Estimating Monetary Policy Rules When Nominal Interest Rates Are Stuck at Zero

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Abstract

Did the Federal Reserve’s response to economic fundamentals change with the onset of the global financial crisis? Answering this question is tricky because the policy interest rate set by the Federal Reserve has been stuck between zero and 25 basis points since late 2008, essentially the beginning of the crisis. Data from surveys of economic forecasts allow us to sidestep the censoring problem and estimate monetary policy rules by regression. We find that the Fed’s inflation response has significantly decreased while its unemployment response has remained strong. In terms of central bank communication, our results can viewed as an argument that the Federal Reserve’s commitment to stable inflation has become weaker in the eyes of the professional forecasters—and probably the financial markets as well.

Keywords: monetary policy, policy rule, survey data, market perceptions, censoring, zero lower bound, Blue Chip survey.

JEL codes: E53, E58.

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1 Introduction

Did the Federal Reserve’s response to economic fundamentals change with the onset of the global financial crisis? Since the 1990s, simple monetary policy rules for the nominal short rate have been widely used to analyze the behavior of central banks—most notably the Fed’s behavior following Taylor (1993). The short rate is critical for the dynamics of the whole economy, not only by representing the reaction of a central bank to the current state of the economy but also by providing information of the economy’s future to financial markets. Estimates of policy rules depend on meaningful covariation between nominal short rates and economic fundamentals such as the inflation and unemployment rates. But since December 2008, around the onset of the global financial crisis, U.S. nominal short rates have been stuck at their zero lower bound (ZLB). Hence, policy rates are censored at zero and observed nominal short rates cannot provide the meaningful variation required to answer our research question.

A textbook way of tackling the censoring problem is to apply a limited dependent variable econometric method to the historical data. However, these methods cannot define a hypothesis test of whether or not policy response parameters changed between the pre-ZLB period (when data is always uncensored) and the ZLB period (when data is always censored). In contrast, this paper avoids the censoring problem by using data on agent expectations—obtained from surveys of economic forecasts—and answers whether or not Fed policy has changed since the global financial crisis. Even while the historical short rates were close to zero, one-year-ahead forecast of the nominal short rate stayed above zero at least until August 2011 when a calendar-based forward guidance was introduced by the Federal Open Market Committee.

1In this paper, we assume that the federal funds rate is effectively at its ZLB as far as it is below 25 basis points. See Bernanke and Reinhart (2004) for reasons why the short-term rate had better not be pushed down all the way to zero.

2This is our censoring problem: Even when both the rate recommended by a simple rule and so the desired target rate are negative, the ZLB-constrained target rate announced by the Fed stays positive. See Rudebusch (2009) and Curdia and Woodford (2011) among others for the argument that target rates would be negative. The censorship is with regards to the targeted policy rate, not the actual nominal short rate.

3Estimation of a policy rule with historical data during the ZLB confronts the problem of limited dependent variables since the left-hand side variable is censored at zero. A maximum likelihood estimator based only on censored observations cannot be consistent (for example, premises of Newey and McFadden 1994 Theorem 2.1 cannot be satisfied) leading to undefined asymptotic distributions and hypothesis tests. Therefore, historical data cannot satisfactorily answer whether or not a policy shift took place in response to the global financial crisis. One probably could expand our definition to partial identification and conduct modified statistical tests. However, this would be a different question than we answer here, which is manifest as a simple parameter break test. Issues of partial identification are interesting but beyond the scope of this paper.

4Using forecasts data has recently been explored in Devereux, Smith and Yetman (2012) and Coibion and Gorodnichenko (2012). These papers do not discuss the relationship of these data to the censoring problem presented by the ZLB. Swanson and Williams (2012) ask whether there has been a change in various maturity U.S. treasury prices’ response to data releases, which avoids this censoring problem when longer maturities are used, however does not answer the policy response question we ask.
Market Committee (FOMC). By using the one-year-ahead forecasts available in various issues of Blue Chip Economic Indicators, we avoid the ZLB-induced censoring problem and estimate the Fed’s monetary policy rule using conventional regression methods. Prior to the global financial crisis (and the ZLB), policy rules estimated with forecast data are comparable to those estimated with historical data: We find a short rate response to inflation just under 2 and to unemployment between -1/2 and -1, which is similar to estimates in previous literature and our own estimates using historical data.

Our results using forecast data show that during the global financial crisis the Federal Reserve’s inflation response has significantly decreased while its unemployment response has remained strong. In particular, the inflation response point estimate significant decreases by half and no longer satisfies the “Taylor principle.” In terms of central bank communication, this can be viewed as an argument that the Federal Reserve’s commitment to stable inflation has become weaker in the eyes of the professional forecasters—and probably the financial markets as well. This result could be interpreted as either bad (the loss of inflation-fighting credibility) or good (successful forward guidance of the Fed’s future “accommodative stance”\(^5\)), and we do not advocate one interpretation over the other.

In addition to avoiding censorship, Blue Chip Economic Indicators includes long-horizon forecasts of inflation, unemployment and the real short rate. We take these forecasts as \textit{data} for respective long-run (or target) values that are unobserved and under-identified by conventional policy rule estimation using historical data. We argue that having data for long-run values (or targets) is a further advantage of using these forecast data, and this advantage allows us to tightly estimate policy response parameters.

A price we pay for our simple solution is that now the object estimated is market participants’ \textit{perception} of how the Federal Reserve sets policy. However, this is often assumed to coincide with the actual policy rule (by rational expectations) or else is interesting on its own as an object (see Hamilton, Pruitt and Borger 2011). It is critical for monetary economists to understand the Fed’s reaction function since the global financial crisis, as well as the financial markets’ perception of its behavior. This paper specifically investigates if market professionals believe the Federal Reserve’s response to real activity and inflation has shifted because of the financial crisis. If one interprets central bank credibility as its perceived response to inflation, our results shed light on whether the Federal Reserve has remained credible since short rates hit the ZLB at the end of 2008.

The plan of the paper is as follows. Sections 2 and 3 describe our data and estimation methods. In Section 4, we present our main results and discuss several sensitivity analyses that support the robustness of our results. Section 5 concludes.

\(^5\)FOMC Statement, October 2012.
2 Data

The main idea of this paper is to turn to data on agents’ *expectations* to answer our research question. A key advantage of such data is that they may forecast future economic conditions that warrant uncensored policy rates, even when current short rates at the time of the forecast are stuck at the ZLB. This means that all the necessary ingredients—dependent and explanatory variables—retain meaningful variability that can be used to estimate the key policy response coefficients. In turn, in the usual way we can then form hypothesis tests that answer our research question.

A brief digression on terminology may be helpful. This paper uses two types of data. The first type of data we refer to as *actual, historical or realized* data are prices or measurements that actually prevailed in the U.S. experience. Examples are the 3-month T-Bill rate during December 2005 or the unemployment rate measured by the Bureau of Labor Statistics for December 2005—these are actual, historical or realized data. The second type of data we refer to as *expectations or forecast* data are one-year-ahead forecasts of future economic conditions. An example is the 3-month T-Bill rate expected to prevail in 2006Q4, as forecast in December 2005.

Our main data are Blue Chip Economic Indicators. Collected monthly since the late 1970s, each issue of Blue Chip Economic Indicators surveys between 30 to 60 professional forecasters in the financial services, consultancy, and academic industries, with emphasis on forecasters in the financial services industry. A median forecast is compiled from this panel for each quarterly forecast horizon, but the full panel of responses are not shown.\(^6\)

Studies such as Ang, Bekaert and Wei (2007) suggest that surveys contain remarkably accurate forecasts, therefore we make the assumption that they accurately represent market expectations. Blue Chip surveys provide medium-horizon (one- to six-quarter ahead) forecasts of CPI inflation, the 3-month T-Bill rate, and the unemployment rate. Biannually the survey also publishes long-horizon forecasts (the five-year average value five years ahead) for these variables, which we convert to a monthly measure by a simple linear interpolation. We make our primary measure of policy the 3-month T-Bill rate so that we have both its medium- and long-horizon forecasts.\(^7\) We use CPI inflation and unemployment to represent

\(^6\)The full panel is shown for forecasts pertaining to the current year and following year of the month in which the survey is conducted. For example, the March 2000 survey shows the full panel of responses of year-over-year forecasts for 2000 and 2001. However, this means that the forecast horizon of the responses shifts by survey month (i.e. the November 2000 survey shows the full panel of response of year-over-year forecasts for 2000 and 2001, too), which would hinder our exercise. Blue Chip Financial Forecasts, a separate but similar survey, does show the full panel of responses for quarterly forecast horizon. See Footnote 7 for further comparison between the two surveys.

\(^7\)It is conventional to measure policy by the federal funds rate, but this variable is not in our data. Instead, we note that over our sample period the correlation between these two series’ actual values is greater than 0.99 and so this data limitation is negligible. Blue Chip Financial Forecasts do contain forecasts for the
Figure 1: **Short-Term Interest Rates**

*Notes:* 3-month T-Bill rates. Actual values from the Federal Reserve Board of Governors, solid line; median one-year-ahead forecast from Blue Chip Economic Indicators, dashed line; long-horizon forecasts from Blue Chip Economic Indicators, dotted line.

Inflation and activity pressures, respectively. We use data starting in 1986 when all variables are observed, and these series are plotted in Figures 1–3.

The long-horizon forecast of the short rate, inflation rate and unemployment rate are plotted by the dotted lines in Figures 1-3, respectively. Taking the difference between the long-horizon short-rate forecast and the long-horizon inflation-rate forecast yields the long-horizon real-rate forecast, by the Fisher identity. Figure 1 shows the long-horizon forecasts of the short rate have fallen steadily over the past three decades; much of this follows the fall in the long-horizon forecasts of inflation plotted in Figure 2. Meanwhile, the long-horizon forecast of the unemployment rate hovers around familiar values of five to six percent. Nevertheless, there has been important variation in this long-horizon unemployment forecast, particularly around 2008 where it appears that professional forecasters draw a connection between the global financial crisis and the long-run level of the unemployment rate in the U.S.

Figure 1 demonstrates the main idea of our paper. Although the realized short rate (the solid line) hits the ZLB at the end of 2008, the one-year-ahead forecast for the short rate (the fed funds rate and the GDP growth rate, but do not include an unemployment rate forecast. Therefore, we opt for using Blue Chip Economic Indicators since it forecasts unemployment and 3-month T-Bill rates, the latter of which are extremely similar to the desired fed funds rates.
dashed line) remains above the ZLB until the middle of 2011. It is rather inconsequential to our exercise that this indicates that market expectations exhibited serially correlated forecast errors during the ZLB. What is important for our analysis is that the forecasts manifest a systematic relationship between short rate and inflation and unemployment.

The end of Figure 1 demonstrates why we end our data sample in August 2011. At the August 2011 FOMC meeting, the Committee announced that “exceptionally low” short rates would likely be held “at least through mid-2013.”

Our surveys—as well as price-based measures of market expectations such as fed funds futures or eurodollar contracts—respond immediately to this policy statement. Surveys published after the August 2011 FOMC statement exhibit medium-horizon short rate forecasts at the ZLB, through the present day. Therefore this paper makes statements about the possible shift of perceived monetary policy during the period of December 2008 to August 2011, relative to previous periods.

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8Earlier, the FOMC introduced the “extended period” language in March 2009. Subsequently, the Committee extended the calendar-based forward guidance to late 2014 in January 2012 and to mid-2015 in September 2012.
Figure 3: Unemployment Rates

Notes: Civilian unemployment rates. Actual values from the Bureau of Labor Statistics, solid line; median one-year-ahead forecast from Blue Chip Economic Indicators, dashed line; long-horizon forecasts from Blue Chip Economic Indicators, dotted line.

3 Estimating Policy Rules

A Taylor-type policy rule specification is

\[ i_\tau = (1 - \rho) \left[ (r^* + \pi^*) + \beta (\pi_\tau - \pi^*) + \delta (u_\tau - u^*) \right] + \rho i_{\tau-1}. \]  

(1)

This form follows Boivin (2006) in using unemployment to measure real activity; it is similar in spirit to rules using the output gap—as in Taylor (1993) among many others. The policy rate \( i_\tau \) responds to the inflation response coefficient \( \beta \) times the deviation of inflation \( \pi_\tau \) from its long-run target \( \pi^* \); the unemployment response coefficient \( \delta \) times the deviation of unemployment \( u_\tau \) from the non-accelerating inflation rate of unemployment (NAIRU) \( u^* \); and the lagged short rate via “gradual adjustment” dictated by \( \rho \). When both inflation and unemployment are at their long-run values of \( \pi^* \) and \( u^* \), the nominal short rate converges to the economy’s equilibrium real rate \( r^* \) plus inflation \( \pi^* \) as implied by the Fisher equation.

Previous implementations of (1) typically use historical data directly as arguments in the rule or to instrument any expectational terms. An obvious problem is that the \( * \)-variables—the inflation target, the NAIRU and the equilibrium real rate—are under-identified by the data. What one conventionally estimates is the reduced form

\[ i_\tau = c_0 + c_1 \pi_\tau + c_2 u_\tau + c_3 i_{\tau-1}. \]
Note that \( c_1 = \beta, c_2 = \delta \) and \( c_3 = \rho \), but \( c_0 = (1 - \rho)(r^* + \pi^*) - \beta \pi^* - \delta u^* \) which highlights the typical identification problem. Note that \( c_0 \) is a function of both the unobserved \( r^*, \pi^*, u^* \) and the response parameters \( \beta, \delta \). Even if \( r^*, \pi^*, u^* \) are unchanged, we expect \( c_0 \) to shift if \( \beta \) or \( \delta \) shift. Therefore, when estimating (1) using historical data, we will need to allow \( c_0 \) to break when we allow either \( \beta \) or \( \delta \) to break.

### 3.1 Rule Using Forecasts Data

Using forecast data instead of historical data provide us at least two distinctive features. First, the forecast data include one-year-ahead forecasts for unemployment, inflation, and the short rate. As mentioned in Section 2, these forecasts are not censored even after the ZLB started to bind. Second, these data include long-horizon forecasts of the short rate, inflation rate and unemployment rate. In macroeconomic models with well-defined steady-state growth paths, long-horizon forecasts are synonymous for the values of the equilibrium nominal interest rate, monetary policy’s inflation target, and the NAIRU. Therefore, the survey of forecasts gives us a unique opportunity to pin down the value of the *-variables using data. We believe this to be a terrific advantage of these data, that they can be used to avoid the identification problem intrinsic to conventional Taylor-type policy rule estimation using historical data.

The forecast-data-based rule we estimate is

\[
i_{q(t)+4,t} = (r^*_t + \pi^*_t) + \beta(\pi_{q(t)+4,t} - \pi^*_t) + \delta(u_{q(t)+4,t} - u^*_t)
\]  

(2)

where \( t \) is measured in months and \( q(t) \) gives the quarter of month \( t \). Equation (2) takes forecasts as data, and we assume that these forecasts are the representative forecaster’s conditional expectation. For example, \( i_{q(t)+4,t} \) is the time \( t \) expectation of the short rate during quarter \( q(t) + 4 \), and \( \pi^*_t \) is the time \( t \) expectation of the average inflation rate over a long horizon.

At least four features of our main specification are worth mentioning. First, we have made a particular horizon choice: Our main results use the the four-quarter-ahead forecast. We do this to prevent, as much as possible, the effect of the ZLB against using expectations data that represent higher, business cycle frequencies. Second, (2) includes no parametric constant to be estimated and instead pins down all movements of the nominal short-rate forecast to the inflation response \( \beta \pi_{q(t)+4,t} \), unemployment response \( \delta u_{q(t)+4,t} \), and shifts in the *-variables \( (r^*_t + \pi^*_t) - \beta \pi^*_t - \delta u^*_t \). As mentioned above, we view this discipline by the data as an advantage of our exercise. Third, (2) is written without a forecaster-specific index: Our benchmark results use the median forecast to stand in for the representative forecaster.

Finally, we have zeroed out the gradual adjustment parameter \( \rho \). Regarding this choice, we point out that Rudebusch (2006) argues that serially-correlated residuals account for the
persistence of short rates, reflecting variables used by policymakers but excluded by our simple analysis. Additionally, Hamilton, Pruitt and Borger (2011) find that responses to past inflation and real activity differ such that gradual adjustment may not be the most useful mechanism for smoothing the short rate’s dynamics. With these previous studies in mind, we decide to abstract from gradual adjustment for three reasons. First, our main results demonstrate that (2) fits the survey data extremely well without it. Second, we wish to prevent, as much as possible, the effect of the ZLB and using a closer horizon’s (i.e. monthly) short-rate forecast in the adjustment term works against that aim. Third, we prefer to explicitly evaluate the policy rule residual in Section 4.2, allowing room to discuss what omitted variables may be imposed by the Taylor-type specification, and connect them to measures of risk in the spirit of Atkeson and Kehoe (2009).

3.2 Rule Using Historical Data

To compare the forecast-data-based rule to conventional policy rule estimates based on historical data, we consider the equation

\[ i_t = c_0 + c_1 \pi_t + c_2 u_t \]  

where \( i_t, \pi_t, u_t \) are the actual values of the policy, inflation and unemployment rates in month \( t \). To maintain consistency with the variables available in the Blue Chip Economic Indicators, we use 3-month T-Bill rates, the civilian over-16 unemployment rate, and annual growth of the CPI excluding food and energy for the short rate, unemployment rate, and inflation rate, respectively.\(^9\)

3.3 Methodology

To answer the paper’s main question we allow parameters to break at certain dates. Our estimation is via maximum likelihood. Using (2) we write the log-likelihood allowing for parameter breaks:

\[ l(X_T; \sigma, \{\beta_k, \delta_k\}_{k=1, \ldots, B}) = -T \log(2\pi\sigma) - \frac{1}{2\sigma^2} \sum_{k=1}^{B} \left( \sum_{t \in T_k} \varepsilon_t^2 \right) \]  

where \( \varepsilon_t \) is the residual from (2) if using forecast data, or (3) if using historical data. \( B \) is the number of parameter regimes in the data sample \( T \) and \( T_k \) is the set of month indices in

\(^9\)CPI excluding food and energy, or “core” inflation, is used to measure policy since it strips out price fluctuations (food and energy prices) that are typically considered temporary. Congruently, core-CPI inflation is a better forecaster of future CPI inflation than headline inflation itself. Hence we find it reasonable to associate realized core inflation (in the rule estimated on historical data) with expected headline inflation (in the rule estimated on forecasts data).
regime \( k \).\(^{10}\) Log-likelihood (4) using (2) is maximized by ordinary least squares run without a constant, using dummy variables for regimes.

When we estimate the rule using historical data (3), the \( r_t^*, \pi_t^*, u_t^* \) combine with \( \beta, \delta \) and become regime-specific constants, while the \( \pi_{q(t)+4,t}, u_{q(t)+4,t} \) become actual values of inflation and unemployment observed at time \( t \), \( \pi_t, u_t \). Log-likelihood (4) using (3) is maximized by ordinary least squares, this time including a constant, using dummy variables for regimes. Note that the constant will be regime-specific if we allow \( \beta \) or \( \delta \) change, as noted above.

4 Estimation Results

Our main results use Blue Chip forecast data. We compare these estimates to results obtained using historical data. We find qualitative agreement between the two estimation approaches on the pre-ZLB sample, supporting the idea that forecast and historical data reveal similar policy response functions when both types of data are uncensored. We then confirm that historical data during the ZLB provides coefficient estimates that are essentially zero because of data censorship. On the other hand, using forecast data during the ZLB we find statistically significant policy response coefficients. We find evidence that the inflation response has diminished, with a point estimate below 1, while the unemployment response is unchanged.

We explore whether our conclusions are robust to including an additional parameter break around the year 2000, which has been found by many previous studies. When we do so, we estimate that the policy response to unemployment was very strong in the pre-ZLB 2000s and therefore that the ZLB unemployment response is significantly weaker; we continue to find that the policy response to inflation has significantly fallen during the ZLB.

We then explore the part of forecasted policy that is not associated with forecasted inflation or unemployment. We find a significant relationship between our estimated short-rate residuals and investors’ perceived risk, suggesting that market participants believe the Fed sets policy in response to financial market risk as well as to fundamentals at the business cycle frequency.

Finally, we use our estimates, realized data, and Blue Chip survey data to analyze the counterfactual implied short rate since the onset of the ZLB.
Table 1: Historical Data Results

<table>
<thead>
<tr>
<th></th>
<th>Full Sample</th>
<th>Break</th>
<th>Break</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>pre-ZLB</td>
<td>ZLB</td>
</tr>
<tr>
<td>Inflation response</td>
<td>1.551</td>
<td>1.943</td>
<td>-0.034</td>
</tr>
<tr>
<td>standard error</td>
<td>(0.108)</td>
<td>(0.108)</td>
<td>(0.035)</td>
</tr>
<tr>
<td>LR test</td>
<td></td>
<td>&lt; 0.001</td>
<td></td>
</tr>
<tr>
<td>Unemployment response</td>
<td>-0.714</td>
<td>-1.277</td>
<td>-0.008</td>
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<tr>
<td>standard error</td>
<td>(0.076)</td>
<td>(0.098)</td>
<td>(0.021)</td>
</tr>
<tr>
<td>LR test</td>
<td></td>
<td>&lt; 0.001</td>
<td></td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.788</td>
<td>0.853</td>
<td></td>
</tr>
</tbody>
</table>

Notes: Monthly data from Federal Reserve Board of Governors and Bureau of Labor Statistics, January 1986 to August 2011. The policy rate is measured by the 3-month T-Bill, inflation is measured by the 12-month growth rate of the core CPI index. In the Break panel a break is allowed at December 2008. Standard errors are provided in parenthesis. Rows marked LR test reports the $p$-value for the likelihood ratio (Chow) test of no break in the indicated parameter, allowing other parameters to break.

4.1 Parameter Estimates

We begin with the estimates obtained from historical data in Table 1. Both the full sample and pre-ZLB estimates agree with typical estimates of Taylor rule parameters. The inflation responses satisfy the Taylor principle, and the unemployment coefficients are significantly negative. The ZLB period estimates are indistinguishable from zero. This result is exactly as expected since the actual short rate contains basically no meaningful variation (on the policy response rule) since hitting the ZLB.

Table 2 reports our main results. We measure nominal short rate forecasts by those for the 3-month T-Bill rate (due to the discussion in Section 2) and measure inflation forecasts by those for the growth rate in the CPI index. Since our inquiry is focused on whether or not policy shifted at a known point in time, we do not search for breakpoints. The breakpoint

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10 We abstract from explicitly accounting for the month-to-month changes in exact forecast horizon. That is, the 4-quarter-ahead forecast for January is for the following year’s second quarter which is 14 months away; the 4-quarter-ahead forecast for February is for the following year’s second quarter which is 13 months away; and for March, 12 months away. Adding this technical complication is trivial but the effects are negligible.

11 We do not report the estimated constant in the Table because it does not uniquely identify a policy parameter of interest. In practice the estimated constant shifts between periods—the pre-ZLB point estimate is 5.10 and ZLB point estimate is 0.25. The historical data constant’s instability and identification problem is a reason that using forecast data is advantageous, since the latter data uniquely pin down all policy variables including long-run values (targets).
<table>
<thead>
<tr>
<th></th>
<th>Full Sample</th>
<th>Break</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>pre-ZLB</td>
</tr>
<tr>
<td>Inflation response</td>
<td>1.712</td>
<td>1.772</td>
</tr>
<tr>
<td>standard error</td>
<td>(0.276)</td>
<td>(0.280)</td>
</tr>
<tr>
<td>LR test</td>
<td></td>
<td>0.067</td>
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<tr>
<td>Unemployment response</td>
<td>−0.658</td>
<td>−0.801</td>
</tr>
<tr>
<td>standard error</td>
<td>(0.063)</td>
<td>(0.147)</td>
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<tr>
<td>LR test</td>
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<td>0.979</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.775</td>
<td>0.792</td>
</tr>
</tbody>
</table>

Notes: Monthly data from Blue Chip Economic Indicators, median forecasts, January 1986 to August 2011. The policy rate is measured by the 3-month T-Bill, inflation is measured by the CPI index. In the Break panel a break is allowed at December 2008. Standard errors are provided in parenthesis. Rows marked LR test reports the $p$-value for the likelihood ratio (Chow) test of no break in the indicated parameter, allowing other parameters to break.

considered by our main results is in December 2008, the month the effective fed funds rate entered the ZLB region of “0 to 1/4 percentage points,” and we label months after this date as the “ZLB” period.

The model fit is quite good. Our exercise is not describing historical short rates; the model fits the market’s forecast of the short rate using a simple linear combination of their forecasts of inflation and unemployment. Evidently these joint expectations are well-explained by a simple Taylor-type rule. We suggest that this has a straightforward explanation: Simple rules are widely used to communicate policy, as described in speeches by Kohn (2007) and Bernanke (2010).

Has the Federal Reserve’s response to inflation and unemployment changed during the ZLB? Our key findings are the likelihood ratio (Chow) test results reported in the rows marked LR test in Table 2. In each, the $p$-value is of the null hypothesis that the indicated parameter is unchanged by the ZLB, while other parameters are permitted to break.

We fail to find evidence that the policy response to unemployment has diminished during the ZLB: The Chow test of a no-break null hypothesis has a $p$-value of 0.98. We estimate that the policy rate falls about 80 basis points for every percentage point rise in the unemployment rate, both prior-to and during the global financial crisis.

However, there is evidence that the policy response to inflation is significantly lower since the onset of the global financial crisis: The Chow test of a no-break null hypothesis
Table 3: Forecasts Data Results with Two Breaks

<table>
<thead>
<tr>
<th></th>
<th>pre-2000</th>
<th>2000-2008</th>
<th>ZLB</th>
</tr>
</thead>
<tbody>
<tr>
<td>Inflation response</td>
<td>1.215</td>
<td>1.821</td>
<td>0.754</td>
</tr>
<tr>
<td>standard error</td>
<td>(0.227)</td>
<td>(0.511)</td>
<td>(0.347)</td>
</tr>
<tr>
<td>LR test</td>
<td></td>
<td></td>
<td>0.019</td>
</tr>
<tr>
<td>Unemployment response</td>
<td>−0.301</td>
<td>−1.683</td>
<td>−0.804</td>
</tr>
<tr>
<td>standard error</td>
<td>(0.093)</td>
<td>(0.407)</td>
<td>(0.090)</td>
</tr>
<tr>
<td>LR test</td>
<td></td>
<td></td>
<td>&lt; 0.001</td>
</tr>
<tr>
<td>$R^2$</td>
<td></td>
<td></td>
<td>0.870</td>
</tr>
</tbody>
</table>

Notes: Monthly data from Blue Chip Economic Indicators, median forecasts, January 1986 to August 2011. The policy rate is measured by the 3-month T-Bill, inflation is measured by the CPI index. A break is allowed at January 2000 and December 2008. Standard errors are provided in parenthesis. Rows marked LR test report the $p$-value for the likelihood ratio (Chow) test of no break in the indicated parameter, allowing other parameters to break, between the 2000-2008 and ZLB periods.

The results of Ang et al. (2011) and Hamilton et al. (2011) suggest that monetary policy (especially as perceived by market participants) changed around 2000. Therefore we consider the effect of controlling for another breakpoint in our estimation around January 2000. Table 3 reports the results.\(^{12}\)

These results are qualitatively similar to our benchmark findings. We estimate that the policy response to unemployment was very strong in the pre-ZLB 2000s—a short rate decrease of 170 basis points for each 1 percentage point increase in unemployment. This contrasts with a smaller short response of 30 basis points in the 1980s and 1990s. Therefore, we now find statistical evidence for a decrease in unemployment response during the

\(^{12}\)It is necessarily the case that the “ZLB” estimates are identical to those in Table 2.
ZLB relative to 2000-2008. However, this response is still strong (80 basis points for each percentage point of unemployment) and higher than the unemployment response prior to 2000.

We continue to find that the policy response to inflation has significantly fallen during the ZLB relative to any previous period, and the statistical significance of the parameter break only increases ($p$-value of 0.02) in this alternative specification. We view this as support for the main result given by our benchmark results in Table 2, which hereafter we discuss in greater detail.

### 4.2 Understanding the Residuals

Turning back to our benchmark results, the solid line of Figure 4 represent the residuals of the one-break regression. In a conventional policy rule regression with realized data, error terms capture policy “shocks.” However, in our benchmark rule using forecast data, the residuals represent a shift of the forecasted short rate that is not accounted for by a linear combination of inflation and unemployment forecasts. We find a few features noteworthy. First, the Greenspan’s chairmanship starting in mid-1987 was characterized by reasonably small residuals until around 2001. This means that forecasters believed that short rates largely reflected economic conditions.

In the first half-decade of the 21st century, the residuals are persistently negative. This implies that market professionals during this time expected more expansionary policy than would otherwise be called for by economic conditions. Starting in 2005 the expected expansionary boost is steadily removed, and by the time of Bernanke’s chairmanship starting in early 2006 the residuals move around zero. However, in late 2007 an expected expansionary boost emerges and spikes in early 2008 as the recession gains momentum. The negative residuals just as quickly vanished around 2009.

What accounts for the residuals’ fluctuation? Atkeson and Kehoe (2009) argue that “business cycle movements of the short rate arise as a result of the Fed’s endogenous policy response to exogenous business cycle fluctuations in risk.” Hence, an important component of these residuals’ evolution may come from forecasters’ perceived risk.

We first proxy for risk by the term premium, the difference between 10-year and 2-year bond yields, which previous literature suggests captures investors’ perception of financial market risk at a business cycle frequency. A regression of the residuals on the term spread

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13. We also ran a regression of the residuals on nowcasts of unemployment and inflation (as well as their forecasts) to check if the shifts reflect economic conditions at the time of the forecasts or dynamic adjustment. The coefficients were statistically insignificant and the variation explained minimal.

14. We are careful here to use the term premium as measured in the month before the survey month, since the economic indicator survey is collected during the first few days of the month it is released, therefore the prior month’s bond prices are what is known to survey respondents at the time they report their forecasts.
shows a strong connection that is both economically and statistically significant: The estimate is $-0.25$ with standard error of 0.03 and $R^2$ of 15%. The coefficient estimate implies that forecasters translate a 100 basis point increase in the term spread with about a 25 basis point decrease in future policy action, irrespective of what inflation and the unemployment rate are forecasted to be.

The dashed line in Figure 4 plots the residuals that are fitted to the term premium. The fit is reasonably good at a business cycle frequency during the 2000s until the crisis begins at the end of 2008. Thereafter, the fitted values move in an opposite direction from the residuals. We find this interesting because one of the Fed’s unconventional policy responses during the ZLB period has been to target medium-maturity treasuries for large-scale asset purchases, directly affecting the measured term premium. If our policy residuals are indeed capturing the market’s perceived risk, then the term-premium fitted values might move contrarily during the crisis because the Fed is directly lowering the measured term premium yet not lowering overall market risk.

Our second proxy for risk is the VIX, which is available only since 1990, and its fitted values are plotted as the dotted line in Figure 4. We find no evidence of a statistically significant connection between the VIX and the estimated policy residuals. This is perhaps

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15 Using robust or HAC standard errors makes very little difference: The $t$-stat remains below $-6$. 
unsurprising given Atkeson and Kehoe’s (2009) argument that the Fed responds to risk at a business cycle frequency, instead of higher frequency financial market risk. Hence, the VIX may largely capture financial market conditions to which the Fed does not respond, in the view of market forecasters.

4.3 Making Use of Estimation Results

In any regression analysis, estimation results are usually used for two purposes: In-sample and out-of-sample analysis. To perform an out-of-sample analysis, it is important to have an in-sample estimation period that has a similar characteristic to the out-of-sample period in mind. Since it is commonly believed that many economic phenomena might have changed with the emergence of the global financial crisis, it is critical to estimate policy rules with a sample after 2008—as we do in this paper.

We make use of estimation results based on forecast data between December 2008 and August 2011 in two ways. First, noting that the target federal funds rate is zero during this period and afterwards, it is of immense interest what the unconstrained nominal short-term interest would be were it not for the zero lower bound (see Rudebusch 2009). As of this writing the target funds rate is at the zero lower bound, and one of the most important topic in financial circles is when the short rate will be lifted from the zero lower bound. The second use of our estimation results aim to answer this question.

4.3.1 Unconstrained Short-Term Rate

In Figure 5, the solid line depicts the actual historical short-term interest rate until December 2008. For the ensuing period until August 2011—which corresponds to the latter part of our estimation sample—the solid line represents the short rate that our benchmark estimation produces when the independent variables are the actual economic outcomes. This line implies the unconstrained interest rate if there were no ZLB and if the estimated policy rule set the short rate to any value. This implied (or recommended) rate decreased until late 2010, when the rate is about $-3\%$; afterwards, the implied rate moved upward as economic conditions bottomed out.

Our estimation sample ends in August 2011, when the FOMC introduced forward guidance that was effective beyond the one-year horizon. However, this does not prevent us from extending the calculation of implied rates after August 2011. The solid line until October 2012 (when this draft is written) is based on the historical outcomes of inflation and unemployment. As economic conditions kept improving, implied rates were on an increasing

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16A one-time intercept shift is allowed at the start of this period, which allows for a smooth transition from actual rates and implied rates. Figure 4, at the December 2008 vertical line, shows that this intercept shift is small.
Figure 5: Short Rates Recommended by a ZLB Policy Rule

Notes: Through October 2012 uses: actual unemployment and inflation data; Blue Chip long-horizon unemployment, inflation and real rate forecasts; and the ZLB parameter estimates in Table 2. From November 2012 onwards, uses: unemployment and inflation forecasts from the November Blue Chip survey; long-horizon unemployment, inflation and real rate forecasts from the October Blue Chip survey; and the ZLB parameter estimates in Table 2. See the text for further details.

4.3.2 Timing of the Lift-Off from the ZLB

Starting with November 2012, the solid line of Figure 5 is based on the economic forecasts that are available in the November 2012 Blue Chip issue. This line indicates that the short rate is projected to be below zero for all of 2013. It is interesting to project when the rate will be higher than 25 basis points; this will be possible using future Blue Chip issues as the end-point of the Fed’s forward-guidance comes close.\footnote{It may also be interesting to compare the contour of the solid line as we update the Blue Chip survey from month to month, which we do not do.}

To investigate the determinants of the lift-off, our analysis provides a decomposition of the lift-off into different components. In particular, our policy rule accounts for changes in recommended policy as due to the forecasted inflation gap (the dashed line) and the forecasted unemployment gap (the dotted line). These two lines of Figure 5 show that economic conditions are projected to return to normal only very slowly. The sideways movement of the unemployment gap (actual and forecast) is largely responsible for the short rate’s sideways movement from 2012 onwards. However, we also have not seen a rise in the inflation gap
(actual or forecast) despite the Fed’s expansionary forward guidance.

5 Conclusion

It is crucial to policymakers that they understand how their behavior is perceived by financial markets. Given the ongoing global financial crisis, it is quite natural to ask: Have policymakers’ behavior, or markets’ perceptions thereof, changed? But the conventional method of empirically answering this question is unavailable because short rates are censored by the zero lower bound. We get around this hurdle by using surveys of forecasts to estimate by simple regression the policy rule that has prevailed, at least in professional forecasters’ perception, since the onset of the global financial crisis.

We find that the Fed’s inflation response has significantly fallen by more than half. This could portend troublesome times ahead when the Fed does start to combat inflation. This could instead reflect the success of the Fed’s expansionary forward guidance that “a highly accommodative stance of monetary policy will remain appropriate for a considerable time after the economic recovery strengthens” (FOMC 2012). Meanwhile, the Fed’s response to unemployment has stayed strong throughout the pre-crisis and crisis periods.

Looking ahead, the Fed’s forward guidance on August 2011 sent even one-year-ahead short rate forecasts to the ZLB. However, as we near the end of the Fed’s “highly accommodative stance,” one-year-ahead short rate forecasts will once again take on meaningful variation. At that time it will be imperative to measure of the Fed’s perceived behavior, and this paper’s approach will ably do so.
References


