ELSEVIER Journal of Financial Economics 47 (1998) 315–337

JOURNAL OF Financial ECONOMICS

Macroeconomic news and bond market volatility[1]

Charles M. Jones[a], Owen Lamont[b,*], Robin L. Lumsdaine[c]

[a] *Graduate School of Business, Columbia University, New York, NY 10027, USA*
[b] *Graduate School of Business, University of Chicago, Chicago, IL 60637, USA*
[c] *Department of Economics, Brown University, Providence, RI 02912, USA*

Received 8 November 1996; received in revised form 28 May 1997

Abstract

We examine the reaction of daily Treasury bond prices to the release of U.S. macroeconomic news. These news releases (of employment and producer price index data) are of interest because they are released on periodic, preannounced dates and because they are associated with substantial bond market volatility. We investigate whether these nonautocorrelated announcements give rise to autocorrelated volatility. We find that announcement-day volatility does not persist at all, consistent with the immediate incorporation of information into prices. We also find a risk premium on these release dates. © 1998 Elsevier Science S.A. All rights reserved.

JEL classification: C32; G12

Keywords: Volatility; Information; News; GARCH; Risk

* Corresponding author. Tel.: 773/702-6414; fax: 773/702-0458; e-mail: owen.lamont@gsb.uchicago.edu.

[1] We thank Walter Toshi Baily, Bob Korajczyk, Jim Poterba, Mark Watson, seminar participants at the University of Chicago, Columbia University, Cornell University, and the University of Montreal, and especially Ludger Hentschel (the referee) for helpful comments. We also thank Mark Mitchell for supplying data, and Amy C. Ko and Sydney Ludvigson for research assistance. Lamont was supported by the FMC Faculty Research Fund at the Graduate School of Business, University of Chicago. A portion of this research was completed while Lumsdaine was a National Fellow at the Hoover Institution. We also thank the Financial Research Center at Princeton University for support. A previous version of this paper circulated as 'Public Information and the Persistence of Bond Market Volatility'.

PII S0304-405X(97)00047-0

1. Introduction

Asset prices are volatile, and this volatility is predictable over time. Asset returns are also predictable over time. Financial economists agree about these facts, but they do not agree on their implications. The sources of day-to-day volatility remain elusive. A large percentage of the variation in prices appears unexplainable, and identifiable news events do not appear to drive much of the volatility of prices (e.g., Roll, 1988; Cutler et al., 1989). If asset prices and expected returns are driven by macroeconomic risks, the identity of these macroeconomic risks remains a mystery (e.g., Cochrane and Hansen, 1992).

A fundamental obstacle to testing asset pricing theories is our inability to accurately measure 'news' and 'macroeconomic risk'. Therefore, in this paper we focus on a type of news and risk that is known, and test whether the volatility of asset prices and movements in risk premiums correspond to the known properties of this news and risk. Specifically, we examine the reaction of daily bond returns to the release of macroeconomic news. We investigate the response of Treasury bond prices to regularly scheduled U.S. government releases of the producer price index (PPI) and employment data. The key features of these announcements are that they are not clustered in time but are exogenously released on periodic, preannounced dates, and they are known to cause substantial bond market volatility (e.g., McQueen and Roley, 1993; Ederington and Lee, 1993). An extensive literature (e.g., Cornell, 1983) also examines money supply announcements, which we do not study here.

These announcements enable us to address two questions in this paper. First, is public information about the macroeconomy immediately reflected in bond prices, or do its effects on volatility persist over the course of several days? Second, do bonds earn positive risk premiums on these dates, when they are exposed to macroeconomic risk? We answer these questions by looking at the second and first moments, respectively, of the time series of Treasury bond returns.

Our main goal is to study how public information about macroeconomic fundamentals moves asset prices. First, we study the sources of persistent volatility in asset prices. It is a well-documented fact that volatility in financial markets is correlated over time (for a review of recent work, see Bollerslev et al., 1992; for an earlier discussion, see Fama, 1970). Although remarkable progress has been made in modeling this process empirically, relatively little is known about *why* financial market volatility is autocorrelated. Bollerslev et al. (1992) note, "While serial correlation in conditional second moments is clearly a property of speculative prices, a systematic search for the causes of this serial correlation has only recently begun" (p. 20).

Since volatility is equivalent to information flow in a large class of models (e.g., Ross, 1989), one possible explanation is that public information arrives in clusters. Such autocorrelation is plausible; we know from everyday experience that publicly observable events do not occur independently over time. For example, suppose the President proposes a new tax bill to Congress on Monday. On Tuesday, the Congress reacts with a counterproposal, and so forth. Thus, if the news-generating process has autocorrelated volatility, we would expect financial market prices also to have autocorrelated volatility. This volatility predictability is perfectly consistent with efficient markets.

As a preliminary exercise, we find some empirical support for this explanation. Mitchell and Mulherin (1994) collect an index of news events based on headline widths on the front page of the *New York Times*, 1983–1990. They construct a daily database defining major news events as those with headlines that are at least three columns wide. Using this database (which the authors kindly provided us), we create a dummy variable equal to one on days on which major news events occur and zero otherwise. This daily variable has a first-order autocorrelation coefficient of about 0.20. This positive autocorrelation is significant. Using a nonparametric runs test, we can reject the hypothesis that news is distributed randomly across weeks (the Z-statistic testing for random runs is about eight). From this example there is evidence that the news-generating process is in fact positively autocorrelated at daily frequencies.

Alternatively, autocorrelated volatility could arise for other reasons. Perhaps things other than news move prices; random changes in liquidity or random 'sentiment' shocks may slow the incorporation of news into prices. Perhaps prices do respond immediately to news, but incorrectly: systematic cognitive error, by investors who are not fully rational, distorts the role of news through under- or overreaction. For example, Klibanoff et al. (1998) use observable news to examine whether there is under- or overreaction in closed-end fund prices. Last, perhaps news is only gradually incorporated into asset prices, as differently informed agents digest the information, observe market volume and prices, and draw inferences about each other's private information. For example, Brock and LeBaron (1996) show how learning can give rise to positively autocorrelated volatility even when fundamentals follow a homoskedastic random walk. Some empirical evidence suggests these alternatives. For example, French and Roll (1986) find that stock market volatility is lower when the market is closed, even if businesses are open.

Our strategy is to focus on information releases that are not autocorrelated. Specifically, we explore whether shocks to bond volatility on macroeconomic announcement days are as persistent as shocks on nonannouncement days. If announcement shocks do not persist, it would suggest that market prices quickly incorporate public information and that the trading process does not inherently generate persistent volatility in response to news. On the other hand, strong persistence of announcement shocks would suggest an alternative explanation: some feature of the trading (or information-gathering) process itself causes volatility to be autocorrelated, regardless of the nature of the news.

The fact that asset prices respond at all to macroeconomic releases of PPI and employment data is evidence against the 'strong' form efficiency of markets. If market prices fully reflected all private information, then these government announcements would not be news at all. For example, the PPI is simply the result of a survey of prices taken by the Bureau of Labor Statistics. If the financial market equilibrium were fully revealing, publishing a survey could not move prices. For example, Huberman and Schwert (1985) study Israeli indexed bonds and estimate that up to 15% of new information about inflation is not incorporated until the announcement of the price index.

The timing of government news releases is exogenous to financial markets. We use a dummy variable equal to one on the day that the government announces macroeconomic news. Since this news is released once a month, this dummy variable is negatively autocorrelated. A key assumption we make is that the news released on announcement dates is a one-time lump of exogenous news.

A possible problem with our identifying assumption is that other agents (e.g., policymakers) may react to this news after some delay. One scenario is that the Federal Reserve, upon observing the release, changes its policy on the subsequent day (even though the Fed typically learns the contents of the report on the night prior to release). To explore this possibility, we examine whether Fed policy changes tend to occur subsequent to announcement dates. Specifically, we compare announcement dates to dates on which the Fed changed its target interest rate, as documented in Rudebusch (1995). We find no evidence that Fed target changes are more likely to occur on days immediately following announcement days (using Rudebusch's sample period of 1974–1979 and 1984–1992). Target changes occur on 5.3% of the days in the sample, but on only 5.0% of the post-announcement days. Naturally, employment and PPI data influence later behavior of the Fed. However, since we focus on days immediately following announcements, this finding is consistent with no delayed reaction, at least on the part of the Fed and for the time horizons we study.

To model volatility, we develop a variant of the autoregressive conditional heteroskedasticity (ARCH) framework of Engle (1982). Engle finds that quarterly U.K. inflation has autocorrelated volatility. Since interest rates contain a component related to expected future inflation, we would naturally expect interest rates also to have autocorrelated volatility. Engle, Lilien, and Robins (1987) document this by estimating a variant of an ARCH model using monthly three- and six-month Treasury bill interest rates. Bollerslev et al. (1988) use quarterly data on 20-year Treasury bonds in their multivariate linear GARCH(p,q) model and find that conditional covariances are a significant determinant of risk premiums. Bollerslev et al. (1992) summarize previous ARCH literature that has used interest rate data. In this paper, we use the GARCH(1,1) model (Bollerslev, 1986) as the starting point for our investigation of the impact of public information releases. We then develop alternative specifications that allow announcement-day and nonannouncement-day shocks to affect conditional variance differently.

Our second goal in this paper is to test whether Treasury bonds earn higher expected returns when exposed to greater macroeconomic risks. Previous studies have estimated time-varying risk premiums as a function of the estimated conditional volatility. Such ARCH-in-mean (ARCH-M) results from Engle et al. (1987) and French et al. (1987) suggest that, in the context of a specific model of time-varying conditional volatility, periods of high ex ante volatility are also periods of high ex ante returns. Unfortunately, inferences drawn on the basis of ARCH-M models have been found to be highly susceptible to model misspecification (e.g., Pagan and Sabau, 1991). Therefore, we do not estimate an ARCH-M model, but instead focus only on what is clearly known in advance by market participants: bond market volatility is high on announcement days. We test whether such days exhibit high expected returns.

The paper is organized as follows. In Section 2, we examine the data and present preliminary ordinary-least squares (OLS) regressions. This section documents that, as in previous research, PPI and employment announcements have large contemporaneous effects on bond market volatility. It also shows that bonds earn significantly higher excess returns on announcement days. In Section 3, we develop models of daily volatility persistence. We document that daily bond market volatility is highly persistent. We test whether announcement-day shocks to volatility are as persistent as nonannouncement-day shocks. We find that announcement-day volatility shocks are significantly different from other shocks in that they exhibit no persistence at all. Section 4 presents conclusions and suggestions for future work.

2. Data and preliminary analysis

In this study we examine daily returns on five-, ten-, and 30-year Treasury bonds. We choose these assets because of our interest in macroeconomic announcements, which we know have a material impact on the Treasury bond market. We calculate excess returns on holding Treasury bonds over the spot rate (assumed equal to the rate on three-month T-bills) for the period 1979–1995. We use a long time period because Monte Carlo evidence in Hong (1987) and Lumsdaine (1995) emphasizes the need for a large amount of data in maximum likelihood estimation of models with conditional heteroskedasticity.

We calculate returns using the Federal Reserve's constant maturity interest rate series. Returns are calculated from the published yields using a hypothetical bond with the stated maturity and a coupon equal to the yield, thus trading at par or face value [similar to the method used by Ibbotson and Associates (1994), for example]. We calculate an end-of-period price on this bond using the next day's yield. Total returns equal capital appreciation plus the excess income over the short rate that accrues over the holding period, which varies from one to four days due to weekends and holidays. Employment announcement dates and PPI announcement dates were supplied to us by the Bureau of Labor Statistics.

Our sample runs from October 9, 1979 (following the adoption of monetary targets by the Federal Reserve) to December 31, 1995. We choose to start our sample in 1979 for two reasons. First, we find evidence of an October 1979 structural break in the time series, and using late 1979 as a structural break point in interest rate data is standard (see, for example, Hardouvelis, 1988; Engle and Ng, 1993a; Klemkosky and Pilotte, 1992). Second, we find no evidence that PPI and employment releases had an effect on bond market volatility prior to 1979 (see below). This relative unimportance of the employment and PPI data prior to 1979 may be due to changes in Fed practices, changes in data quality, learning by financial markets, or other changes in macroeconomic structure, information, and the financial system. Krueger (1996) provides evidence on the changes in the quality of the employment data and the amount of news coverage it received, beginning in 1979.

Table 1 gives summary statistics for daily excess returns, and compares nonannouncement and announcement days. Tables 2 and 3 examine these same data and test whether announcement days are significantly different from nonannouncement days. As can be seen in the first row of Table 1, excess returns are 0.01–0.02% per trading day in this period, or about 2.7–4.6% per year. The magnitude of daily returns is sometimes quite large, with returns for the ten-year bond as high as 4.69% (on October 20, 1987, the day after the stock market crash) and as low as − 3.67% (on February 19, 1980). Neither of these two dates is an announcement date. There is also evidence of first-order autocorrelation in excess returns. In addition, excess returns are significantly positively skewed and significantly fat-tailed.

The upper half of Table 1 motivates our use of the ARCH class of models. The first-order autocorrelation coefficient (ρ) for the absolute value of excess returns ranges from 0.07 to 0.18, and ranges from 0.04 to 0.16 for the squared value. Like stock returns and foreign exchange returns, bond returns exhibit autocorrelated volatility.

The bottom half of Table 1 focuses on the announcement dates. Since the results for PPI and employment are so similar, hereafter we pool announcement dates. For both the employment release and the PPI release, bond market volatility is far higher on release dates than on nonrelease dates. For example, absolute returns for the ten-year bond average 0.558% on announcement days and 0.388% on other days. In contrast, the first-order autocorrelation coefficients of volatility following announcement days are not very different from average, rising from 0.12 to 0.19 for the ten-year bond squared excess returns but falling slightly for the analogous absolute value of excess returns.

Table 1 also shows that excess returns are much higher on announcement dates, averaging 0.104% for the ten-year bond. In fact, during this period, most of the excess returns earned by bonds are accounted for by the 9% of trading days with an employment or PPI announcement. On nonannouncement days, the other 91% of the sample, bonds earned much smaller excess returns (0.005% for the ten-year bond). It is worth noting that the ex ante excess return on government bonds is not necessarily positive in theory or in fact. For example, Campbell (1995) finds using monthly data that ten-year Treasury bonds earned negative excess returns over the period 1952–1991.

To put the numbers in perspective, the mean excess returns on announcement days imply annualized excess returns of more than 22% per year for all three maturities. Thus it seems that exposure to macroeconomic risk earns a high risk premium.

Another way to characterize the risk premium is in terms of a Sharpe or reward-to-risk ratio. On nonannouncement days, bondholders bear substantial risk in return for a very low excess return, for a Sharpe ratio of 0.01. On employment and PPI release days, the Sharpe ratio for ten-year bonds over Treasury bills is $0.104/0.718 = 0.145$. Over a similar period, we find using data from the Center for Research in Securities Prices that the value-weighted index of all NYSE, AMEX, and Nasdaq stocks has a daily Sharpe ratio of about 0.03. This ratio is substantially below the reward-to-risk ratio for bonds on announcement days. Perhaps the reward-to-risk ratio on release days is so high because the type of risk on important macroeconomic announcement days is fundamentally different from the risk on other days. Financial market volatility associated with fundamental news–news that pertains to future consumption, rates of return, and business conditions – earns a high return.

We turn next to simple OLS regressions to explore the relationship of the announcement dates to both risk and return. Table 2 documents the volatility of daily excess returns using day-of-the-week and announcement indicator variables. Again, we measure volatility in two ways, using both squared and absolute values of excess returns. We discuss results for absolute excess returns; the table shows that results using squared returns are similar.

The results indicate that there are day-of-the-week effects for return volatility. We see that bond market volatility is generally highest on Mondays and Fridays and lowest on Wednesdays, exhibiting a U-shaped curve over the week. This pattern contrasts somewhat with the stock market, in which return variances decline over the course of the week (e.g., French, 1980). Note that, because we also include announcement-day dummy variables, we are able to separate the Friday and announcement effects (83% of the announcements were on Fridays, and 37% of the Fridays had announcements).

Table 2 also confirms that, controlling for day-of-the-week effects, announcement days have significantly higher volatility than average. Across the three bonds studied, volatility increases by more than one-third on event days, as measured by the absolute value of excess returns. These differences in volatility are highly statistically significant.

Table 1
Summary statistics: Treasury bond daily excess returns

XR_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from October 9, 1979 to December 31, 1995.

	5-yr			10-yr			30-yr		
	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$
Full Sample (N = 4051)									
Mean	0.011[a]	0.153[c]	0.265[c]	0.014	0.329[c]	0.404[c]	0.018	0.613[c]	0.564[c]
Std Dev	0.391	0.440	0.288	0.573	0.848	0.407	0.783	1.505	0.544
Min	− 2.487			− 3.669			− 3.941		
Max	3.014			4.692			7.252		
ρ (autocorr.)	0.105	0.157	0.183	0.075	0.124	0.159	0.048	0.039	0.068
Kurtosis	6.257			4.644			4.012		
Skewness	0.259			0.230			0.207		
Employment report announcement dates (N = 189)									
Mean	0.091[bx]	0.310[cz]	0.425[cz]	0.112[ax]	0.621[cz]	0.613[cz]	0.120	1.085[c]	0.831[c]
Std Dev	0.551	0.534	0.361	0.782	0.967	0.496	1.037	1.475	0.629
ρ (t to $t+1$)	0.081	0.164	0.125	0.074	0.151	0.151	0.065	0.085	0.054
PPI announcement dates (N = 189)									
Mean	0.086[cy]	0.211[cz]	0.341[cz]	0.119[by]	0.495[cz]	0.533[cz]	0.160[by]	0.884[cz]	0.741[cz]
Std Dev	0.453	0.382	0.309	0.696	0.846	0.461	0.929	1.321	0.580
ρ (t to $t+1$)	0.121	0.234	0.132	0.070	0.166	0.165	0.101	0.101	0.120

Announcement dates ($N = 364$)									
Mean	0.082[cz]	0.249[cz]	0.375[cz]	0.104[cy]	0.525[cz]	0.558[cz]	0.133[cy]	0.956[cz]	0.776[cz]
Std Dev	0.493	0.443	0.329	0.718	0.842	0.463	0.970	1.356	0.595
ρ (t to $t+1$)	0.092	0.222	0.144	0.064	0.187	0.157	0.063	0.106	0.083
Nonannouncement dates ($N = 3687$)									
Mean	0.004	0.143[c]	0.254[c]	0.005	0.309[c]	0.388[c]	0.006	0.580[c]	0.543[c]
Std Dev	0.379	0.438	0.281	0.556	0.846	0.398	0.761	1.515	0.534
ρ (t to $t+1$)	0.109	0.151	0.192	0.078	0.119	0.162	0.046	0.033	0.068

[a]Different from zero at the 10% significance level (two-tailed).
[b]Different from zero at the 5% significance level (two-tailed).
[c]Different from zero at the 1% significance level (two-tailed).
[x]Different from nonannouncement mean value at the 10% level (one-tailed).
[y]Different from nonannouncement mean value at the 5% level (one-tailed).
[z]Different from nonannouncement mean value at the 1% level (one-tailed).

Table 2
Treasury bond return volatility by day of week and event day

Mean values of the volatility of the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill, estimated using an OLS regression with dummy variables for weekdays and announcement days. Returns are expressed in percent, i.e., multiplied by 100. Announcement is a dummy variable which equals one on the PPI and employment report announcement dates. The sample extends from October 9, 1979 to December 31, 1995. Heteroskedasticity-consistent standard errors are given in parentheses (White, 1980).

	5-yr	10-yr	30-yr
Panel A: Absolute value of excess returns			
Monday	0.293	0.432	0.602
	(0.015)	(0.021)	(0.028)
Tuesday	0.252	0.385	0.545
	(0.010)	(0.014)	(0.019)
Wednesday	0.221	0.345	0.486
	(0.008)	(0.012)	(0.016)
Thursday	0.261	0.412	0.576
	(0.011)	(0.016)	(0.022)
Friday	0.285	0.429	0.572
	(0.012)	(0.016)	(0.020)
Announcement ($t+1$)	−0.043	−0.071	−0.089
	(0.015)	(0.023)	(0.030)
Announcement	0.097	0.136	0.208
	(0.019)	(0.027)	(0.034)
Announcement ($t-1$)	−0.022	−0.022	−0.021
	(0.020)	(0.028)	(0.037)
Panel B: Squared excess returns			
Monday	0.211	0.407	0.773
	(0.026)	(0.045)	(0.079)
Tuesday	0.150	0.321	0.615
	(0.017)	(0.036)	(0.075)
Wednesday	0.098	0.231	0.456
	(0.009)	(0.023)	(0.041)
Thursday	0.140	0.319	0.603
	(0.015)	(0.028)	(0.049)
Friday	0.166	0.343	0.575
	(0.018)	(0.029)	(0.039)
Announcement ($t+1$)	−0.038	−0.075	−0.157
	(0.023)	(0.044)	(0.063)
Announcement	0.088	0.189	0.381
	(0.026)	(0.048)	(0.075)
Announcement ($t-1$)	−0.038	−0.045	−0.084
	(0.033)	(0.067)	(0.100)

Table 3
Treasury bond daily excess returns by day of week and event day

Mean values of the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill, estimated using an OLS regression with dummy variables for weekdays and announcement days. Returns are expressed in percent, i.e., multiplied by 100. Announcement is a dummy variable which equals one on the PPI and employment report announcement dates. The sample extends from October 9, 1979 to December 31, 1995. Heteroskedasticity-consistent standard errors are given in parentheses (White, 1980).

	5-yr	10-yr	30-yr
Monday	− 0.054	− 0.075	− 0.087
	(0.020)	(0.028)	(0.039)
Tuesday	0.030	0.049	0.063
	(0.013)	(0.020)	(0.027)
Wednesday	0.004	− 0.004	− 0.007
	(0.011)	(0.017)	(0.023)
Thursday	0.010	0.009	0.032
	(0.015)	(0.023)	(0.032)
Friday	0.020	0.029	0.024
	(0.016)	(0.024)	(0.031)
Announcement ($t + 1$)	0.001	0.007	− 0.023
	(0.021)	(0.032)	(0.043)
Announcement	0.063	0.077	0.108
	(0.028)	(0.041)	(0.055)
Announcement ($t - 1$)	0.022	0.036	0.036
	(0.026)	(0.039)	(0.053)

There is no evidence that this effect is present prior to 1979. When we estimate the model in Table 2 using 2659 daily observations from February 1, 1969 to October 5, 1979, the effect of announcement days on volatility is minute and we are unable to reject the hypothesis that news dates have no effect on volatility. The estimated announcement-day effect (with standard error) for the three bonds, respectively, is 0.007 (0.012), 0.013 (0.015), and 0.049 (0.033) for absolute returns and 0.003 (0.010), 0.005 (0.015), and 0.020 (0.034) for squared returns.

The financial press often claims that financial markets are particularly quiet on the days prior to these announcements. For example, a typical headline in the *Wall Street Journal* reads 'Treasurys Decline in Light Trading as Market Awaits Today's Report on Employment in July' (8/5/94). This finding might be called the 'calm before the storm' effect. We investigate this claim by including a lead of the announcement dummy in the regression. For all three bonds, absolute returns are lower than average on days preceding macroeconomic announcements. For example, ten-year bond absolute returns on the day before an announcement are 0.071 percentage points below the average volatility. This estimate represents an 18% fall from the full-sample average volatility of 0.404%. Like the announcement-day effect, the magnitude of the 'calm before the storm' effect is monotonically increasing with the maturity of the bond. The effect is statistically significant for absolute returns at all maturities. We conclude, therefore, that there is some evidence for the 'calm before the storm' effect. Volume data (as in French et al., 1989) might cast further light on this question.

Last, Table 2 includes a one-day lag of the announcement-day dummy. If the shocks to volatility on announcement days generate persistent volatility, we would expect that the day after an announcement day would have higher than average volatility. Table 2 shows that for all three bonds, volatility on the day after an announcement is *lower* than average. Table 2 thus shows no evidence that the trading process itself generates autocorrelated volatility in response to a one-time piece of news. Of course, Table 2 does not attempt to model conditional heteroskedasticity formally; we proceed to this estimation in Section 3.

Table 3 further examines the risk premiums earned by bonds on announcement days. The first five rows show day-of-the-week effects. Most strikingly, excess returns are consistently negative on Mondays. Since Table 2 shows that Mondays are high-volatility days, we find no obvious relation between volatility and mean excess returns across days of the week.

The seventh row of Table 3 shows the effect of including the announcement-day dummy. Holding constant day-of-the-week effects, excess returns are 0.077 percentage points higher on announcement days for the ten-year bond. This higher return is significant with a t-statistic of between 1.9 and 2.3 across the three bonds we study. Of course, the estimation in Table 3 has limited power, since from an econometric standpoint, failure to account for conditional heteroskedasticity is inefficient. Thus the more complex modeling in Section 3 can provide more precise estimates of the risk premiums earned on announcement days. Note that these excess returns are smaller than those associated with news releases in Table 1 (e.g., 0.104% for the ten-year bond) because excess returns appear to be higher on Fridays (0.029% for the ten-year bond Friday dummy).

In contrast, the sixth and eighth rows show that the days before and after announcement dates (which we know from Table 2 do not have high volatility) do not have high returns.

The pattern of the announcement-day effect across different maturities is consistent with the idea that higher exposure to macroeconomic risk results in higher expected returns. Tables 1 and 2 show that longer-maturity bonds have both higher average volatility and higher announcement-day volatility. Table 3 shows that excess returns on announcement days are also monotonically increasing in the maturity of the bond, as one would expect if news about future inflation and interest rates affects long bonds more than short bonds.

The results from Table 3 are not caused by a small number of large outliers. We drop the bottom and top 1% of the observations on excess returns and reestimate Table 3 for each of the three bonds (we drop the extreme 82 out of 4050 observations for each of the three bonds, which results in dropping between 11 and 13 of the announcement dates). The results are similar, although the announcement-day coefficient drops somewhat: for the three bonds, the coefficient on the announcement-day dummy (with standard error) is 0.063 (0.028), 0.052 (0.025), and 0.044 (0.037).

The fact that returns are significantly higher on announcement dates is similar to results that might be obtained from an ARCH-M specification. We consider the results in Tables 2 and 3 (and the more powerful risk premium tests in the next section) important because they supply evidence of a time-series risk/return tradeoff without reliance on an ARCH-M specification.

We also investigate the sensitivity of the results to changes in measuring returns. We replicate all our empirical tests using raw returns instead of returns in excess of the Treasury bill rate. We find nearly identical results for the size and statistical significance of announcement-day effects on average returns and on volatility. In addition, excluding the weekday dummies has no effect on the results.

In summary, the simple OLS regressions presented here provide straightforward answers to our two questions. First, days following announcements do not exhibit high unconditional volatility, which suggests that the trading process itself does not generate autocorrelated volatility in response to a one-time piece of macroeconomic news. Second, bonds earn much higher returns when they are more exposed to macroeconomic risk. We next turn to more complex ways of modeling volatility, which allow us to estimate jointly both the pattern of conditional heteroskedasticity and the size of the risk premium.

3. Modeling changing variance

Perhaps the most commonly used model of financial asset return volatility is the GARCH(1,1) model proposed by Bollerslev (1986). Though it is not necessarily the correct specification of the return-generating process, it is an important benchmark, because the same model has been estimated across a number of asset classes and sampling frequencies. In addition, theoretical results are available for quasi-maximum likelihood estimators of this model (e.g., Lumsdaine, 1996). Finally, GARCH models provide approximate descriptions of conditional volatility for a wide variety of volatility processes (see Nelson, 1990b; Nelson and Foster, 1994).

For these reasons, we begin by estimating a univariate GARCH(1,1) model of daily bond returns adjusted for announcement-day effects. We initially accommodate these volatility effects using the procedure outlined in Andersen and

Bollerslev (1997). Dummy variables allow us to measure the (contemporaneous) impact of an announcement day on both conditional mean returns and conditional volatility. Specifically, we assume returns are of the form

$$R_t = \mu + \theta I_t^{\mathrm{A}} + \phi_1 R_{t-1} + s_t^{1/2}\varepsilon_t, \tag{1}$$

$$s_t = 1 + \delta_0 I_t^{\mathrm{A}}, \tag{2}$$

$$h_t = \omega + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1}, \tag{3}$$

where I_t^{A} is the announcement indicator dummy variable, s_t is the volatility seasonal for time t, δ_0 measures the volatility effect of an announcement on day t, and ε_t is a random variable with conditional mean zero and conditional variance h_t, independent of s_t. That is, we estimate a (multiplicative) seasonal dummy for volatility on the day of the announcement. The specification (1)–(3) allows the conditional variance to differ on announcement days; that is, excess return innovations have conditional variance h_t on nonannouncement days and $h_t(1 + \delta_0)$ on announcement days.

This specification can also easily accommodate day-of-the-week effects via the volatility seasonal. However, since the day-of-the-week effects contribute little and do not affect the results, to conserve on space we report regressions with no day-of-the-week effects. The coefficient θ measures changes in mean returns on announcement dates. We estimate an autoregressive model for the first moment because we find small but highly significant positive autocorrelation in Treasury bond returns. This autocorrelation could be due to microstructure effects in measuring prices, or it could be due to equilibrium partial adjustment. We also estimate a model with higher-order autoregressive terms, but the higher-order terms are insignificant and thus are excluded.

Reported parameter estimates for Eqs. (1)–(3) are obtained by quasi-maximum likelihood estimation using a normal likelihood function; starting values for the conditional variance process h_t are set equal to the unconditional variance, as suggested in Engle and Bollerslev (1986). The results are reported in Table 4 for the five-, ten-, and 30-year bond returns. Robust standard errors, as in Bollerslev and Wooldridge (1992), are also provided.

We begin with results for the first moment of returns. The autoregressive coefficient for excess returns is below 0.1. Most importantly, we confirm the earlier evidence of a risk premium on announcement days; returns on such days average 0.07, 0.09, and 0.12 percentage points higher for the three bonds. All three are significant.

Estimation of (3) reveals a statistically significant impact of the announcement; volatility increases by 115% on announcement days for the ten-year bond. However, we are more interested in the persistence of shocks to volatility on release and nonrelease days. While there is no unique characterization of volatility persistence (see, e.g., Nelson, 1990a), one standard measure of

Table 4
Benchmark AR(1)-GARCH(1, 1) model of Treasury bond daily returns

Quasi-maximum likelihood estimates of the model

(1) $R_t = \mu + \theta I_t^A + \phi R_{t-1} + s^{1/2}\varepsilon_t$,

(2) $s = (1 + \delta_0 I_t^A)$,

(3) $h_t = \omega + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1}$,

where R_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill rate, ε_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from October 9, 1979 to December 31, 1995. Robust standard errors are given in parentheses (Bollerslev and Wooldridge, 1992).

	First moment parameters				Second moment parameters		
	5-yr	10-yr	30-yr		5-yr	10-yr	30-yr
μ	0.004	0.008	0.015	ω	0.0008	0.0021	0.0031
	(0.005)	(0.008)	(0.012)		(0.0002)	(0.0004)	(0.0007)
ϕ	0.090	0.083	0.047	δ_0	1.388	1.150	1.015
	(0.016)	(0.015)	(0.015)		(0.206)	(0.178)	(0.166)
θ	0.066	0.088	0.122	α	0.050	0.051	0.037
	(0.024)	(0.035)	(0.052)		(0.011)	(0.009)	(0.007)
				β	0.938	0.937	0.955
					(0.014)	(0.010)	(0.008)
log L	2464.61	756.29	− 686.31				

persistence comes from the j-step-ahead forecast of conditional variance (see, e.g., Andersen and Bollerslev, 1997):

$$ {}_t h_{t+j} - \sigma^2 = (\alpha + \beta)^j (h_t - \sigma^2), \tag{4}$$

where ${}_t h_{t+j}$ is the expectation at time t of conditional volatility at time $t+j$ and σ^2 is the unconditional variance. If $\alpha + \beta < 1$, this can be used to compute a half-life, that is, the time it takes on average for conditional variance h_t to revert halfway to its unconditional value. This is computed by setting $(\alpha + \beta)^j = \frac{1}{2}$; the half-life j is thus defined as $-[\ln(2)]/\ln(\alpha + \beta)$. For each of the bonds, $\alpha + \beta$ is about 0.99 but statistically different from unity; the three bonds have corresponding half-lives of between 57 and 86 trading days (three to four months). Thus, by using this metric we would conclude that in all three assets, there is a high degree of volatility persistence.

We are principally interested in the persistence properties of the estimated conditional volatility following announcements. On one hand, if markets immediately incorporate information into prices, announcement news would exhibit little persistence. On the other hand, if volatility is caused by some feature of the trading or information-gathering process itself, persistence following announcement news may be higher than on nonannouncement days. Hence, we supplement the benchmark GARCH(1,1) model with an alternative specification that allows the degree of volatility persistence to vary.

Are announcement-day volatility shocks more or less persistent than shocks on other days? To answer this question, we estimate a model in which the conditional variance is a regime-switching GARCH process, similar to Gray (1996). Unlike standard regime-switching processes such as Hamilton (1989) or Gray (1996), in which the timing of the regime shifts must be estimated from the data, we treat the regime shift as occurring deterministically at announcement dates.[2] Specifically, the model consists of (1) and the following specification of the announcement seasonal and the conditional variance:

$$s_t = (1 + \delta_0 I_t^{\mathrm{A}})(1 + \delta_1 I_{t-1}^{\mathrm{A}}), \tag{2'}$$

$$h_t = \omega + [\alpha_0 + \alpha_{\mathrm{A}} I_{t-1}^{\mathrm{A}}]\,\varepsilon_{t-1}^2 + [\beta_0 + \beta_{\mathrm{A}} I_{t-1}^{\mathrm{A}}]\,h_{t-1}. \tag{3'}$$

On the day after an announcement, conditional volatility changes regime. First, we allow a deterministic change in volatility, δ_1, on the day immediately following an announcement day. More importantly, announcement-day shocks are allowed to feed differently into conditional volatility via α_{A}. Last, the importance of lagged conditional volatility on these days can also vary via β_{A}. This very general specification allows announcement shocks and nonannouncement shocks to have different effects.

Note that if $\alpha_{\mathrm{A}} = \beta_{\mathrm{A}} = \delta_1 = 0$, (2′)–(3′) reduces to model (2)–(3). Thus, the regime-switching model nests the benchmark model as a special case. This model also nests the special case in which announcement shocks do not persist at all. In particular, if $\alpha_{\mathrm{A}} = -\alpha_0$ and $\beta_{\mathrm{A}} = \delta_1 = 0$, then announcement-day shocks to volatility do not have any effect on future volatility; h evolves as though those shocks were equal to zero.

Results for this model are given in Table 5. Using robust standard errors, we compute Wald statistics as in Lumsdaine (1995) for the joint null hypothesis that $\alpha_{\mathrm{A}} = 0$, $\beta_{\mathrm{A}} = 0$, and $\delta_1 = 0$ and strongly reject the benchmark model in favor of the regime-switching alternative. The value of the statistic is 17.64, 17.22, and 15.64 for the five-, ten-, and 30-year bonds, respectively; the $\chi^2(3)$ 5% critical value is 7.81.

The regime-switching model of (1), (2′), and (3′) leads to quite different conclusions about announcement persistence than simpler models. Using the

[2] We thank Ludger Hentschel, the referee, for suggesting this specification. In a previous version of this paper, we estimated several alternative volatility specifications. All models give similar results over the 1979–1995 period, namely that announcement-day volatility does not appear to persist at all.

Table 5
GARCH(1,1) with differing parameters following announcements

Quasi-maximum likelihood estimates of the model

(1) $R_t = \mu + \theta I_t^A + \phi R_{t-1} + s^{1/2}\varepsilon_t$,

(2) $s = (1 + \delta_1 I_{t-1}^A)(1 + \delta_0 I_t^A)$,

(3) $h_t = \omega_0 + [\alpha_0 + \alpha_A I_{t-1}^A]\,\varepsilon_{t-1}^2 + [\beta_0 + \beta_A I_{t-1}^A]\,h_{t-1}$,

where R_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill rate, ε_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from October 9, 1979 to December 31, 1995. Robust standard errors are given in parentheses (Bollerslev and Wooldridge, 1992).

	First moment parameters				Second moment parameters		
	5-yr	10-yr	30-yr		5-yr	10-yr	30-yr
μ	0.004	0.008	0.015	ω_0	0.0011	0.0028	0.0051
	(0.005)	(0.007)	(0.011)		(0.0002)	(0.0004)	(0.0008)
ϕ	0.085	0.081	0.046	δ_0	1.379	1.163	1.050
	(0.015)	(0.015)	(0.015)		(0.224)	(0.193)	(0.182)
θ	0.056	0.080	0.114	δ_1	0.173	0.178	0.139
	(0.026)	(0.037)	(0.053)		(0.152)	(0.156)	(0.155)
				α_0	0.061	0.060	0.045
					(0.015)	(0.011)	(0.010)
				α_A	− 0.055	− 0.053	− 0.047
					(0.020)	(0.018)	(0.018)
log L	2486.91	772.62	− 670.93	β_0	0.937	0.935	0.947
					(0.021)	(0.016)	(0.017)
				β_A	− 0.035	− 0.015	0.015
					(0.104)	(0.103)	(0.117)

regime-switching model, we find that announcement-day shocks are strikingly different in terms of their persistence. For announcement shocks, the estimates of α_A are all negative and very close to $-\alpha_0$. In fact, we cannot reject the hypothesis that announcement shocks do not persist at all (i.e., $\alpha_A = -\alpha_0$). Wald statistics for this hypothesis are 0.37, 0.37, and 0.03 for the three bonds, all well below the 5% critical value of 3.84. This is not due to large standard errors, as the coefficient estimates are quite precise. Further, the estimates of β_A are all small and statistically indistinguishable from zero, which means that on the day after announcement days, conditional volatility continues to decay at the regular rate β_0. Last, we cannot reject the hypothesis that the days after announcement days do not have higher than average unconditional volatility, since δ_1 is indistinguishable from zero.

Table 6
Restricted model: announcement-day shocks have no further effect

Quasi-maximum likelihood estimates of the model

(1) $R_t = \mu + \theta I_t^A + \phi R_{t-1} + s^{1/2}\varepsilon_t$,

(2) $s = (1 + \delta_0 I_t^A)$,

(3) $h_t = \omega_0 + \alpha[1 - I_{t-1}^A]\,\varepsilon_{t-1}^2 + \beta h_{t-1}$,

where R_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill, ε_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from October 9, 1979 to December 31, 1995. Robust standard errors are given in parentheses (Bollerslev and Wooldridge, 1992).

	First moment parameters				Second moment parameters		
	5-yr	10-yr	30-yr		5-yr	10-yr	30-yr
μ	0.003	0.007	0.014	ω_0	0.0011	0.0030	0.005
	(0.005)	(0.007)	(0.012)		(0.0002)	(0.0004)	(0.001)
ϕ	0.084	0.080	0.047	δ_0	1.373	1.134	1.010
	(0.015)	(0.015)	(0.015)		(0.206)	(0.178)	(0.166)
θ	0.054	0.078	0.115	α	0.063	0.061	0.045
	(0.025)	(0.037)	(0.052)		(0.013)	(0.010)	(0.008)
log L	2484.63	769.98	−672.58	β	0.934	0.934	0.949
					(0.013)	(0.009)	(0.008)

The results from Table 5 suggest the further restriction that $\alpha_A = -\alpha_0$ and $\beta_A = \delta_1 = 0$. That is, we constrain the model such that announcement-day shocks have no effect on future volatility, that deseasonalized conditional volatility continues to decay at the usual rate, and that there is no deterministic change in conditional volatility on the day after an announcement. Using robust Wald statistics, we do not reject this restriction. Wald statistics are 1.75 for the five-year, 2.12 for the ten-year, and 1.49 for the 30-year bond, all well below the 7.81 critical value. For clarity, we also estimate this restricted model explicitly in Table 6.

In summary, the restricted model in Table 6 has the following implications. Announcement-day shocks to volatility do not persist at all. Volatility is higher on announcement days, but this increase in volatility is purely transitory. Equivalently, there is no information in announcement-day returns that is helpful in forecasting return variance on the day after an announcement.

Our results concur with those of Ederington and Lee (1993, 1995) that market prices quickly incorporate the information in these macroeconomic announcements, and that volatility quickly returns to preannouncement levels. At least for this subset of public information releases, near-zero estimates of $\alpha_0 + \alpha_A$ are consistent with the hypothesis that the trading or information-gathering process does not generate volatility on succeeding days. Thus, we provide a counterexample to the hypothesis that something in the trading or information-gathering process turns all information arrivals into positively autocorrelated volatility. Our results suggest that, in general, researchers should focus on autocorrelated public or private information release as the source of persistent volatility.

In summary, the autoregressive conditional heteroskedasticity models presented in this section give answers that are quite similar to the simple OLS regressions. First, announcement shocks have no effect on volatility in the days following announcements, which suggests that the trading process itself does not generate autocorrelated volatility. Second, across all the different models estimated, bonds earn higher returns when they are exposed to higher risk. This risk premium is always monotonically increasing in the maturity of the bond, and the estimate is always statistically significant.

4. Conclusions

Understanding the determinants of asset prices and the way that markets process new information about these determinants is a central question of finance. To study this question, economists have examined both expected returns and unexpected (realized) returns. One approach to this question is simply to assume that any unexpected movement in returns reflects new information about fundamentals, and any expected movement in returns reflects predictable fundamental risk. Thus, for example, Engle and Ng (1993b) define 'news' as the unpredictable component of realized returns, and go on to examine its time series properties. Fama and French (1992) find predictable differences in expected returns across different types of stocks, and interpret the results as suggesting different exposures to economic risk factors. A second approach has been to attempt to identify ex post the news that moves asset prices, and to connect ex ante movements in expected returns with movements in fundamental risk. Roll (1988) and Cutler et al. (1989) try to explain realized returns with measures of news, with very limited success. The consumption-CAPM literature (see Cochrane and Hansen, 1992 for references) attempts to explain expected returns with aggregate consumption risk, again with little success.

In this paper, we take a different route. Instead of trying to find ex post explanations for all asset price movements, we focus on a subset of asset price movements that are associated with observable news about macroeconomic conditions. Standard asset pricing theory predicts that a one-time shock of public news about economic fundamentals should cause a one-time movement in asset prices. When the volatility of news is serially uncorrelated, the volatility of returns should also be serially uncorrelated. Standard theory also predicts that if this increase in macroeconomic risk occurs at a predictable time, assets exposed to the risk should earn higher expected returns at the time of higher expected risk.

The data confirm these predictions. Unlike nonannouncement shocks, shocks to volatility that occur on announcement days have no subsequent impact on daily volatility. Further, the predictable risk that bonds bear on announcement days is compensated with higher expected excess returns. We conclude that markets set expected returns to reflect risk, that the release of public information does not generate autocorrelated volatility, and that markets quickly incorporate public information into prices. The result that volatility on subsequent days is not raised by announcement-day shocks is consistent with the intraday evidence presented in Ederington and Lee (1993).

Why is return volatility autocorrelated in general? Our results suggest that the documented volatility autocorrelation in asset returns is partially caused by volatility autocorrelation of the news-generating process. Most news is clustered over time (though not the announcements we study), so volatility is clustered over time. Why are asset returns predictable? Our results are consistent with the idea that expected returns are predictable because fundamental risk is predictable. What determines asset prices? Our results are that expected returns vary with macroeconomic risk.

Of course, we cannot conclude that prices only reflect fundamental risk, or that only genuine news moves prices, since we are focusing on a subset of news events. This subset of events explains only a small fraction of the variance of returns. For example, McQueen and Roley (1993) find from changes in ten-year Treasury bond yields and a larger set of news events 1977–1988 that macroeconomic surprises explain about 9% of the variation in yields. We also have nothing to say about how markets incorporate private information, since we have only examined public information. Finally, our inferences depend critically on the assumption that announcement days actually do contain a one-time piece of fundamental news.

Nevertheless, it is comforting that when we can find clean, observable macroeconomic shocks, asset prices respond the way theory says they should. Our findings also shed economic light on existing empirical research on financial volatility. First, the results suggest that conditional heteroskedasticity is in part a description of the time series properties of the news-generation process. Second, the results corroborate the basic ARCH-in-mean hypothesis: higher conditional volatility is accompanied by higher expected returns.

On a more technical level, our results have direct implications for modeling conditional heteroskedasticity in asset returns. In particular, we find strong evidence that return shocks vary in their persistence depending on their sources. Ignoring this variation results in model misspecification and is likely to lead to inferior estimates of conditional volatility. This pattern has practical considerations as well, most obviously in the area of option pricing.

Much remains for future research. We find evidence of a 'calm before the storm' effect. When markets know that a large shock is forthcoming, return volatility decreases. This effect might suggest interesting interactions of volume, private information, preferences towards risk, or portfolio rebalancing. This question deserves further study. Also, our results are limited in that they only examine the univariate response of returns to risk. The news announcements used here might yield insights about the shape of the term structure and about the covariance of bond returns with other assets.

Specifically, it remains puzzling that, in general, stock prices seem less affected by macroeconomic news than bond prices (see McQueen and Roley, 1993 for references and some evidence on this issue). If asset returns are driven by fundamental risk, and if (as we assume) announcement days are days of high risk, how can it be that bonds are affected more than stocks? Since bonds but not stocks have high return variance on announcement days, we might expect that the conditional covariance of stock and bond returns falls on announcement days. Future work, perhaps along the lines of Bollerslev et al. (1988), could address this issue.

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