Measuring the Effect of the Zero Lower Bound on Medium- and Longer-Term Interest Rates

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Abstract

The zero lower bound on nominal interest rates has constrained the Federal Reserve's setting of the federal funds rate since December 2008. According to many macroeconomic models, this should have greatly reduced the effectiveness of monetary policy and increased the efficacy of fiscal policy. However, standard macroeconomic theory also implies that private-sector decisions depend on the entire *path* of expected future short-term interest rates, not just the current level of the federal funds rate. Thus, interest rates with a year or more to maturity are arguably more relevant for assessing the effects of the zero lower bound on the economy, and it is unclear to what extent the zero bound has constrained those yields. In this paper, we propose a novel approach to measure the effects of the zero lower bound on interest rates of any maturity. We compare the sensitivity of interest rates of various maturities to macroeconomic news during periods when short-term interest rates are very low to that during normal times. We find that yields on Treasury securities with six months or less to maturity were strongly affected by the zero bound during most or all of the period when the federal funds rate was near zero. In sharp contrast, yields with more than two years to maturity continued to respond to economic news in their usual way. One- and two-year Treasury yields represent an intermediate case, being only partially attenuated by the zero bound over part of the period when the funds rate was near zero. We discuss the implications of these findings for monetary and fiscal policy. Finally, we offer two explanations for our results: First, market participants often expected the zero bound to constrain policy for only a few quarters, minimizing its effects. Second, the Fed's unconventional policy actions may have helped offset the effects of the zero bound on medium- and longer-term rates.

Keywords: monetary policy, zero lower bound, fiscal policy, fiscal multiplier

JEL Classification: E43, E52, E62

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1 Introduction

The zero lower bound on nominal interest rates has clearly constrained the federal funds rate—the Federal Reserve's traditional monetary policy instrument—since December 2008, when it was lowered to a floor of essentially zero. However, standard textbook macroeconomic models (e.g., Clarida, Galí, and Gertler 1999, Woodford 2003) imply that monetary policy affects the economy through its effects on *longer*-term interest rates, not just the current level of the overnight rate, and it is not clear whether the zero lower bound has substantially constrained the Fed's ability to affect these longer-term interest rates. Theoretically, if a central bank has the ability to commit to future values of the 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). Empirically, Gürkaynak, Sack, and Swanson (2005b) found that the effects of Federal Reserve policy announcements on asset prices are driven primarily by the Federal Open Market Committee (FOMC) statement's effects on financial market expectations of *future* monetary policy, rather than changes in the current federal funds rate target. And Figure 1 shows that yields on U.S. Treasury securities with a year or more to maturity remained well above zero throughout much of 2008–11, suggesting that those yields were not *directly* constrained by the zero bound during this period. Indeed, on several occasions, the Federal Reserve was able to generate a decline in medium- and longer-term Treasury yields of as much as 20 basis points by managing monetary policy expectations and by asset purchases.[1] These actions and their effects suggest that monetary policy still had room to affect the economy, despite the constraint on the current level of the federal funds rate.

In addition to monetary policy considerations, the extent to which the zero lower bound affects interest rates of different maturities has important implications for fiscal policy. Numerous authors have emphasized that the macroeconomic effects of fiscal policy are much larger when the zero lower bound is binding, because in that case interest rates do not rise in response to higher output, and private investment is not "crowded out" (e.g., Christiano, Eichenbaum, and Rebelo 2011, Erceg

[1]For example, on August 9, 2011, the FOMC stated, "The Committee currently anticipates that economic conditions... are likely to warrant exceptionally low levels for the federal funds rate at least through mid-2013." In response to this announcement, the 2-year Treasury yield fell 8 basis points (bp), while the 5- and 10-year Treasury yields each fell 20 bp. Note that, in normal times, it would take a surprise change in the federal funds rate of about 100 bp to generate a fall of 8 to 20 bp in intermediate-maturity yields (Gürkaynak et al. 2005b).

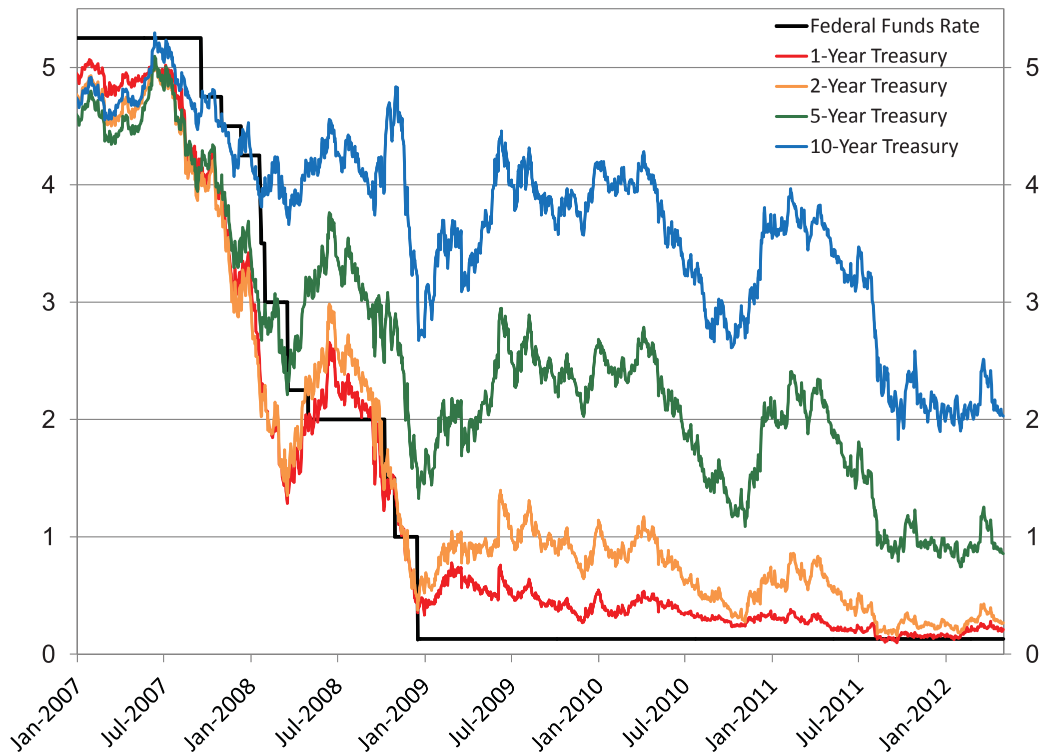

Figure 1. Federal funds rate target and 1-, 2-, 5-, and 10-year zero-coupon Treasury yields from January 2007 through April 2012. (Data are from the Gürkaynak, Sack, and Wright (2007) online dataset.)

and Lindé 2010, Eggertsson and Krugman 2011, DeLong and Summers 2012). However, standard macroeconomic theory implies that private-sector spending is determined by the expected future path of short-term interest rates and not just the current level of the overnight rate. As emphasized by Christiano et al. (2011) and Erceg and Lindé (2010), the fact that the one-period interest rate is at the zero lower bound is largely irrelevant for measuring the fiscal multiplier. Instead, the extent to which private spending is crowded out by fiscal stimulus depends on the *length of time* that monetary policy is constrained by the zero lower bound. In other words, the fiscal multiplier depends on the extent to which medium- and longer-term yields are affected by the zero bound.

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 those yields were unconstrained by the zero bound (taken to be 1990–2000). In periods in which the yield is about as sensitive to news as the benchmark sample, we say that the yield is unaffected by the zero bound. In periods when the yield responds very little or not at all to news, we say that the yield is completely constrained. Intermediate cases are measured by the degree of the yield's sensitivity to news relative to the benchmark period, and the severity and statistical significance of the effects can be assessed using standard econometric techniques. To our knowledge, this is the first empirical study to shed light on how relevant the zero bound constraint has been for intermediate- and longer-maturity yields, and thus to what extent the zero bound has hindered the effectiveness of monetary policy and amplified the effectiveness of fiscal policy.

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 basis points (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, and this "effective" lower bound may vary across countries or over time.[2] For example, the Federal Reserve has held the target federal funds rate at a floor of 0 to 25 bp from December 2008 through at least April 2012, 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 nominal rates 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.[3] 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

[2] See Bernanke and Reinhart (2004) for a discussion of the institutional barriers to lowering the policy rate all the way to zero.

[3] As another example, in 2003, the Federal Reserve lowered the funds rate to 1 percent, at which point it began to use forward guidance, such as the phrase "policy accommodation can be maintained for a considerable period," to try to lower longer-term interest rates without cutting the funds rate any further (see, e.g., Bernanke and Reinhart 2004; the quotation is from the FOMC statement dated August 12, 2003). Thus, one can make a good case that the effective lower bound on the funds rate in 2003–04 was 100 bp rather than zero.

bounds that may be greater than zero or change over time.

Third, the sensitivity of interest rates to news is more relevant than the level of yields for the fiscal multiplier. As discussed in Christiano et al. (2011) and others, what is crucial for the fiscal multiplier is whether or not 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. If we estimate that the interest rates relevant for private-sector spending are half as sensitive to news as in "normal" times, that would suggest a fiscal multiplier that is roughly halfway between the normal value of 1 or a little less estimated, e.g., by Christiano et al. (2011) and the much larger, constant-interest-rate multiplier estimated by those same authors.

To preview our results, we find that Treasury yields with six months or less to maturity were severely constrained by the zero lower bound from the spring of 2009 through April 2012 (the end of our sample). In contrast, Treasury yields with more than two years to maturity were substantially unaffected by the zero bound over this same period. One- and two-year Treasury yields represent an intermediate case, with the one-year yield's sensitivity to news showing attenuation beginning in 2010, and the two-year yield's sensitivity showing attenuation beginning only in late 2011. Importantly, our method provides a quantitative measure of the *degree* to which the zero bound affects each yield, as well as the periods during which it was affected.

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

The remainder of the paper proceeds as follows. Section 2 lays out a simple New Keynesian model that illustrates three important points we will use in our empirical analysis. Section 3 describes our empirical framework. Our main results are reported in Section 4. Section 5 considers various extensions, additional results, and discussion related to our main findings. Section 6 concludes.

2 An Illustrative Model

In this section, we use a simple macroeconomic model to illustrate the effects of the zero lower bound on the responsiveness of yields to economic news. In particular, we use the model to illustrate three important points that we will employ in our empirical analysis, below. First, we show 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 responsiveness of yields to news is greatest at short maturities and is smaller for longer-term yields. Second, the effects of the zero bound on the sensitivity of yields to news is essentially *symmetric*—that is, the responsiveness of yields to both positive and negative announcements falls by about the same amount when the zero bound is strongly binding on short-term rates. Third, the zero bound dampens the sensitivity of yields to news by similar amounts for different types of shocks, so long as the persistence of those shocks are not too different. Readers who are willing to take these three points for granted can skip ahead to the next section.

We conduct the analysis in this section using a standard, simple, three-equation New Keynesian model (cf. Clarida, Galí, and Gertler (1999) and Woodford (2003), among others) that describes the evolution over time t of the output gap, y_t, inflation rate, π_t, and one-period risk-free nominal interest rate, i_t. The purpose of this exercise is to illustrate qualitatively how the zero lower bound affects the sensitivity of bond yields to news, so the model is deliberately simplistic and not intended to capture the quantitative effects we estimate below.

The model's output gap equation is derived from the household's consumption Euler equation, and relates the output gap this period to the expected output gap next period and the difference between the current ex ante real interest rate, $i_t - E_t\pi_{t+1}$, and natural rate of interest, r_t^*:

$$y_t = -\alpha(i_t - E_t\pi_{t+1} - r_t^*) + E_t y_{t+1}. \tag{1}$$

Solving this equation forward, assuming $\lim_{k\to\infty} E_t y_{t+k} = 0$, we have:

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

This equation makes clear that the current level of the output gap depends on the entire expected future *path* of short-term interest rates, inflation, and the natural rate of interest. As emphasized by Woodford (2003) and Erceg et al. (2000), among others, the quantity $E_t \sum_{j=0}^{\infty} i_{t+j}$ can be interpreted as a nominal long-term interest rate in the model.

We model shocks to the output gap as shocks to the natural interest rate, r_t^*. We assume that the natural interest rate follows a stationary AR(1) process,

$$r_t^* \;=\; (1-\rho)\bar{r}^* + \rho r_{t-1}^* + e_t, \tag{3}$$

where $\rho \in (-1, 1)$ and $\bar{r}^*$ denotes the unconditional mean of the natural rate.

The equation for inflation is derived from profit maximization by monopolistically competitive firms with Calvo price contracts, and is given by:

$$\pi_t \;=\; \gamma y_t + \beta E_t \pi_{t+1} + \mu_t, \tag{4}$$

where μ_t can be thought of as a markup shock, assumed to follow a stationary AR(1) process:

$$\mu_t \;=\; \delta \mu_{t-1} + v_t, \tag{5}$$

where $\delta \in (-1, 1)$.

The one-period interest rate in the model is set according to a Taylor (1993) Rule, subject to the constraint that i_t must be nonnegative:

$$i_t \;=\; \max\{\,0,\; \pi_t + r_t^* + 0.5(\pi_t - \bar{\pi}) + 0.5 y_t\}, \tag{6}$$

where $\bar{\pi}$ denotes the central bank's inflation target, taken to be 2 percent. 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 shocks to the natural interest rate 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)–(5), we assume that long-term bond yields in the model are determined by the expectations hypothesis. Thus, the M-period yield to maturity on a zero-coupon nominal bond, i_t^M, is given by:

$$i_t^M \;=\; E_t \sum_{j=0}^{M-1} i_{t+j} \;+\; \phi^M, \tag{7}$$

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

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 lower bound is not a binding constraint on the setting of 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 the model using a nonlinear perfect foresight algorithm, as in Reifschneider and Williams (2000), which solves for the impulse response functions of the model to an output or inflation shock under the assumption that the private sector assumes that realized values of all future innovations will be zero. In each scenario, impulse 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.

We set the model parameters $\alpha = 1.59$ and $\gamma = 0.096$, based on Woodford (2003), and choose illustrative values for the shock persistences of $\rho = 0.85$ and $\delta = 0.5$. We calibrate the magnitude of the shocks to r_t^* and μ_t so that they each generate a 5 basis point 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 news surprise typically moves shorter-term yields by only a few basis points.[4]

The top panels of Figure 2 report the impulse response functions of the one-period nominal interest rate, i_t, to a shock to output and to inflation, achieved through shocks to r_t^* and μ_t, respectively. In each of these panels, the solid black line depicts the impulse response function to a positive shock to output or inflation in the case where the zero lower bound is not binding—i.e., the standard impulse response function to an output or inflation shock in a textbook New Keynesian model. The dashed red line in each panel depicts the impulse response function for i_t to the same shock when the zero lower bound is binding—that is, when the initial value of r_t^* is set equal to -4 percent. In each panel, the dashed red impulse responses are computed relative to a baseline in which r_t^* begins at -4 percent but is returning toward $\bar{r}^*$, so that the zero bound ceases to bind the short-term interest rate i_t in the fourth period.

[4]See our empirical results, below, for discussion. Gürkaynak, Sack, and Swanson (2005a) and Gürkaynak, Levin, and Swanson (2010) also discuss this fact.

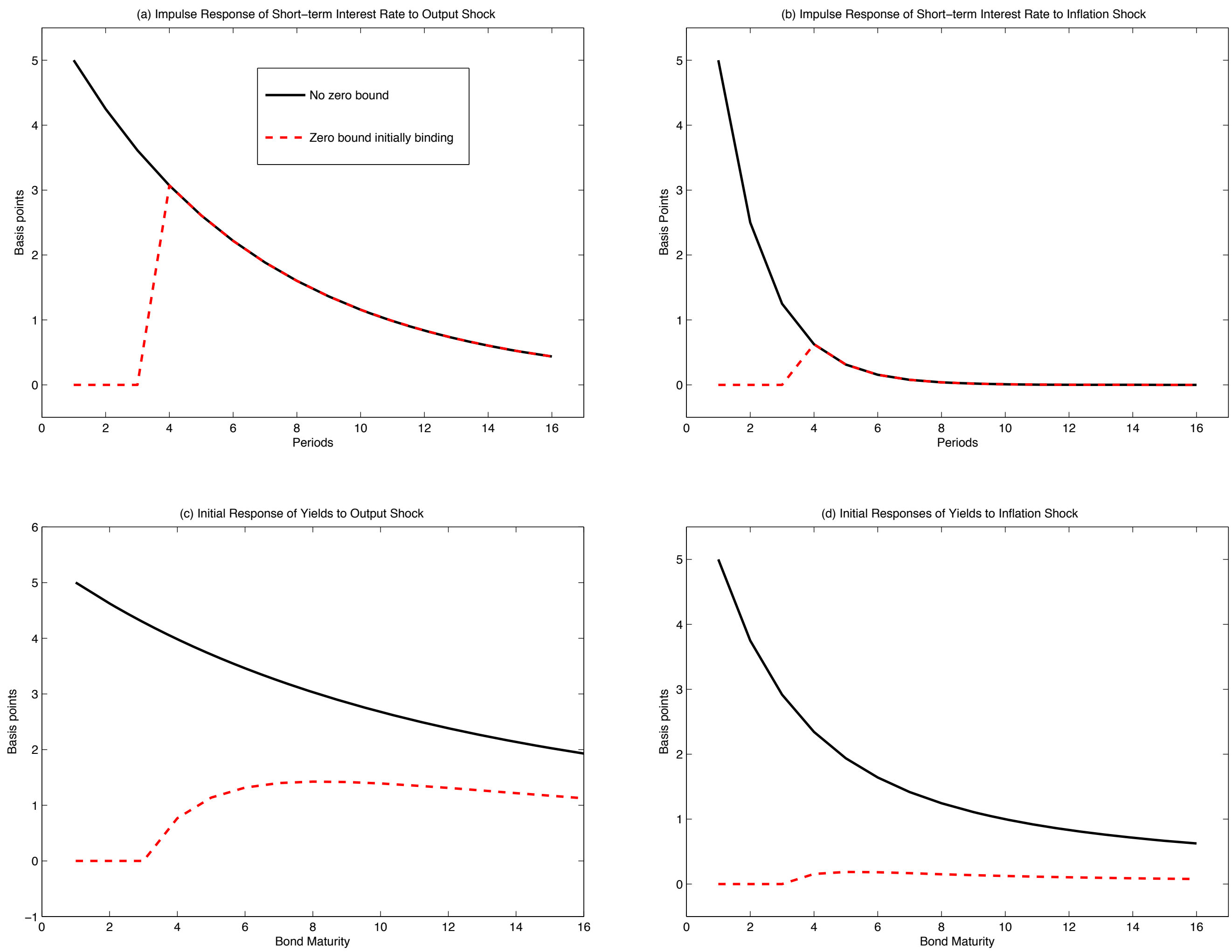

Figure 2. Response of short-term interest rate and the yield curve to output and inflation shocks in a simple New Keynesian model, with and without the zero bound constraint on monetary policy. Shocks are normalized to produce a 5 basis point effect on the one-period nominal interest rate on impact. (a) impulse response function of one-period interest rate to an output shock and (b) an inflation shock. (c) initial response in period 1 of the yield curve to an output shock and (d) an inflation shock; x-axis in the bottom panels denotes bond yield maturity rather than periods after the shock. See text for details.

Note that, once the zero bound ceases to bind in Figure 2, the behavior of the interest rate i_t is identical to what would occur absent the zero bound—that is, the red and black lines in the top panels 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, in turn, lead to a more persistent difference in the path of interest rates.

The bottom panels of Figure 2 depict the responses on impact of the yield curve to an output or inflation shock in period 1, the period when the shock hits. Thus, the bottom panels of Figure 2 are not impulse response functions, but rather plot the instantaneous response of the entire yield curve at a single point in time.

The first main point to take away from the model is that the response of the yield curve to shocks is attenuated when the zero bound constrains policy, and the degree of attenuation declines with the maturity of the bond. This can be seen clearly in the bottom panels of Figure 2. For the shortest maturities, there is a total lack of responsiveness of the yield curve to an output or inflation shock when the zero bound is binding, whereas for the longest-maturity bonds, the response of the yield curve to an output or inflation shock becomes closer to the normal, unconstrained response. Intermediate-maturity bonds are constrained by the zero bound to an intermediate extent. The intuition for these results is clear and holds more generally than the simple illustrative model of this section.

The second point to take away from the model is that the responses of yields to shocks are essentially symmetric to positive and negative shocks. Figure 2 plots the response of the model to small positive shocks, but the results for small negative shocks of the same size are exactly the same in absolute value. This symmetry holds perfectly as long as the number of periods that policy is constrained at the zero bound does not change, which is the case for small shocks.

Even for larger shocks, the responses of yields are essentially symmetric. Figure 3 plots the absolute value of the impulse responses of the model to a positive (dashed red line) and negative (solid blue line) output shock that are each ten times larger than in Figure 2 for the case where the zero bound is binding. These are truly gigantic shocks, relative to the typical macroeconomic data release surprise in our sample. Yet the two lines in the first panel of Figure 3 are still almost identical, except that the dashed red line lifts off from the zero bound one period sooner than the blue line, because the positive shock increases policymakers' desired interest rate above the zero bound in that period. When the zero bound ceases to bind in either model in period 4, both lines are identical for the same reasons as in Figure 2. The second panel of Figure 3 reports the corresponding absolute value response of the yield curve on impact.

The fact that the zero bound causes the yield curve to be damped almost symmetrically to positive and negative announcements can be counterintuitive at first, since the zero bound is a

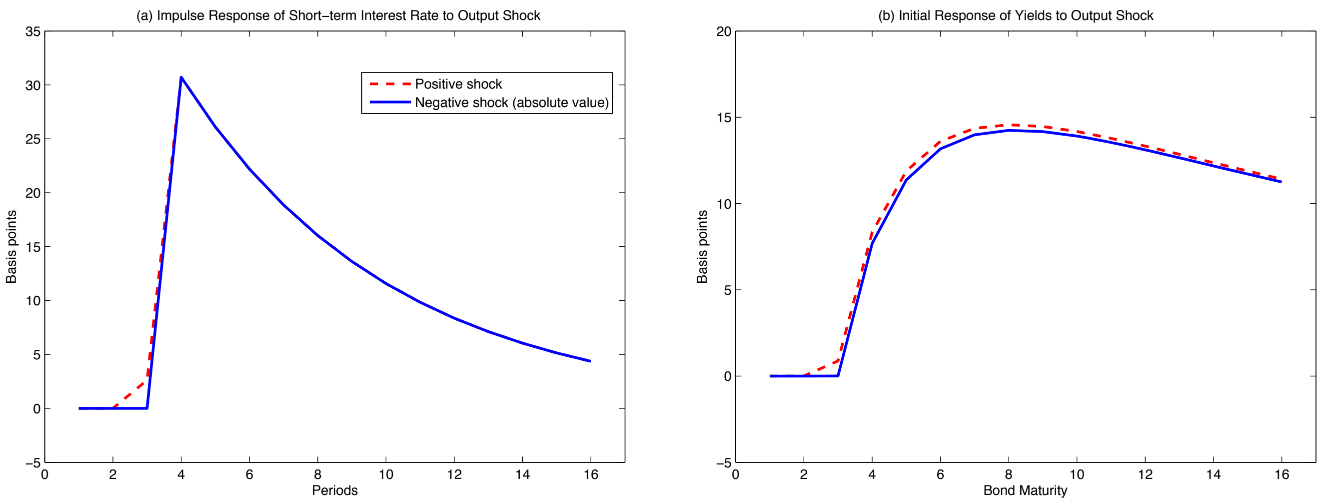

Figure 3. Absolute value of responses of short-term interest rate and yield curve to large positive and negative output shocks in the simple New Keynesian model when the zero bound is binding. Shocks are normalized to produce a 50 bp effect on the one-period nominal interest rate on impact, ten times as large as in Figure 2. (a) impulse response function of one-period interest rate to the output shock; (b) initial response of the yield curve to an output shock. See text for details.

one-sided constraint. Nevertheless, the intuition is clear and holds much more generally than in just the simple model of this section: When the zero bound is a severe constraint on policy—that is, policymakers would like to set the one-period nominal interest rate far 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 when the zero bound is binding, 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 is typically 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 third and final point to take away from the model is that the dampening effects of the zero bound on the sensitivity of yields is qualitatively the same regardless of the specific nature of the shock. However, the quantitative effects of the zero bound do depend on the persistence of the shock. That is, the dampening of the sensitivity of yields is greater for short-lived shocks that have their main effects on the short-term interest rate during the period that the zero bound is a binding constraint, as in the case of the inflation shock in Figure 2. The more persistent output shock continues to have large effects on interest rates in periods when the zero bound is not binding. As a result, there is less dampening of the sensitivity of longer-term yields in response to those shocks.

Note that this difference across the two shocks is not because of the type of shock, but because of the difference in persistence. If the degree of persistence in the two shock processes were the same, then the attenuation across maturities would be essentially identical for the output and inflation shocks. In models with more complicated dynamics, the effects of the zero bound would differ more substantially across the two types of shocks, but even in those models it remains true that the degree of attenuation across maturities is determined primarily by the length of time the zero lower bound is expected to bind, and not by the type of shock.

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 our theoretical model, this would only be exactly true if all of the shocks have identical persistence characteristics in terms of their effects on the short-term interest rate. Empirically, these persistences are unlikely to be exactly the same, but we view this assumption as a reasonable approximation that can be tested, which we do 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.[5] 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. (2005a), 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. 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

[5] Kuttner (2001) tests and confirms this hypothesis for the case of monetary policy announcements.

regarding these data are provided in Gürkaynak, Sack, and Swanson (2005a), Gürkaynak, Levin, and Swanson (2010), and in Section 5.5, below.

3.2 The Sensitivity of Treasury Yields to Macro 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. (2005a) 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.[6]

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.[7]

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

[6]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.

[7]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	1.66	(2.90)	2.08	(4.23)	1.47	(2.52)
Consumer Confidence	0.29	(0.58)	2.68	(5.83)	2.70	(5.39)
Core CPI	0.81	(2.57)	2.35	(4.39)	1.72	(3.38)
GDP (advance)	0.34	(0.72)	−0.19	(−0.19)	−0.68	(−0.62)
Initial Claims	−0.29	(−1.36)	−0.64	(−2.42)	−0.40	(−1.44)
ISM Manufacturing	0.99	(1.47)	3.46	(7.23)	2.63	(4.98)
Leading Indicators	0.81	(1.56)	1.19	(2.35)	0.69	(1.09)
New Home Sales	1.48	(3.57)	2.01	(4.84)	2.07	(4.30)
Nonfarm Payrolls	2.46	(4.40)	4.61	(7.02)	2.90	(4.03)
Core PPI	0.52	(1.40)	0.88	(1.78)	1.34	(2.61)
Retail Sales ex. autos	1.20	(3.35)	1.84	(2.84)	1.19	(1.83)
Unemployment rate	−1.53	(−2.19)	−1.97	(−2.59)	−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 nonlinear 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 errors; 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.

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 3-month Treasury yields is nonfarm payrolls, for which a one-standard-deviation surprise causes yields 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. Thus, the high-frequency regressions in Table 1 provide powerful, highly statistically significant measures of the sensitivity of these different Treasury yields to news.[8]

[8] 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 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

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 more structural specifications for the time-varying sensitivity coefficients δ in Section 5, 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 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

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.

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 April 2012.[9] When τ corresponds to the midpoint of a given calendar year $i \in \{1990, 1991, \ldots, 2011\}$, 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 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).[10]

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 April 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

[9] 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 April 2012.

[10] From July 2011 through April 2012, we also perform this interpolation using the estimated $\delta^{\tau_{2012}}$, recognizing that the rolling δ^τ do not exactly equal $\delta^{\tau_{2012}}$ as τ approaches April 2012.

	Treasury yield maturity					
	3-month		2-year		10-year	
Capacity Utilization	0.72	(1.52)	1.48	(2.89)	0.83	(2.48)
Consumer Confidence	0.76	(2.90)	1.37	(3.72)	0.88	(2.50)
Core CPI	0.40	(1.91)	1.91	(5.01)	1.27	(3.82)
GDP (advance)	0.93	(3.17)	1.44	(2.41)	0.98	(1.70)
Initial Claims	−0.30	(−1.81)	−1.10	(−5.35)	−0.98	(−5.08)
ISM Manufacturing	1.24	(3.23)	2.74	(7.09)	2.02	(5.97)
Leading Indicators	0.20	(0.61)	0.28	(0.86)	0.26	(0.91)
New Home Sales	0.84	(2.63)	0.66	(1.99)	0.52	(1.96)
Nonfarm Payrolls	3.06	(7.67)	4.84	(9.55)	2.96	(6.73)
Core PPI	0.22	(0.80)	0.53	(1.54)	0.87	(3.13)
Retail Sales ex. autos	0.84	(3.77)	1.87	(4.91)	1.60	(4.18)
Unemployment rate	−1.23	(−3.51)	−1.26	(−2.77)	−0.35	(−0.88)
# Observations	2747		2747		2747	
R^2	.08		.17		.10	
$H_0 : \beta$ constant, p-value	1.000		1.000		1.000	
$H_0 : \delta$ symmetric, p-value	.089		.280		.365	
$H_0 : \delta$ constant, p-value	$< 10^{-16}$		$< 10^{-10}$		.016	

Table 2. Coefficient estimates β for nonlinear regression $\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i}\beta X_t + \varepsilon_t$ at daily frequency from Jan. 1990 to Apr. 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.

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, there is essentially no loss in fit from using (9) rather than (11), relative to the degrees of freedom of the restriction: the p-values are equal to 1 to at least three decimal places. The assumption of a constant β in (9) is thus very consistent with the data.

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 $i = 1990, \ldots, 2012$.[11] 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 .089, that result is entirely driven by a large outlier on October 20, 2008, when the 3-month T-bill yield jumped by 41 bp.[12] Excluding that single observation, the p-value for the 3-month Treasury yield hypothesis test is .699. 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} and 10^{-10}, respectively. 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 .016. Although the 10-year yield's sensitivity to macroeconomic 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 4 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 ± 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

[11] 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.

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

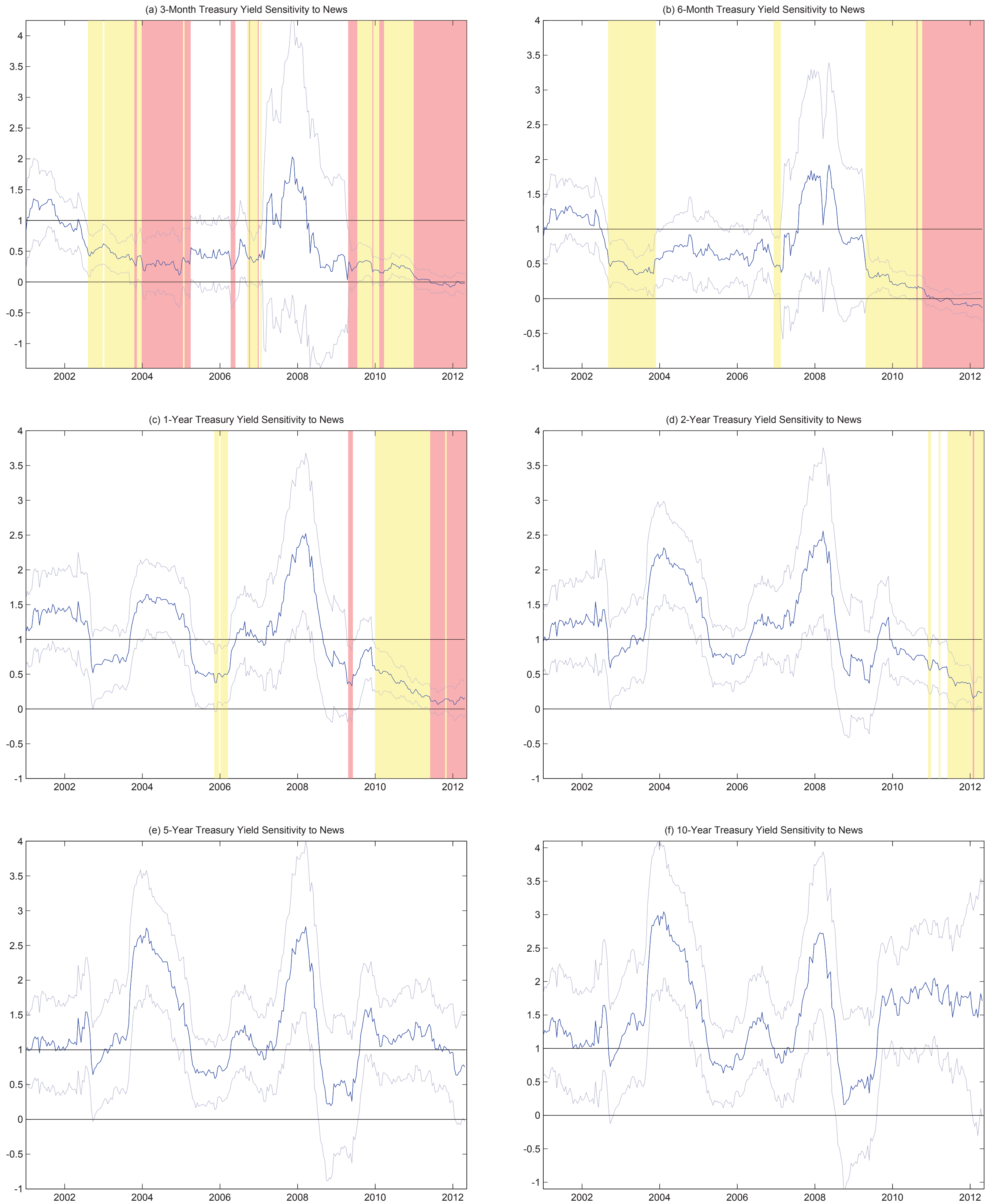

Figure 4. Time-varying sensitivity coefficients δ^τ from regression (10) for (a) 3-month, (b) 6-month, (c) 1-year, (d) 2-year, (e) 5-year, and (f) 10-year Treasury yields. Dotted gray lines depict heteroskedasticity-consistent $\pm$2-standard-error bands, adjusted for two-stage sampling uncertainty in (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.

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.[13] 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 4 shows that the sensitivity of the 3-month Treasury yield to macroeconomic news has varied between about 0 and 2 from 2001 through 2012. From the spring of 2009 through the end of our sample in April 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 might appear more surprising in the first panel of Figure 4 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."[14] 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,

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

[14] 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.

is a noteworthy feature of our approach.

Panel (b) of Figure 4 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 April 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 4. The 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. To the extent that crowding out is determined by yields with a year or more to maturity, we would thus conclude that crowding out in 2003–04 would have been no different from normal. Similarly, to the extent that the Fed can influence monetary policy expectations at a horizon of a year or more, we would conclude that monetary policy was relatively unconstrained by the zero bound during 2003–04.

What is perhaps more surprising in the middle panels of Figure 4 is how little and how late the zero bound seems to have affected these intermediate-maturity yields between December 2008 and April 2012. 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 mid-2011. Only in late 2011 and 2012 does the 1-year Treasury yield cease responding to news. The sensitivity of the 2-year yield to news was not significantly less than unity until 2011, and even through 2012 remains partially responsive to news. To the extent that the Fed can influence monetary policy expectations at a horizon of two years or more, we would thus conclude that monetary policy was not substantially constrained by the zero bound until 2011, and even through April 2012 had substantial ability to influence 2-year Treasury yields. Similarly, to the extent that crowding out is related to intermediate-maturity yields such as the 2-year Treasury, we find little evidence that crowding out in 2008–10 would have been any less than normal. (Although in 2011–12, our estimates imply that crowding out would have been about half as large as normal, based on the 2-year Treasury yield's sensitivity to news.)

The last two panels of Figure 4 report results for yields with maturities of five and ten years, which are also remarkable. There are no red or yellow shaded regions in these panels, because nowhere in the sample is their sensitivity to news significantly less than unity. Thus, even in 2012, when Treasury yields out to two years are at least partially attenuated by the zero lower bound, the 5- and 10-year yields appear to remain unconstrained.

5 Discussion

The results of the previous section raise several important questions that we now investigate in greater detail. First, we check the robustness of our results by comparing them to private-sector expectations of the time of federal funds rate "liftoff" from the zero bound. Second, we provide evidence that the Federal Reserve can in fact manage monetary policy expectations at horizons out to several quarters. Third, we explore the implications of the Federal Reserve's large-scale purchases of long-term bonds for our results. Fourth, we investigate why the sensitivity of interest rates to news sometimes exceeds unity, and why that sensitivity varies over time even when monetary policy is far away from the zero bound. And finally, we show that the distribution of our macroeconomic surprise data from 2008–12 is not very different from the distribution of those surprises before the financial crisis and recession.

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 the zero bound for only a single quarter, then medium- and longer-term interest rates should be nearly unaffected by the zero bound, but if the funds rate is expected to be at the zero bound for several years, then even the 10-year Treasury yield should be noticeably affected.

Figure 5 plots the number of quarters that the private sector expected the funds rate to remain

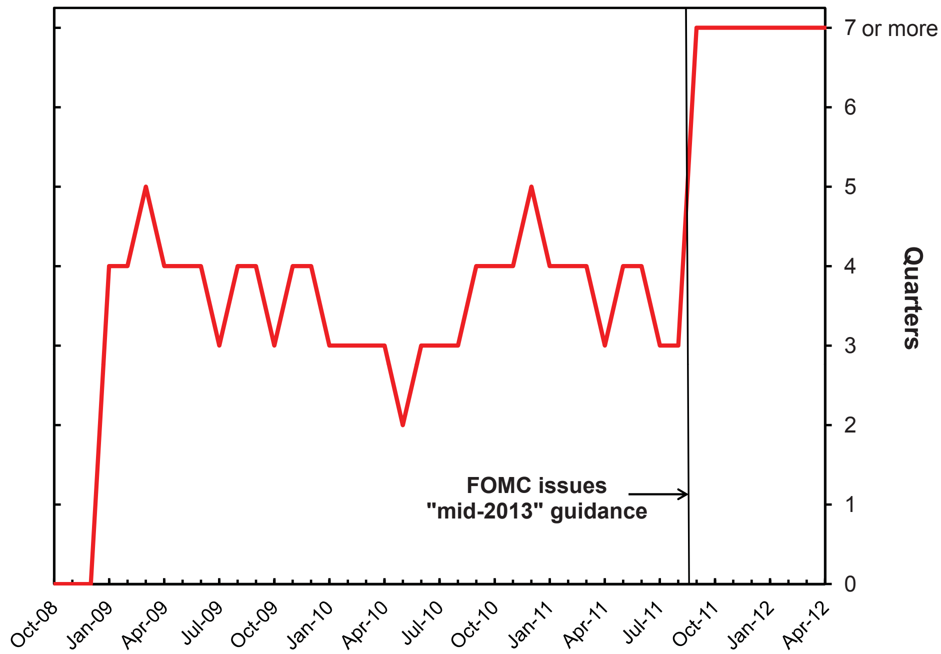

Figure 5. Expected number of quarters until the first federal funds rate increase above 25 bp, from the monthly Blue Chip survey of forecasters.

below 25 bp, 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 for any length of time. After the FOMC cut the target funds rate to near 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. After the FOMC announced on August 9, 2011, that it expected to keep the funds rate at zero until at least "mid-2013," private-sector expectations of the time until liftoff jumped to seven or more quarters (the Blue Chip forecast horizon extends forward only six quarters).

The implication of the forecasts underlying Figure 5 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 a more substantial attenuation with respect to news. And in fact, this corresponds closely to our time-varying sensitivity results in Figure 4.

Figure 6 provides an additional perspective on these results from the interest rate options

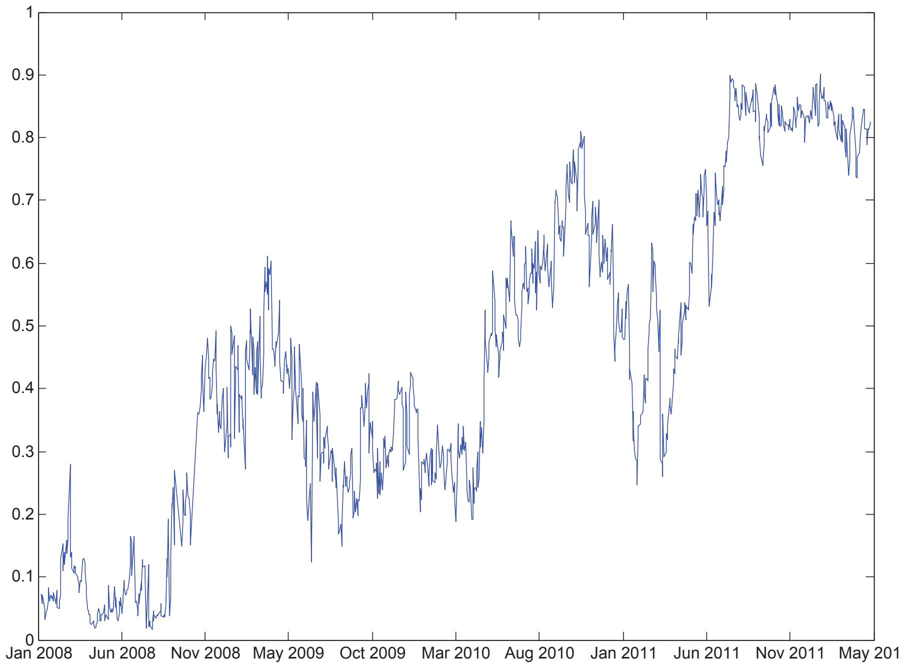

Figure 6. Probability the federal funds rate would be less than 50 bp in five quarters, estimated from options data. See text for details.

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.[15] 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 6 from January 2008 to April 2012. This probability was generally less than ten percent from January through October 2008, and rises sporadically over time to about 85 percent from September 2011 through April 2012.

The implied probabilities in Figure 6 corroborate the survey results in Figure 5 and our sensitivity estimates in Figure 4. 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

[15]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 April 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 very flexible functional form. Eurodollar options are the most liquid options available on a short-term interest rate and thus provide us with 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.

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 around a central tendency of about 50 to 60 percent. Given the relatively low level and substantial movements in these option-implied probabilities of being at the zero bound, it is not surprising that 2-year or even 1-year Treasury yields would respond to news almost normally throughout much of this period. Only beginning in late 2011 do we see the probabilities in Figure 6 increase to around 85 percent, corresponding to a more reduced sensitivity of the 2-year Treasury yield to news.

Figure 7 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.[16] Thus, a Eurodollar future with one quarter to expiration is closely related to market expectations about the federal funds rate from 3 to 6 months ahead, and a Eurodollar future with 3 quarters to expiration closely reflects market expectations about monetary policy from 9 to 12 months ahead.

The results in Figure 7 confirm those for Treasury yields in Figure 4. 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 almost zero in 2010–12. And just like the 1-year Treasury yield, the sensitivity to news of Eurodollar futures with 2–4 quarters to expiration remained close to normal all the way up until 2010, at which point they fell.

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 in August that it expected to keep the funds rate near zero until "mid-2013". This mirrors very closely the behavior of the 2-year Treasury yield in Figure 4. Even longer-maturity Eurodollar futures, with 5 to 7 quarters to expiration, continued to respond normally to news until the FOMC announced in

[16]See Gürkaynak, Sack, and Swanson (2007) for additional details regarding Eurodollar futures. They compare the forecasting performance of Eurodollar futures for the federal funds rate to other market-based measures of monetary policy expectations and econometric forecasts. They find that Eurodollar futures perform as well as or better than any other measure at horizons of six months or more, the horizon which is most interesting for our present analysis. Also note that Figure 7 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.

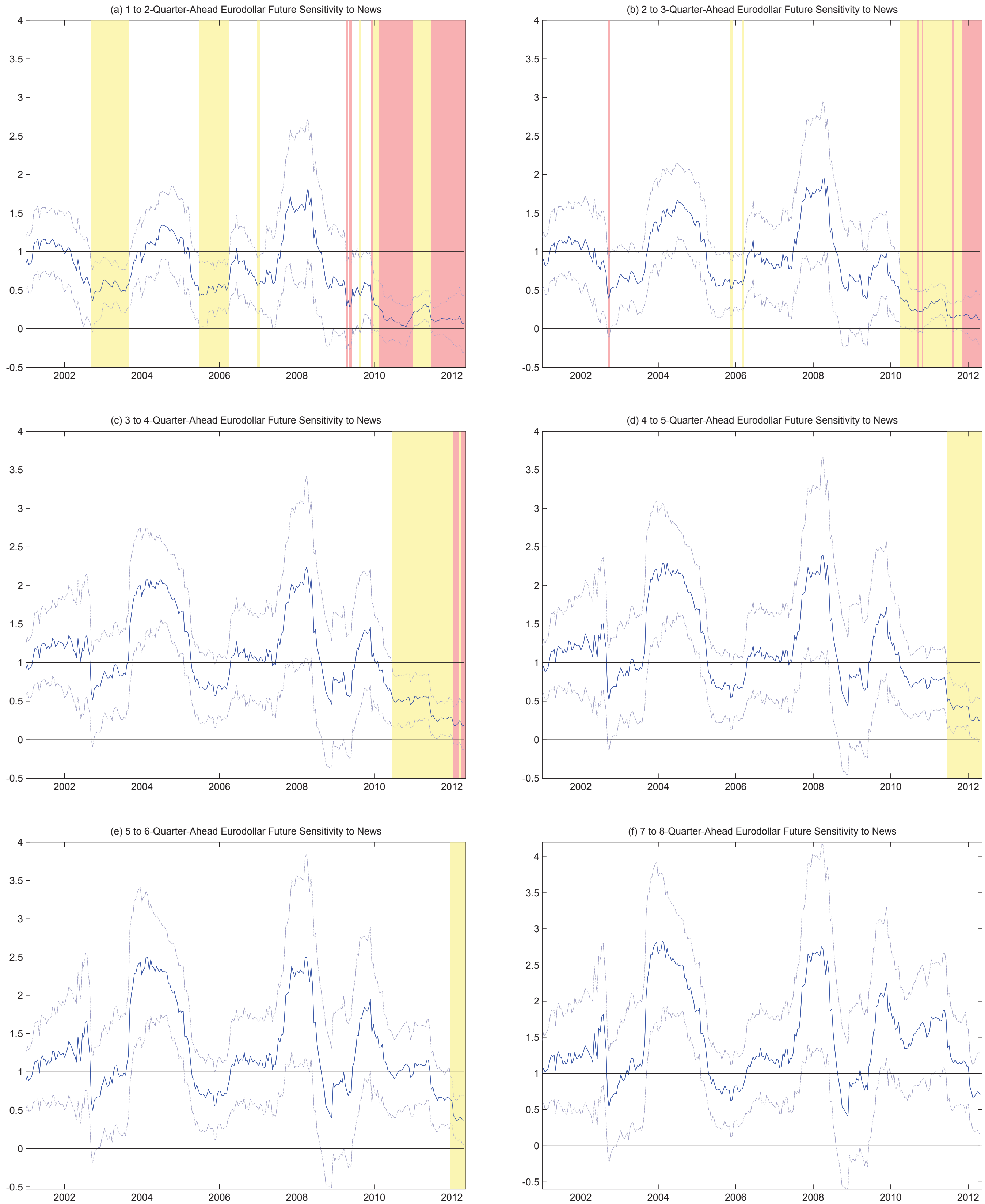

Figure 7. Time-varying sensitivity coefficients δ^{τ} from regression (10) for Eurodollar futures contracts with (a) 1–2 quarters, (b) 2–3 quarters, (c) 3–4 quarters, (d) 4–5 quarters, (e) 5–6 quarters, and (f) 7–8 quarters to expiration. Eurodollar futures settle based on the spot 3-month Eurodollar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

January 2012 that it expected to keep the funds rate near zero until "late 2014".

These results corroborate our findings for Treasury yields and, like the Blue Chip survey data, suggest that financial markets did not expect the zero bound to constrain the federal funds rate for more than a few quarters until about August 2011. Only then do we see interest rate expectations more than four quarters ahead begin to behave in an attenuated fashion.

5.2 Can the Fed Manage Monetary Policy Expectations?

Our findings above suggest that interest rate expectations more than a few quarters ahead were largely unaffected by the zero bound until at least August 2011. Thus, even though the federal funds rate was severely constrained by the zero lower bound beginning in December 2008, monetary policy more broadly defined might not have been substantially constrained to the extent that the Federal Reserve can affect monetary policy expectations. In this section, we briefly review the evidence on the Federal Reserve's ability to influence these expectations by any means other than changes in the current level of the federal funds rate. Note that, even if the Fed had no ability to influence monetary policy expectations, our findings above would still have interesting implications for fiscal policy and the degree to which fiscal stimulus crowded out private-sector investment.

In theory, a central bank can influence private-sector expectations of future monetary policy if the bank has some ability to at least partially commit to its future policy actions. Schaumburg and Tambalotti (2007) and Debortoli and Nuñes (2010) define a continuum of partial commitment technologies that lie between perfect commitment and perfect discretion, in the sense of Kydland and Prescott (1977). Since perfect discretion is a limiting case along this continuum and implies no ability to commit whatsoever, it seems likely—or at least theoretically possible—that monetary policymakers would have some ability to influence private-sector expectations of future monetary policy actions at least a few periods into the future. Nevertheless, the Federal Reserve's ability to manipulate expectations about monetary policy several quarters into the future is ultimately an empirical question.

Empirically, there are several studies of the Federal Reserve's ability to influence longer-term interest rates through its communications, such as through the statements released by the FOMC after each monetary policy meeting. Gürkaynak, Sack, and Swanson (2005b) separately identify

	Treasury yields					
	3-month	6-month	1-year	2-year	5-year	10-year
FOMC drops "considerable period" language on Jan. 28, 2004						
Jan. 27, 2004	0.91	0.98	1.17	1.694	3.082	4.391
Jan. 28, 2004	0.94	1.00	1.295	1.86	3.221	4.494
change (bp)	3	2	12.5	16.6	13.9	10.3
FOMC projects zero funds rate "at least through mid-2013" on Aug. 9, 2011						
Aug. 8, 2011	0.05	0.07	0.173	0.271	1.133	2.591
Aug. 9, 2011	0.03	0.06	0.13	0.172	0.928	2.362
change (bp)	−2	−1	−4.3	−9.9	−20.5	−22.9

Table 3. Response of Treasury yields to significant changes in FOMC statements on Jan. 28, 2004, and Aug. 9, 2011. In both cases, there was no change in the current federal funds rate target, but the statement described a substantial change in the outlook for the funds rate relative to market expectations. See text for details.

the impact of FOMC actions (that is, changes in the federal funds rate target) and statements, and find that FOMC statements have highly statistically significant effects on Treasury yields out to maturities of 10 years. In fact, more than half of the explainable variation in the response of two-year Treasury yields (and almost 90 percent of the variation in the response of 10-year yields) to FOMC announcements is attributable to the FOMC's statements, rather than to changes in the current federal funds rate target. The authors' interpretation of this finding is not that statements have some mysterious independent power over longer-term interest rates, but rather that statements affect longer-term yields by changing financial market expectations about the future path of the funds rate. Bernanke, Reinhart, and Sack (2004) review these results and come to very similar conclusions using slightly different methods.[17] More recently, Campbell et al. (2012) extend the Gürkaynak et al. (2005b) analysis through the end of 2011, and find that the FOMC's statements continued to have similarly large effects on longer-term bond yields throughout the financial crisis and its aftermath in 2007–11.

Table 3 highlights two examples of this effect. From August 2003 until January 2004, the FOMC stated after each of its meetings that the accommodative stance of monetary policy "can be maintained for a considerable period." On January 28, 2004, in response to the strengthening

[17] Kohn and Sack (2004) also find that FOMC statements and Congressional testimony by the Fed Chairman have significant effects on longer-term interest rates, and present evidence that changes in monetary policy expectations are the primary driver of these changes.

economic outlook, the FOMC dropped this phrase from its statement and replaced it with the phrase "the Committee believes it can be patient in removing its policy accommodation."[18] Even though the funds rate target itself was unchanged on that date, the change in the statement was read by financial markets as indicating that the FOMC would begin raising the funds rate sooner than previously expected.[19] The result was that longer-term Treasury yields responded dramatically to the announcement, rising by about 10 to 16 bp at maturities of one to ten years.

Similarly, on August 9, 2011, in response to the weakening economic outlook, the FOMC announced that "economic conditions… are likely to warrant exceptionally low levels for the federal funds rate at least through mid-2013." Because the federal funds rate was already at an effective lower bound of 0 to 25 bp, there was no change in the FOMC's current federal funds rate target. Analogous to the previous example, financial markets read the change in statement as signaling the FOMC would likely begin raising the funds rate later than previously expected.[20] As a result, longer-term Treasury yields fell substantially, about 10 to 23 bp at maturities of two to ten years. The Blue Chip survey evidence presented in Figure 5 also suggests that this announcement had a large effect on monetary policy expectations.

These two examples are representative of the more comprehensive and systematic evidence in Gürkaynak et al. (2005b), Kohn and Sack (2004), Bernanke et al. (2004), and Campbell et al. (2012). These authors all find statistically and economically significant effects of FOMC communications on longer-term bond yields, above and beyond any effects of changes in the current level of the funds rate. We conclude from these studies that the FOMC does have the ability to influence monetary policy expectations for at least the next few years, and thereby affect the level of intermediate-maturity Treasury yields and Eurodollar futures with several quarters to expiration.

[18] The statements released after each FOMC meeting are available from the Federal Reserve Board's public web site.

[19] For example, the front page of *The Wall Street Journal* reported the following morning that "investors interpreted the omission of 'considerable period' as a signal that the Fed is closer to raising rates than many thought." ("Fed Clears Way for Future Rise in Interest Rates," *The Wall Street Journal,* Jan. 29, 2004, Greg Ip, p. A1.) Note that 10–15 bp is the typical response of 5- and 10-year Treasury yields to a 100-bp surprise change in the federal funds rate target (Gürkaynak et al. 2005b), so the changes in Table 3 are large.

[20] The front page of *The Wall Street Journal* the following morning noted that, in response to the FOMC statement, financial markets "lowered their expectations for when the Fed will start tightening policy." ("Markets Sink Then Soar after Fed Speaks," *The Wall Street Journal,* Aug. 10, 2011, Sudeep Reddy and Jonathan Cheng, p. A1.)

5.3 The Federal Reserve's Purchases of Long-Term Bonds

The Federal Reserve may also be able to affect longer-term interest rates by changing the supply of long-term bonds available to the private sector through large-scale open-market purchases. Although standard representative-agent asset pricing models do not allow for the quantity of a security to have any effect on its price, Vayanos and Vila (2009) provide a modern, arbitrage-free foundation for the earlier "portfolio balance" and "preferred habitat" models of Tobin (1958) and Modigliani and Sutch (1966).[21] Intuitively, if private-sector investors have heterogeneous preferences for different bond maturities, and arbitrage across maturities is limited, then the supply of longer-term bonds in the market can affect longer-term bond yields.

Empirically, Bernanke et al. (2004), Krishnamurthy and Vissing-Jorgensen (2007, 2011), Gagnon et al. (2011), Swanson (2011), and others find that large changes in the supply of Treasury securities have had appreciable effects on the yields of those securities. Between 2008 and 2011, the FOMC announced several rounds of large-scale purchases of longer-term Treasury and agency-backed securities, amounting to over $2.3 trillion in total.[22] These purchases represented a substantial fraction of the quantity of longer-term Treasury bonds in the hands of the private sector (Gagnon et al. 2011), and thus would be expected to have appreciable effects on longer-term bond yields based on the findings of the studies cited above.

Indeed, these purchases may help to explain a surprising feature of our results in Figure 4, namely, that 5- and 10-year Treasury yields were no less sensitive to news between 2008 and 2012 than in normal times. On the one hand, a finding of relatively little attenuation in the sensitivity of longer-term yields before August 2011 is not surprising, given how quickly market participants expected the funds rate to lift off from zero. But the continued lack of any visible attenuation for 5- and 10-year Treasuries in late 2011 and 2012 is surprising. For example, our illustrative model

[21] See also Hamilton and Wu (2012), who relate the Vayanos-Vila model to a standard arbitrage-free affine term structure model to estimate quantity effects.

[22] On November 25, 2008, the FOMC announced that it would purchase $500 billion of mortgage-backed securities and $100 billion of debt directly issued by the housing-related government-sponsored enterprises (GSEs); on March 18, 2009, the FOMC announced it would purchase an additional $750 billion of mortgage-backed securities, an additional $100 billion of GSE debt, and $300 billion of longer-term Treasury securities; on November 3, 2010, the FOMC announced that it would purchase an additional $600 billion of longer-term Treasury securities; and on September 21, 2011, the FOMC announced it would exchange an additional $400 billion of short-term Treasury securities for an equal amount of long-term Treasury bonds, which did not increase the FOMC's total holdings of Treasury securities but substantially altered the composition of those holdings.

in Section 2 would predict that the sensitivity of 5-year Treasury yields to news would be at least slightly attenuated after August 2011 given that the sensitivity of 2-year Treasury yields to news was only half as large as normal during that time.

A natural explanation for this finding is that market participants expected the Fed to adjust its purchases of long-term bonds in response to changing economic conditions. Since the stuides cited above found these purchases had substantial effects on long-term yields, this could have offset any dampening effect from the zero bound on those yields. The net result would be less attenuation in the sensitivity of 5- and 10-year Treasury yields to news between 2008 and 2012, perhaps even resulting in no net attenuation, as we found in Figure 4.

5.4 Why Are Interest Rates Sometimes More Sensitive to News?

Another interesting feature of Figures 4 and 7 is that the interest rate sensitivity coefficients δ^τ are sometimes estimated to be significantly *higher* than normal. For example, Treasury yields with two or more years to maturity and Eurodollar futures with three or more quarters to expiration were all more than twice as sensitive to news in 2004 and from mid-2007 to mid-2008 as in our benchmark sample from 1990–2000.

A natural explanation for why interest rates might have been more sensitive to news at those times is that financial markets were unusually uncertain about the outlook for monetary policy. In an environment of higher monetary policy uncertainty, each data release can have larger effects on financial market expectations of the future path of monetary policy; this would be the case, for example, if financial markets use methods similar to Kalman filtering to form expectations of future monetary policy. As a result, any given data surprise would tend to have a larger effect on medium- and longer-term interest rates, and our interest rate sensitivity regressions (9) and (10) would show a heightened sensitivity of those yields to news.

We investigate the importance of this effect by considering a more structural specification for the time-varying sensitivity coefficients δ^τ in regressions (9) and (10). In particular, we consider nonlinear regressions 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 δ, including a measure of monetary policy

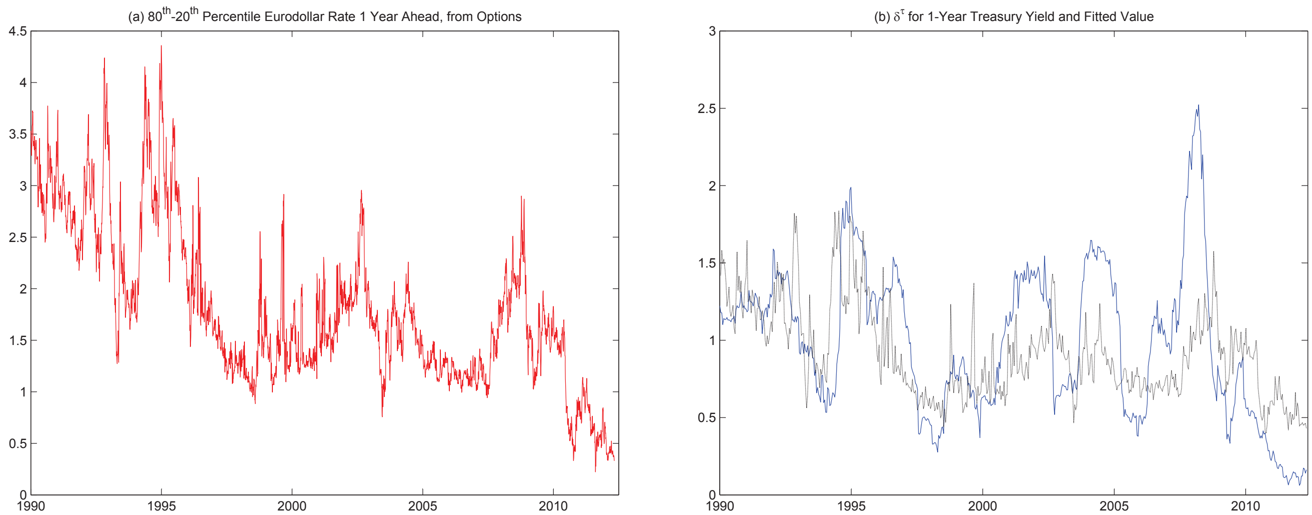

Figure 8. (a) Solid red line depicts the difference between the 80^{th} and 20^{th} percentiles of the 1-year-ahead Eurodollar rate distribution, derived from Eurodollar options. (b) Solid blue line depicts time-varying sensitivity coefficients δ^τ from regression (10) for the 1-year Treasury yield; dotted black line plots the fitted value $f(Z_t)$ from regression (12)–(13) where Z_t includes the width of the interquintile range from panel (a). See text for details.

uncertainty, and f denotes the nonlinear functional form

$$f(Z_t) = \max\{\, 0,\ \theta + \phi Z_t \,\}, \tag{13}$$

where θ and ϕ are parameters to be estimated along with γ and β in (12). The nonnegativity constraint in (13) reflects the view that, although an interest rate's sensitivity to news may rise and fall, it should never flip sign. (In practice, the estimated values of θ and ϕ almost never produce negative values for $f(Z_t)$, so the nonlinearity in (13) can essentially be ignored.) Thus, instead of estimating δ^{τ_i} or δ^τ in an unrestricted way in (9) or (10), with annual dummies or rolling regressions, regression (12) allows the sensitivity of Treasury yields to vary at daily frequency along with variations in Z_t.

We measure monetary policy uncertainty using Eurodollar options, which are the most liquid options available on a shorter-term interest rate and thus represent the best measure of monetary policy uncertainty we have over the next several quarters. Using these options, we can estimate the distribution of the spot Eurodollar rate at different points in the future, such as in one year's time; we then use the distance between the 80^{th} and 20^{th} percentiles of this distribution as our measure of monetary policy uncertainty.[23] We plot this series in the first panel of Figure 8. The

[23]On each day from 1996–2012, we compute the distance between the 80^{th} and 20^{th} percentiles for general dis-

	Treasury yield maturity					
	1-year		2-year		10-year	
Monetary policy uncertainty	.458	(4.05)	.447	(3.98)	.318	(1.89)
time trend ($\times 10^{-3}$)	.063	(1.13)	.108	(1.80)	.089	(1.90)
# Observations	2747		2747		2747	
R^2	.19		.19		.11	

Table 4. Coefficient estimates ϕ from regression $\Delta y_t = \gamma + f(Z_t)\beta X_t + \varepsilon_t$, where $f(Z_t) = \max\{0, \theta + \phi Z_t\}$, at daily frequency from Jan. 1990 to Apr. 2012. Heteroskedasticity-consistent t-statistics in parentheses. See text for details.

distance between the 80$^{\text{th}}$ and 20$^{\text{th}}$ 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 35 bp in 2012. This general downward trend in monetary policy uncertainty (discussed in Swanson, 2006) is punctuated by increases in 1994 (when the Fed began to raise interest rates after the 1991 recession), 2002, 2004, and 2007–08.

Results for the coefficients ϕ in regression (12)–(13) for the 1-, 2-, and 10-year Treasury yields are reported in Table 4.[24] In each regression, a constant θ is included and is normalized so that the average value of $f(Z_t)$ equals unity over the benchmark 1990–2000 sample, in order to ensure that θ, ϕ, and β are separately identified. A time trend is also included in Z_t to allow the downward trend in monetary policy uncertainty and the fluctuations around that trend to have possibly differing effects on each yield's sensitivity to news.

In Table 4, the estimated coefficient on monetary policy uncertainty is positive for each yield, highly statistically significant for the 1- and 2-year yields, and significant at the 5 percent level for a one-sided test in the positive direction for the 10-year yield. The R^2 for each regression is essentially the same as in regression (9) with annual dummies, implying no change in fit from using the alternative specification (12)–(13). The coefficient point estimates of about 0.4 imply that, if the width of the interquintile range for the one-year-ahead Eurodollar rate widens by one

tributions for the Eurodollar spot rate estimated using Eurodollar options with one year to expiration and a range of strike prices, as discussed in our computation of Figure 6. Detailed options data are not available to us prior to 1996, so from 1990–1995, we compute the width of this interquintile range using the implied volatility on Eurodollar options with one year to expiration, assuming a lognormal distribution for the one-year-ahead spot Eurodollar rate, computed by staff at the Federal Reserve Board.

[24]Results for the 3-month Treasury yield are not statistically significant and are not reported. This result is not surprising given that the 3-month yield's sensitivity to news over much of the 2000–12 period was constrained by the zero lower bound and thus responded very little to news.

percentage point, the Treasury yield's sensitivity to news increases by about 0.4, or 40 percent of the sensitivity during the benchmark 1990–2000 sample. For the 2- and 10-year Treasury yields, the estimated time trend is slightly positive, implying that the fluctuations in monetary policy uncertainty around its trend have a larger effect on the sensitivity of those yields to news than the downward trend in that uncertainty.

Fitted values from regression (12)–(13) for the 1-year Treasury yield's sensitivity to news are reported in the second panel of Figure 8. The solid blue line in the figure depicts the time-varying sensitivity coefficient δ^{τ} from regression (10) for the 1-year Treasury yield, and the dotted black line plots the fitted values $f(Z_t)$ from regression (12)–(13). There is a close correspondence between the increases in monetary policy uncertainty in 1994, 2002, 2004, and 2007–08 with the increases in interest rate sensitivity over those same periods. Monetary policy uncertainty was also relatively low in 1993, 1997–98, and 2010–11, helping to explain the lower levels of interest rate sensitivity in those periods.

Nevertheless, monetary policy uncertainty fails to explain some notable features of the variation in Treasury yield sensitivity in the figure. For example, the very large spike in 2007–08 is only partially 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.[25] Also of particular interest is the observation that the 1-year yield's sensitivity to news is lower in 2011–12 than can be explained by the drop in monetary policy uncertainty alone. We interpret this fact as reflecting the additional constraining effects of the zero lower bound on the sensitivity of the 1-year Treasury yield to news. That is, not only was monetary policy less uncertain at that time, it was also very close to zero, and the latter fact is likely to have had additional attenuating effects on the sensitivity of the 1-year yield to news.[26]

[25] For example, financial market expectations of large-scale bond purchases or other market interventions by the Federal Reserve may have contributed to the sensitivity of Treasury yields over this period, as discussed in the preceding section. Alternatively, financial markets may have been concerned about the risks of a prolonged "lost decade" scenario, as in Japan in the 1990s, which could have also raised the sensitivity of monetary policy expectations and Treasury yields to news.

[26] For example, the zero lower bound makes the current level of the federal funds rate completely unresponsive to news, which is not the case away from the zero bound, regardless of the level of monetary policy uncertainty. Of course, it is also important to note that one of the main reasons monetary policy uncertainty fell to such low levels in 2010–11 in the first place was because of the zero bound. Moreover, it was the zero bound constraint that led the FOMC to declare its expectation that it would keep the federal funds rate at zero until "at least mid-2013," which reduced uncertainty about the path of monetary policy through mid-2013 even further. Thus, even the component of the decline in Treasury yield sensitivity in 2011–12 that can be explained by decreases in monetary policy uncertainty

5.5 Post-2007 Distribution of Macroeconomic Data Release Surprises

It is important to note that the macroeconomic data release surprises X_t can be regarded as strictly exogenous in regressions (9) and (10). This is because the surprise component of each data release is independent of all past values of the interest rates on the left-hand side of these regressions under the assumption that our survey expectations data incorporate all relevant information as of the day before the release. To the extent that this assumption is true, and 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 fairly similar across these samples. This can be seen in Figure 9, 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 the degree to which interest rates of any maturity are affected by the presence of the zero lower bound. Our method uses the high-frequency sensitivity of yields to macroeconomic news to measure and statistically test the extent to which the zero bound affected any given yield's behavior. Importantly, our method provides a quantitative

can still be attributed to the zero lower bound.

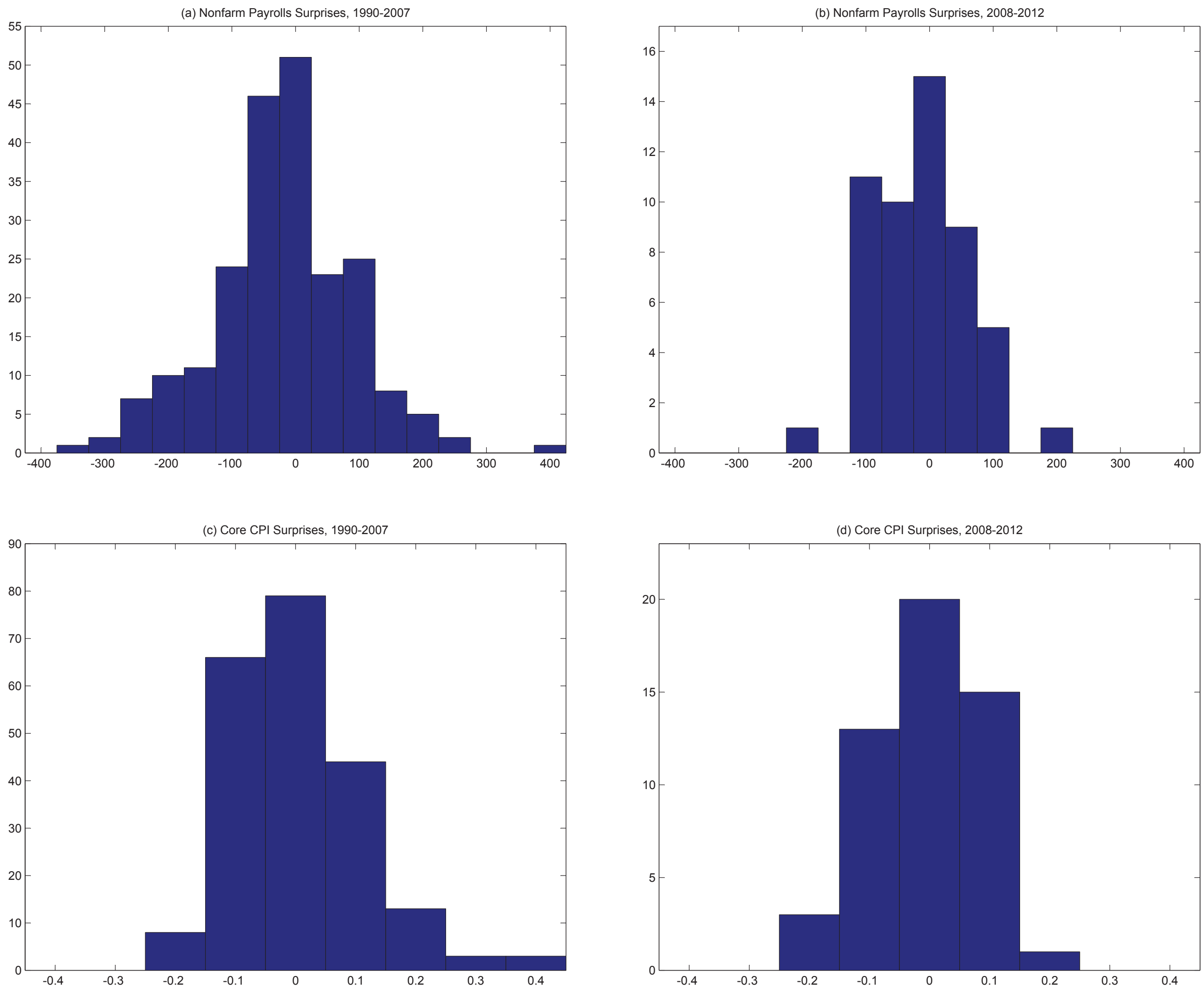

Figure 9. Top panels depict empirical distribution of the surprise component of nonfarm payrolls announcements from (a) 1990–2007 and (b) 2008–12, rounded to the nearest 50 thousand workers. Bottom panels depict the distribution of core CPI surprises from (c) 1990–2007 and (d) 2008–12, rounded to the 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.

measure of the severity of the zero bound constraint on each yield as well as a statistical test for the periods during which each yield was affected.

We find that interest rates with six months or less to maturity were substantially constrained by the zero bound between 2009 and April 2012, but interest rates with more than two years to maturity were largely unaffected. One- and two-year Treasury yields represent an intermediate case, being only partly constrained for part of this period. Over much of this period, financial markets expected the funds rate to remain at the zero bound for only a few quarters. Only beginning in late 2011—around the time of the FOMC's "mid-2013" guidance—do we see financial markets begin to expect short-term rates to remain at the zero bound for a longer period of time.

Our results have important implications for both fiscal and monetary policy. To the extent that medium- and longer-term interest rates are more relevant than very short-term rates for private-sector spending, we find little evidence that crowding out would have been any less than normal over much of 2008–12. Only beginning in late 2011 do medium-term Treasury yields show significant attenuation in their responsiveness to macroeconomic news.

For monetary policy, our results imply that policymakers had substantial room to affect medium- and longer-term interest rates throughout much of 2008–12, 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 Treasury and Agency-backed bonds.

The methods we have developed in this paper can be generalized beyond the United States and applied to any economy for which financial markets are sufficiently well developed. In particular, it would be very interesting to see our methods applied to other economies that have faced the zero lower bound in recent years, such as Japan, the U.K., Canada, Sweden, and the Euro area.

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