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MEASURING THE EFFECT OF THE ZERO LOWER BOUND ON MEDIUM- AND LONGER-TERM INTEREST RATES

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ABSTRACT

The federal funds rate has been at the zero lower bound for over four years, since December 2008. According to standard macroeconomic models, this should have greatly reduced the effectiveness of monetary policy and increased the efficacy of fiscal policy. However, these models also imply that asset prices and private-sector decisions depend on the entire path of expected future short-term interest rates, not just the current level of the overnight rate. Thus, interest rates with a year or more to maturity are arguably more relevant for asset prices and the economy, and it is unclear to what extent those yields have been affected by the zero lower bound. In this paper, we measure the effects of the zero lower bound on interest rates of any maturity by comparing the sensitivity of those interest rates to macroeconomic news when short-term interest rates were very low to that during normal times. We find that yields on Treasury securities with a year or more to maturity were surprisingly responsive to news throughout 2008–10, suggesting that monetary and fiscal policy were likely to have been about as effective as usual during this period. Only beginning in late 2011 does the sensitivity of these yields to news fall closer to zero. We offer two explanations for our findings: First, until late 2011, market participants expected the funds rate to lift off from zero within about four quarters, minimizing the effects of the zero bound on medium- and longer-term yields. Second, the Fed's unconventional policy actions seem to have helped offset the effects of the zero bound on medium- and longer-term rates.

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1 Introduction

The federal funds rate—the Federal Reserve's traditional monetary policy instrument—has been at a lower bound of essentially zero for over four years, since December 2008. According to many macroeconomic models, the binding constraint of the zero lower bound on nominal interest rates should have greatly reduced the effectiveness of monetary policy and increased the efficacy of fiscal policy during this period (e.g., Christiano, Eichenbaum, and Rebelo, 2011; Woodford, 2011).[1] However, standard macroeconomic theory, such as Clarida, Galí, and Gertler (1999) and Woodford (2003), as well as the papers cited above, imply that asset prices and the economy are affected by the entire *path* of expected future short-term interest rates, not just the current level of the overnight federal funds rate. Thus, interest rates with a year or more to maturity are arguably more relevant for asset prices and the economy, and it is not clear whether the zero bound has substantially constrained the Fed's ability to affect these longer-term yields.

Theoretically, if a central bank has the ability to commit to future values of its policy rate, it can work around the zero bound constraint by promising monetary accommodation in the future once the zero bound ceases to bind (Reifschneider and Williams, 2000; Eggertsson and Woodford, 2003).[2] Empirically, Gürkaynak, Sack, and Swanson (2005a) show that the Federal Reserve's monetary policy announcements affect asset prices primarily through their effects on financial market expectations of *future* monetary policy, rather than changes in the current federal funds rate. Thus, there are both theoretical and empirical reasons to think that monetary policy can remain effective even when the overnight interest rate is zero. In fact, 1- and 2-year Treasury yields remained substantially above zero throughout much of 2008–10 (Figure 1), suggesting that monetary policy still had room to affect the economy despite the constraint on the funds rate. And indeed, the Federal Reserve generated a drop in medium- and longer-term Treasury yields of as much as 20 basis points (bp) on several occasions between 2008 and 2012 by managing monetary policy expectations or purchasing assets (Gagnon et al., 2011; Wright, 2012).[3]

[1] See also Eggertsson (2009), Erceg and Lindé (2010), Eggertsson and Krugman (2012), and DeLong and Summers (2012). Intuitively, the macroeconomic effects of fiscal policy are larger when the zero bound is binding, because then interest rates do not rise in response to higher output, so private investment and consumption are not "crowded out".

[2] The central bank does not need a perfect commitment technology for this to be true; a partial commitment technology, as in Schaumburg and Tambalotti (2007), is sufficient.

[3] Note that, in normal times, it would take a surprise change in the federal funds rate of more than 100 bp to generate a fall of 20 bp in medium- and longer-term Treasury yields (Gürkaynak et al., 2005).

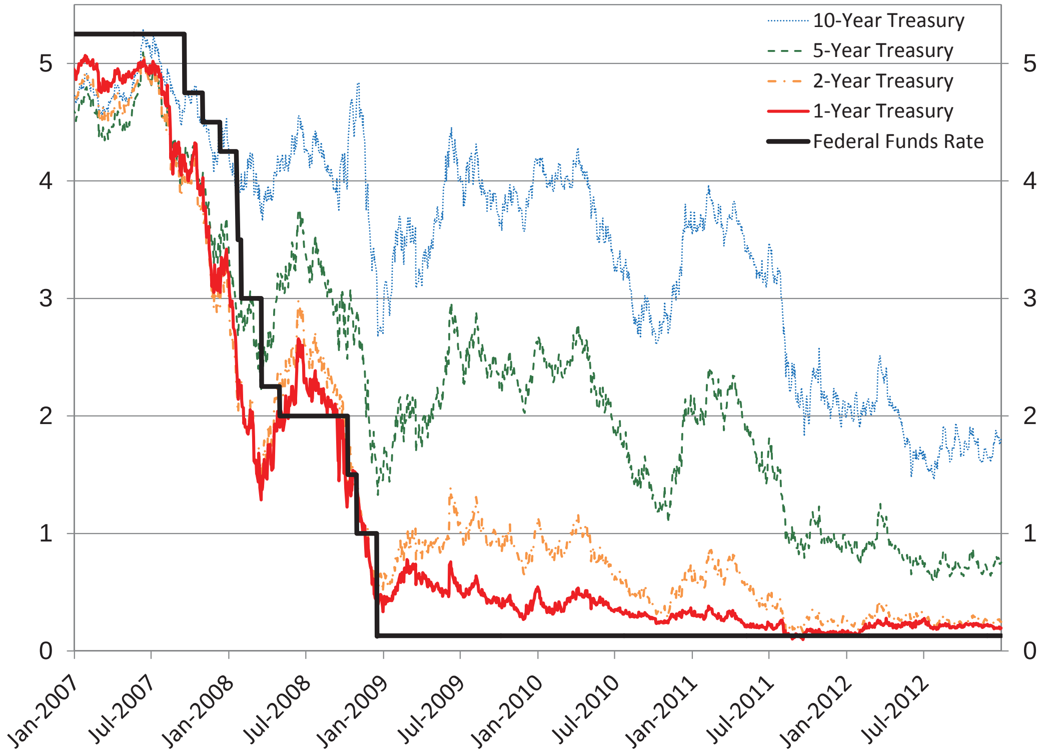

Figure 1. Federal funds rate target and 1-, 2-, 5-, and 10-year zero-coupon Treasury yields from January 2007 through December 2012. Data are from the Federal Reserve Board and the Gürkaynak, Sack, and Wright (2007) online dataset.

In this paper, we propose a novel method of measuring the extent to which interest rates of any maturity—and hence monetary policy, more broadly defined—are constrained by the zero lower bound. In particular, we estimate the time-varying sensitivity of yields to macroeconomic announcements using high-frequency data and compare that sensitivity to a benchmark period in which the zero bound was not a concern (taken to be 1990–2000). In periods when a given yield is about as sensitive to news as in the benchmark sample, we say that yield is unconstrained. In periods, if any, where a yield responds very little or not at all to news, we say that yield is largely or completely constrained. Intermediate cases are measured by the degree of the yield's sensitivity to news relative to the benchmark sample, and the severity and statistical significance of the constraint can be assessed using standard econometric techniques.

We emphasize that the level of a yield alone is not a useful measure of whether that yield is constrained by the zero lower bound, for at least three reasons. First, there is no way to quantify the severity of the zero bound constraint or its statistical significance using the level of the yield alone. For example, if the one-year Treasury yield is 50 bp, there is no clear way to determine whether that yield is severely constrained, mildly constrained, or even unconstrained. By contrast, the method we propose in this paper provides an econometrically precise answer to this question.

Second, the lower bound on nominal interest rates may be above zero for institutional reasons (Bernanke and Reinhart, 2004), and this "effective" lower bound may vary across countries or over time. For example, the Federal Reserve has kept the federal funds rate at a floor of about 0 since December 2008, but the Bank of England has maintained a floor of 50 bp for its policy rate over the same period while conducting unconventional monetary policy on a similar scale to the Federal Reserve. Evidently, the effective floor on the monetary policy rate in the U.K. is about 50 bp rather than zero. As a result, a 50 or even 100 bp gilt yield in the U.K. might be substantially constrained by the effective U.K. lower bound of 50 bp, while a similar 50–100 bp yield in the U.S. might be only mildly constrained or unconstrained. The approach in this paper relies on the sensitivity of interest rates to news rather than the level of rates, and thus can accommodate effective lower bounds that may be greater than zero or change over time.

Third, the sensitivity of yields to news is more relevant than the level of yields for the fiscal multiplier. As emphasized by Christiano et al. (2011), Woodford (2011), and others, what is crucial for the fiscal multiplier is whether interest rates *respond* to a government spending shock; the level of yields by itself is largely irrelevant. Although the zero lower bound motivates the analysis in those studies, their results are all derived in a "constant interest rate" environment in which nominal yields can be regarded as fixed at any absolute level.

To our knowledge, this paper represents the first attempt to directly measure the effects of the zero lower bound constraint on the behavior of intermediate- and longer-maturity yields, and thus the extent to which the zero bound has hindered the effectiveness of monetary policy and amplified the effectiveness of fiscal policy. Previous studies of the effects of the zero bound on longer-term yields have focused on how to *model* the yield curve when rates are near zero (e.g., Gorovoi and Linetsky, 2004; Kim and Singleton, 2012). In contrast, we directly *measure* rather than model the zero bound's effects, so our estimates are model-free. This is an important advantage of our approach, since there are a variety of nonnegative dynamic term structure models (NDTSMs) that lead to a variety of predictions regarding the behavior of longer-term yields near the zero bound

(Kim and Singleton, 2012).

To preview our results, we find that Treasury yields with one or two years to maturity were surprisingly responsive to news throughout 2008–10, despite the federal funds rate being essentially zero over this period. Contrary to conventional wisdom, this suggests that the efficacy of monetary and fiscal policy were likely close to normal in 2008–10. Only beginning in late 2011 do we see the sensitivity of the 2-year Treasury yield to news fall significantly below normal. We also show that 5- and 10-year Treasury yields were essentially unconstrained by the zero bound throughout our sample, while Treasuries with six months or less to maturity have been severely constrained since the spring of 2009. Importantly, our method provides a quantitative measure of the *degree* to which the zero bound affects each yield, as well as a statistical test for the periods during which it was constrained.

We provide two explanations for our findings. First, up until August 2011, market participants expected the zero bound to constrain policy for only a few quarters, minimizing the zero bound's effects on medium- and longer-term yields. Second, the Federal Reserve's large-scale purchases of long-term bonds and management of monetary policy expectations may have helped offset the effects of the zero bound on medium- and longer-term interest rates.

Our estimates are consistent with, and can help inform, the NDTSM literature mentioned above. Although most NDTSMs to date have been estimated using Japanese data, Bauer and Rudebusch (2013) apply one such model to the U.S.[4] They estimate that, prior to August 2011, financial markets expected the zero bound to constrain U.S. short-term rates for only a few quarters; after that date, they estimate the funds rate was expected to lift off much later. Our findings corroborate and support these features of their model, which allow it to fit the relatively high level and volatility of intermediate-maturity yields prior to August 2011 without large term premia.

The remainder of our paper proceeds as follows. Section 2 lays out a simple New Keynesian model that helps to motivate our empirical analysis and provide intuition for our results. Section 3 describes the details of our high-frequency empirical methodology. Our main results are reported

[4]We do not consider Japanese data in this paper because we do not have access to the surprise component of Japanese macroeconomic announcements or daily bond yield data that go back far enough to estimate the sensitivity of Japanese yields to news when the zero bound is not binding. (This is partly due to data availability and partly because the zero bound has been a potential constraint in Japan for so long, back to the 1990s.) Swanson and Williams (2013b) apply the methods of the present paper to yields and exchange rates in the U.K. and Germany.

in Section 4. Section 5 considers the broader implications of our results for monetary and fiscal policy, and various extensions and robustness checks. Section 6 concludes.

2 An Illustrative Model

A simple theoretical model helps to motivate our empirical analysis and provide intuition for our results. The purpose of this section is to illustrate qualitatively how the zero bound might be expected to affect the sensitivity of bond yields to news, so the model is deliberately simplistic and not intended to capture all the details of the effects we estimate below.

We follow the standard three-equation New Keynesian framework of Clarida, Galí, and Gertler (1999) and Woodford (2003), among others. The output gap, $\tilde{y}_t$, satisfies a forward-looking IS curve,

$$\tilde{y}_t = -\alpha(i_t - E_t\pi_{t+1} - r_t^*) + E_t\tilde{y}_{t+1}, \tag{1}$$

where i_t denotes the one-period nominal interest rate at time t, π_t the inflation rate, r_t^* a "natural" rate of interest, E_t the mathematical expectation conditional on information at time t, and α is a parameter. We model shocks to output as coming through shocks to r_t^*.

Equation (1) can be solved forward, assuming $\lim_{j\to\infty} E_t\tilde{y}_{t+j} = 0$, to get

$$\tilde{y}_t \;=\; -\alpha\, E_t \sum_{j=0}^{\infty} \{i_{t+j} - \pi_{t+j+1} - r_{t+j}^*\}. \tag{2}$$

Equation (2) makes it clear that the output gap at time t depends on the entire expected *future path* of short-term interest rates (as well as inflation and the natural rate of interest), rather than just the current short-term interest rate. Thus, even if the current one-period interest rate is constrained by the zero lower bound, the effect of that constraint on the economy may be negligible if expectations of future short-term interest rates are unconstrained.

Inflation, π_t, satisfies a standard New Keynesian Phillips curve,

$$\pi_t \;=\; \gamma\tilde{y}_t + \beta E_t\pi_{t+1} + \mu_t, \tag{3}$$

where μ_t can be interpreted as a markup shock, and γ and β are parameters.

The one-period nominal interest rate is set by the central bank according to a Taylor-type (1993) monetary policy rule, subject to the constraint that i_t must be nonnegative:

$$i_t \;=\; \max\{\,0,\; \pi_t + r_t^* + 0.5(\pi_t - \bar{\pi}) + 0.5\tilde{y}_t\}, \tag{4}$$

where $\bar{\pi}$ denotes the central bank's inflation target. Note that monetary policy is assumed to respond to the current level of the natural interest rate. This implies that, absent the zero lower bound, monetary policy perfectly offsets the effects of natural interest rate shocks on the output gap and inflation. Of course, the presence of the zero lower bound implies that, in certain circumstances, monetary policy will be unable to offset such shocks.

Consistent with the log-linearized structure of the economy implicit in equations (1)–(3), long-term bond yields in the model are determined by the expectations hypothesis.[5] Thus, the M-period yield to maturity, i_t^M, on a zero-coupon nominal bond is given by

$$i_t^M \; = \; E_t \sum_{j=0}^{M-1} i_{t+j} \; + \; \phi^M, \tag{5}$$

where ϕ^M denotes a term premium that may vary with maturity M but is constant over time.

We set $\alpha = 1.59$, $\beta = .99$, and $\gamma = .096$, based on Woodford (2003), and we set $\bar{\pi}$ to 2 percent. The shocks r_t^* and μ_t are assumed to follow AR(1) processes,

$$r_t^* \;\; = \;\; \rho r_{t-1}^* + e_t, \tag{6}$$

$$\mu_t \;\; = \;\; \delta \mu_{t-1} + v_t, \tag{7}$$

with persistences $\rho = 0.85$ and $\delta = 0.5$. We calibrate the magnitude of the shocks to r_t^* and μ_t so that they each generate a 5 bp response of the one-period interest rate i_t on impact in the absence of the zero lower bound. This calibration is consistent with our empirical finding, below, that any given macroeconomic announcement typically moves short-term rates by only a few basis points.

We solve for the impulse response functions of the model under two scenarios: First, a scenario in which the initial value of r_t^* is substantially greater than zero, so that the zero bound is not a binding constraint on the short-term interest rate; and second, a scenario in which the initial value of r_t^* is -4, which is sufficient for the zero bound to constrain the short-term nominal rate i_t for several periods. In the latter case, we solve for the impulse response functions of the model using a nonlinear perfect foresight algorithm, as in Reifschneider and Williams (2000), in which the private sector assumes the realized values of all future innovations will be zero. In each scenario, impulse

[5] Intuition for the behavior of long-term bond yields is most easily obtained by focusing on the expectations component of those yields rather than the term premium, since risk premia on long-term bonds are not well understood. Moreover, our empirical results can be understood without resorting to a story about risk premia, which also motivates this modeling choice. We discuss the possible effects of time-varying term premia in Section 5.3, below.

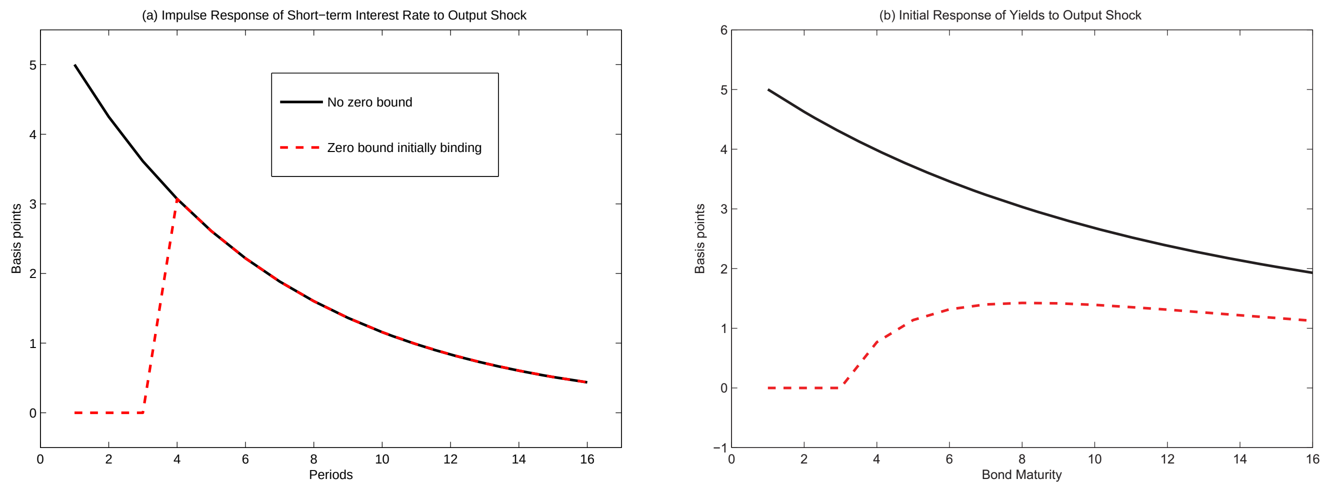

Figure 2. (a) Impulse response function of one-period interest rate i_t and (b) instantaneous response in period 1 of the yield curve to an output shock. Shock is normalized to produce a 5 bp effect on the one-period nominal interest rate on impact. See text for details.

responses are computed as the difference between the path of the economy after the shock and the baseline path of the economy absent the shock.

Figure 2 reports the response of the short-term interest rate and yield curve to a shock to output, achieved through an innovation to r_t^*. (Responses to an inflation shock are reported in Swanson and Williams, 2013a, and are very similar to those in Figure 2.) In the first panel, the solid black line depicts the impulse response function of the one-period nominal interest rate i_t when the zero lower bound is not binding—i.e., the standard impulse response function to an output shock in a New Keynesian model. The dashed red line depicts the impulse response function for i_t to the same shock when the zero bound is binding—that is, when the initial value of r_t^* is set equal to -4 percent. Note that the dashed red impulse response is computed relative to a baseline in which r_t^* begins at -4 percent but is returning toward 0, so that the zero bound ceases to bind i_t in the fourth period. From that point onward, the behavior of i_t is identical to what would occur absent the zero bound—that is, the red and black lines in the left panel of Figure 2 are identical. This is because output and inflation in this particular model are purely forward-looking. In more general models with output or inflation inertia, the zero bound would have more persistent effects on output and inflation, which would lead to a more persistent difference in the path of i_t.

The second panel of Figure 2 plots the instantaneous response of the yield curve in the period when the shock hits (i.e., period 1). Thus, the second panel of Figure 2 is not an impulse response function, but rather the instantaneous response of the entire yield curve at a single point in time.

Figure 2 shows that, when short-term interest rates are constrained by the zero lower bound, yields of *all* maturities respond less to economic announcements than if the zero bound were not present; moreover, the reduction in the sensitivity of yields to news is greatest at short maturities and is smaller for longer-term yields. This result is intuitive: for the shortest maturities, there is a total lack of responsiveness to an output or inflation shock when the zero bound is binding. According to equation (5), longer-term yields are an average of expected future short-term interest rates over the life of the bond. Since the expected path of future short-term interest rates in the model converges back toward steady state over time, short-term interest rates are expected to be unconstrained by the zero lower bound at some point in the future. As a result, longer-term bond yields are not as attenuated as shorter-term yields in their sensitivity to news.

The second point to take away from the model is that the responses of yields to shocks are essentially *symmetric*—that is, the sensitivity of yields to positive and negative announcements falls by about the same amount when the zero bound is strongly binding on short-term rates. Figure 2 plots the response of the model to small positive shocks, but the results for small negative shocks are virtually identical in absolute value.[6] This implication of the model can be surprising at first, since the zero bound is a one-sided constraint. Nevertheless, the intuition is clear: when the zero bound is strongly binding—that is, policymakers would like to set i_t substantially below zero for several periods—then short-term yields are completely unresponsive to *both* positive and negative shocks, as long as those positive shocks are not large enough to bring short-term rates above the zero bound. Longer-term yields are also about equally damped in response to positive and negative shocks because: (a) longer-term yields are an average of current and expected future short-term rates, (b) current short-term rates do not respond to either positive or negative shocks, and (c) expected future short-term rates respond symmetrically to positive and negative shocks in periods in which the zero bound is not binding. There are very few periods in which expected future short-term rates are unconstrained by the zero bound for the positive shock but still constrained for the negative shock, and even in those periods the interest rate differential between the two cases

[6] This symmetry is perfect if the number of periods that policy is constrained by the zero bound does not change, which is the case for small shocks. Even for shocks that are very large by empirical standards, Swanson and Williams (2013a) show that the response of yields in the model is essentially symmetric, for the reasons discussed below.

is typically very small. These small differences are negligible compared to the response of the yield curve as a whole, so the result is almost perfectly symmetric. We also test this restriction in our empirical work below, and find that it is not rejected by the data.

The final point to take away from the model is that the zero bound attenuates the sensitivity of yields to news by similar amounts for different types of shocks, as long as the effects of those shocks on short-term interest rates have similar persistence. Intuitively, the degree of attenuation across maturities is determined primarily by the length of time the zero bound is expected to bind, and not by the type of shock. If two different shocks in the model have similarly persistent effects on short-term interest rates, then the attenuation across maturities is also approximately the same. In our empirical work below, we assume that the zero bound attenuates the sensitivity of the yield curve to news by the same amount for all shocks. In the model above, this would only be exactly true if the shocks had identically persistent effects on the short-term interest rate, but we view this assumption as a reasonable empirical approximation. We also test this assumption below, and find that it is not rejected by the data.

3 Empirical Framework

We now seek to estimate the extent to which Treasury securities of different maturities have been more or less sensitive to macroeconomic announcements over time. We do this in three steps: First, we identify the surprise component of major U.S. macroeconomic announcements. Second, we estimate the average sensitivity of Treasury securities of each maturity to those announcements over a benchmark sample, 1990–2000, during which the zero bound was not a constraint on yields. Third, we compute the sensitivity of each Treasury yield in subsequent periods and compare it to the benchmark sample to determine when and to what extent each yield was affected by the presence of the zero lower bound. Periods in which the zero bound was a significant constraint on a given Treasury yield should appear in this analysis as periods of unusually low sensitivity of that security to macroeconomic news. We describe the details of each of these three basic steps in turn.

3.1 The Surprise Component of Macroeconomic Announcements

Financial markets are forward-looking, so the expected component of macroeconomic data releases should have essentially no effect on interest rates (Kuttner, 2001, tests and confirms this hypothesis for the case of monetary policy announcements). To measure the effects of these announcements on interest rates, then, we first compute the unexpected, or surprise, component of each release.

As in Gürkaynak et al. (2005b), we compute the surprise component of each announcement as the realized value of the macroeconomic data release on the day of the announcement less the financial markets' expectation for that realized value. We obtain data on financial market expectations of major macroeconomic data releases from two sources: Money Market Services (MMS) and Bloomberg Financial Services. Both MMS and Bloomberg conduct surveys of financial market institutions and professional forecasters regarding their expectations for upcoming major data releases, and we use the median survey response as our measure of the financial market expectation. An important feature of these surveys is that they are conducted just a few days prior to each announcement—historically, the MMS survey was conducted the Friday before each data release, and the Bloomberg survey can be updated by participants until the night before the release—so these forecasts should reflect essentially all relevant information up to a few days before the release. Anderson et al. (2003) and other authors have verified that these data pass standard tests of forecast rationality and provide a reasonable measure of ex ante expectations of the data release, which we have verified over our sample as well.

Data from MMS for some macroeconomic series are available back to the mid-1980s, but are only consistently available for a wider variety of series starting around mid-1989, so we begin our sample on January 1, 1990.[7] Bloomberg survey data begin in the mid-1990s but are available to us more recently. When the two survey series overlap, they agree very closely, since they are surveying essentially the same set of financial institutions and professional forecasters. Additional details regarding these data are provided in Gürkaynak, Sack, and Swanson (2005b), Gürkaynak, Levin, and Swanson (2010), and in Section 5.5, below.

[7]MMS survey data for two of our series, consumer confidence and initial claims, begin in 1991. All of our results are essentially unchanged if we begin our analysis in 1991 rather than 1990.

3.2 The Sensitivity of Treasury Yields to Macroeconomic Announcements

In normal times, when Treasury yields are far away from the zero lower bound, those yields typically respond to macroeconomic news. To measure this responsiveness, Gürkaynak et al. (2005b) estimate daily-frequency regressions of the form

$$\Delta y_t = \alpha + \beta X_t + \varepsilon_t \,, \tag{8}$$

where t indexes days, Δy_t denotes the one-day change in the Treasury yield over the day, X_t is a vector of surprise components of macroeconomic data releases that took place that day, and ε_t is a residual representing the influence of other news and other factors on the Treasury yield that day. Note that most macroeconomic data series, such as nonfarm payrolls or the consumer price index, have data releases only once per month, so on days for which there is no news about a particular macroeconomic series, we set the corresponding element of X_t equal to zero.[8]

Table 1 reports estimates of regression (8) for the 3-month, 2-year, and 10-year Treasury yields from January 1990 through December 2000, a period in which we assume the zero lower bound did not constrain these yields. We exclude days on which no major macroeconomic data releases occurred, although the results are very similar whether or not these non-announcement days are included. To facilitate interpretation of the coefficients in Table 1, each macroeconomic data release surprise is normalized by its historical standard deviation, so that each coefficient in the table is in units of basis points per standard-deviation surprise in the announcement.[9]

The first column of Table 1 reports results for the 3-month Treasury yield. Positive surprises in output or inflation cause the 3-month Treasury yield to rise, on average, consistent with a Taylor-type reaction function for monetary policy, while positive surprises in the unemployment rate or initial jobless claims (which are countercyclical economic indicators) cause the 3-month Treasury yield to fall. The data release that has the largest effect on the 3-month Treasury yield is nonfarm

[8] Thus, if we write X as a matrix with columns corresponding to macroeconomic series and rows corresponding to time t, each column of X will be a vector consisting mostly of zeros, with one nonzero value per month corresponding to dates on which news about the corresponding macroeconomic series was released. (For GDP, there is one nonzero value per quarter, and for initial claims, one nonzero value per week.)

[9] The historical standard deviations of these surprises are as follows: capacity utilization, 0.34 percentage points; consumer confidence, 5.1 index points; core CPI, 0.11 percentage points; real GDP, 0.76 percentage points; initial claims for unemployment insurance, 18.9 thousand workers; NAPM/ISM survey of manufacturers, 2.04 index points; leading indicators, 0.18 index points; new home sales, 60.6 thousand homes; nonfarm payrolls, 102.5 thousand workers; core PPI, 0.26 percentage points; retail sales excluding autos, 0.43 percentage points; and the unemployment rate, 0.15 percentage points.

	Treasury yield maturity					
	3-month		2-year		10-year	
Capacity Utilization	1.68	(2.93)	2.10	(4.23)	1.47	(2.51)
Consumer Confidence	0.29	(0.59)	2.67	(5.84)	2.69	(5.40)
Core CPI	0.79	(2.55)	2.33	(4.38)	1.71	(3.38)
GDP (advance)	0.33	(0.71)	−0.18	(−0.19)	−0.67	(−0.62)
Initial Claims	−0.29	(−1.36)	−0.63	(−2.42)	−0.39	(−1.43)
ISM Manufacturing	0.98	(1.47)	3.44	(7.23)	2.61	(4.99)
Leading Indicators	0.83	(1.58)	1.20	(2.36)	0.69	(1.09)
New Home Sales	1.46	(3.56)	1.98	(4.84)	2.04	(4.30)
Nonfarm Payrolls	2.44	(4.41)	4.56	(7.02)	2.86	(4.03)
Core PPI	0.52	(1.40)	0.87	(1.78)	1.33	(2.60)
Retail Sales ex. autos	1.19	(3.35)	1.83	(2.84)	1.18	(1.83)
Unemployment rate	−1.54	(−2.19)	−1.98	(−2.60)	−0.96	(−1.23)
# Observations	1303		1303		1303	
R^2	.07		.19		.09	
$H_0 : \beta = 0,\ p$-value	$< 10^{-8}$		$< 10^{-16}$		$< 10^{-16}$	

Table 1. Coefficient estimates β from linear regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$ at daily frequency on days of announcements from Jan. 1990 to Dec. 2000. Change in yields Δy_t is in basis points; surprise component of macroeconomic announcements X_t are normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics in parentheses. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. See text for details.

payrolls, for which a one-standard-deviation surprise causes the yield to move by about 2.5 bp on average, with a t-statistic of about 4.5. Taken together, the twelve data releases in Table 1 have a highly statistically significant effect on the 3-month Treasury yield, with a joint F-statistic above 5 and a p-value of less than 10^{-8}. The results for the 2- and 10-year Treasury yields in the second and third columns are similar, with joint statistical significance levels that are even higher than for the 3-month yield.[10] Thus, the high-frequency types of regressions conducted in Table 1 provide a great deal of power and information with which to estimate time-variation in the sensitivity of these yields to news in the next section.[11]

[10] The response of the 2-year Treasury yield to news is often larger than the response of the 3-month yield. In other words, the response of the yield curve to news tends to be hump-shaped. This is consistent with the standard result in monetary policy VARs that the federal funds rate has a hump-shaped response to output and inflation shocks (e.g., Sims and Zha 1999), and the finding in estimated monetary policy rules that the federal funds rate has inertia, so that the central bank responds only gradually to news (e.g., Sack and Wieland 2000). For simplicity, we considered a noninertial monetary policy rule in the previous section, but the key observations from that model are essentially unchanged if an inertial policy rule is used instead.

[11] Although the magnitudes of the coefficients in Table 1 are only a few basis points per standard deviation and the R^2 less than 0.2, these results should not be too surprising given the low signal-to-noise ratio of any single monthly data release for the true underlying state of economic activity and inflation. There are several reasons for this. For one, our surprise data cover only the headline component of each announcement, while the full releases are much

3.3 Measuring the Time-Varying Sensitivity of Treasury Yields

In principle, one can measure the time-varying sensitivity of Treasury yields to news by running regressions of the form (8) over one-year rolling windows. However, this approach suffers from small-sample problems because most macroeconomic series have data releases only once per month, providing just twelve observations per year with which to identify each element of the vector β.

We overcome this small-sample problem by imposing that the *relative* magnitude of the elements of β are constant over time, so that only the overall magnitude of β varies as the yield in question becomes more or less affected by the presence of the zero lower bound. Intuitively, if a Treasury security's sensitivity to news is reduced because its yield is starting to bump up against the zero bound, then we expect that security's responsiveness to all macroeconomic data releases to be damped by a roughly proportionate amount. This assumption is supported by the illustrative model in Section 2 and by empirical tests we conduct below.

Thus, for each given Treasury yield, we generalize regression (8) to a nonlinear least squares specification of the form:

$$\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i} \beta X_t + \varepsilon_t, \tag{9}$$

where the parameters γ^{τ_i} and δ^{τ_i} are scalars that are allowed to take on different values in each calendar year $i = 1990, 1991, \ldots, 2012$. (The reason for the notation γ^{τ_i}, δ^{τ_i} rather than γ^i, δ^i will become clear shortly.) The use of annual dummies in (9) is deliberately atheoretical in order to "let the data speak" at this stage; we will consider higher-frequency and more structural specifications for the time-varying sensitivity coefficients δ in Section 5.4, below. Note that regression (9) greatly reduces the small-sample problem associated with allowing every element of β to vary across years, because in (9) there are about 140 observations of βX_t per year with which to estimate each scalar δ^{τ_i}. Regression (9) also brings about twice as much data to bear in the estimation of β relative to the 1990–2000 sample considered in Table 1.

We must choose a normalization to separately identify the coefficients β and δ^{τ_i} in (9). We

richer: e.g., the employment report includes not just nonfarm payrolls and the unemployment rate, but also how much of the change in payrolls is due to government hiring, how much of the change in unemployment is due to workers dropping out of the labor force, and revisions to the previous two nonfarm payrolls announcements. The situation is very similar for all of the other releases in Table 1, and details such as these typically have a substantial effect on the markets' overall interpretation of a release. The important point to take away from Table 1 is that the large number of observations and extraordinary statistical significance of the regressions implies that they are extremely informative about the sensitivity of Treasury yields to economic news.

normalize the δ^{τ_i} so that they have an average value of unity from 1990–2000, which we take to be a period of relatively "normal" or unconstrained Treasury yield behavior. An estimated value of δ^{τ_i} close to unity thus represents a year in which the given Treasury yield behaved normally in response to news, while an estimated value of δ^{τ_i} close to zero corresponds to a year in which the given Treasury yield was completely unresponsive to news. Intermediate values of δ^{τ_i} correspond to years in which the Treasury yield's sensitivity to news was partially attenuated.

To provide a finer estimate of the periods during which each Treasury yield's sensitivity was attenuated, we also estimate daily rolling regressions of the form

$$\Delta y_t = \gamma^\tau + \delta^\tau \hat{X}_t + \varepsilon_t^\tau, \tag{10}$$

where $\hat{X}_t \equiv \hat{\beta} X_t$ denotes a "generic surprise" regressor defined using the estimated value of $\hat{\beta}$ from (9), and (10) is estimated over one-year rolling windows centered around each business day τ from January 1990 through December 2012.[12] When τ corresponds to the midpoint of a given calendar year $i \in \{1990, 1991, \ldots, 2012\}$, the estimated value of the attenuation coefficient δ^τ agrees exactly with δ^{τ_i} from regression (9). But we can also estimate (10) for any business day τ in our sample, and plot the coefficients δ^τ over time τ to provide a finer estimate of the periods during which each Treasury yield's sensitivity to news was attenuated. When we plot the standard errors in regression (10) around the point estimates for δ^τ, we account for the two-stage sampling uncertainty by using the estimated standard errors of the δ^{τ_i} from regression (9) as benchmarks and interpolating between them using the standard errors estimated in (10).[13]

As an alternative to the sensitivity of Treasury yields, δ^τ, one could also consider the time-varying unconditional volatility of those yields—that is, the six-month or one-year rolling standard deviation of yield changes Δy_t. Looking at the unconditional volatility of yields is simpler and requires less data than our sensitivity-based approach above. The primary disadvantage of unconditional volatility is that it can vary due to changes in the variance of the ε_t in regressions (9) and

[12] Toward either end of our sample, the regression window gets truncated and thus becomes smaller and less centered, approaching a six-month leading window in January 1990 and a six-month trailing window in December 2012.

[13] Nonlinear least squares regression (9) is a one-stage regression whose standard errors are estimated consistently and efficiently under standard econometric assumptions. Thus, we use these standard errors for δ^{τ_i} as benchmarks. Let σ^τ denote the OLS standard error for δ^τ on date τ from daily rolling regression (10) (which ignores the sampling uncertainty for $\hat{\beta}$). Let ς^{τ_i} denote the benchmark standard errors for δ^{τ_i} from NLS regression (9). For τ between benchmark dates τ_i and τ_{i+1}, we scale up σ^τ by a factor of $\frac{\tau_{i+1}-\tau}{\tau_{i+1}-\tau_i} \cdot \frac{\varsigma^{\tau_i}}{\sigma^{\tau_i}} + \frac{\tau-\tau_i}{\tau_{i+1}-\tau_i} \cdot \frac{\varsigma^{\tau_{i+1}}}{\sigma^{\tau_{i+1}}}$.

(10) because of factors such as liquidity and risk premia. If the unconditional volatility of a given Treasury yield is greater than zero, that tells us little about whether monetary policy can affect that yield or not, because it could be that the residuals ε_t for that yield are large. In contrast, if the *sensitivity* of a given Treasury yield to macroeconomic news, δ^τ, is substantially greater than zero, then the case that monetary policy can affect that yield is much stronger. Thus, the sensitivity of yields to macroeconomic news is more relevant than unconditional volatility for monetary and fiscal policy.

4 Main Results

Table 2 reports nonlinear least squares estimates for β in regression (9) for the 3-month, 2-year, and 10-year Treasury yields over the sample January 1990 through December 2012. The results in Table 2 are generally similar to those in Table 1, although the number of observations in Table 2 is more than twice as large as in Table 1, owing to the longer sample.

At the bottom of Table 2, we report results for three specification tests. First, we test the hypothesis that the relative response coefficients β in regression (9) are constant over time—and only the scalar attenuation coefficients δ^{τ_i} vary—against an alternative in which every element of β is permitted to vary independently across calendar years, that is:

$$\Delta y_t = \gamma^{\tau_i} + \beta^{\tau_i} X_t + \varepsilon_t. \tag{11}$$

As can be seen in Table 2, the restricted specification (9) is consistent with the data, with p-values above .25 for each regression.[14] As an additional robustness check on this restriction, we have also run regression (9) with a smaller, more homogeneous set of announcements X_t, such as nonfarm payrolls alone, or nonfarm payrolls, initial claims, and the unemployment rate. In the interest of space, these results are reported in an online Appendix, but the estimated time-varying sensitivity coefficients δ^{τ_i} are very similar to those reported in this section (although the estimates are much noisier when the number of regressors is reduced in this way).

[14]In Swanson and Williams (2013a), we performed a Wald test of this restriction, which produced p-values of 1 to three decimal places. The results reported here are for a GMM J-test of overidentifying restrictions, which was suggested by a referee as being a more robust test. Regression (9) has 57 coefficients while there are 297 moment conditions of the form $Ex_t\varepsilon_t = 0$ corresponding to the unrestricted model (11). (Recall that we do not have data for two series in 1990.) The J-test p-values reported in Table 2 are for the continuous-updating GMM estimator, which Hansen, Heaton, and Yaron (1996) find gives the most reliable test statistics in finite samples.

	Treasury yield maturity					
	3-month		2-year		10-year	
Capacity Utilization	0.73	(1.56)	1.49	(2.89)	0.68	(2.02)
Consumer Confidence	0.75	(2.90)	1.37	(3.71)	0.84	(2.43)
Core CPI	0.39	(1.88)	1.89	(5.00)	1.17	(3.60)
GDP (advance)	0.92	(3.15)	1.42	(2.40)	0.95	(1.69)
Initial Claims	−0.30	(−1.82)	−1.10	(−5.35)	−0.95	(−5.02)
ISM Manufacturing	1.23	(3.24)	2.72	(7.09)	1.98	(5.96)
Leading Indicators	0.20	(0.62)	0.28	(0.85)	0.28	(1.01)
New Home Sales	0.83	(2.65)	0.65	(1.99)	0.50	(1.93)
Nonfarm Payrolls	3.03	(7.67)	4.79	(9.54)	2.95	(6.79)
Core PPI	0.22	(0.79)	0.52	(1.54)	0.85	(3.14)
Retail Sales ex. autos	0.83	(3.76)	1.86	(4.92)	1.62	(4.31)
Unemployment rate	−1.24	(−3.53)	−1.26	(−2.78)	−0.41	(−1.07)
# Observations	2829		2829		2829	
R^2	.08		.17		.10	
$H_0 : \beta$ constant, p-value	.598		.265		.630	
$H_0 : \delta$ symmetric, p-value	.095		.310		.319	
$H_0 : \delta$ constant, p-value	$< 10^{-16}$		$< 10^{-16}$		.015	

Table 2. Coefficient estimates β from nonlinear regression $\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i}\beta X_t + \varepsilon_t$ at daily frequency from Jan. 1990 to Dec. 2012. Coefficients indexed τ_i may take on different values in different calendar years. Δy_t and X_t are as in Table 1. Heteroskedasticity-consistent t-statistics in parentheses. $H_0 : \beta$ constant p-value is for the test that β is fixed over time and only the δ^{τ_i} vary. $H_0 : \delta$ symmetric tests whether δ^{τ_i} is the same for positive and negative surprises βX_t. $H_0 : \delta$ constant tests whether $\delta^{\tau_i} = 1$ for all years i. See text for details.

Second, we test the hypothesis that the δ^{τ_i} in (9) are the same for positive and negative surprises βX_t, against an alternative in which we allow separate attenuation coefficients $\delta_+^{\tau_i}$ and $\delta_-^{\tau_i}$ for positive and negative values of βX_t in each calendar year i. In other words, we separate the data into two groups—those announcements that have positive implications for Treasury yields, and those that have negative implications—and test whether the attenuation coefficients $\delta_+^{\tau_i} = \delta_-^{\tau_i}$ for each $i = 1990, \ldots, 2012$.[15] As can be seen in Table 2, this restriction is also not rejected by the data, with p-values typically substantially above ten percent. Although the symmetry restriction appears to be marginally rejected for the 3-month Treasury yield, with a p-value of .095, that result is entirely driven by a large outlier on October 20, 2008, when the 3-month T-bill yield jumped by 41 bp.[16] Excluding that single observation, the p-value for the 3-month Treasury yield hypothesis

[15]The first group consists of all of the unemployment rate and initial claims surprises that are less than zero, and all of the positive surprises in the other statistics. The second group consists of all of the unemployment rate and initial claims surprises that are greater than zero, and all of the negative surprises in the other statistics.

[16]The only macroeconomic data released that day was leading indicators, which had a positive surprise of about

test is .709. We conclude that this restriction is also consistent with the data.

Third, we test the hypothesis that the time-varying sensitivity coefficients δ^{τ_i} in (9) are constant over time. That is, we test whether $\delta^{\tau_i} = 1$ for each calendar year $i = 1990, \ldots, 2012$. In contrast to the previous two tests, here the data strongly reject the restriction for the 3-month and 2-year Treasury yields, with p-values less than 10^{-16}. Clearly, the sensitivity of these two yields to macroeconomic news has varied substantially over time. The constant-δ restriction for the 10-year yield is also rejected, but less strongly, with a p-value of .015. Although the 10-year yield's sensitivity to news does appear to have varied over time, the assumption of constant sensitivity for this yield is not nearly as inconsistent with the data as for the shorter-maturity yields.

Figure 3 plots the time-varying sensitivity coefficients δ^{τ} from regression (9) as a function of time τ, using the daily rolling regression specification (10). The six panels of the figure depict results for the 3-month, 6-month, and 1-, 2-, 5-, and 10-year Treasury yields. The solid blue line in each panel plots the estimated value of δ^{τ} on each date τ, while the dotted gray lines depict heteroskedasticity-consistent $\pm$2-standard-error bands, adjusted for the two-stage estimation procedure as described in the preceding section. In each panel, horizontal black lines are drawn at 0 and 1 as benchmarks for comparison, corresponding to the cases of complete insensitivity to news and normal sensitivity to news, respectively.

In each panel, the yellow shaded regions denote periods during which the estimated value of δ^{τ} is significantly less than unity at the one percent level. We use a conservative threshold here so that the shaded regions represent periods in which the yield was clearly less sensitive to news than normal. In addition, if the hypothesis $\delta^{\tau} = 0$ cannot be rejected, then the region is shaded red.[17] Thus, red shaded regions correspond to periods in which the Treasury yield was essentially insensitive to news, while yellow shaded regions correspond to periods in which the yield was partially—but not completely—unresponsive to news.

Panel (a) of Figure 3 shows that the sensitivity of the 3-month Treasury yield to macroecon-

2 standard deviations. According to *The Wall Street Journal,* the major news that day was that J.P. Morgan Chase and other large banks lent billions of dollars to their counterparts in Europe, which "spurred improvement in the commercial paper market... With more appetite for risk came an exodus out of government debt. Treasury bills suffered the most... Bills were also pressured by more than \$80 billion in bill supply... from the Treasury department." ("Credit Markets: Bonds and Stocks Show Signs of Healing," *The Wall Street Journal,* October 21, 2008, Romy Varghese and Emily Barrett, p. C1.)

[17]We use a standard five percent threshold here. A one percent threshold would result in the red shaded regions being slightly larger.

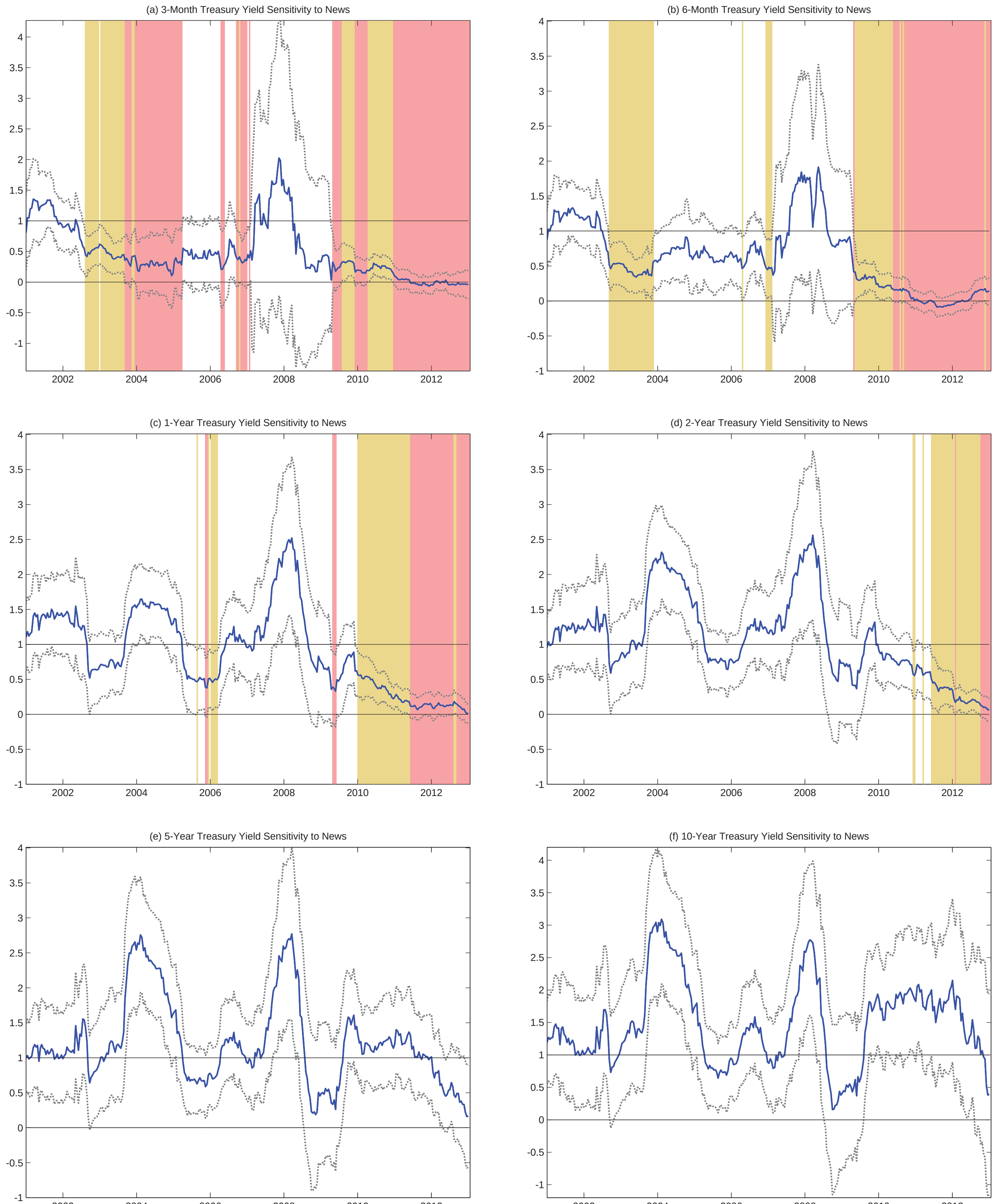

Figure 3. Time-varying sensitivity coefficients δ^τ from regression (10) for (a) 3-month, (b) 6-month, (c) 1-year, (d) 2-year, (e) 5-year, and (f) 10-year Treasury yields. Dotted gray lines depict heteroskedasticity-consistent ± 2-standard-error bands, adjusted for two-stage sampling uncertainty in (10). $\delta^\tau = 1$ corresponds to normal Treasury sensitivity to news; $\delta^\tau = 0$ to complete insensitivity. Yellow shaded regions denote δ^τ significantly less than 1; red shaded regions denote δ^τ significantly less than 1 and not significantly different from 0. See text for details.

omic news has varied between about 0 and 2 from 2001 through 2012. From the spring of 2009 through the end of 2012, the 3-month Treasury yield was either partially or completely insensitive to news. It is natural to interpret this insensitivity as being driven by the zero lower bound, since the federal funds rate and 3-month Treasury yields were both essentially zero from December 2008 through the end of our sample. At the shortest end of the yield curve, at least, Treasury yields appear to have been substantially constrained by the zero bound from the spring of 2009 onward.

What is perhaps more surprising in the first panel of Figure 3 is that the 3-month Treasury yield was also partially or completely insensitive to news throughout 2003 and 2004, a period during which the federal funds rate target and 3-month Treasury yield never fell below 1 percent. However, the Fed had recently lowered the funds rate to 1.25 percent in November 2002 and again to 1 percent in June 2003, and at the time, a level of the funds rate below 1 percent was regarded as costly for institutional reasons (Bernanke and Reinhart, 2004). Rather than try to lower the funds rate below 1 percent, the FOMC opted instead to switch to a policy of managing monetary policy expectations, using phrases such as "policy accommodation can be maintained for a considerable period."[18] Thus, even though the funds rate was not constrained by a floor of zero in 2003 and 2004, our results show that the 3-month Treasury yield behaved *as if* it had been constrained by a floor of 1 percent. The fact that our empirical method picks up the constraints faced by monetary policy in 2003–04, and the potential absence of crowding out of fiscal policy over the same period, is a noteworthy feature of our approach.

Panel (b) of Figure 3 reports analogous results for the 6-month Treasury yield, which are generally similar to those for the 3-month yield: the sensitivity to macroeconomic news ranges between 0 and 2, and from the spring of 2009 through the end of 2012, the 6-month yield was either partially or completely unresponsive to news. In contrast to the 3-month yield, however, the 6-month yield's sensitivity to news was much less attenuated in 2003–04. Thus, to the extent that the effective lower bound of 1 percent was a substantial constraint on monetary policy in 2003–04, that constraint did not appear to extend out to maturities beyond 3 months in 2004.

Results for 1- and 2-year Treasury yields are reported in the middle panels of Figure 3. The

[18]The "considerable period" language was introduced into the FOMC statement on August 12, 2003, and continued until the end of January 2004, at which point it was replaced with the phrase, "the Committee believes that it can be patient in removing its policy accommodation." The funds rate was finally raised on June 30, 2004.

sensitivity of these intermediate-maturity yields to news is less attenuated than that of 3- and 6-month yields throughout our sample. For example, both the 1- and 2-year yields behaved close to normally throughout 2003–04, implying that they were relatively unaffected by the FOMC's implicit floor of about 1 percent during this period. Thus, to the extent that the FOMC can affect yields with a year or more to maturity, we would conclude that the effectiveness of monetary and fiscal policy were very close to normal in 2003–04.

What is perhaps most surprising in the middle panels of Figure 3 is how little and how late the zero bound seems to have affected these intermediate-maturity yields after 2008. The sensitivity of the 1-year yield to news was only significantly less than unity beginning in 2010, and even then is partially responsive to news until late 2011. Only beginning in late 2011 does the 1-year Treasury yield cease responding to news. The 2-year Treasury yield's sensitivity to news was generally not significantly attenuated until late 2011, and even then remained partially responsive to news until late 2012. Thus, to the extent that the Fed can influence monetary policy expectations over a horizon out to two years, we conclude that monetary and fiscal policy were likely to have been about as effective as usual until at least late 2011.

The bottom two panels of Figure 3 report results for 5- and 10-year Treasury yields, which are also remarkable. There are essentially no red or yellow shaded regions in these panels, because the sensitivity of these yields to news is never significantly less than one until the last few weeks of 2012. Even in late 2011 and 2012, when Treasury yields out to two years were becoming substantially constrained, 5- and 10-year yields remained largely unconstrained. However, the sensitivity of the 5-year yield to news declined throughout 2012, and given the substantial decline in the 5- and 10-year yields' sensitivity to news toward the end of 2012, it would be very interesting to see how this sensitivity evolves going forward.

5 Discussion

We now discuss the broader implications of our empirical findings and perform several extensions and robustness checks. First, we compare our results to private-sector expectations of the time until federal funds rate "liftoff" from the zero bound. Second, we discuss the implications of our

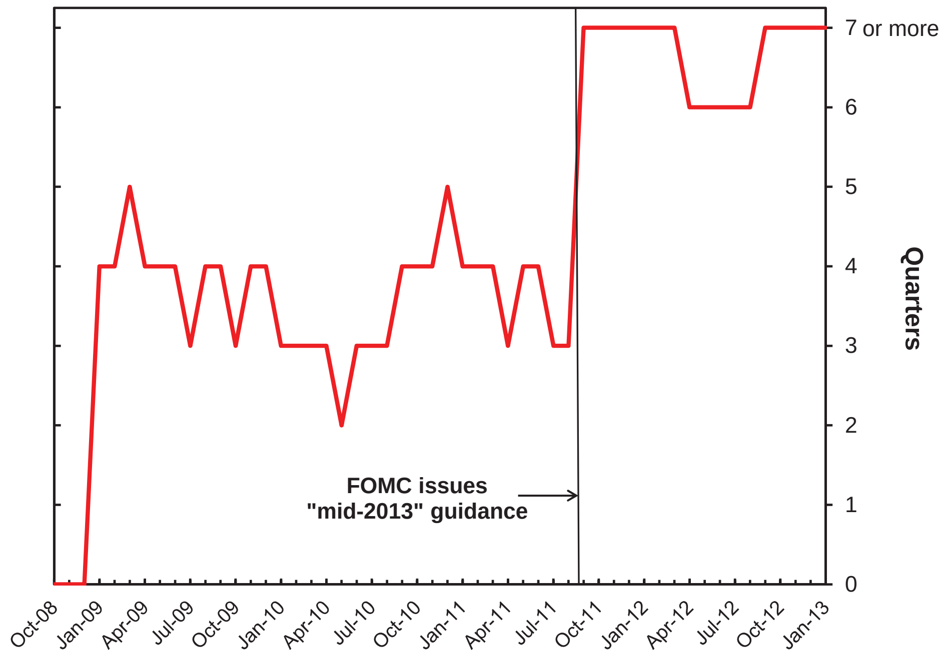

Figure 4. Expected number of quarters until the first federal funds rate increase above 25 bp, from the monthly Blue Chip survey of forecasters. Data are top-coded at "7 or more" quarters due to the forecast horizon length published by Blue Chip.

findings for the fiscal multiplier. Third, we show that time-varying term premia would not affect the interpretation of our main results. Fourth, we investigate to what extent a reduced sensitivity of Treasury yields to news can be explained mechanically by a lower level of yields or by changes in monetary policy uncertainty. Finally, we show that the distribution of our macroeconomic surprise data from 2008 onward is not very different from the distribution of those surprises before 2008.

5.1 Private-Sector Expectations of Federal Funds Rate "Liftoff" from Zero

Our illustrative model in Section 2 implies that the sensitivity of medium- and longer-term Treasury yields to news is closely related to the length of time that the federal funds rate is expected to be at the zero lower bound. For example, if the funds rate is expected to be at zero for just one quarter, then medium- and longer-term interest rates should be nearly unaffected by the zero bound, whereas if the funds rate is expected to be at the zero bound for several years, then even 5- or 10-year Treasury yields should be noticeably affected.

Figure 4 plots the number of quarters until the private sector expected the funds rate to be 25 bp or higher, as measured by the median, "consensus" response to the monthly Blue Chip survey of professional forecasters. Prior to December 2008, the FOMC was not expected to lower the funds rate below 25 bp. After the FOMC cut the target funds rate to about zero in December 2008, the Blue Chip consensus expectation of the length of time until the first funds rate increase then fluctuated between two and five quarters until August 2011. On August 9, 2011, the FOMC announced that it expected to keep the funds rate near zero "at least through mid-2013," and private-sector expectations of the time until liftoff jumped to seven or more quarters (the Blue Chip forecast horizon extends forward only six quarters).[19]

The implication of the forecasts underlying Figure 4 is that, from about January 2009 until August 2011, the sensitivity of Treasury yields with a year or less to maturity should have fallen close to zero, while that for maturities of two years or more should have been only partially attenuated. Only beginning in August 2011 would we expect to see yields with two years to maturity show substantial attenuation with respect to news. And in fact, this corresponds closely to our time-varying sensitivity results in Figure 3.

Figure 5 provides an additional perspective on these results from the interest rate options market. Using daily options data with a range of strike prices and five quarters to expiration, we can estimate the entire implied distribution of the federal funds rate in five quarters' time at daily frequency.[20] We can then use these estimated distributions to back out the implied probability that the federal funds rate would be less than 50 bp in five quarters' time, which we plot in Figure 5 from January 2008 to December 2012.

The implied probabilities in Figure 5 corroborate the survey results in Figure 4 and our sensitivity estimates in Figure 3. Before September 2008, options traders apparently viewed the probability of the funds rate being less than 50 bp in five quarters' time as very low—less than about

[19]The Federal Reserve Bank of New York also surveys primary dealers in the Treasury market, with survey results since January 2011 made available at http://www.newyorkfed.org/markets/primarydealer_survey_questions.html. The results of this survey show a similar jump in the median primary dealer forecast, from 5 quarters in August 2011 to 9 quarters in September 2011.

[20]We do not need to assume normality for these distributions because we observe option prices for multiple different strikes. On each day from January 2008 through December 2012, we use the range of available Eurodollar option put and call prices with five quarters to expiration to estimate the implied distribution of the spot 3-month Eurodollar rate in five quarters' time, using a flexible functional form. Eurodollar options are the most liquid options on a short-term interest rate and thus provide the best measure of the distribution of possible short-term interest rate outcomes. We use the spread between overlapping federal funds futures and Eurodollar futures rates at a one-year horizon to convert these implied distributions for the 3-month Eurodollar rate into an implied distribution for the federal funds rate. These probability estimates ignore risk premia and thus represent implied risk-neutral probabilities.

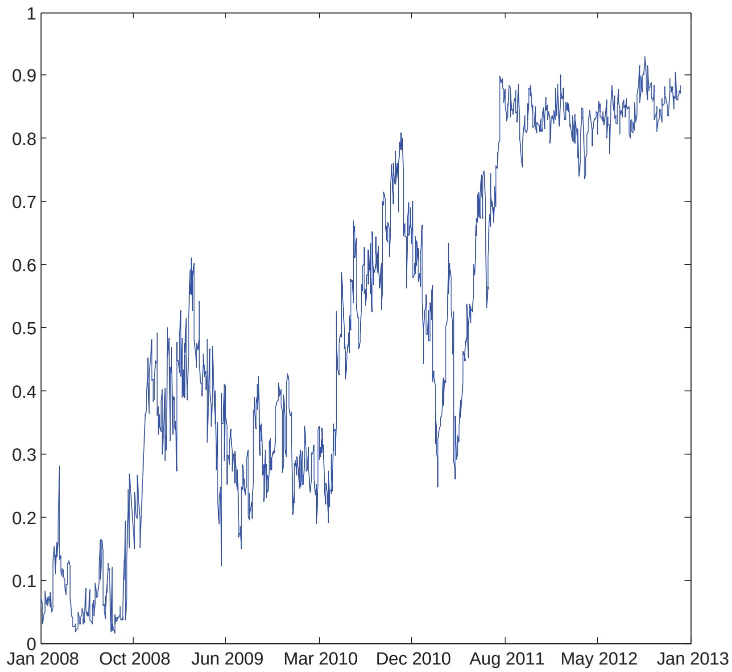

Figure 5. Probability the federal funds rate would be less than 50 bp in five quarters, estimated from options data. See text for details.

10 percent. Between September 2008 and mid-2010, this probability rose modestly to somewhere between 20 and 45 percent—larger than before, but still less likely than not. From mid-2010 to mid-2011, this probability fluctuated more widely between about 30 and 70 percent. Given the relatively low level and substantial movements in these probabilities, it's not surprising that 2-year or even 1-year Treasury yields responded almost normally to news throughout much of this period. Only beginning in August 2011 do we see the probabilities in Figure 5 increase to around 85 percent, corresponding to a more reduced sensitivity of the 2-year Treasury yield to news.

Figure 6 provides a final robustness check on these results by applying regression (10) to Eurodollar futures rather than Treasury yields. Eurodollar futures are the most heavily traded futures contracts in the world and settle at expiration based on the spot 3-month term Eurodollar deposit rate in London.[21] Thus, a Eurodollar future with one quarter to expiration is closely related

[21]See Gürkaynak, Sack, and Swanson (2007) for additional details regarding Eurodollar futures. They compare the ability of a variety of financial market instruments and econometric models to forecast the federal funds rate and find that Eurodollar futures perform as well as or better than any other measure at horizons of six months or more, which is the most relevant for our present analysis. Figure 6 lists the expiration of each contract as 1–2 quarters ahead, 2–3 quarters ahead, etc., because contracts expire in March, June, September, and December of each year; thus the number of quarters to expiration can lie anywhere between n and $n+1$ quarters, depending on whether the current date t is closer to the beginning or the end of the current quarter.

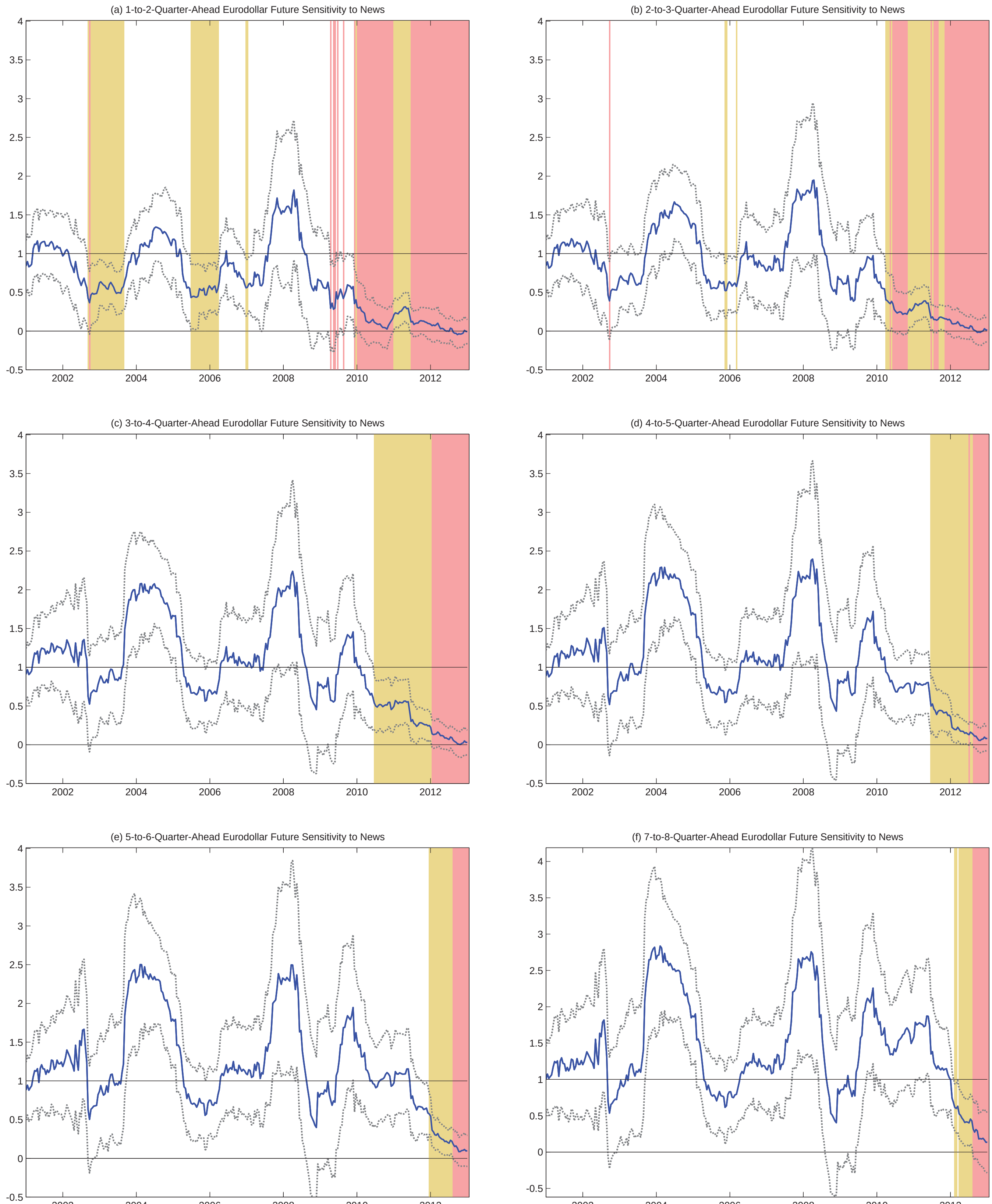

Figure 6. Time-varying sensitivity coefficients δ^{τ} from regression (10) for Eurodollar futures contracts with (a) 1–2 quarters, (b) 2–3 quarters, (c) 3–4 quarters, (d) 4–5 quarters, (e) 5–6 quarters, and (f) 7–8 quarters to expiration. Eurodollar futures settle based on the spot 3-month Eurodollar deposit rate at expiration, and thus correspond to forward interest rates beginning at expiration and ending 1 quarter after expiration. See notes to Figure 3 and text for details.

to market expectations about the federal funds rate from 3 to 6 months ahead, a Eurodollar future with 2 quarters to expiration closely reflects market expectations about future monetary policy from 6 to 9 months ahead, and so on.

The results in Figure 6 confirm those for Treasury yields in Figure 3. Just like 3- and 6-month Treasury yields, the sensitivity to news of Eurodollar futures with 1 to 2 quarters to expiration was attenuated in 2003, and fell to essentially zero in 2010–12. Similarly, with 4 to 5 quarters to expiration, Eurodollar futures' sensitivity to news remained near normal levels until late 2011, around the time of the FOMC's announcement that it expected to keep the funds rate near zero "at least through mid-2013." This mirrors very closely the behavior of the 2-year Treasury yield in Figure 3. Even longer-maturity Eurodollar futures, with 5 to 8 quarters to expiration, continued to respond normally to news until the FOMC announced in January 2012 that it expected to keep the funds rate near zero "at least through late 2014."

The results in this section suggest that financial markets did not expect the zero bound to constrain the funds rate for more than a few quarters until about August 2011, around the time of the FOMC's "mid-2013" forward guidance. Only then do we see interest rate expectations more than four quarters ahead begin to behave in an attenuated fashion. An interesting question, then, is to what extent the decline in Treasury yield sensitivity represents a voluntary, endogenous commitment by the FOMC to keep interest rates low for a long time, rather than an exogenous constraint? (As an example of the latter, the state of the economy could be so weak that the central bank is simply not expected to raise rates above zero for a very long time, which could be true even under perfect discretion.)

Our empirical methods do not discriminate between these two different ways in which monetary policy expectations might be fixed at zero. Nevertheless, the distinction between an exogenous constraint and endogenous commitment by the FOMC matters very little for our results and conclusions. First, from the point of view of the fiscal multiplier (discussed below), the distinction does not matter; the only relevant issue is whether or not monetary policy expectations respond to changes in the economic outlook, regardless of the reason. Second, in periods when we estimate a given Treasury yield *is* sensitive to news—including the case of partial sensitivity $0 < \delta^\tau < 1$—our results tell us that *neither* an exogenous constraint nor endogenous commitment is fully binding, so monetary policy can still try to operate through that yield. Third, in periods when we estimate a given Treasury yield is *not* sensitive to news ($\delta^\tau \approx 0$), it is still useful for policymakers to know that that yield has come to the point where it can no longer be manipulated.

Finally, we note that the distinction between an exogenous constraint and endogenous commitment is not so sharp in practice. After all, it is precisely *because of* the zero lower bound that the FOMC issued its "mid-2013", "late 2014", and "mid-2015" forward guidance.[22] Even though this forward guidance was in some sense voluntary, it was clearly driven by the (exogenous) presence of the zero lower bound.

5.2 Implications for the Fiscal Multiplier

Our empirical results have important implications for the growing literature on the fiscal multiplier at the zero lower bound, such as Christiano et al. (2011), Woodford (2011), and the other studies cited in footnote 1. An important finding of that literature is that the fiscal multiplier is larger the greater the fraction of the change in government spending that is expected to take place while the short-term interest rate is at zero.[23] Put differently, for a given path of fiscal stimulus, the multiplier is larger the longer the zero bound is expected to constrain the short-term interest rate. Intuitively, when the zero lower bound is expected to constrain the short-term rate for a longer period of time, monetary policy is more constrained. The more constrained is monetary policy, the less there is any "crowding out" of fiscal stimulus by changes in the private sector's expected path of future short-term interest rates.

Our empirical results suggest that, until late 2011, financial markets expected the federal funds rate to rise above zero relatively quickly, within the next four quarters. By contrast, Christiano et al. (henceforth CER) consider a scenario in which the zero bound is expected to bind for a longer period of time—either 8 or 12 quarters—in their quantitative simulations using the Altig et al.

[22] After its September 2012 meeting, the FOMC announced it expected to keep the funds rate near zero "at least through mid-2015."

[23] For example, "Our basic result is that the multipliers are higher the larger the percentage of the spending that comes on line when the nominal interest rate is zero" (Christiano et al. 2011, p. 112); "A key lesson from this analysis is that... it is critical that the spending come on line when the economy is actually in the zero bound. Spending that occurs after that yields very little bang for the buck and actually dulls the impact of the spending that comes on line when the zero bound binds" (ibid., p. 112); and "Hence, while there is a positive effect on output during the crisis of increased government purchases at date $t < T$, an anticipation of increased government purchases at dates $t \geq T$ has a *negative* effect on output prior to date T" (Woodford 2011, p. 22, emphasis original).

(2011) model. When CER consider a scenario in which the zero bound binds for only 4 quarters, they find that the fiscal multiplier is essentially no different from normal—that is, no greater than when the short-term interest rate is unconstrained (CER, footnote 12).

Based on the results in CER and our own estimates of the length of time markets expected the zero bound to constrain short-term interest rates, we conclude that the fiscal multiplier was likely close to normal throughout 2008–10. Only beginning in late 2011, when markets began to expect the federal funds rate would be stuck at zero for a longer period of time, would we expect the fiscal multiplier to approach the larger values estimated by those authors.[24]

More generally, our results suggest that the sensitivity of intermediate-maturity bond yields to economic news is a good indicator of the relative size of the fiscal multiplier. Intermediate-maturity yields are less sensitive to news when the zero bound is expected to constrain the federal funds rate for a longer period of time. Thus, as a general rule, periods when the fiscal multiplier is larger are also periods when intermediate-maturity bond yields are less sensitive to news, consistent with the standard IS-LM intuition of a smaller degree of crowding out.

5.3 Time-Varying Term Premia and Treasury Yield Sensitivity

The simple, linearized New Keynesian model in Section 2 helps to motivate our analysis and facilitate the interpretation of our results. However, it also implies that longer-term bond yields satisfy the expectations hypothesis; that is, long-term yields equal the average expected short-term rate over the lifetime of the bond, as in equation (5). While that equation allows for the existence of a term premium ϕ^M that may vary with maturity M, the premium is assumed to be constant over time. Contrary to this assumption, many empirical studies in finance find that term premia vary substantially over time (e.g., Campbell and Shiller, 1991; Campbell, Lo, and MacKinlay, 1997).

The effects of any time-variation in term premia on the *sensitivity* of Treasury yields to major macroeconomic announcements is unclear, however, and could be positive, negative, or zero.[25]

[24]It is interesting that financial markets' expectation of a quick liftoff from the zero bound in 2008–10 turned out to be incorrect ex post. Nevertheless, as is clear from Woodford's (2011) analysis, it is the private sector's *expectations* at time t regarding the future path of short-term interest rates and government spending that is crucial for determining the effect on output at time t.

[25]For example, risk premia should increase when the economy is weak (e.g., Campbell and Cochrane, 1999; Piazzesi and Swanson, 2008; and Lettau and Ludvigson, 2010). If risk aversion was higher during the financial crisis, then good news about the economy would tend to produce a larger decline in term premia to offset the expected increase in the path of the federal funds rate, producing a *decreased* Treasury yield sensitivity to news. Alternatively, the

Regardless, the interpretation of our main results is essentially independent of any such effects. Our main empirical findings are all consistent with the simple model of Section 2, and do not represent a "sensitivity puzzle" that would require searching outside the model for an explanation. Although we find that Treasury yields with a year or more to maturity were surprisingly responsive to news in 2008–10, that finding is consistent with survey-based expectations of a rapid liftoff from the zero bound, depicted in Figure 4; thus, even though that finding is surprising, it agrees with measures of monetary policy expectations at the time and does not require an explanation based on term premia. After the FOMC's "mid-2013" forward guidance in August 2011, expectations of funds rate liftoff shifted out dramatically (Figure 4) and the sensitivity of intermediate-maturity bond yields fell close to zero, again consistent with the simple model of Section 2 and not very suggestive of an explanation based solely on term premia.

Although time-varying term premia do not seem able to explain our main empirical results on their own, they may help to explain some features of the data, such as why 5- and 10-year Treasury yields were about as sensitive to news from 2008–12 as in normal times. This finding is particularly surprising after August 2011, given that the sensitivity of the 2-year Treasury yield was so low, and that the simple model of Section 2 predicts *all* yields should be attenuated by the zero bound to some extent. An answer to this puzzle is that, if term premia on 5- and 10-year Treasuries varied along with changes in market expectations about Fed purchases of long-term bonds, the increased sensitivity of 5- and 10-year term premia to news could have offset any dampening effect from the zero bound on those yields.[26] The net result would be less attenuation in 5- and 10-year Treasury yield sensitivity to news between 2008 and 2012, perhaps even no net attenuation.

5.4 Other Explanations for Time-Varying Treasury Yield Sensitivity

The sensitivity of Treasury yields to news could vary for reasons other than the zero lower bound or time-varying term premia. For example, it is well known that interest rate volatility—and thus,

financial crisis could have led to an increased demand for safe, liquid assets. In that case, bad news about the economy might cause a "flight to quality" that would drive Treasury yields below the expected decrease in the path of the federal funds rate, producing an *increased* sensitivity of Treasury yields to news.

[26]Between 2008 and 2012, the FOMC announced several rounds of long-term bond purchases, amounting to over $3.2 trillion in total. Many authors, such as Krishnamurthy and Vissing-Jorgensen (2011, 2012), Gagnon et al. (2011), and Swanson (2011), estimate that such large-scale changes in the supply of long-term bonds affect the yields on those securities. Hamilton and Wu (2012) show how these purchases likely affect term premia in particular.

presumably, interest rate sensitivity to news—declines along with the overall level of yields (e.g., Chan et al., 1992; Kim and Singleton, 2012). Thus, part of our estimated decline in Treasury yield sensitivity since 2009 could simply reflect the decline in Treasury yields over that period.

Changes in financial market uncertainty about future short-term interest rates could also be an important determinant of bond yield sensitivity to news. Suppose that financial markets use a Kalman filter or Bayesian updating to revise their expectations about the path of future short-term interest rates. Then the financial market sensitivity on day t to a macroeconomic data release depends on the variance of the surprise component of that data release and the day $t-1$ prior variance of the variable being forecast—in this case, the future short-term interest rate. If the market's prior variance is very small, then market participants have a great deal of confidence in their expectation and will respond relatively little to any data release on day t. On the other hand, if the market's date $t-1$ prior variance is very large, then markets will respond much more strongly to any news on day t. This effect could help to explain why our estimates of δ^τ in Figures 3 and 6 are sometimes significantly *higher* than normal as well as lower than normal; for example, the 2-year Treasury yield's sensitivity to news was more than twice as high in 2004 and from mid-2007 to mid-2008 as in the benchmark sample, 1990–2000.

We investigate the importance of these explanations by considering a more structural specification for the time-varying sensitivity coefficients δ^τ in regressions (9) and (10). In particular, we consider a nonlinear regression of the form

$$\Delta y_t \;=\; \gamma + f(Z_t)\beta X_t + \varepsilon_t, \tag{12}$$

where Z_t denotes a vector of explanatory variables for Treasury yield sensitivity, such as the level of yields or a measure of monetary policy uncertainty, and

$$f(Z_t) \;=\; \theta + \phi Z_t, \tag{13}$$

where θ and ϕ are parameters to be estimated along with γ and β in (12). Thus, instead of estimating δ^{τ_i} or δ^τ in an unrestricted way in (9) or (10), with annual dummies or rolling regressions, the specification (12)–(13) allows the sensitivity of Treasury yields to vary at daily frequency along with variations in Z_t. The constant θ in (13) is normalized so that the average sensitivity $f(Z_t)$ is equal to unity over the benchmark sample, 1990–2000, as in regression (9).

Regressions with Treasury Yield Sensitivity a Function of Explanatory Variables

	(1)		(2)		(3)	
(A) 3-month Treasury Yield						
own interest rate level	.130	(3.49)			.051	(0.93)
monetary policy uncertainty			.526	(4.73)	.433	(2.58)
time trend						
R^2	.11		.11		.11	
(B) 2-year Treasury Yield						
own interest rate level	.048	(2.09)			−.023	(−0.72)
monetary policy uncertainty			.292	(5.90)	.332	(4.58)
time trend						
R^2	.19		.19		.19	
(C) 10-year Treasury Yield						
own interest rate level	−.020	(−0.38)			−.120	(−1.32)
monetary policty uncertainty			.074	(0.81)	.248	(1.51)
time trend						
R^2	.11		.11		.11	

Table 3. Coefficient estimate ϕ from nonlinear regression $\Delta y_t = \gamma + (\theta + \phi Z_t)\beta X_t + \varepsilon_t$ at daily frequency from Jan. 1990 to Dec. 2012. Heteroskedasticity-consistent t-statistics in parentheses. See text for details.

Table 3 reports results for regression (12)–(13) for the 3-month, 2-year, and 10-year Treasury yields. In the first column, Z_t includes just the level of that same yield. The sensitivity of the 3-month and 2-year yields to news is significantly positively related to the levels of those yields, as expected, although the relationship for the 10-year yield is insignificant and negative. A decline in the 3-month Treasury yield of 1 percentage point is associated, on average, with a decrease in the yield's sensitivity to news of about 13 percent, relative to the benchmark sample. The R^2 for these regressions are a bit higher than for regression (9), indicating an improvement in fit from the higher-frequency variation in sensitivity allowed by the more structural specification (13).

The second column of Table 3 reports results when Z_t includes a measure of monetary policy uncertainty. As described in Section 5.1, we use Eurodollar options to estimate the implied probability distribution for the Eurodollar rate one year ahead, and take the distance between the 80^{th} and 20^{th} percentiles of that distribution as a measure of monetary policy uncertainty.[27] This

[27] Detailed options data are not available to us prior to 1996, so from 1990–1995, we compute the width of the

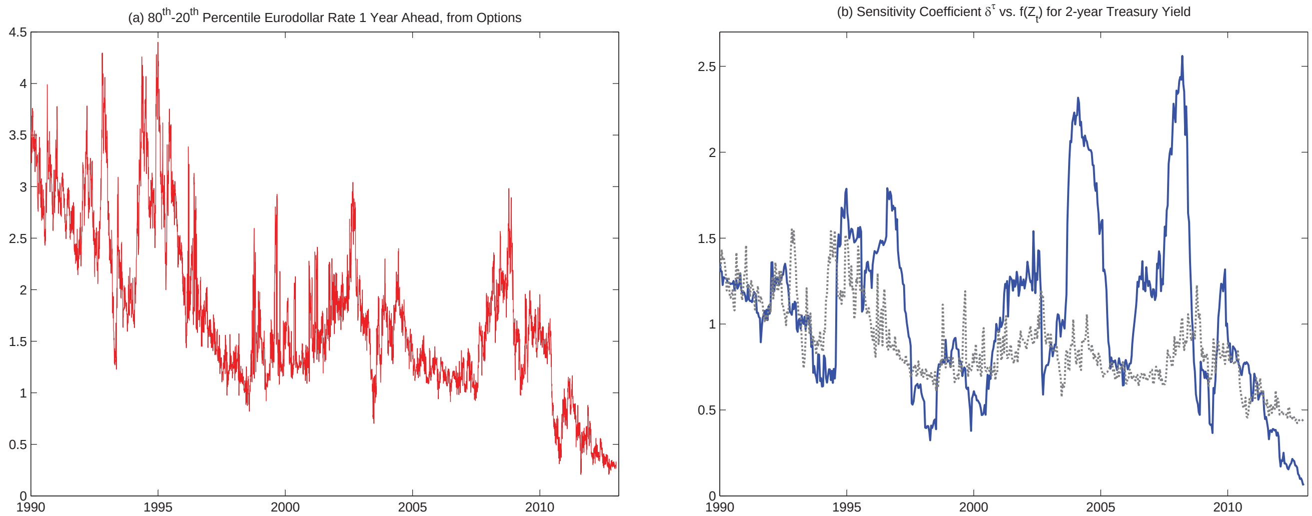

Figure 7. (a) Difference between 80th and 20th percentiles of the 1-year-ahead Eurodollar rate distribution, derived from Eurodollar options; (b) Solid blue line and dashed black line plot time-varying sensitivity coefficients δ^τ from regression (10) and $f(Z_t)$ from (12)–(13), respectively, for the 2-year Treasury yield. Estimates for $f(Z_t)$ are from column (2) of Table 3. See text and notes to Table 3 for details.

interquintile range is depicted in the first panel of Figure 7. The distance between the 80th and 20th percentiles of the one-year-ahead Eurodollar rate is about 3.5 percentage points at the beginning of 1990 and declines over time to about 25 bp by the end of 2012. This general downward trend in uncertainty (discussed in Swanson, 2006) is punctuated by increases in 1994 (when the Fed began to raise rates after the 1991 recession), 2002, 2004, and 2008–09.

Monetary policy uncertainty is significantly and positively related to the 3-month and 2-year Treasury yields' sensitivities to news in Table 3, as expected, although the coefficient for the 10-year yield is again insignificant. A decrease of 1 percentage point in the interquintile range is associated with a decrease in the 3-month Treasury's sensitivity to news of about 53 percent, relative to the benchmark sample.

The third column of Table 3 includes both the level of yields and monetary policy uncertainty as explanatory variables. In these regressions, the latter variable remains highly statistically significant while driving out the former, indicating that the level of yields is not highly correlated with sensitivity once the effect of monetary policy uncertainty is taken into account. The R^2 from these regressions are likewise no higher than those from the first two columns, suggesting little or no

interquintile range using the implied volatility on Eurodollar options with one year to expiration, computed by staff at the Federal Reserve Board assuming a lognormal distribution for the one-year-ahead spot Eurodollar rate.

additional gain to including interest rate levels in the regression once monetary policy uncertainty is included.[28]

Thus, an important takeaway from Table 3 is that bond yield sensitivity to news can vary for reasons other than the zero lower bound. It's even possible for our econometric test to reject normal sensitivity to news when the zero lower bound is not a factor (e.g., Swanson and Williams (2013b) find one such example of a "false positive" for long-term gilts in the U.K. in 2005). However, this does not reduce the validity of our test or imply that it is flawed. It just means that—as with any econometric test—we must inspect the results carefully and not leap to conclusions.

In this spirit, Figure 7(b) compares the estimated sensitivity coefficient δ^τ from regression (10) to $f(Z_t)$ from regression (12)–(13) for the 2-year Treasury yield, using the estimates from column (2) of Table 3 for $f(Z_t)$. The solid blue line in the figure plots δ^τ, while the dashed black line depicts $f(Z_t)$. As can be seen in the figure, monetary policy uncertainty is highly correlated with δ^τ in the 1990s, but has been much less so since 2000. For example, the very large spike in δ^τ in 2007–08 is only partly explained by the rise in monetary policy uncertainty during this period, suggesting that the extraordinary sensitivity of Treasury yields to news during that episode was attributable to other factors as well. More recently, the 2-year yield's sensitivity δ^τ falls much more in 2011–12 than would normally be attributable to the drop in monetary policy uncertainty alone. We interpret this last fact as reflecting the additional constraining effects of the zero lower bound and the Federal Reserve's forward guidance on the sensitivity of the 2-year Treasury yield to news.

Of course, we emphasize that the main reason monetary policy uncertainty fell to such low levels in 2010–12 in the first place is precisely because of the zero lower bound. In particular, the presence of the zero bound constraint was what led the FOMC to issue its "mid-2013", "late 2014", and "mid-2015" forward guidance, discussed above, which greatly reduced uncertainty about the near-term path of monetary policy. Thus, even the declines in Treasury yield sensitivity in 2011–12 that can be explained by decreases in monetary policy uncertainty can still be attributed, in a deeper sense, to the zero lower bound.

[28] We also experimented with including a time trend in these regressions, since the steady decline in monetary policy uncertainty over time in Figure 7(a) might be expected to have a different effect on yield curve sensitivity than variations in uncertainty due to the business cycle or other factors. However, the time trend had essentially no effect on the results.

5.5 Post-2007 Distribution of Macroeconomic Data Release Surprises

In our main empirical regressions (9) and (10), the surprise component of each data release in X_t can be regarded as strictly exogenous, under the assumption that our survey expectations data incorporate all relevant information as of the day before the release. (Under this assumption, the surprise component of each data release is independent of all past and future values of the interest rate changes on the left-hand side of these regressions.) To the extent that regressions (9) and (10) are correctly specified, strict exogeneity implies that the empirical distribution of the macroeconomic surprise data X_t is irrelevant for our estimates of the relative response coefficients β or time-varying sensitivity coefficients δ.

Nevertheless, one might be concerned that regression specifications (9) and (10) are simplifications that assume a linear structure with respect to X_t. As a result, it would be reassuring if the distribution of data surprises X_t in 2008–12 was not dramatically different from our benchmark sample 1990–2000, or the pre-crisis sample 1990–2007.

In fact, the distribution of these macro data surprises is similar across these samples. This can be seen in Figure 8, which plots the surprise component of nonfarm payrolls and core CPI announcements over the 1990–2007 and 2008–12 periods. Results for other macroeconomic data releases and the 1990–2000 period are similar. This finding might seem puzzling at first given the severity of the 2007–09 recession, but one should bear in mind that financial markets were quick to realize the severity of the downturn, so financial market expectations of the data fell about in line with the decline in the data itself. As a result, the surprises in the data releases, relative to the one-day-ahead expectations, do not look very different from earlier periods.

6 Conclusions

In this paper, we have developed a novel method to measure whether and to what extent interest rates of any maturity are affected by the presence of the zero lower bound. Our method provides both a quantitative measure of the severity of the zero bound constraint on each yield and a statistical test for the periods during which that yield was affected.

We find that interest rates with a year or more to maturity were surprisingly responsive to

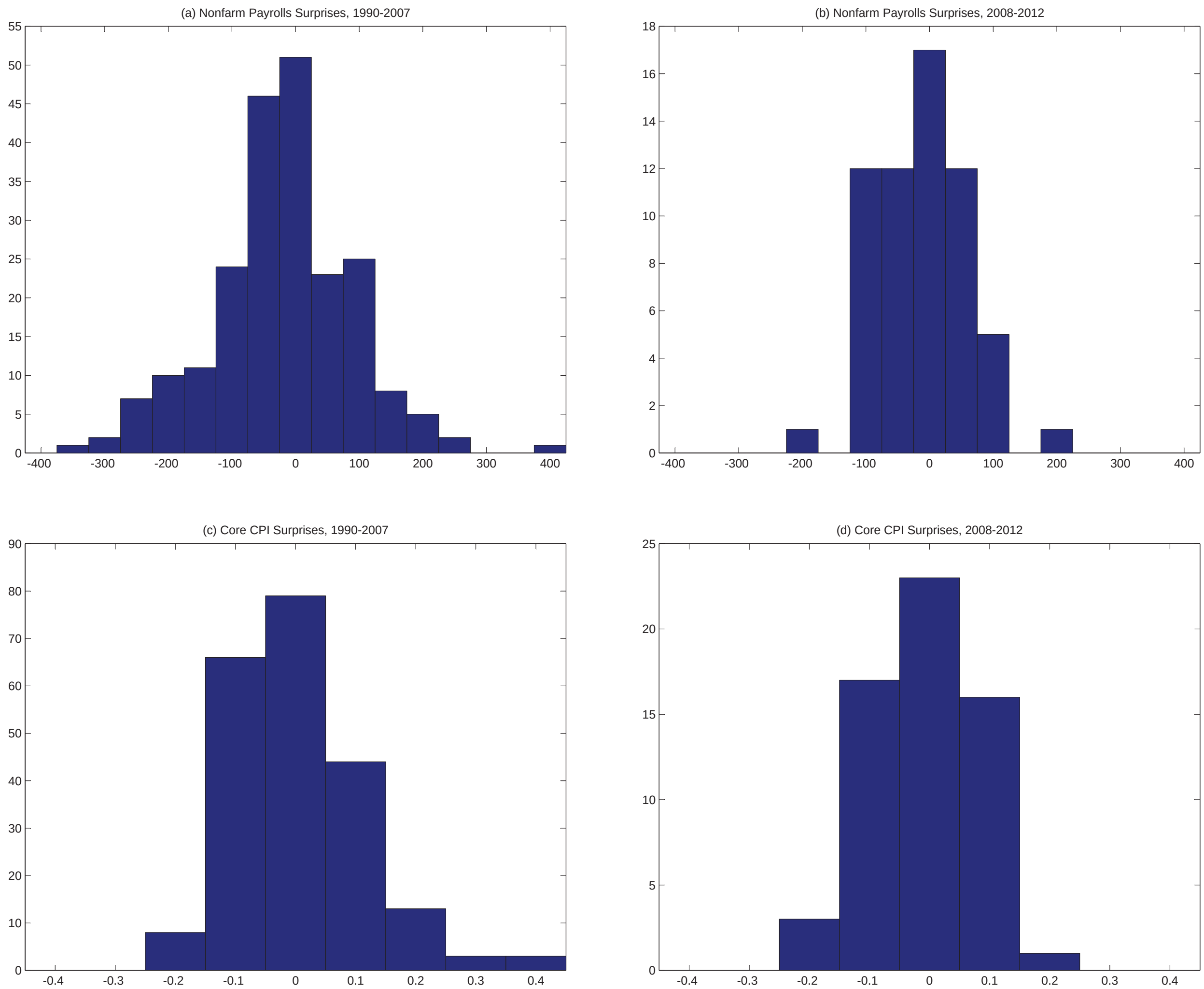

Figure 8. Top panels depict empirical distribution of the surprise component of nonfarm payrolls announcements from (a) 1990–2007 and (b) 2008–12, rounded to the nearest 50 thousand workers. Bottom panels depict the distribution of core CPI surprises from (c) 1990–2007 and (d) 2008–12, rounded to the nearest 0.1 percent. The surprise distributions of these and other macroeconomic data releases are relatively similar pre- and post-crisis. See text for details.

news throughout 2008–10. Only beginning in late 2011—around the time of the FOMC's "mid-2013" forward guidance—do we see the sensitivity of intermediate-maturity Treasury yields fall closer to zero. There appear to be two main explanations for this finding: First, up until late 2011, financial markets consistently expected the federal funds rate to lift off from zero within about four quarters. Second, the FOMC's forward guidance and large-scale asset purchases—and private-sector expectations of Federal Reserve policies along these lines—continued to move medium- and longer-term interest rates even when short-term rates were stuck at zero.

Our results have important implications for both monetary and fiscal policy. For monetary policy, our findings imply that policymakers had substantial room to affect medium- and longer-term interest rates until at least late 2011, despite the federal funds rate being at the zero lower bound. Indeed, on several occasions, the FOMC appears to have directly affected those longer-term yields by managing expectations of future monetary policy and conducting large-scale purchases of longer-term bonds.

For fiscal policy, taking the results in Christiano et al. (2011) as given, our findings suggest that the fiscal multiplier was probably close to normal throughout 2008–10, because financial markets at that time expected the zero bound constraint to last only about four quarters or less. Only beginning in late 2011, when the sensitivity of two-year Treasury yields to news was reduced—and the expected time until the first federal funds rate increase was lengthened to seven quarters or more—would our results suggest that the fiscal multiplier approached the larger values estimated by Christiano et al. (2012) and other authors.

More generally, the methods we have developed in the present paper can be extended beyond the United States and applied to any economy for which financial markets are sufficiently well developed. For example, Swanson and Williams (2013b) consider bond yields and exchange rates in the U.K. and Germany and find that: 1) USD/GBP and USD/EUR exchange rates have been essentially unaffected by the zero lower bound, 2) yields on German bunds were essentially unconstrained by the zero bound until late 2012, and 3) yields on U.K. gilts were substantially constrained by the zero bound in 2009 and 2012, but were surprisingly responsive to news in 2010–11. It would be very interesting to see these methods applied to other economies that have faced the zero lower bound in recent years, such as Japan, Canada, Sweden, and other members of the Euro area.

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