-1.4048	1.87	1.87	10.07
15	2002	-1.12	0.32	0.05	-0.0033	2.45		
		-2.18	7.30	2.93	-2.7824	4.54		
30	1998	12.57	1.08	0.13	-0.0050	8.48		
		7.17	1.02	1.09	-1.7641	4.29		
30	1999	22.91	1.47	0.54	-0.0211	4.58		
		6.36	0.56	2.06	-3.4145	1.70		
30	2000	4.00	0.51	0.11	-0.0034	4.85		
		6.36	4.50	6.34	-5.8386	1.26		
30	2001	8.89	-1.01	-0.02	0.0027	-4.71	17.89	51.97
		9.08	-3.66	-0.36	1.9994	-6.53	2.72	22.73
30	2002	1.73	0.05	0.02	-0.0008	0.49		
		2.86	0.77	3.57	-1.7641	0.92		
60	1998	12.23	1.58	-0.04	0.0002	6.13		
		5.49	2.51	-1.27	0.6707	4.99		
60	1999	23.51	1.44	-0.01	-0.0003	0.76		
		4.77	1.11	-0.09	-0.4062	0.36		
60	2000	1.49	0.78	0.01	-0.0002	8.33		
		1.51	5.82	1.08	-1.3811	2.66		
60	2001	6.48	0.36	-0.02	0.0003	-5.86	25.18	48.29
		2.92	0.69	-0.85	0.7255	-4.77	3.55	33.48
60	2002	20.95	-2.31	0.08	-0.0007	-1.35		
		13.58	-8.77	6.37	0.6707	-1.72		
90	1998	5.85	-0.01	0.01	-0.0001	2.30		
		1.92	-0.06	3.28	-4.0165	1.87		
90	1999	29.24	2.68	-0.08	0.0005	-5.14		
		5.43	2.94	-2.66	2.3860	-1.21		
90	2000	4.44	0.17	0.02	-0.0002	10.13		
		2.87	1.24	3.80	-4.0860	2.52		
90	2001	14.93	0.18	-0.01	0.0001	-5.41	28.83	51.95
		3.91	0.90	-2.22	2.1320	-3.17	2.74	20.17
90	2002	17.85	1.03	-0.04	0.0004	-5.45		
		6.43	3.13	-4.40	-4.0165	-3.11		

Table A.2: Dummy Variable Coefficient Estimates for Tests of the Expectations Hypothesis, Controlling for Time-Varying Term Premia,Transactions Data, A2/P2-Rated Firms

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$	$\alpha_{m,6}$	$\alpha_{m,7}$
7	1998	4.84	23.99	4.57	-1.1071	1.86		
		0.51	9.19	5.34	-7.7745	0.30		
7	1999	53.70	24.41	0.04	-0.7402	-0.66		
		9.58	6.67	0.08	-4.7307	-0.39		
7	2000	24.21	17.70	0.81	-0.5492	5.72		
		1.81	8.39	1.43	-4.1092	0.47		
7	2001	5.58	0.36	1.67	0.7253	3.63	-16.81	82.98
		2.08	0.50	37.93	10.2709	1.31	-3.28	9.13
7	2002	12.19	0.08	-0.57	0.1053	-0.85		
		6.98	0.34	-6.64	-7.7745	-0.71		
15	1998	24.51	5.76	1.62	-0.1289	-1.12		
		4.95	1.32	1.56	-2.3771	-0.33		
15	1999	44.17	20.47	0.40	-0.1349	-1.13		
		10.55	12.34	1.17	-6.1224	-0.90		
15	2000	11.07	10.36	1.11	-0.1039	-16.36		
		4.98	12.76	4.63	-6.1429	-1.59		
15	2001	8.92	5.69	-0.06	-0.0258	12.81	0.48	49.28
		5.98	8.60	-0.42	-1.4789	5.03	0.10	5.93
15	2002	-4.17	3.25	0.07	-0.0152	-1.34		
		-4.33	11.88	1.73	-2.3771	-1.59		
30	1998	46.17	3.95	-0.17	0.0013	10.98		
		15.40	3.55	-1.61	0.5043	6.70		
30	1999	33.66	4.12	0.27	-0.0141	11.24		
		7.07	1.13	0.75	-1.6416	2.47		
30	2000	-0.53	5.18	0.12	-0.0065	-28.58		
		-0.43	15.40	2.59	-4.7210	-4.84		
30	2001	21.34	3.63	-0.32	0.0080	2.69	8.11	39.20
		14.33	6.06	-4.24	3.5781	1.86	1.94	9.95
30	2002	10.47	0.75	0.01	-0.0009	6.12		
		5.86	1.06	0.10	0.5043	4.97		
60	1998	45.58	4.41	-0.13	0.0010	8.46		
		19.75	6.01	-4.28	2.6102	4.75		
60	1999	31.21	3.08	-0.06	0.0002	4.65		
		5.64	2.33	-0.93	0.2066	1.84		
60	2000	-12.85	1.92	-0.04	0.0010	-24.00		
		-6.57	5.40	-2.47	3.9108	-5.64		
60	2001	33.55	1.79	-0.07	0.0006	-8.97	18.20	36.05
		19.27	4.96	-3.73	2.6894	-4.65	3.73	9.38
60	2002	43.89	-3.17	0.12	-0.0011	-1.17		
		19.66	-9.98	8.44	2.6102	-0.63		
90	1998	-8.82	3.08	-0.02	-0.0001	13.77		
		-1.19	5.51	-1.35	-0.7041	3.22		
90	1999	37.83	3.41	-0.08	0.0005	-3.36		
		5.88	2.77	-2.07	1.6526	-0.54		
90	2000	-11.78	0.63	-0.01	0.0002	-15.33		
		-3.69	2.24	-1.06	2.9103	-2.85		
90	2001	10.60	2.83	-0.05	0.0002	-12.15	20.37	37.36
		1.57	5.97	-4.89	2.8729	-3.59	2.23	5.63
90	2002	21.58	1.34	-0.03	0.0002	6.21		
		5.74	4.33	-3.39	-0.7041	2.73		

Table A.3: Dummy Variable Coefficient Estimates for Tests of the Expectations Hypothesis, Controlling for Time-Varying Term Premia, Dealer-Quote Data

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$
7	1989	-35.82	10.66	4.45	-0.4876	3.27
		-5.74	3.06	10.08	-1.1031	1.71
7	1990	-20.99	61.01	9.03	-2.0625	9.15
		-1.44	11.69	7.85	-9.6117	1.62
7	1991	-7.07	49.01	4.78	-1.5415	2.99
		-1.67	10.43	11.22	-9.5760	1.09
7	1992	-26.53	6.94	2.73	-0.5276	16.93
		-16.47	6.93	18.15	-13.9015	15.81
7	1993	-8.65	4.61	0.82	-0.1586	6.19
		-4.99	6.99	8.28	-7.1808	4.17
7	1994	-7.03	7.18	0.48	-0.2235	10.44
		-5.68	5.58	3.61	-3.8704	11.31
7	1995	9.98	6.02	-0.27	-0.2164	3.44
		4.69	5.00	-5.93	-4.4006	4.34
7	1996	-7.26	7.25	1.78	-0.1833	8.31
		-2.68	2.51	4.97	-1.5691	3.58
15	1989	2.98	-0.28	-0.23	0.0420	1.59
		1.45	-0.74	-1.19	0.2203	1.23
15	1990	16.06	14.58	0.91	-0.0923	22.98
		2.83	11.70	2.59	-4.0051	10.04
15	1991	-7.77	5.13	1.25	-0.0833	3.06
		-2.79	7.85	6.33	-6.3428	1.16
15	1992	12.41	6.95	-1.34	0.0577	12.21
		5.08	3.89	-3.18	2.6861	5.37
15	1993	1.56	2.07	-0.25	0.0121	5.44
		1.45	7.94	-3.29	2.9951	4.52
15	1994	3.81	2.56	-0.18	0.0043	13.70
		1.71	5.77	-1.00	0.3634	6.09
15	1995	7.99	1.22	-0.42	0.0295	3.25
		4.96	5.90	-3.51	3.3557	3.45
15	1996	6.01	2.99	-0.15	0.0146	8.04
		2.58	7.85	-0.90	1.3917	6.43
30	1989	-13.51	0.86	-0.01	0.0017	-3.30
		-9.58	2.26	-0.19	0.0290	-2.53
30	1990	23.19	-2.28	0.18	0.0004	32.25
		6.26	-1.31	0.84	0.0649	18.75
30	1991	25.32	0.23	-0.07	0.0036	2.85
		16.83	0.29	-0.75	1.6264	1.14
30	1992	25.85	5.07	-0.49	0.0101	13.14
		10.65	7.78	-8.12	7.4137	9.72
30	1993	9.33	2.57	-0.26	0.0059	3.04

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Table A.3: (continued)

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$
		7.19	4.63	-5.08	5.3451	2.62
30	1994	18.33	4.04	-0.34	0.0072	13.28
		11.82	7.03	-7.32	7.5515	8.38
30	1995	-3.48	0.53	0.02	-0.0010	4.18
		-4.20	3.05	0.78	-1.2961	5.00
30	1996	1.02	-0.10	0.03	0.0009	3.96
		0.82	-0.31	0.85	0.8156	2.80
60	1989	1.27	-0.91	0.01	0.0002	-6.22
		0.75	-5.98	1.82	0.0231	-3.87
60	1990	7.56	3.29	-0.10	0.0011	46.31
		2.66	3.94	-2.32	2.0314	16.54
60	1991	23.05	0.71	-0.02	0.0002	-6.94
		9.85	0.99	-0.50	0.4289	-2.50
60	1992	4.56	2.39	-0.05	0.0001	8.74
		2.66	7.59	-3.39	0.6062	5.60
60	1993	2.42	2.07	-0.08	0.0007	-9.23
		1.59	11.85	-9.62	7.5475	-5.31
60	1994	-18.72	1.31	0.03	-0.0008	13.06
		-4.80	3.97	2.06	-4.4086	6.60
60	1995	-6.27	-0.01	0.02	-0.0002	10.85
		-4.13	-0.05	3.16	-3.1137	6.71
60	1996	-9.35	0.33	-0.02	0.0004	4.24
		-6.09	1.42	-1.18	2.1919	2.79
90	1989	15.42	-0.91	0.01	0.0001	-8.30
		3.95	-1.93	0.34	0.0048	-2.78
90	1990	14.00	0.31	0.02	-0.0002	55.88
		4.41	1.03	1.74	-2.0219	19.71
90	1991	32.12	1.10	-0.03	0.0002	-9.22
		11.13	3.28	-2.78	2.1752	-2.85
90	1992	-5.10	-0.28	0.04	-0.0004	9.50
		-1.47	-1.24	7.27	-8.7713	4.61
90	1993	-1.32	0.43	-0.01	0.0000	-13.63
		-0.50	1.45	-0.83	0.4195	-4.72
90	1994	-5.33	-1.47	0.06	-0.0004	13.15
		-1.15	-3.46	4.30	-4.1200	3.89
90	1995	-3.83	-0.52	0.02	-0.0002	5.27
		-1.43	-3.45	4.53	-4.2905	2.25
90	1996	-4.42	-0.17	-0.00	0.0001	-0.93
		-1.43	-0.59	-0.27	1.2788	-0.42

Table A.4: Dummy Variable Coefficient Estimates for Tests of the Expectations Hypothesis, Controlling for Time-Varying Term Premia, Dealer-Quote Data Prior to February 4, 1994

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$
7	1989	-38.83	9.47	4.57	-0.4325	3.99
		-5.53	2.57	10.13	10.1282	2.04
7	1990	-26.19	56.73	8.79	-1.9341	9.41
		-1.74	8.84	8.21	8.2078	1.68
7	1991	-9.31	45.36	4.51	-1.4234	3.88
		-1.93	7.51	8.75	8.7497	1.35
7	1992	-27.35	6.05	2.65	-0.4991	17.68
		-14.74	4.51	13.58	13.5799	14.96
7	1993	-8.87	4.05	0.77	-0.1433	6.45
		-5.34	4.59	6.42	6.4164	4.24
15	1989	3.46	-0.60	-0.29	0.0497	2.36
		1.58	-1.31	-1.45	-1.4534	1.72
15	1990	13.86	13.48	0.88	-0.0847	23.29
		2.28	8.91	2.53	2.5253	10.12
15	1991	-7.36	4.68	1.18	-0.0765	4.07
		-2.64	6.37	5.80	5.7950	1.47
15	1992	12.08	6.40	-1.31	0.0578	13.19
		4.75	3.60	-3.20	-3.2009	5.63
15	1993	1.75	1.95	-0.26	0.0130	5.86
		1.39	7.28	-3.52	-3.5215	4.50
30	1989	-12.11	0.87	-0.06	0.0037	-3.00
		-7.21	1.81	-0.84	-0.8397	-1.92
30	1990	21.47	-2.59	0.14	0.0022	30.46
		6.25	-1.75	0.78	0.7767	14.43
30	1991	23.82	-0.01	-0.07	0.0042	3.50
		13.38	-0.01	-0.95	-0.9530	1.37
30	1992	22.24	4.26	-0.43	0.0092	13.79
		7.80	6.30	-7.41	-7.4077	8.38
30	1993	7.88	1.84	-0.20	0.0048	2.22
		4.85	3.23	-4.00	-4.0044	1.56
60	1989	0.73	-0.76	0.00	0.0004	-5.99
		0.32	-2.26	0.10	0.0999	-3.01
60	1990	5.14	3.58	-0.13	0.0015	43.44
		1.42	3.80	-2.62	-2.6181	12.27
60	1991	21.78	0.73	-0.03	0.0004	-6.20
		8.07	0.95	-0.69	-0.6901	-2.05
60	1992	0.47	1.91	-0.04	0.0001	8.89
		0.18	4.98	-2.55	-2.5461	4.54
60	1993	-1.16	1.83	-0.07	0.0007	-11.40
		-0.47	8.55	-7.98	-7.9831	-4.86
90	1989	14.87	-0.77	-0.00	0.0001	-8.08

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Table A.4: (continued)

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$
		3.13	-1.40	-0.13	-0.1321	-2.20
90	1990	11.41	0.58	0.00	-0.0001	53.54
		2.46	1.21	0.27	0.2705	13.23
90	1991	28.56	1.30	-0.04	0.0003	-8.14
		7.85	3.21	-2.94	-2.9418	-2.13
90	1992	-2.07	-0.69	0.04	-0.0003	8.89
		-0.46	-2.19	6.02	6.0165	3.40
90	1993	-4.12	0.29	-0.01	0.0001	-17.36
		-1.16	0.83	-0.82	-0.8218	-4.44

Table A.5: Dummy Variable Coefficient Estimates for Tests of the Expectations Hypothesis, Controlling for Time-Varying Term Premia, Dealer-Quote Data After February 4, 1994

Days to Maturity	Year	$\alpha_{m,yr,1}$	$\alpha_{m,yr,2}$	$\alpha_{m,yr,3}$	$\alpha_{m,yr,4}$	$\alpha_{m,yr,5}$
7	1994	-5.54	9.94	0.70	-0.3386	9.95
		-5.31	10.35	6.75	6.7534	11.58
7	1995	14.45	8.74	-0.39	-0.3186	1.54
		8.23	9.00	-6.58	-6.5841	1.50
7	1996	-5.46	13.02	2.33	-0.4017	8.30
		-2.17	5.98	8.65	8.6514	4.18
15	1994	4.35	3.16	-0.15	-0.0008	12.78
		1.76	6.15	-0.80	-0.7989	5.16
15	1995	5.93	1.47	-0.31	0.0210	1.25
		2.92	5.91	-2.00	-1.9973	1.09
15	1996	4.28	3.71	0.06	-0.0034	7.11
		1.65	8.04	0.35	0.3464	5.32
30	1994	20.76	5.43	-0.44	0.0087	14.58
		12.05	7.60	-7.61	-7.6141	8.59
30	1995	-4.50	0.56	0.02	-0.0009	2.29
		-4.66	3.86	0.92	0.9216	2.12
30	1996	-1.50	0.24	0.07	-0.0009	5.50
		-1.04	0.90	2.33	2.3348	4.33
60	1994	-11.24	0.89	0.05	-0.0010	16.79
		-3.78	3.06	3.94	3.9426	10.78
60	1995	-5.05	0.01	0.02	-0.0002	10.99
		-4.00	0.06	3.00	3.0010	7.12
60	1996	-8.07	0.18	-0.00	0.0002	6.70
		-6.32	1.25	-0.41	-0.4083	5.84
90	1994	1.47	-1.47	0.06	-0.0004	18.22
		0.45	-3.05	3.83	3.8291	7.08
90	1995	0.03	-0.49	0.02	-0.0002	8.06
		0.01	-3.49	4.69	4.6876	3.56
90	1996	0.84	-0.30	0.00	0.0000	3.20
		0.40	-1.40	0.34	0.3434	2.08

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