

On the Volatility and Comovement of U.S. Financial Markets around Macroeconomic News Announcements

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Abstract

The objective of this paper is to provide a deeper insight into the links between financial markets and the real economy. To that end, we study the short-term anticipation and response of U.S. stock, Treasury, and corporate bond markets to the first release of surprise U.S. macroeconomic information. Specifically, we focus on the impact of these announcements not only on the level, but also on the volatility and comovement of those assets' returns. We do so by estimating several extensions of the parsimonious multivariate GARCH-DCC model of Engle (2002) for the excess holding-period returns on seven portfolios of these asset classes. We find that both the process of price formation in each of those financial markets and their interaction appear to be driven by fundamentals. Yet our analysis reveals a statistically and economically significant dichotomy between the reaction of the stock and bond markets to the arrival of unexpected fundamental information. We also show that the conditional mean, volatility, and comovement among stock, Treasury, and corporate bond returns react asymmetrically to the information content of these surprise announcements. Overall, the above results shed new light on the mechanisms by which new information is incorporated into prices within and across U.S. financial markets.

I. Introduction

There seems to be little doubt among academics and practitioners that the release of macroeconomic news has a significant impact on the prices of securities

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within as diverse asset classes as stocks, Treasury bonds, and corporate bonds. For instance, most asset pricing models provide a snapshot of the cross-sectional relationship between asset returns (or prices) and risk factors at a given point in time. A change in one or more of these factors should therefore affect asset returns, with the dynamic nature of these changes dictated by the dynamics of new information arriving in the market.

The dominant paradigm regarding the response of asset prices to new information (first articulated by Fama (1971)) is that, since markets are efficient, asset prices should react immediately and in an unbiased manner to *new* information. Accordingly, in this paper we explore the functioning of the process of price formation in all three of the main U.S. financial markets—stocks, government bonds, and corporate bonds—around the release dates of the *unexpected* portion of four important macroeconomic news items: the consumer price index (CPI), the unemployment rate, nonfarm payroll employment, and the target federal funds rate. Specifically, we intend to address four basic sets of questions.

First, what is the impact of these macroeconomic announcements on asset returns and asset return volatility in the proximity of their first release? Are the markets where these assets are traded more volatile before the news event and less volatile afterward? This would be the case, for example, if the arrival of information leads to the resolution of uncertainty and/or disagreement among market participants (e.g., Pasquariello (2007)). However, the analysis of Ross (1989) may lead to the opposite inference. Ross argues that, in an arbitrage-free economy, the volatility of prices should be related to the arrival of information in an efficient market. Accordingly, Foster and Viswanathan (1993) and Pasquariello and Vega (2007) show that, *ceteris paribus*, both the availability and the realizations of a public signal increase price volatility.

Second, does the release of macroeconomic news surprises affect the markets for different asset classes in different ways? The rationale for heterogeneity in the fluctuations of stock and (government and corporate) bond returns in response to macroeconomic news is intuitive. For instance, higher inflation may increase only the volatility of the bond market, since stocks (and, to a lesser extent, corporate bonds) offer a natural hedge against inflation.

Third, do macroeconomic announcements affect the existing degree of covariance between different asset classes? Covariance shifts may stem from information spillovers across markets (e.g., Shiller (1989), King and Wadhvani (1990), Pindyck and Rotemberg (1990), (1993), Karolyi and Stulz (1996), and Connolly and Wang (2003)), from more intense portfolio rebalancing activity across stocks and bonds (e.g., Fleming, Kirby, and Ostdiek (1998), Kodres and Pritsker (2002)), from financial constraints (e.g., Kyle and Xiong (2001)), or from greater dispersion of beliefs among speculators (e.g., Pasquariello (2007), Kallberg and Pasquariello (2008)) before and after their arrivals.

Fourth, is the impact of unexpected macroeconomic news releases instantaneous or protracted over time? Are the observed shocks to volatility and comovement of a transitory or persistent nature? For instance, Jones, Lamont, and Lumsdaine (1998) observe that clustering of information arrivals, market sentiment, or gradual learning may explain why an increase in return volatility would persist over time. Therefore, lack of persistence in announcement shocks would

suggest that additional information gathering or the trading process do not intrinsically increase return volatility.¹

To tackle these questions, we combine a database of daily excess holding-period returns on seven asset portfolios—stocks traded on the NYSE-AMEX and NASDAQ; 5-, 10-, and 30-year Treasury bonds; and Aaa- and Baa-rated long-term corporate bonds—with data on those four macroeconomic announcements and their consensus expectations between 1986 and 2002. The choice of daily spot data is not casual. First, higher-frequency data are unavailable for two of the asset classes under consideration (i.e., corporate and government bonds) over our entire sample period. Since one important aspect of our study is to compare the reaction of all U.S. financial markets to the release of U.S. macroeconomic news, full sample coverage for all asset classes represents a necessary prerequisite. Second, using higher-frequency data over shorter sample periods, recent studies find a significant intraday reaction by some of these markets to macroeconomic news, often within minutes of their release (e.g., Balduzzi, Elton, and Green (2001), Andersen, Bollerslev, Diebold, and Vega (2003), (2007), Green (2004), and Fleming and Piazzesi (2005)). Thus, the daily frequency of our data is most likely to bias our estimates of that reaction downward. Yet, as we soon discuss, our evidence is strong. Lastly, higher-frequency data are afflicted by microstructure frictions (e.g., bid-ask bounce, quote clustering, price staleness, inventory effects) that may bias inferences drawn upon them (e.g., Bai, Russell, and Tiao (2004), Andersen, Bollerslev, and Meddahi (2005), and references therein). These frictions generally become immaterial over longer horizons.²

The expectation data, from the International Money Market Services (MMS) survey (for macroeconomic variables) and from futures prices (for federal funds rate decisions), are customarily assumed to represent unbiased estimates of the anticipated portion of these announcements.³ Hence, they allow us to identify the unexpected component in those information arrivals when released to the public. Then we estimate the impact of these events on conditional returns, return volatility, and return covariance using several extensions of the multivariate generalized autoregressive conditional heteroskedasticity-dynamic conditional correlation (GARCH-DCC) integrated moving average (IMA) model of Engle (2002). This specification, while more flexible and parsimonious than most available multivariate models, has been shown to perform equally well in a variety of situations.

The objective of our study is to contribute to the empirical literature relating macroeconomic fundamentals, and macroeconomic news in particular, to asset pricing dynamics. Studies in this literature differ on multiple grounds: their choice of news, their choice of market (bonds, stocks, or currencies), the moments of the return distribution they examine, and the statistical methodology they employ.⁴

¹Jones et al. (1998) provide some supporting evidence for this argument in the U.S. Treasury market using actual (rather than unexpected) employment and producer price announcements.

²See, for example, the discussion in Hasbrouck (2006).

³The main properties of these data sets and references to their use in the literature are given in Section II.

⁴An incomplete list of the most recent research includes Jones et al. (1998), Li and Engle (1998), Fleming and Remolona (1999), Bomfim and Reinhart (2000), Balduzzi et al. (2001), Kuttner (2001),

Only a few of these studies focus on the simultaneous impact of macroeconomic news releases on both stocks and government bonds or their futures and none on the corporate bond market. Some of these studies use the *actual* news releases, i.e., do not separate the expected component of the released information from the unexpected one (e.g., Jones et al. (1998), Fleming and Remolona (1999)). Most of these studies concentrate on the first and second moments of asset returns. Fleming et al. (1998) consider the volatility linkages of stocks, bonds, and money market instruments, while Connolly, Stivers, and Sun (2005), (2007) investigate the impact of stock uncertainty (measured by the implied volatility from equity index options) on the time-varying comovement between government bond and stock returns within and across countries. Yet neither of these studies specifically relates the estimated spillovers to macroeconomic news.

Our study contributes to the aforementioned literature on several dimensions. First, we investigate the impact of important U.S. macroeconomic news on the joint distribution of the returns in three important U.S. financial markets: equity, Treasury bonds, and corporate bonds. This is arguably a reasonable course of action, since none of these markets exists in a vacuum and investors can, and often do, hold and trade many of these securities at the same time. Second, we use survey and futures data to extract the unexpected components of these news announcements. Third, we analyze the impact of their releases not only on the returns of the three asset classes but also on their volatility and covariation. This effort allows us to provide a more comprehensive picture of the effect of surprise macroeconomic information on the behavior of asset prices in the U.S. capital markets.

We find that the process of price formation in the U.S. financial markets appears to be driven by fundamentals, albeit often heterogeneously so. Specifically, our evidence reveals a statistically and economically significant dichotomy, to our knowledge novel to the literature, between the reaction of the stock and bond markets to the release of unanticipated fundamental information. After initially falling, conditional stock return volatility increases on the day that these surprise macroeconomic news items are released. Conditional bond return volatility instead first rises and then declines, the more so the shorter the maturity of the bond portfolio and the lower its likelihood of default. We also show that conditional return comovement within and across stock and bond markets most often decreases—rather than increasing, as commonly believed—in correspondence with those releases, even after controlling for volatility shifts. Finally, we find that the conditional dynamics of stock, Treasury, and corporate bond returns react asymmetrically to the *specific* information content of surprise announcements (e.g., especially, but not exclusively, to the release of “bad” news). Overall, these results shed new light on the mechanisms by which information is incorporated into prices.

The remainder of the paper is organized as follows. Section II describes our database. Section III develops our econometric approach to estimating the impact

Flannery and Protopapadakis (2002), Andersen et al. (2003), (2007), Bomfim (2003), Connolly and Wang (2003), Bernanke and Kuttner (2005), and Boyd, Hu, and Jagannathan (2005). Some recent studies also analyze the impact of macroeconomic news releases on stock and bond market liquidity (e.g., Brandt and Kavajecz (2004), Green (2004), Vega (2006), and Pasquariello and Vega (2007)).

of unexpected news releases on asset returns. Section IV discusses the first set of results for conditional returns, return volatility, and return comovement. Section V tests for the differential effect of positive and negative news on stock and bond return dynamics. Section VI concludes.

II. Data

The basic data set we use in this paper consists of a variety of daily, continuously compounded excess holding-period returns (over 3-month Treasury bills) on three asset classes—stocks, Treasury bonds, and corporate bonds—whose prices are expected to be affected by four macroeconomic announcements: the *total* CPI, the unemployment rate, and nonfarm payroll employment, exogenously released on a monthly basis; and the target federal funds rate, potentially endogenous to the former variables. Our choice of macroeconomic announcements is motivated by several considerations. First, there is broad agreement that the level and dynamics of employment and inflation within a country represent the main concern of that country's monetary authorities when setting their interest rate policy. Second, there is ample evidence that financial markets are most sensitive to *direct* news about unemployment and inflation (e.g., Fleming and Remolona (1999), Krueger and Fortson (2003), Boyd et al. (2005), and Pasquariello and Vega (2007)). Third, the impact of other, *indirect* announcements about unemployment and inflation (such as retail sales or consumer confidence) on any of the asset classes that we study can only weaken rather than enhance the significance of our results.

We analyze seven time series of asset returns: two for the stock market, three for the Treasury bond market, and two for the corporate bond market. Excess stock holding-period returns (including distributions) are computed for the Center for Research in Security Prices value-weighted portfolios made up of NYSE and AMEX stocks (r_t^{NYX}) and of NASDAQ stocks (r_t^{NAQ}). We calculate holding-period returns on Treasury bonds using the constant-maturity 5-, 10-, and 30-year time series of interest rates constructed by the U.S. Treasury from yields on actively traded issues. These rates are converted into excess 5-year (r_t^{5Y}), 10-year (r_t^{10Y}), and 30-year (r_t^{30Y}) holding-period returns on hypothetical par bonds with the corresponding maturity, as in Jones et al. (1998). A similar procedure is applied to two portfolios of corporate bonds, one made of Aaa-rated (r_t^{Aaa}) and the other made of Baa-rated (r_t^{Baa}) U.S. corporate bonds, with maturity equal to or greater than 20 years, from Moody's Investors Service.⁵ Aaa-rated bonds carry the smallest degree of investment risk. Baa-rated bonds are neither highly protected nor poorly secured, and are generally deemed to have some speculative characteristics. Our sample covers a period of roughly 16 years, the longest period over which data on all assets under examination are *continuously* available: from January 3, 1986 (the first day for which daily Baa portfolio yields become available), to February 14, 2002 (the last day for which constant-maturity 30-year

⁵For these two portfolios, we assume that the hypothetical par corporate bonds have a maturity of 20 years.

Treasury bond yields are available).⁶ This period also largely overlaps with the tenure of Alan Greenspan as the chairman of the Board of Governors of the Federal Reserve System.⁷

Summary statistics for these series are reported in Table 1. Returns are in percentage (i.e., multiplied by 100). Not surprisingly, given the growth experienced by the U.S. stock market in the past three decades, the mean excess returns r_t^{NYSE} and r_t^{NASDAQ} are positive, significant, and the highest among the asset classes under examination. Excess bond returns are positive as well and increase with maturity and likelihood of default. Daily excess returns are also characterized by little or no skewness, strong leptokurtosis, and small positive autocorrelation ($\hat{\rho}_1 > 0$). Finally, the Ljung-Box portmanteau test for up to the fifth-order serial correlation (LB(5)) strongly rejects the null hypothesis that excess holding-period returns on stocks, Treasury, highest-grade (Aaa), and medium-grade (Baa) bonds are white noise.

TABLE 1
Descriptive Statistics: Excess Returns

Table 1 reports summary statistics for the time series of daily continuously compounded excess holding-period returns for the value-weighted portfolios of NYSE and AMEX stocks (r_t^{NYSE}), NASDAQ stocks, (r_t^{NASDAQ}), constant-maturity 5-year Treasury bonds (r_t^{5Y}), constant-maturity 10-year Treasury bonds (r_t^{10Y}), constant-maturity 30-year Treasury bonds (r_t^{30Y}), Aaa-rated corporate bonds (r_t^{Aaa}), and Baa-rated corporate bonds (r_t^{Baa}) described in Section III, over the 3-month Treasury bill, for the sample from 01/03/1986 to 02/14/2002 (the last day for which r_t^{30Y} is available). Returns are expressed in percentages (i.e., multiplied by 100). N is the number of available observations for each series, μ is the mean, σ is the standard deviation, and Min and Max are the minimum and maximum value for the series over the sample. Skew is the coefficient of skewness, and Kurt is the excess kurtosis; their standard errors for asymptotic normal distributions are $(6/N)^{1/2}$ and $(24/N)^{1/2}$, respectively. $\hat{\rho}_1$ is the estimated first-order autocorrelation, while LB(5) is the value of the Ljung-Box test of randomness for up to the fifth-order autocorrelation, asymptotically distributed as $\chi^2_5[5]$ under the null hypothesis that the series is white noise. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Daily Excess Holding-Period Returns						
	r_t^{NYSE}	r_t^{NASDAQ}	r_t^{5Y}	r_t^{10Y}	r_t^{30Y}	r_t^{Aaa}	r_t^{Baa}
μ	0.053***	0.056**	0.030***	0.034***	0.039**	0.040***	0.043***
σ	0.940	1.442	0.280	0.451	0.685	0.424	0.393
Min	-18.233	-11.453	-1.721	-2.774	-3.882	-2.991	-3.039
Max	8.799	14.234	2.948	4.629	7.191	3.405	1.973
$\hat{\rho}_1$	0.085***	0.082***	0.063***	0.054***	0.036**	0.073***	0.059***
LB(5)	45.37***	41.67***	24.77***	25.59***	15.10***	37.37***	25.18***
Skew	-2.181***	-0.123***	0.132***	-0.031	0.156***	-0.304***	-0.511***
Kurt	41.565***	9.952***	5.427***	4.908***	5.147***	4.846***	3.803***
N	4,069	4,069	4,069	4,069	4,069	4,069	4,069

We assemble a database of significant macroeconomic announcements over our sample interval. We collect information on all meetings of the Federal Open Market Committee (FOMC) from Bloomberg. In particular, we focus on target federal funds rate decisions, which represent the most explicit public disclosure by the Federal Reserve of its stance on monetary policy over the sample period.⁸

⁶The U.S. Treasury suspended issuing 30-year bonds in October 2001 and resumed issuing them only in February 2006. Consequently, the 30-year Treasury constant-maturity series was discontinued on February 18, 2002, and reintroduced only on February 9, 2006.

⁷Alan Greenspan was sworn in as chairman of the Federal Reserve in August 11, 1987, succeeding Paul Volcker.

⁸Recent studies (e.g., Boukus and Rosenberg (2006)) find that U.S. financial markets respond not only to such *quantitative* information as FOMC rate decisions but also to such *qualitative* information as official statements, speeches, or FOMC minutes released by the Federal Reserve to the public.

By law, the FOMC must meet at least four times a year, but it has held a minimum of eight scheduled meetings per year since 1981. Nonetheless, Bomfim and Reinhart (2000) observe that more than 75% of the changes in the intended federal funds rate from 1989 to 1993 occurred between those meetings. Furthermore, until the end of 1993 the Federal Reserve made the vast majority of its interest rate decisions in the afternoon, when the federal funds market in New York was virtually closed, and declared them by conducting open-market operations the next day. However, since March 28, 1994, the federal funds rate has been released regularly at 2:15PM Eastern Standard Time (EST). We control for this delay by shifting the “effective” federal funds announcement date by 1 day until then.⁹

Monthly real-time (*actual*) announcements (i.e., as reported in the original press releases) on total CPI (CPI_t , in percentage in Table 2), nonfarm payroll changes (PAY_t , in thousands in Table 2), and unemployment (UNE_t , in percentage in Table 2) and their corresponding releases dates are obtained from the Bureau of Labor Statistics (BLS).¹⁰ We classify these economic announcements according to whether their content was anticipated by the market. For that purpose, we use the database of forecasts compiled by MMS. This database is available to us for the period 1986–2002.¹¹ Over the past two decades, MMS has surveyed dozens of economists and money managers, collecting their forecasts for a broad range of macroeconomic variables. MMS surveys are conducted via telephone every last Friday prior to each news announcement, and the resulting median forecasts are released during the following week. According to several studies of their information content (e.g., Urich and Wachtel (1984), Balduzzi et al. (2001)), MMS expectations are generally unbiased and less noisy estimates of the corresponding realized macroeconomic variables than those generated by extrapolative models (e.g., Figlewski and Wachtel (1981)). Balduzzi et al. (2001) also show that MMS expectations are not stale, i.e., are likely to be based upon all information available at the time of the survey (including other, most recent macroeconomic announcements).¹²

Consistent with the existing literature on macroeconomic news arrivals (e.g., Balduzzi et al. (2001), Andersen et al. (2003), (2007)), we define an announcement of type e on day t , A_t^e , as a *surprise* if its absolute difference with respect to the “market consensus,” the corresponding median forecast F_t^e , is “large”

This evidence, albeit consistent with the FOMC’s increased commitment to transparency over the last few years of our sample, is sensitive to the specific algorithm that extracts common themes from multi-page, often ambiguous statements.

⁹For more on the timing discrepancy between FOMC announcements and the end of trading in the federal funds market, see Kuttner (2001) and Bernanke and Kuttner (2005).

¹⁰The monthly news releases containing CPI data are usually made available to the public at 8:30AM EST on a day in the second full week of each month. News releases of monthly changes in nonfarm payroll and unemployment instead come (again at 8:30AM EST) about 3 weeks into the month.

¹¹Extending our sample beyond 2002 is problematic. Indeed, since being acquired by Informa in 2003, MMS has stopped providing its survey services. Recently, another company, Action Economics, has resumed collecting economists’ forecasts of U.S. macroeconomic variables, although not from the date when MMS ended its survey. Furthermore, it is not clear whether Action Economics has been using exactly the same survey techniques previously employed by MMS.

¹²For a more detailed description of the MMS database and its properties, see Andersen et al. (2003), (2007).

TABLE 2
Descriptive Statistics: Macroeconomic Events

Table 2 reports summary statistics for the following monthly preliminary news releases: Federal Reserve target rates (FED, in percentage), target rate increases (FED⁺), target rate decreases or unchanged (FED⁻), Consumer Price Index (CPI) inflation (CPI, in percentage) from the Bureau of Labor Statistics (BLS), CPI increases (CPI⁺), CPI decreases or unchanged (CPI⁻), unemployment rate change (UNE, in percentage) from the BLS, unemployment rate increase (UNE⁺), decreased or unchanged unemployment rate (UNE⁻), nonfarm payroll employment change (PAY, in thousands) from the BLS, increase in nonfarm payroll employment (PAY⁺), and decreased or unchanged nonfarm payroll employment (PAY⁻). The sample is from 01/03/1986 to 02/14/2002. *N* is the number of available observations for each series; μ is the mean, Med is the median, and σ is the standard deviation of the announced variable; and Min and Max are the minimum and maximum value for the announced variable over the sample. Market expectations for these announcements are from MMS; announcement surprises are computed with respect to market consensus according to a procedure described in Section III. Here, μ_{SUR} , Med_{SUR}, σ_{SUR} , and *N*_{SUR} are the mean, median, standard deviation, and number of available observations, respectively, of the surprise difference (when positive, +, negative, -, or both) between the announced and expected amount. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Macroeconomic Events											
	FED			CPI			UNE			PAY		
	FED	FED ⁺	FED ⁻	CPI	CPI ⁺	CPI ⁻	UNE	UNE ⁺	UNE ⁻	PAY	PAY ⁺	PAY ⁻
μ	-0.034*	0.342***	-0.117***	0.256***	0.294***	-0.084***	-0.006	0.179***	-0.105***	146.1***	216.8***	-118.9***
Med	0.000	0.250	0.000	0.200	0.300	0.000	0.000	0.100	-0.100	167.5	209.0	-96.00
σ	0.246	0.149	0.175	0.201	0.168	0.146	0.170	0.106	0.102	180.3	122.8	92.92
Min	-0.500	0.055	-0.500	-0.400	0.100	-0.400	-0.400	0.100	-0.400	-415.0	2.000	-415.0
Max	0.750	0.750	0.000	1.100	1.100	0.000	0.600	0.600	0.000	705.0	705.0	-1.000
<i>N</i>	161	29	132	189	170	19	190	66	124	190	150	40
μ_{SUR}	-0.011	0.271***	-0.256***	-0.015	0.152***	-0.125***	-0.040**	0.162***	-0.166***	-10.09	110.3***	-112.2***
Med _{SUR}	-0.060	0.180	-0.145	-0.100	0.100	-0.100	-0.100	0.100	-0.200	-31.00	90.00	-91.00
σ_{SUR}	0.380	0.296	0.255	0.149	0.071	0.053	0.183	0.108	0.075	133.6	74.99	73.63
<i>N</i> _{SUR}	71	33	38	128	51	77	140	54	86	146	67	79

(i.e., greater than a predetermined threshold). In this study, we use a threshold of 5 basis points (bp) for CPI_t and UNE_t , and of 20,000 jobs for PAY_t ; the results that follow are not meaningfully affected by alternative thresholds, since the vast majority of expected announcements occurred at or very close to the median forecast. We use a similar approach (and the same 5 bp threshold) for the target rate decisions by the Federal Reserve (FED_t , in percentage in Table 2). Nonetheless, we measure the corresponding market expectations using the 30-day interest rate futures contract traded on the Chicago Board of Trade (CBOT), since changes in the federal funds futures rate, an intuitive aggregation of market-wide policy expectations, have been shown to represent more efficient predictors of FOMC's target rate changes.¹³ We employ the algorithm devised by Kuttner (2001) to estimate the 1-day rate surprise from the 1-day change in the spot-month futures rate around FOMC announcements. Because CBOT futures prices are available only from October 3, 1988, we use MMS forecasts for the first 3 years of the sample.

According to Table 2, there were 161 FOMC meetings over our sample period 1986–2002. During that period, the Federal Reserve decided to increase the federal funds rate much less frequently rather than either leaving it unchanged or cutting it. About 44% of all rate decisions were unexpected by the market. The CPI series suggests an annual inflation rate of 3% between 1986 and 2002. Not surprisingly, given the behavior of the U.S. economy over the last two decades,

¹³Krueger and Kuttner (1996), Bomfim and Reinhart (2000), and Kuttner (2001) explore the properties of various proxies for market expectations of FOMC rate changes implied by federal funds future rates. Piazzesi and Swanson (2008) find that forecasts from 1-day changes in federal funds futures rates are the least affected by biases induced by time-varying risk premia.

CPI deflation was much less common ($CPI_t \leq 0$ only 19 times), and almost always went undetected by the market. Recessions were of shorter length ($UNE_t > 0$ in 66, and $PAY_t \leq 0$ in 40, of 190 BLS press releases) and generally exhibited greater per month intensity than expansions, but both were equally difficult to predict on average (both A_t^{UNE} and A_t^{PAY} were sufficiently different from F_t^{UNE} and F_t^{PAY} more than 70% of the times, regardless of their sign).

III. The Empirical Model

We now study the short-term anticipation and response of the U.S. stock, Treasury, and corporate bond markets to the arrival of relevant U.S. macroeconomic news. The goal of our multimarket analysis is to shed light not only on the effects of these announcements on conditional mean excess returns and return volatility, but also on the comovement and interaction between those markets in the proximity of their occurrence.

The GARCH specification proposed by Bollerslev (1986) and its many univariate and multivariate extensions are among the most widely adopted models that describe time-varying volatility and covariances.¹⁴ However, in most cases, multivariate GARCH models are not flexible enough, and the number of parameters in them too large, to introduce complex forms of conditional comovement among asset returns. In this paper, we use the DCC multivariate model of Engle (2002) to analyze the short-term behavior of the U.S. financial market in proximity to the release of macroeconomic news. The DCC specification has the flexibility of univariate GARCH models without the complexity of traditional multivariate GARCH specifications. We begin by proposing the following GARCH(1,1) model to describe the evolution of daily excess holding-period asset returns r^i :

$$\begin{aligned}
 (1) \quad r_t^i &= \mu_t^e + \rho_t^e r_{t-1}^i + \gamma_t^e(0) I_t^e(0) S_t^e(0) + \varepsilon_t^i, \\
 \varepsilon_t^i &= \sqrt{h_t^i} e_t^i \quad e_t^i | F_{t-1} \sim N(0, 1), \quad \text{and} \\
 h_t^i &= s_t^i \left[\omega_t^e + \alpha_t^e (e_{t-1}^i)^2 + \beta_t^e h_{t-1}^i \right],
 \end{aligned}$$

where F_{t-1} denotes the information set at time $t - 1$ and $s_t^i = 1 + \sum_{k=-1}^{+1} \delta_i^e(k) I_t^e(k) |S_t^e(k)|$. As standard in the finance literature, equation (1) specifies a first-order autocorrelation model for excess holding-period returns, to control for nonsynchronicity in prices, microstructure effects, and gradual convergence to equilibrium. Yet the above specification is novel to the literature, for it allows for both the *arrival* and the *extent* of macroeconomic news surprises of type e (either FED_t , CPI_t , UNE_t , or PAY_t) to affect not only conditional mean asset returns (e.g., as in Andersen et al. (2003), (2007), Bomfim (2003)) but also the conditional variance of excess return innovations ε_t^i .

Specifically, in equation (1), $I_t^e(k)$ is an event dummy variable equal to 1 if a surprise macroeconomic event of type e occurred at time $t + k$, and 0 otherwise

¹⁴See, for example, Nelson and Foster (1994), Engle (1995), and Kroner and Ng (1998) for a review.

(e.g., as in Jones et al. (1998)), while S_t^e are news surprises standardized by their sample standard deviation $\hat{\sigma}^e$ (i.e., $S_t^e = (A_t^e - F_t^e) / \hat{\sigma}^e$) to control for differences in units of measurement across announcements (as in Balduzzi et al. (2001), among others), and $S_t^e(k) = S_{t+k}^e$ if $I_t^e(k) = 1$, and 0 otherwise. Thus, the coefficient $\gamma_i^e(0)$ captures the average impact of the *actual* surprise announcements $S_t^e(k)$ on the mean excess return on the announcement dates, while the dummy coefficients $\delta_i^e(k)$ proxy for the sequential impact of *absolute* standardized macroeconomic news surprises $|S_t^e(k)|$ on conditional return variance. The modeling choice to interact $|S_t^e(k)|$, rather than $S_t^e(k)$, with $I_t^e(k)$ in the expression for h_t^i in equation (1) is motivated by existing evidence (e.g., Jones et al. (1998), Bomfim (2003)) and by preliminary unreported analysis indicating that the mere occurrence of surprise macroeconomic events, regardless of their sign, affects conditional return volatility.¹⁵ Hence, allowing for signed news surprises in S_t^i would weaken both the magnitude and significance of their estimated impact on h_t^i .¹⁶ We can interpret the resulting coefficient $\delta_i^e(1)$ as a measure of *anticipation*, i.e., as the percentage impact of the release of a unit absolute news surprise of type e on the conditional variance of the excess return r^i , before that release actually occurs at time $t + 1$. The coefficient $\delta_i^e(0)$ is a proxy for the additional, *contemporaneous* percentage impact of a unit absolute news surprise released at time t on h_t^i . Finally, the coefficient $\delta_i^e(-1)$ is a measure of *persistence*, i.e., of the additional percentage impact of the arrival of a unit absolute news surprise on h_t^i , after the information has already been revealed at time $t - 1$.¹⁷

As mentioned earlier, we also study the impact of news arrivals on the structure of conditional comovement among asset classes. To that purpose, measurement issues are the object of considerable debate. In particular, Forbes and Rigobon (2002) argue that shocks to the conditional covariance among asset returns in proximity to certain macroeconomic events may be due to shocks to return volatility. We tackle this problem by assuming that the conditional covariance between any two standardized residuals

$$\eta_t^i = \frac{\varepsilon_t^i}{\sqrt{s_t^i h_t^i}} \quad \text{and} \quad \eta_t^j = \frac{\varepsilon_t^j}{\sqrt{s_t^j h_t^j}}$$

at time t , given F_{t-1} , q_t^{ij} , is described accurately by the following exponential smoothing function:

$$(2) \quad q_t^{ij} = s_t^{i*} \left[\lambda^e q_{t-1}^{ij} + (1 - \lambda^e) \eta_{t-1}^i \eta_{t-1}^j \right],$$

¹⁵See, for example, the evidence on the estimation of a GARCH(1,1) model similar to equation (1) but in which $I_t^e(k)|S_t^e(k)|$ is replaced by $I_t^e(k)$ in a previous version of this paper (available at http://papers.ssrn.com/sol3/papers.cfm?abstract_id=676145) as well as our analysis of *signed* announcement surprises in Section V.

¹⁶The ensuing inference for h_t^i is unaffected by replacing $S_t^e(0)$ with $|S_t^e(0)|$ in the expression for r^j in equation (1).

¹⁷Equivalently, we can interpret the model of equation (1) as the amended, conditional version of a simple unconditional event study of the behavior of the second moment of asset returns around the official dates of the release of unexpected macroeconomic information.

where $s_t^{i*} = 1 + \sum_{k=-1}^{+1} d_{i,j}^e(k) I_t^e(k) |S_t^e(k)|$. We add $\eta_t^i \eta_t^j$ to both sides of equation (2) and rearrange terms to obtain an IMA process with no intercept,

$$(3) \quad \eta_t^i \eta_t^j = s_t^{i*} \eta_{t-1}^i \eta_{t-1}^j + \left(q_t^{ij} - \eta_t^i \eta_t^j \right) - \lambda^e s_t^{i*} \left(q_{t-1}^{ij} - \eta_{t-1}^i \eta_{t-1}^j \right),$$

where the errors $q_t^{ij} - \eta_t^i \eta_t^j$, with a conditional mean of 0, are Martingale differences by construction and can be thought of as white noise sequences. Equation (3) is an extension of the DCC IMA model of Engle (2002) with common cross-asset adjustments (λ^e). More importantly, the above model allows us to measure the anticipated ($d_{i,j}^e(1)$), contemporaneous ($d_{i,j}^e(0)$), and persistent ($d_{i,j}^e(-1)$) percentage impact of the release of a unit absolute news surprise of type e on the conditional covariance between any pair of standardized residuals η_t^i and η_t^j in a parsimonious way, while controlling for the impact of these arrivals on conditional volatility and mean excess returns, as advocated by Forbes and Rigobon (2002). Indeed, any common fundamental information shock stemming from the news arrival at time t directly affects the processes for r_t^i and h_t^i in equation (1) via the parameters $\gamma_i^e(0)$, $\delta_i^e(-1)$, $\delta_i^e(0)$, and $\delta_i^e(1)$.¹⁸

The model of equations (1) and (3) provides a basic representation of the behavior of asset returns in proximity to macroeconomic news announcements. As such, it is designed to capture only the first-order, symmetric impact of news arrivals on the U.S. equity and bond markets. Therefore, evidence from its estimation in support of any of the arguments discussed in Section I can be interpreted as strong evidence. We investigate second-order effects, in particular the possibly asymmetric impact of “good” and “bad” news on return dynamics, in Section V.

As first suggested by Engle (2002), we estimate the model of equations (1) and (3) in two steps. In the first step, we estimate equation (1) separately for each asset i by the quasi-maximum likelihood (QML) procedure described in Bollerslev and Wooldridge (1992). Hence, the resulting estimates are consistent and asymptotically efficient. In the second step, we estimate the parametric model of equation (3) jointly for all assets i , again by QML, using the parameters obtained in the first step. Under some mild regularity conditions, consistency of those parameters ensures consistency, although not efficiency, of the estimates stemming from the second step. Engle and Sheppard (2001), Engle (2002), and Cappiello, Engle, and Sheppard (2006) show that these estimates perform well in a variety of situations and provide reasonable empirical results with a minimal loss of efficiency.¹⁹

¹⁸In a recent paper, Andersen, Bollerslev, Diebold, and Labys (2003) argue that model-free estimates of daily variances from intraday return data perform well in comparison with standard GARCH models of daily volatility. Along those lines, Christiansen and Rinaldo (2007) use futures intraday data to measure realized Treasury bond-U.S. stock return correlations in proximity to macroeconomic news announcements. However, the GARCH(1,1)-DCC IMA model of equations (1) and (3) allows us to control for the contemporaneous impact of these news arrivals on conditional return levels when estimating realized volatility, as well as on conditional return volatility when estimating realized return correlations.

¹⁹Efficiency of the second-step DCC IMA parameters could be achieved by *jointly* estimating both equations (1) and (3). Yet convergence is often problematic, since the resulting log-likelihood function is often “extremely flat” (Cajigas and Urga (2009)). Joint estimation of our basic model and its extensions by QML (when successful, in unreported analysis) leads to virtually identical inference.

IV. The Basic Results

We start our analysis by estimating the model of equations (1) and (3) for each of the four macroeconomic announcements in our sample according to the two-step procedure described in Section III. We report the resulting QML estimates (and robust t -statistics) of selected parameters of equations (1) and (3) in Tables 3 and 4, respectively, when $e = \text{FED, CPI, UNE, or PAY}$. In unreported analysis, the inclusion of dummy variables to control for day-of-the-week effects or the day-level clustering of other macroeconomic announcements (e.g., Jones et al. (1998), Kim, McKenzie, and Faff (2004)) does not significantly affect our inference, either qualitatively or quantitatively.²⁰

Table 3 reveals a dichotomy between the reaction of the U.S. stock and bond markets to economic news. Conditional stock return volatility is somewhat *lower* the day before ($\delta_i^e(1) < 0$) and significantly higher the day of their arrival ($\delta_i^e(0) > 0$). In particular, our estimates indicate that between 1986 and 2002, upon the release of a unit absolute standardized macroeconomic news surprise (with the exception of $e = \text{FED}$), conditional return volatility for the NYSE-AMEX (NASDAQ) is on average no less than 50% (35%) and almost 100% (60%) higher than during nonannouncement days. This is significant, since those surprise announcement windows are frequent, accounting for about 21% of all days in our sample (see Tables 1 and 2). In contrast, we find that the coefficients $\delta_i^e(1)$ are always *positive* and mostly statistically significant for conditional government and corporate bond return volatility. For instance, the day before the release of a unit absolute unemployment news ($e = \text{UNE or PAY}$), return volatility increases by a minimum of 62% ($\delta_i^{\text{PAY}}(1) = 0.621$) for Aaa-rated long-term corporate bonds to a maximum of 115% ($\delta_i^{\text{PAY}}(1) = 1.152$) for 5-year Treasury bonds. This suggests that there is a considerable increase in uncertainty among bond market participants in anticipation of the release of macroeconomic news. This effect is only partially short-lived, since the contemporaneous coefficients $\delta_i^e(0)$, albeit often significantly negative (at the 1% level), are of much lower magnitude (e.g., $\delta_i^{\text{PAY}}(0) = -0.249$ for 5-year Treasury bonds), while $\delta_i^e(-1)$ is either small, negative, and weakly significant, or 0. Stock return volatility instead declines appreciably after the announcements, especially for NYSE-AMEX ($\delta_i^{\text{CPI}}(-1) = -0.189$, $\delta_i^{\text{UNE}}(-1) = -0.119$, and $\delta_i^{\text{PAY}}(-1) = -0.221$ in Table 3). This evidence indicates that in the U.S. bond market only, any additional information stemming from economic news does not appear to augment price fluctuations; i.e., this news appears to induce (at least partial) resolution of uncertainty and/or disagreement only among bond market participants.

Accordingly, qualitatively similar inference can also be drawn from estimating equation (3) and its extensions with a generalized method of moments procedure in the spirit of that proposed by Ang, Piazzesi, and Wei (2006) to account for the sampling uncertainty of the first-step GARCH parameters.

²⁰Furthermore, the timing of the macroeconomic releases in our sample is not clustered and (with the exception of the few unscheduled FOMC meetings over our sample period) can be deemed exogenous to financial markets. Yet our model does not allow assessing whether the clustering of various unexpected macroeconomic news releases in the same direction over a short period of time may have larger effects on U.S. financial markets than the average effect of each news release *separately*, as stemming from the estimation of equations (1) and (3). In these circumstances, our approach is likely to bias such effects downward. We thank the referee for pointing this out.

TABLE 3
GARCH Model for Excess Returns: Surprise Events

Table 3 reports quasi-maximum likelihood (QML) estimates and robust *t*-statistics (Bollerslev and Wooldridge (1992), in parentheses) over a sample from 01/03/1986 to 02/14/2002 (4,069 observations) for the GARCH(1, 1) model:

$$\begin{aligned}
 (1) \quad r_t^i &= \mu_i^e + \rho_i^e r_{t-1}^i + \gamma_i^e(0)I_t^e(0)S_t^e(0) + \varepsilon_t^i, \\
 \varepsilon_t^i &= \sqrt{s_t^i}e_t^i \quad e_t^i|F_{t-1} \sim N(0, h_t^i), \quad \text{and} \\
 h_t^i &= \omega_i^e + \alpha_i^e (e_{t-1}^i)^2 + \beta_i^e h_{t-1}^i,
 \end{aligned}$$

where F_{t-1} denotes the information set, $s_t^i = 1 + \sum_{k=-1}^{+1} \delta_i^e(k)I_t^e(k)|S_t^e(k)|$, r_t^i is the daily continuously compounded excess return on asset *i* (in percentage), and S_t^e are actual standardized news. In equation (1), $I_t^e(k) = 1$ and $S_t^e(k) = S_t^e$ if a surprise Federal Reserve rate change ($e = \text{FED}$), a surprise CPI ($e = \text{CPI}$), unemployment ($e = \text{UNE}$), or payroll announcement ($e = \text{PAY}$) was made on day $t+k$, and 0 otherwise. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	r_t^{NYX}	r_t^{NAO}	r_t^{5Y}	r_t^{10Y}	r_t^{30Y}	r_t^{Aaa}	r_t^{Baa}
$\gamma_i^{\text{FED}}(0)$	-0.125 (-1.54)	-0.135* (-1.71)	-0.004 (-0.14)	-0.015 (-0.34)	0.018 (0.30)	0.002 (0.05)	-0.009 (-0.30)
$\delta_i^{\text{FED}}(-1)$	-0.134*** (-3.91)	-0.127*** (-3.81)	-0.049 (-0.79)	-0.030 (-0.47)	-0.026 (-0.45)	-0.018 (-0.35)	-0.027 (-0.54)
$\delta_i^{\text{FED}}(0)$	0.048 (0.24)	0.099 (0.56)	-0.067 (-1.18)	0.021 (0.22)	0.070 (0.75)	0.007 (0.11)	0.036 (0.44)
$\delta_i^{\text{FED}}(1)$	0.138 (0.65)	0.062 (0.42)	0.450*** (2.67)	0.129 (1.09)	0.021 (0.27)	0.075 (0.89)	0.173 (1.54)
$\gamma_i^{\text{CPI}}(0)$	0.089 (1.28)	0.040 (0.63)	0.012 (0.67)	0.012 (0.40)	0.033 (0.71)	0.013 (0.59)	0.022 (0.89)
$\delta_i^{\text{CPI}}(-1)$	-0.189*** (-3.74)	0.002 (0.03)	0.157 (1.51)	0.119 (1.18)	0.039 (0.44)	-0.074 (-1.04)	-0.121** (-2.10)
$\delta_i^{\text{CPI}}(0)$	0.988*** (4.53)	0.594*** (3.15)	-0.178*** (-2.85)	-0.178*** (-2.89)	-0.154** (-2.31)	-0.120 (-1.62)	-0.107 (-1.55)
$\delta_i^{\text{CPI}}(1)$	-0.104 (-1.64)	-0.199*** (-3.83)	0.177* (1.71)	0.213** (2.12)	0.224** (2.23)	0.107 (1.25)	0.178* (1.94)
$\gamma_i^{\text{UNE}}(0)$	-0.092 (-1.57)	-0.142** (-2.44)	0.012 (0.52)	0.011 (0.29)	-0.004 (-0.08)	0.020 (0.76)	-0.013 (-0.47)
$\delta_i^{\text{UNE}}(-1)$	-0.119*** (-2.91)	-0.098 (-1.33)	-0.155*** (-4.10)	-0.166*** (-5.13)	-0.129*** (-3.38)	-0.137*** (-2.96)	-0.087 (-1.51)
$\delta_i^{\text{UNE}}(0)$	0.543*** (3.57)	0.352*** (2.09)	-0.203*** (-14.86)	-0.183*** (-9.25)	-0.168*** (-7.05)	-0.160*** (-4.36)	-0.172*** (-6.23)
$\delta_i^{\text{UNE}}(1)$	-0.091 (-1.52)	-0.081 (-0.97)	1.035*** (6.81)	0.912*** (6.53)	0.643*** (5.54)	0.624*** (5.23)	0.597*** (4.83)
$\gamma_i^{\text{PAY}}(0)$	-0.104 (-1.51)	-0.066 (-0.91)	0.016 (0.73)	0.036 (1.00)	0.038 (0.71)	0.013 (0.49)	0.018 (0.67)
$\delta_i^{\text{PAY}}(-1)$	-0.221*** (-7.67)	-0.182*** (-4.19)	-0.148*** (-3.12)	-0.123** (-2.26)	-0.112** (-2.12)	-0.104* (-1.69)	-0.093 (-1.62)
$\delta_i^{\text{PAY}}(0)$	0.853*** (4.38)	0.465** (2.46)	-0.249*** (-11.99)	-0.236*** (-7.99)	-0.210*** (-6.40)	-0.163*** (-3.13)	-0.198*** (-5.35)
$\delta_i^{\text{PAY}}(1)$	-0.170*** (-3.76)	-0.088 (-1.04)	1.152*** (6.24)	0.943*** (5.77)	0.680*** (5.33)	0.621*** (4.73)	0.565*** (4.60)

The arrival of employment news has the greatest impact on the government bond market, especially before its release, while CPI news has the lowest; the corresponding effect on corporate bonds, although of the same sign, is of somewhat smaller scale. Interestingly, nonfarm payroll surprises, although released simultaneously with unemployment numbers, are preceded and accompanied by relatively more pronounced volatility shocks. This evidence is consistent with some recent studies (e.g., Andersen et al. (2007), Pasquariello and Vega (2007)) and

TABLE 4
DCC IMA Model: Surprise Events

Table 4 reports intra- and across-asset class averages of quasi-maximum likelihood (QML) estimates and their robust t-statistics (Bollerslev and Wooldridge (1992), in parentheses) for the following DCC IMA model:

$$(3) \quad \eta_t^i \eta_t^j = s_t^{i*} \eta_{t-1}^i \eta_{t-1}^j + (q_t^{ij} - \eta_t^i \eta_t^j) - \lambda^e s_t^{i*} (q_{t-1}^{ij} - \eta_{t-1}^i \eta_{t-1}^j),$$

where $s_t^{i*} = 1 + \sum_{k=-1}^{t-1} d_{i,j}^e(k) I_t^e(k) |S_t^e(k)|$, $\eta_t^i = \varepsilon_t^i / \sqrt{s_t^i h_t^i}$ is the daily standardized residual from the GARCH(1, 1) model of equation (1) for r_t^i , the excess return on asset i ; S_t^e are actual standardized news; and $I_t^e(k) = 1$ and $S_t^e(k) = S_{t-k}^e$ if a surprise Federal Reserve rate change ($e = \text{FED}$), a surprise CPI ($e = \text{CPI}$), unemployment ($e = \text{UNE}$), or a payroll announcement ($e = \text{PAY}$) was made on day $t + k$, and 0 otherwise. Equation (3) is estimated between 01/03/1986 and 02/14/2002 (i.e., over 4,069 observations). The STOCK category is made of $i = \text{NYSE, NAQ}$; the GOVT category is made of $i = 5Y, 10Y, 30Y$; and the CORP category is made of $i = \text{Aaa, Baa}$.

	$\overline{d_{i,j}^{\text{FED}}(-1)}$			$\overline{d_{i,j}^{\text{FED}}(0)}$			$\overline{d_{i,j}^{\text{FED}}(1)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.874 (-25.03)	-0.881 (-16.35)	-0.783 (-11.67)	0.181 (1.28)	-0.320 (-8.84)	-0.696 (-12.42)	-0.274 (-37.41)	-0.652 (-15.65)	-1.111 (-13.06)
GOVT		-0.330 (-7.31)	-0.912 (-14.85)		-0.429 (-7.75)	-0.685 (-28.32)		-0.248 (-7.73)	-0.012 (-2.51)
CORP			-0.151 (-11.39)			-0.794 (-38.15)			0.580 (4.05)
	$\overline{d_{i,j}^{\text{CPI}}(-1)}$			$\overline{d_{i,j}^{\text{CPI}}(0)}$			$\overline{d_{i,j}^{\text{CPI}}(1)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	0.700 (7.48)	0.348 (3.36)	-1.555 (-25.14)	0.564 (6.57)	-1.192 (-19.16)	-1.072 (-18.29)	-0.938 (-18.20)	-0.763 (-12.55)	-0.852 (-14.40)
GOVT		-2.685 (-22.64)	-1.152 (-15.12)		-0.389 (-11.96)	-0.466 (-13.90)		-0.323 (-7.58)	-0.657 (-17.92)
CORP			-0.106 (-2.46)			-0.255 (-6.92)			-0.808 (-21.42)
	$\overline{d_{i,j}^{\text{UNE}}(-1)}$			$\overline{d_{i,j}^{\text{UNE}}(0)}$			$\overline{d_{i,j}^{\text{UNE}}(1)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.206 (-4.99)	-0.727 (-14.50)	-0.844 (-17.74)	-1.138 (-37.20)	-1.033 (-15.85)	-0.934 (-17.29)	-1.267 (-30.54)	-1.096 (-20.94)	-0.992 (-18.91)
GOVT		-0.495 (-7.93)	-1.077 (-22.84)		-0.866 (-20.37)	-1.504 (-30.78)		0.091 (1.41)	-0.427 (-14.98)
CORP			-1.161 (-29.35)			-1.700 (-40.82)			-0.073 (-1.39)
	$\overline{d_{i,j}^{\text{PAY}}(-1)}$			$\overline{d_{i,j}^{\text{PAY}}(0)}$			$\overline{d_{i,j}^{\text{PAY}}(1)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.948 (-33.18)	-0.881 (-11.35)	-0.746 (-9.51)	-0.389 (-24.25)	-0.935 (-13.44)	-1.059 (-13.89)	-0.314 (-9.16)	-1.016 (-16.80)	-1.055 (-15.02)
GOVT		-0.545 (-13.29)	-0.362 (-11.39)		-1.191 (-27.96)	-1.224 (-31.40)		0.001 (0.02)	-0.032 (-0.69)
CORP			-0.300 (-4.52)			-1.294 (-34.37)			-0.009 (-0.17)

suggests that nonfarm payroll has the greatest information content among all public signals of the state of the U.S. economy available to investors and speculators. The milder reaction of both government and corporate bonds to CPI surprises may instead stem from the relative stability of inflation expectations and the Federal Reserve's significant credibility in fighting inflation over our sample period 1986–2002. Yet the evidence in Table 3 also suggests that the behavior of stock

and bond markets close to target rate decisions by the Federal Reserve is both less economically and less statistically significant than any other event in our sample.

The absolute magnitude of $\delta_i^e(1)$ and $\delta_i^e(0)$ is decreasing in the maturity of the bond portfolios and in the likelihood of default, as proxied by the Moody's ratings. The latter is somewhat surprising, given the greater sensitivity of corporate bonds of poorer quality to the business cycle (e.g., Gertler and Lown (1999)). The notion that 5-year Treasury bonds are the most liquid (e.g., Fleming (2003), Brandt and Kavajecz (2004)) may instead explain the former, since we would expect the intensity of information gathering and portfolio rebalancing induced by the news arrivals to be greater for more liquid securities.²¹ Accordingly, target rate decisions have a positive and statistically significant impact on the conditional return volatility exclusively of 5-year Treasury bonds ($\delta_i^{\text{FED}}(0) = 0.450$). Alternatively, mean reversion in short-term interest rates (e.g., Chapman and Pearson (2000) and references therein) may lead to a weaker impact of an information shock on longer-term bonds, since any impact of such shocks on the short end of the yield curve is expected to die out for its long end. U.S. stock markets instead display a remarkable homogeneity of responses to most surprise macroeconomic announcements. Specifically, the impact of their arrival on NYSE-AMEX conditional return and volatility is qualitatively and quantitatively similar to that estimated for NASDAQ, despite the fact that most "new economy" companies are usually deemed less sensitive to employment or inflation news than to the state of their specific industry. For instance, the conditional return volatility for both NYSE-AMEX and NASDAQ declines by roughly 13% on average ($\delta_i^{\text{FED}}(-1) = -0.134$ for r_i^{NYX} and $\delta_i^{\text{FED}}(-1) = -0.127$ for r_i^{NAQ} in Table 3) following unit absolute surprise Federal Reserve rate decisions.

At this stage of the analysis, we find weak or no evidence of a relation between the conditional mean excess holding-period returns of each of the asset classes we study and the release of macroeconomic news surprises. In particular, Table 3 shows that the corresponding estimated coefficient $\gamma_i(0)$ is insignificant for all but (NASDAQ) stocks ($\gamma_i^{\text{FED}}(0) = -0.135$ for r_i^{NAQ}), as in Bomfim (2003). Table 4 provides evidence of a much greater dichotomy between the stock and the bond markets when we analyze our estimates for the parameters of the DCC IMA covariance model of equation (3). There we report intra- and across-asset class averages (i.e., within and among stocks, S , government bonds, G , and corporate bonds, C) of both QML estimates of the dummy coefficients $d_{i,j}^e(k)$ in equation (3) and their robust t -statistics. These average estimates suggest that the release of the surprise economic announcements in our sample is preceded by sharply lower comovement among U.S. stock markets (e.g., $\overline{d_{S,S}^{\text{UNE}}}(1) = -1.267$), but by much less so among government and corporate bond markets. Declining comovement among NYSE-AMEX and NASDAQ stocks is often magnified on the day of news arrivals and the day afterward (e.g., $\overline{d_{S,S}^{\text{UNE}}}(0) = -1.138$ and $\overline{d_{S,S}^{\text{UNE}}}(1) = -0.206$), even after controlling for their impact on conditional stock return and return

²¹For example, see Chowdhry and Nanda (1991). Yet shorter-maturity bonds are also less sensitive to fluctuations of interest and inflation rate expectations potentially induced by that news. According to Table 3, this effect is weaker than the maturity (or liquidity) effect.

volatility. It is, however, during those days that comovement among government bond returns and among corporate bond returns drops most significantly for all announcement types. The dynamics of intra-asset class comovement of excess holding-period returns are relatively homogeneous in intensity and significance, except when $e = \text{CPI}$. For instance, it is only in those circumstances that the estimated preannouncement decoupling among NYSE-AMEX and NASDAQ returns is partially reversed ($\overline{d_{S,S}^{\text{CPI}}(1)} = -0.938$ but $\overline{d_{S,S}^{\text{CPI}}(0)} = 0.564$ and $\overline{d_{S,S}^{\text{CPI}}(-1)} = 0.700$), while the estimated announcement-day decline in comovement among government bond returns is the most persistent ($\overline{d_{G,G}^{\text{CPI}}(-1)} = -2.685$) and the weakest among Moody's portfolios of Aaa- and Baa-rated corporate bonds ($\overline{d_{C,C}^{\text{CPI}}(0)} = -0.255$ and $\overline{d_{C,C}^{\text{CPI}}(-1)} = -0.106$).

Finally, comovement among all asset classes decreases in proximity to all macroeconomic news releases. For example, we find that comovement between stocks and government bonds, stocks and corporate bonds, and government and corporate bonds declines on average by 103%, 93%, and 150%, respectively, in correspondence with the release of a unit absolute standardized unemployment news surprise ($\overline{d_{S,G}^{\text{UNE}}(0)} = -1.033$, $\overline{d_{S,C}^{\text{UNE}}(0)} = -0.934$, and $\overline{d_{G,C}^{\text{UNE}}(0)} = -1.504$ in Table 4). CPI and nonfarm payroll surprises lead to negative comovement shocks of similar magnitude. Federal Reserve rate decisions are instead accompanied by the smallest (yet still economically significant) drop in across-asset class covariance, consistent with the results reported in Table 3. This evidence is important, for it suggests that any multiasset trading activity due to portfolio rebalancing, information spillover, financial constraints, or shifts to the degree of information asymmetry and heterogeneity among traders induced by macroeconomic news, translates into either more negative or less positive—but not more positive, as commonly thought—comovement among the excess holding-period returns in our sample.

V. Asymmetric News Impact

In the previous section, we showed that the arrival and absolute magnitude of surprise macroeconomic announcements significantly affected the U.S. equity and bond markets between 1986 and 2002. The sign and magnitude of the impact of their release on the dynamics of asset prices should also depend on their *specific* information content. Indeed, it is reasonable to conjecture that the effect of “good” or “bad” news on the process of price formation in the U.S. financial markets may be asymmetric.²² For example, Boyd et al. (2005) find that in the short run, the reaction of stock and bond prices to unemployment news, measured with respect to a statistical model for final release numbers, is state-dependent. The equity market usually responds positively to surprisingly rising unemployment, while Treasury bond prices react to it only during expansions. Similarly, Veronesi (1999) argues that “good” or “bad” news may increase uncertainty when

²²In light of earlier discussion, we use the terms “good” and “bad” to refer exclusively to the content of the news, rather than to its implications for stock and bond valuations.

released in “bad” or “good” times, respectively.²³ Furthermore, it is possible that the various forms of market frictions and financial constraints described in the literature (e.g., borrowing, short-selling, and wealth constraints) are more binding on investors’ behavior following “negative” announcements. These constraints could then affect not only the level of asset returns (e.g., Diether, Malloy, and Scherbina (2002)) but also their volatility and comovement (e.g., Kyle and Xiong (2001)) differentially in proximity to news arrivals.

We explore these issues by amending the model of equations (1) and (3) to allow for asymmetric effects of news releases. Specifically, we introduce the following GARCH(1,1) model for conditional return volatility:

$$(4) \quad \begin{aligned} r_t^i &= \mu_i + \rho_i r_{t-1}^i + [\gamma_i^e(0, +) I_t^e(0, +) + \gamma_i^e(0, -) I_t^e(0, -)] S_t^e + \varepsilon_t^i, \\ \varepsilon_t^i &= \sqrt{s_t^{i\pm}} e_t^i \quad e_t^i | F_{t-1} \sim N(0, h_t^i), \quad \text{and} \\ h_t^i &= \omega_i + \alpha_i (e_{t-1}^i)^2 + \beta_i h_{t-1}^i, \end{aligned}$$

where $s_t^{i\pm} = 1 + \sum_{k=-1}^{+1} \delta_i^e(k, +) I_t^e(k, +) | S_t^e(k, +) | + \sum_{k=-1}^{+1} \delta_i^e(k, -) I_t^e(k, -) | S_t^e(k, -) |$, and then model comovement among excess return residuals as

$$(5) \quad \eta_t^i \eta_t^j = s_t^{i\pm*} \eta_{t-1}^i \eta_{t-1}^j + \left(q_{t-1}^{ij} - \eta_{t-1}^i \eta_{t-1}^j \right) - \lambda^e s_t^{i\pm*} \left(q_{t-1}^{ij} - \eta_{t-1}^i \eta_{t-1}^j \right),$$

where $s_t^{i\pm*} = 1 + \sum_{k=-1}^{+1} d_{i,j}^e(k, +) I_t^e(k, +) | S_t^e(k, +) | + \sum_{k=-1}^{+1} d_{i,j}^e(k, -) I_t^e(k, -) | S_t^e(k, -) |$, $\eta_t^i = \varepsilon_t^i / \sqrt{s_t^{i\pm}} h_t^i$, $I_t^e(k, +)$, and $S_t^e(k, +)$ are dummy variables equal to 1 and S_{t+k}^e , respectively, if the Federal Reserve announced either a surprisingly large rate increase or a surprisingly small rate cut on day $t+k$ ($S_{t+k}^{\text{FED}} > 0$), or if the CPI or the unemployment rate was reported at day $t+k$ to have either increased surprisingly much or decreased surprisingly little with respect to the previous month ($S_{t+k}^{\text{CPI}} > 0$ or $S_{t+k}^{\text{UNE}} > 0$), or if the nonfarm payroll was reported on day $t+k$ to have either decreased surprisingly much or increased surprisingly little with respect to the previous month ($S_{t+k}^{\text{PAY}} < 0$), and 0 otherwise; vice versa, $I_t^e(k, -)$ and $S_t^e(k, -)$ are dummy variables equal to 1 and S_t^e , respectively, if the Federal Reserve announced either a surprisingly large rate cut or a surprisingly small rate increase on day $t+k$ ($S_{t+k}^{\text{FED}} < 0$), or if the CPI or the unemployment rate was reported at day $t+k$ to have either decreased surprisingly much or increased surprisingly little with respect to the previous month ($S_{t+k}^{\text{CPI}} < 0$ or $S_{t+k}^{\text{UNE}} < 0$), or if the nonfarm payroll was reported on day $t+k$ to have either increased surprisingly much or decreased surprisingly little with respect to the previous month ($S_{t+k}^{\text{PAY}} > 0$), and 0 otherwise. Hence, $S_t^{\text{FED}}(k, +) (I_t^{\text{FED}}(k, +))$, $S_t^{\text{CPI}}(k, +) (I_t^{\text{CPI}}(k, +))$, $S_t^{\text{UNE}}(k, +) (I_t^{\text{UNE}}(k, +))$, and $S_t^{\text{PAY}}(k, -) (I_t^{\text{PAY}}(k, -))$ are “bad” news (dummy variables), while $S_t^{\text{FED}}(k, -) (I_t^{\text{FED}}(k, -))$, $S_t^{\text{CPI}}(k, -) (I_t^{\text{CPI}}(k, -))$, $S_t^{\text{UNE}}(k, -) (I_t^{\text{UNE}}(k, -))$, and $S_t^{\text{PAY}}(k, +) (I_t^{\text{PAY}}(k, +))$ are “good” news (dummy variables).²⁴ According to Table 2, “bad”

²³Further, Andersen et al. (2007) report that, during the 1990s, intraday stock, bond, and currency futures returns experienced heterogeneous conditional mean and volatility jumps in presence of “good” or “bad” news.

²⁴As in Section III, we interact $|S_t^e(k, \pm)|$ with $I_t^e(k, \pm)$ in the expressions for h_t^i and $\eta_t^i \eta_t^j$ in equations (4) and (5), respectively, to allow for a meaningful comparison of the resulting estimates for $\delta_i^e(k, \pm)$ and $d_i^e(k, \pm)$ with the corresponding dummy coefficients $\delta_i^e(k)$ and $d_i^e(k)$ from equations (1) and (3), reported in Tables 3 and 4.

news so defined is generally less frequent but of greater absolute magnitude than “good” news over our sample period 1986–2002.

We estimate the model of equations (4) and (5) according to the two-stage QML procedure described in Section III and report the ensuing coefficients in Table 5 for conditional returns and volatility and in Table 6 for return covariances. Our results reveal a significant degree of asymmetry in the response of the U.S. financial markets to the arrival of macroeconomic news of positive versus negative information content. Table 5 shows that the absolute magnitude of the effects of news releases on return volatility described in Section IV is generally, although not homogeneously, greater when the news is “bad” (especially “bad” CPI and nonfarm payroll news: $S_t^{\text{CPI}}(k, +) > 0$ and $S_t^{\text{PAY}}(k, -) < 0$): Conditional stock (bond) return volatility rises more sharply the day of (before), and drops more sharply the day after (of) their release. For instance, we find that the conditional volatility of 30-year Treasury bond returns increases by 77% ($\delta_i^{\text{PAY}}(1, -) = 0.766$) the day before, and then declines by 22% ($\delta_i^{\text{PAY}}(0, -) = -0.215$) the day of the release of surprisingly “bad” nonfarm payroll numbers ($S_t^{\text{PAY}}(k, -) < 0$), versus an increase of 59% ($\delta_i^{\text{PAY}}(1, +) = 0.589$) and subsequent decline by 21% ($\delta_i^{\text{PAY}}(0, +) = -0.205$) in correspondence with unexpectedly “good” nonfarm payroll announcements ($S_t^{\text{PAY}}(k, +) > 0$). Target rate decisions by the Federal Reserve represent a noteworthy exception, for their absolute impact on the dynamics of conditional stock and bond return volatility is greater when more expansionary (or less tightening) occurs than expected ($S_t^{\text{FED}}(k, -) < 0$). Overall, our evidence suggests that negative macroeconomic news is more likely to induce greater uncertainty in the U.S. stock and bond markets.

According to Table 3, neither the occurrence nor the absolute magnitude of important macroeconomic news affects conditional mean excess holding-period returns for U.S. stocks and bonds between 1986 and 2002. In contrast, Table 5 provides evidence of significant asymmetric effects of signed news surprises on those returns. In particular, the estimates reported in Table 5 suggest that most surprisingly “good” macroeconomic announcements significantly increase conditional mean stock returns but decrease conditional mean Treasury and corporate bond returns. Yet most unexpectedly “bad” announcements have no meaningful contemporaneous impact on any r_t^i . For example, we find that estimates for the contemporaneous impact of unexpectedly expansionary rate decisions by the Federal Reserve ($S_t^{\text{FED}}(k, -) < 0$) on mean excess stock returns are negative and both statistically and economically significant (-19 bp and -25 bp, respectively, i.e., $\gamma_i^{\text{FED}}(0, -) = -0.190$ for r_t^{NYX} and $\gamma_i^{\text{FED}}(0, -) = -0.246$ for r_t^{NAQ} in Table 5). Better than expected inflation news ($S_t^{\text{CPI}}(k, -) < 0$), which generally accompanies a slowing economy and may lead to future target rate cuts, has no impact on mean excess stock returns but leads to a small decrease in conditional 5-year Treasury and Baa-rated corporate bond returns (by 4 bp and 6 bp, respectively, i.e., $\gamma_i^{\text{CPI}}(0, -) = 0.044$ for r_t^{5Y} and $\gamma_i^{\text{CPI}}(0, -) = 0.058$ for r_t^{Baa} in Table 5). Accordingly, worse than expected nonfarm payroll numbers ($S_t^{\text{PAY}}(k, -) < 0$) also lead to lower mean excess government bond returns. However, “good” employment news ($S_t^{\text{UNE}}(k, -) < 0$ and $S_t^{\text{PAY}}(k, +) > 0$), which usually accompanies a growing economy and may lead to future target rate increases, significantly decreases only mean excess NASDAQ and NYSE-AMEX returns, respectively

(by 17 bp and 19 bp, i.e., $\gamma_i^{\text{UNE}}(0, -) = -0.172$ for r_t^{NAQ} and $\gamma_i^{\text{PAY}}(0, +) = -0.194$ for r_t^{NYX} in Table 5).

Finally, the differential information content of macroeconomic news has little or no impact on the (negative) direction of the resulting comovement shocks across asset classes (see Section IV). However, the estimates of the DCC IMA model of equation (5) in Table 6 reveal a significant heterogeneity in the intensity of the impact of the release of “good” or “bad” news on intra- and across-asset class covariances, especially in proximity to the release of CPI data. For example,

TABLE 5
GARCH Model for Excess Returns: Asymmetric Impact

Table 5 reports quasi-maximum likelihood (QML) estimates and robust *t*-statistics (Bollerslev and Wooldridge (1992), in parentheses), over a sample from 01/03/1986 to 02/14/2002 (4,069 observations), for the GARCH(1,1) model:

$$\begin{aligned}
 (4) \quad r_t^i &= \mu_i + \rho_i r_{t-1}^i + [\gamma_i^g(0, +)I_t^g(0, +) + \gamma_i^b(0, -)I_t^b(0, -)] S_t^g e_t^i, \\
 \varepsilon_t^i &= \sqrt{s_t^{i\pm}} e_t^i \quad e_t^i | F_{t-1} \sim N(0, h_t^i), \quad \text{and} \\
 h_t^i &= \omega_i + \alpha_i (e_{t-1}^i)^2 + \beta_i h_{t-1}^i,
 \end{aligned}$$

where $s_t^{i\pm} = 1 + \sum_{k=-1}^{+1} \delta_i^g(k, +) |I_t^g(k, +)| S_t^g(k, +) + \sum_{k=-1}^{+1} \delta_i^b(k, -) |I_t^b(k, -)| S_t^b(k, -)$, F_{t-1} denotes the information set, r_t^i is the daily continuously compounded excess return on asset *i*, S_t^g are actual standardized news, and $I_t^g(k, \pm)$ and $S_t^b(k, \pm)$ are the surprisingly “good” and “bad” macroeconomic news event dummy variables defined in Section V. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	r_t^{NYX}	r_t^{NAQ}	r_t^{5Y}	r_t^{10Y}	r_t^{30Y}	r_t^{Aaa}	r_t^{Baa}
$\gamma_i^{\text{FED}}(0, +)$	-0.023 (-0.19)	-0.026 (-0.28)	0.004 (0.10)	-0.004 (-0.07)	0.025 (0.31)	-0.004 (-0.10)	-0.013 (-0.27)
$\gamma_i^{\text{FED}}(0, -)$	-0.190* (-1.70)	-0.246** (-2.03)	-0.009 (-0.23)	-0.017 (-0.27)	0.005 (0.05)	0.013 (0.28)	-0.003 (-0.07)
$\delta_i^{\text{FED}}(-1, +)$	-0.084 (-0.34)	0.007 (0.04)	-0.030 (-0.28)	0.032 (0.27)	0.024 (0.29)	0.017 (0.24)	-0.011 (-0.16)
$\delta_i^{\text{FED}}(0, +)$	-0.046 (-0.10)	0.037 (0.15)	0.390 (1.27)	0.244 (1.07)	0.069 (0.65)	0.101 (0.85)	0.205 (1.18)
$\delta_i^{\text{FED}}(1, +)$	0.326 (0.76)	0.018 (0.10)	-0.039 (-0.34)	-0.036 (-0.35)	0.011 (0.13)	-0.003 (-0.04)	0.083 (0.62)
$\delta_i^{\text{FED}}(-1, -)$	-0.189*** (-6.39)	-0.193*** (-8.59)	-0.062 (-0.76)	-0.076 (-0.97)	-0.122** (-1.91)	-0.073 (-0.92)	-0.056 (-0.68)
$\delta_i^{\text{FED}}(0, -)$	0.304 (1.07)	0.166 (0.59)	-0.124** (-2.15)	-0.057 (-0.60)	0.129 (0.55)	-0.092 (-1.18)	-0.072 (-0.87)
$\delta_i^{\text{FED}}(1, -)$	-0.067 (-0.46)	0.051 (0.24)	0.562*** (2.72)	0.219 (1.36)	0.058 (0.31)	0.274 (1.63)	0.280* (1.65)
$\gamma_i^{\text{CPI}}(0, +)$	0.132 (1.57)	0.094 (1.19)	-0.022 (-1.01)	-0.032 (-0.87)	-0.036 (-0.57)	-0.022 (-0.63)	-0.020 (-0.51)
$\gamma_i^{\text{CPI}}(0, -)$	0.014 (0.14)	-0.039 (-0.41)	0.044* (1.77)	0.054 (1.30)	0.101 (1.55)	0.047 (1.64)	0.058* (1.90)
$\delta_i^{\text{CPI}}(-1, +)$	0.076 (0.61)	0.188 (1.27)	0.510** (2.53)	0.402** (2.19)	0.183 (1.19)	-0.159* (-1.85)	-0.225*** (-4.12)
$\delta_i^{\text{CPI}}(0, +)$	1.039*** (3.03)	0.600** (2.26)	-0.293*** (-6.36)	-0.279*** (-5.63)	-0.251*** (-4.45)	-0.166* (-1.67)	-0.087 (-0.81)
$\delta_i^{\text{CPI}}(1, +)$	-0.277*** (-6.87)	-0.280*** (-5.57)	0.307* (1.75)	0.299* (1.93)	0.382** (2.34)	0.414** (2.48)	0.372** (2.28)
$\delta_i^{\text{CPI}}(-1, -)$	-0.315*** (-9.73)	-0.143* (-1.77)	-0.025 (-0.24)	-0.028 (-0.26)	-0.035 (-0.33)	0.074 (0.62)	0.072 (0.64)
$\delta_i^{\text{CPI}}(0, -)$	0.809*** (3.39)	0.599*** (2.65)	-0.166** (-1.99)	-0.139 (-1.47)	-0.106 (-1.05)	-0.068 (-0.63)	-0.142* (-1.66)
$\delta_i^{\text{CPI}}(1, -)$	0.049 (0.43)	-0.143* (-1.77)	0.076 (0.63)	0.121 (0.95)	0.104 (0.85)	-0.148* (-1.91)	-0.007 (-0.07)

(continued on next page)

TABLE 5 (continued)
 GARCH Model for Excess Returns: Asymmetric Impact

	r_t^{NYX}	r_t^{NAQ}	r_t^{5Y}	r_t^{10Y}	r_t^{30Y}	r_t^{Aaa}	r_t^{Baa}
$\gamma_i^{UNE}(0, +)$	-0.100 (-0.96)	-0.111 (-1.06)	-0.006 (-0.13)	-0.021 (-0.31)	-0.016 (-0.17)	0.036 (0.77)	-0.016 (-0.34)
$\gamma_i^{UNE}(0, -)$	-0.092 (-1.24)	-0.172** (-2.50)	0.007 (0.27)	0.009 (0.23)	-0.008 (-0.14)	-0.006 (-0.22)	-0.020 (-0.68)
$\delta_i^{UNE}(-1, +)$	-0.111* (-1.74)	-0.171*** (-3.86)	-0.146** (-2.17)	-0.170*** (-4.32)	-0.119** (-2.48)	-0.182*** (-5.55)	-0.082 (-0.65)
$\delta_i^{UNE}(0, +)$	0.603** (2.34)	0.498* (1.87)	-0.194*** (-6.93)	-0.148*** (-3.28)	-0.117** (-2.35)	-0.072 (-0.62)	-0.144* (-1.74)
$\delta_i^{UNE}(1, +)$	-0.012 (-0.10)	0.032 (0.19)	1.224*** (5.28)	0.961*** (4.78)	0.578*** (3.92)	0.759*** (3.20)	0.766*** (3.71)
$\delta_i^{UNE}(-1, -)$	-0.221*** (-3.22)	0.056 (0.40)	-0.063 (-0.70)	0.000 (0.00)	0.017 (0.16)	0.170 (1.42)	0.061 (0.56)
$\delta_i^{UNE}(0, -)$	0.840*** (3.24)	0.134 (0.65)	-0.348*** (-15.05)	-0.341*** (-10.21)	-0.323*** (-12.98)	-0.345*** (-16.54)	-0.324*** (-11.34)
$\delta_i^{UNE}(1, -)$	-0.226*** (-2.77)	-0.133 (-1.23)	1.095*** (4.89)	0.982*** (3.44)	0.808*** (4.21)	0.538*** (3.44)	0.614*** (3.80)
$\gamma_i^{PAY}(0, +)$	-0.194* (-1.85)	-0.112 (-1.17)	-0.042 (-1.24)	-0.032 (-0.58)	-0.040 (-0.50)	-0.008 (-0.22)	-0.025 (-0.67)
$\gamma_i^{PAY}(0, -)$	-0.038 (-0.41)	-0.029 (-0.25)	0.066** (2.37)	0.091* (1.88)	0.096 (1.33)	0.025 (0.62)	0.057 (1.57)
$\delta_i^{PAY}(-1, +)$	-0.212*** (-5.74)	-0.167*** (-2.99)	-0.175*** (-3.43)	-0.151** (-2.22)	-0.103 (-1.43)	-0.028 (-0.28)	-0.130** (-2.12)
$\delta_i^{PAY}(0, +)$	0.643*** (2.91)	0.206 (1.00)	-0.238*** (-9.05)	-0.235*** (-7.93)	-0.205*** (-5.37)	-0.217*** (-6.35)	-0.196*** (-4.81)
$\delta_i^{PAY}(1, +)$	-0.154** (-2.36)	-0.040 (-0.30)	1.101*** (4.70)	0.970*** (4.32)	0.589*** (3.75)	0.430*** (2.78)	0.448*** (3.18)
$\delta_i^{PAY}(-1, -)$	-0.225*** (-3.72)	-0.225*** (-2.95)	-0.040 (-0.38)	-0.112 (-1.38)	-0.121 (-1.49)	-0.158** (-2.36)	-0.024 (-0.22)
$\delta_i^{PAY}(0, -)$	1.143*** (3.50)	0.832** (2.46)	-0.305*** (-7.46)	-0.210*** (-3.35)	-0.215*** (-3.59)	-0.063 (-0.62)	-0.215*** (-3.27)
$\delta_i^{PAY}(1, -)$	-0.218*** (-3.23)	-0.128 (-1.19)	1.175*** (4.68)	0.857*** (4.18)	0.766*** (4.08)	0.661*** (3.45)	0.644*** (3.51)

the comovement between government bond holding-period returns declines by 66% on the day following better-than-expected inflation news, $S_t^{CPI}(k, -) < 0$ ($\overline{d_{G,G}^{CPI}}(-1, -) = -0.656$), in line with estimates for other news releases in our sample, but by much more ($\overline{d_{G,G}^{CPI}}(-1, +) = -3.375$) immediately after worse-than-expected inflation news, $S_t^{CPI}(k, +) > 0$. Similarly, although lower comovement among NYSE-AMEX and NASDAQ stock returns precedes surprisingly expansionary federal funds rate decisions as well ($S_t^{FED}(k, -) < 0$), the estimated decline in intra-asset class return covariances is again more pronounced and persistent in proximity to a tighter-than-expected monetary policy stance by the Federal Reserve ($S_t^{FED}(k, +) > 0$). Comovement within and across all asset classes also displays more pronounced downward dynamics when news about worsening employment conditions ($S_t^{UNE}(k, +) > 0$ and $S_t^{PAY}(k, -) < 0$) is released. This suggests that surprisingly “bad” news arrivals are more likely to induce a wave of rebalancing activity among the asset classes in our sample. Such asymmetry is consistent with recent studies (e.g., Kyle and Xiong (2001)) conjecturing that the multiasset trading activity of informed and uninformed market participants

is more likely to respond to financial constraints in those circumstances. Accordingly, target rate increases and “bad” nonfarm payroll news induce lower comovement among high-rated and low-rated corporate bonds than rate cuts and “good” employment numbers (i.e., are more likely to induce investors to shift their portfolios away from low-quality issuers).

TABLE 6
DCC IMA Model: Asymmetric Covariance Shifts

Table 6 reports intra- and across-asset class averages of quasi-maximum likelihood (QML) estimates and their robust t-statistics (Bollerslev and Wooldridge (1992), in parentheses) for the DCC IMA model:

$$(5) \quad \eta_i^j \eta_i^k = s_i^{j\pm*} \eta_{i-1}^j \eta_{i-1}^k + (q_i^{jj} - \eta_i^j \eta_i^j) - \lambda^e s_i^{\pm*} (q_{i-1}^{jj} - \eta_{i-1}^j \eta_{i-1}^j),$$

where $s_i^{j\pm*} = 1 + \sum_{k=-1}^{+1} d_{i,j}^e(k, +) I_i^e(k, +) |S_i^e(k, +)| + \sum_{k=-1}^{+1} d_{i,j}^e(k, -) I_i^e(k, -) |S_i^e(k, -)|$, $\eta_i^j = \epsilon_i^j / \sqrt{s_i^j h_i^j}$ is the daily standardized residual from the GARCH(1, 1) model of equation (1) for r_i^j , the excess return on asset i ; S_i^e are actual standardized news; and $I_i^e(k, \pm)$ and $S_i^e(k, \pm)$ are the surprisingly “good” and “bad” macroeconomic news event dummy variables defined in Section V. Equation (5) is estimated between 01/03/1986 and 02/14/2002 (i.e., over 4,069 observations). The STOCK category is made of $i = NYX, NAQ$; the GOVT category is made of $i = 5Y, 10Y, 30Y$; and the CORP category is made of $i = Aaa, Baa$.

	$\overline{d_{i,j}^{FED}(-1, +)}$			$\overline{d_{i,j}^{FED}(0, +)}$			$\overline{d_{i,j}^{FED}(1, +)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.999 (-17.75)	-0.772 (-9.91)	-0.718 (-10.22)	-0.194 (-6.24)	-0.863 (-12.31)	-0.881 (-11.33)	-1.270 (-50.09)	-1.251 (-18.71)	-0.476 (-9.71)
GOVT		-0.150 (-8.82)	-0.960 (-10.43)		-0.750 (-8.67)	-0.601 (-5.92)		-0.303 (-9.68)	-0.314 (-10.13)
CORP			-0.154 (-13.73)			0.072 (1.14)			-0.251 (-13.86)
	$\overline{d_{i,j}^{FED}(-1, -)}$			$\overline{d_{i,j}^{FED}(0, -)}$			$\overline{d_{i,j}^{FED}(1, -)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.902 (-45.29)	-0.989 (-23.54)	-0.936 (-19.84)	0.170 (1.19)	-0.258 (-6.58)	-0.652 (-11.36)	-0.265 (-54.25)	-1.105 (-23.39)	-0.938 (-21.29)
GOVT		-0.553 (-7.89)	-0.733 (-11.78)		-0.261 (-6.60)	-0.743 (-45.61)		-0.202 (-8.08)	0.132 (-0.85)
CORP			-0.703 (-11.08)			-0.803 (-47.49)			0.712 (3.47)
	$\overline{d_{i,j}^{CPI}(-1, +)}$			$\overline{d_{i,j}^{CPI}(0, +)}$			$\overline{d_{i,j}^{CPI}(1, +)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	0.606 (5.28)	0.441 (3.07)	-1.035 (-11.40)	-0.105 (-1.38)	-0.629 (-7.49)	-0.751 (-9.35)	-0.277 (-5.07)	-0.649 (-9.23)	-0.758 (-11.32)
GOVT		-3.375 (-12.73)	-1.587 (-12.23)		-0.392 (-8.42)	-0.445 (-8.71)		-1.028 (-11.03)	-0.595 (-14.43)
CORP			-2.751 (-14.01)			-0.263 (-5.35)			-0.081 (-1.30)
	$\overline{d_{i,j}^{CPI}(-1, -)}$			$\overline{d_{i,j}^{CPI}(0, -)}$			$\overline{d_{i,j}^{CPI}(1, -)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.591 (-37.93)	-0.665 (-11.73)	-1.058 (-8.62)	0.805 (6.21)	-1.077 (-16.75)	-0.920 (-9.50)	-1.088 (-16.20)	-0.918 (-10.27)	-0.984 (-11.82)
GOVT		-0.656 (-9.89)	-0.928 (-11.48)		-0.327 (-6.22)	-0.363 (-5.23)		-0.209 (-3.39)	-0.815 (-12.62)
CORP			-1.156 (-19.08)			-0.257 (-4.22)			-1.017 (-16.72)

(continued on next page)

TABLE 6 (continued)
DCC IMA Model: Asymmetric Covariance Shifts

	$\overline{d_{i,j}^{\text{UNE}}(-1,+)}$			$\overline{d_{i,j}^{\text{UNE}}(0,+)}$			$\overline{d_{i,j}^{\text{UNE}}(1,+)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.172 (-3.21)	-0.639 (-12.52)	-0.828 (-14.46)	-0.118 (-1.89)	-1.112 (-14.22)	-1.257 (-19.23)	-1.077 (-21.75)	-1.124 (-22.06)	-0.957 (-14.75)
GOVT		-0.850 (-18.64)	-1.298 (-21.10)		-1.672 (-21.29)	-1.394 (-19.66)		0.020 (0.23)	-0.438 (-10.04)
CORP			-1.165 (-33.12)			-1.757 (-32.51)			-0.206 (-4.44)
	$\overline{d_{i,j}^{\text{UNE}}(-1,-)}$			$\overline{d_{i,j}^{\text{UNE}}(0,-)}$			$\overline{d_{i,j}^{\text{UNE}}(1,-)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.252 (-4.32)	-0.737 (-12.15)	-0.946 (-15.57)	-0.089 (-1.39)	-0.984 (-18.71)	-1.069 (-22.51)	-1.356 (-30.23)	-1.097 (-18.82)	-1.107 (-19.33)
GOVT		-0.680 (-13.47)	-1.106 (-18.27)		-0.752 (-28.11)	-0.728 (-28.22)		0.071 (0.88)	0.043 (0.57)
CORP			-0.654 (-16.45)			-0.641 (-33.93)			-2.703 (-23.58)
	$\overline{d_{i,j}^{\text{PAY}}(-1,+)}$			$\overline{d_{i,j}^{\text{PAY}}(0,+)}$			$\overline{d_{i,j}^{\text{PAY}}(1,+)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-0.178 (-7.02)	-0.888 (-13.67)	-0.857 (-11.88)	-1.107 (-30.25)	-0.857 (-8.97)	-1.048 (-12.03)	-0.447 (-7.98)	-0.821 (-14.67)	-0.664 (-13.07)
GOVT		-0.922 (-23.39)	-0.708 (-15.87)		-1.161 (-19.37)	-1.222 (-24.75)		0.233 (1.94)	0.279 (2.19)
CORP			-0.175 (-3.26)			-1.271 (-28.64)			0.356 (2.69)
	$\overline{d_{i,j}^{\text{PAY}}(-1,-)}$			$\overline{d_{i,j}^{\text{PAY}}(0,-)}$			$\overline{d_{i,j}^{\text{PAY}}(1,-)}$		
	STOCK	GOVT	CORP	STOCK	GOVT	CORP	STOCK	GOVT	CORP
STOCK	-1.013 (-29.96)	-0.736 (-12.37)	-0.583 (-13.26)	-0.291 (-5.63)	-0.924 (-11.46)	-0.986 (-12.32)	-0.167 (-2.63)	-1.137 (-13.88)	-0.986 (-12.45)
GOVT		-0.476 (-17.01)	-0.202 (-13.82)		-1.267 (-22.70)	-1.230 (-23.42)		-0.156 (-3.06)	-0.238 (-5.35)
CORP			0.330 (1.87)			-1.246 (-23.26)			-0.205 (-3.92)

VI. Conclusions

The analysis of the extent to which prices in financial markets incorporate fundamental information is central to the theoretical and empirical finance literature. The traditional notion of market efficiency requires that *new* information about asset payoffs should be quickly and fully reflected in asset prices and drive their dynamics. Prior research has examined the links between financial and real variables by studying the effects of the disclosure of macroeconomic information (often without first identifying its surprise content) on stock and bond markets (often separately).

Our paper contributes to this debate by providing a comprehensive analysis of the impact of the *unexpected* component of important U.S. macroeconomic

news releases on the process of price formation in the three most important U.S. financial markets: equity, government bonds, and corporate bonds. For that purpose, we develop an extension of the multivariate GARCH-DCC IMA model of Engle (2002) that allows us to simultaneously (yet parsimoniously) identify the effects of news arrivals on conditional mean returns, return volatility, and return covariance. We estimate different versions of this model over our sample of announcements of target rate decisions by the Federal Reserve, CPI, unemployment, and nonfarm payroll data from the BLS between 1986 and 2002.

We find that the arrival of surprise macroeconomic news has a statistically and economically significant impact on the U.S. financial markets, but also that this impact varies greatly across asset classes. Conditional stock return volatility decreases on the trading day before, increases on the day when the announcements are made, and subsequently decreases. Conditional bond return volatility instead increases before the news is released and declines afterward. This effect is stronger for shorter maturity bond portfolios (typically, the most liquid and sensitive to the mean-reverting nature of the short rate drift) or less likely to default. The estimated shifts in volatility appear to be persistent in the short run (i.e., do not offset each other completely over a 3-day event window around the announcements). These effects are also asymmetric: Their absolute magnitude is generally greater when the macroeconomic information released represents “bad” news. Conditional mean excess holding-period returns for stock and bonds are instead mostly sensitive to the release of unexpectedly “good” news. Finally, our estimates paint a complex picture of the interaction between asset returns in proximity to macroeconomic news releases. Yet they offer little or no support for the commonly held notion that the arrival of this news is accompanied by greater comovement among asset returns. Indeed, return comovement within and across stock and bond markets most often decreases in correspondence with those announcements, especially in proximity to “bad” news.

Ultimately, this study reports a wide array of novel empirical evidence, stemming from a robust yet manageable methodology, on the effects of the release of macroeconomic news on the moments of returns in U.S. equity, government bond, and corporate bond markets. We suggest that some of this evidence is consistent with the existing theoretical arguments in the literature (e.g., those emphasizing the role of investors’ trading activity). Nonetheless, we hope that our work may stimulate further research on a unifying theory that explains all of these fascinating stylized facts.

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