

# High Frequency Identification of Monetary Non-Neutrality

Emi Nakamura and Jón Steinsson\*

Columbia University

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## Abstract

We provide new evidence on the responsiveness of real interest rates and inflation to monetary shocks. Our identifying assumption is that the increase in the volatility of interest rate news in a 30-minute window surrounding scheduled Federal Reserve announcements arises from news about monetary policy. Real and nominal yields and forward rates at horizons out to 3 years move close to one-for-one at these times implying that changes in expected inflation are small. At longer horizons, the response of expected inflation grows. Accounting for “background noise” in interest rates on FOMC days is crucial in identifying the effects of monetary policy on interest rates, particularly at longer horizons. We show that in conventional business cycle models with nominal rigidities our estimates imply that monetary non-neutrality is large. Our estimates also imply substantial inflation inertia.

Keywords: Real Interest Rates, Term Structure, Identification by Heteroskedasticity.

JEL Classification: E30, E40, E50

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# 1 Introduction

A fundamental question in macroeconomics is how monetary policy affects the economy. The key empirical challenge in answering this question is that most changes in interest rates happen for a reason. For example, the Fed might lower interest rates to counteract the effects of an adverse shock to the financial sector. In this case, the effect of the Fed's actions are confounded by the financial shock, making it difficult to identify the effects of monetary policy. Two sources of existing evidence are structural vector autoregressions (e.g., Christiano, Eichenbaum, and Evans, 1999) and Romer and Romer's (2004) approach of looking at the effects of changes in the intended federal funds rate that are orthogonal to the Fed's information set as measured by its staff forecast. The concern remains, however, that not all endogenous variation has been purged from these measures of monetary shocks.

An alternative approach—the one we pursue in this paper—is to focus on movements in bond prices in a narrow window around scheduled Federal Open Market Committee (FOMC) meetings. This approach was pioneered by Cook and Hahn (1989), Kuttner (2001), and Cochrane and Piazzesi (2002). Key to this method is that, while interest rates are continually being affected by many factors, monetary news is revealed in a lumpy fashion, with a disproportionate fraction of news revealed at the time of FOMC announcements. Since bond prices adjust in real-time to news about the macroeconomy, movements in bond prices at the time of an FOMC announcement reflect the effect of news about current and future monetary policy. This is important for identification since it strips out endogenous variation in interest rates associated with other shocks than monetary shocks. For example, a positive employment announcement that occurs several days or even hours before an FOMC announcement will already have been factored into bond prices when the Fed makes its announcement.

While few would dispute the Fed's impact on nominal interest rates, considerable debate exists about the magnitude of the Fed's impact on the real economy. Using data from the Treasury Inflation Protected Securities (TIPS) market, we show that the monetary policy shocks we identify have large and persistent effects on the real yield curve. A monetary shock that raises the 2-year nominal yield on Treasuries by 105 basis points, raises the 2-year real TIPS yield by 100 basis points. The effect of this shock on the 2-year instantaneous real forward rate is 86 basis points. The impact of the shock then falls monotonically at longer horizons to 72 basis points at 3 years, 39 basis points at 5 years,

and 9 basis point at 10 years.<sup>1</sup> We can infer the response of market expectations about inflation by taking the difference between the response of nominal and real rates. At horizons of 2 and 3 years, the response of inflation to our monetary shock is small relative to the response of real interest rates. At longer horizons, the response of inflation grows to become significantly negative.

Our identification is based on a heteroskedasticity-based estimator developed by Rigobon (2003) and Rigobon and Sack (2004). Early work using FOMC announcements to identify monetary shocks assumed that no other shocks occur on FOMC announcement days. Our identification assumption is weaker. We allow for the possibility that some movements in interest rates in a narrow window around FOMC announcements are associated with non-monetary shocks (we consider a 30-minute window and a 1-day window). Our identifying assumption is that the increase in volatility of interest rates at the time of FOMC announcements is due to monetary news. We use a control sample of equally sized windows of time to adjust for the effect of other shocks which are assumed to have equal variance at the time of FOMC announcements as in the control sample. We show that accounting for “background noise” in interest rates in this way has important implications for the identification of monetary shocks, especially when one uses a 1-day window.<sup>2</sup>

We next investigate the implications of our estimates for the degree of monetary non-neutrality. We show that in a textbook New Keynesian model, key parameters of the model are identified by the relative magnitude of the response of inflation and the response of real interest rates to a monetary shock. Intuitively, there are two forces at play here. First, the Euler equation implies that an increase in the real interest rate leads to a decrease in output. The strength of this force is governed by the intertemporal elasticity of substitution (IES). Second, the resulting decrease in output leads firms to reduce their prices, generating a fall in inflation. The strength of this force is governed by the extent of nominal and real rigidities—i.e., the extent of price adjustment frictions. If the response of inflation to a monetary shock is small relative to the response of real interest rates, this implies that output does not respond much to the real interest rate (small IES) or prices do not respond

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<sup>1</sup>Hanson and Stein (2012) employ a similar high-frequency approach to study the impact of monetary shocks on long-term real interest rates. Our results differ significantly from theirs in that their measure of monetary shocks has a significant effect on instantaneously real forwards even at the 10-year horizon. A key difference is the heteroskedasticity-based estimation approach we employ (discussed below), which accounts for “background noise” in interest rates.

<sup>2</sup>Gurkaynak, Sack, and Swanson (2005) and Fleming and Piazzesi (2005) use intra-day data to assess the impact of FOMC actions on nominal interest rates. This sharply reduces the amount of background noise in interest rates. Wright (2012) uses Rigobon’s identification by heteroskedasticity approach to identify the effects of unconventional monetary policy on interest rates during the recent period over which short-term nominal interest rates have been at their zero lower bound.

much to output (large nominal and real rigidities) or both.

In addition, textbook New Keynesian models imply that inflation is purely forward looking. This means that the largest response of inflation should be immediately following the monetary shock, when all the high real interest rates are in the future. The response of inflation should then dissipate as the response of real interest rates dies out. In contrast to this, the response of inflation that we estimate in the data builds over time. This suggests a model with a substantial degree of inflation inertia—i.e., a lagged inflation term in the Phillips curve.

We next use our empirical evidence on the response of nominal and real interest rates and inflation to a monetary shock to formally estimate the workhorse business cycle model proposed by Christiano, Eichenbaum and Evans (2005, CEE) and further developed by Altig et al. (2011, ACEL) using a simulated method-of-moments approach. Here, we follow in the tradition of Rotemberg and Woodford’s (1997) and CEE’s seminal work on estimating key parameters of business cycle models by matching model-generated impulse responses to those generated by a structural VAR. A key difference is that our empirical impulse responses are estimates using the high-frequency identification approach described above.

Our estimation yields stark conclusions regarding the magnitude of monetary non-neutrality. The summary measure of monetary non-neutrality we use is the relative magnitude of cumulative output and inflation responses following our monetary shock. This measure is 16.2 for our estimation of the CEE/ACEL model, with a confidence interval of [6.8, 146.7]. For comparison, ACEL’s and CEE’s parameter estimates imply values of 10.7 and 1.4 respectively. In other words, our estimates imply a comparable (somewhat larger) degree of monetary non-neutrality than ACEL, and a much more monetary non-neutrality than CEE. The CEE/ACEL model builds in the inflation inertia we discuss above. Our estimation of this model therefore succeeds in matching the delayed response of inflation following a contractionary monetary policy shock.

An important question is whether some of the effects on longer-term real interest rates we estimate reflect risk premia as opposed to changes in expected future short-term real interest rates. A key point is that constant risk premia do not affect our results, since our identification is based on changes in bond yields at the time of FOMC announcements. To address the possibility that risk premia may change at the time of FOMC announcements, we study the effect of our monetary shocks on expected real rates using direct measures of expectations from the *Blue Chip Economic Indicators*, which surveys professional forecasters on their beliefs about future interest rates and inflation. Since

these data are direct measures of expectations, they are immune from risk premium effects. While our estimates based on this approach are less precise than those based on asset prices, they support a similar time-pattern of effects on real interest rates and a small inflation response. Furthermore, we find no evidence that the interest rate effects we identify dissipate quickly after the announcement, as would be predicted by some models of liquidity premia.<sup>3</sup>

In recent years, FOMC announcements have revealed more surprise news about the future path of nominal interest rates than about contemporaneous changes in the Fed Funds rate, which have, to a substantial extent, been anticipated in advance (Gurkaynak, Sack, and Swanson, 2005). The Fed affects longer-term interest rates via its effects on expectations of the future Fed Funds rate. In identifying the effects of monetary shocks, we therefore focus on a “policy news shock” equal to the first principal component of unexpected changes at the time of FOMC announcements in nominal interest rates over the year following an FOMC meeting.<sup>4</sup> We show that analyzing this shock has advantages in terms of distinguishing among alternative models of the economy, since real interest rates far into the future have particularly large effects on consumer and firm behavior.

In the above discussion, we have implicitly assumed that FOMC announcements change the private sector’s beliefs about current and future monetary policy without providing the private sector with new information about the state of the economy. In other words, we have been assuming that the Fed does not have an informational advantage over the private sector. In this case, FOMC announcements may provide the private sector with information about the preferences of the Fed—i.e., how tough they are on inflation—but it may also provide the private sector with information about the Fed’s beliefs about the current and future state of the economy. Even if the Fed and the private sector have the same information set, they may hold different beliefs about the future path of the natural rate of output if they interpret the information differently (perhaps due to believing in somewhat different models).

A potential alternative interpretation of our results to the one we emphasize above is that FOMC announcements reveal information to the private sector about the state of the economy. If this is the case and the private sector believes that the Fed will conduct monetary policy in such a way as to make sure the real interest rate tracks the “natural rate of interest”—i.e., the real interest that

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<sup>3</sup>Hanson and Stein (2012) present a behavioral model in which “search for yield” generates significant risk premium effects of monetary shocks.

<sup>4</sup>Our policy news shock is closely related to the “path factor” studied by Gurkaynak, Sack, and Swanson (2005).

would prevail if prices were fully flexible—then FOMC announcements may affect the expected path of real interest rates without affecting output and inflation. Our empirical evidence is consistent with this interpretation at the short end of the term structure, but not at the long end. In addition, for this effect to be important, the Fed must have a considerable informational advantage relative to private markets arising either from superior data or superior analytical capacities.

The paper proceeds as follows. Section 2 describes the data we use in our analysis. Section 3 describes the construction of our policy news shock and presents our main empirical results regarding the response nominal and real interest rates and inflation to the policy news shock. Section 4 shows what structural parameters our empirical evidence provides information on in the context of a textbook New Keynesian model and quantitatively assess the degree of monetary non-neutrality implied by our empirical evidence by estimating the CEE/ACEL model using indirect inference. Section 5 concludes.

## 2 Data

We use data on interest rates from several sources. First, we use tick-by-tick data on Fed Funds futures and Eurodollar futures from the CME Group (owner of the Chicago Board of Trade and Chicago Mercantile Exchange). Fed Fund futures have been traded since 1989, while Eurodollar futures began trading in the early 1980's. For each month, we make use of the current month's Fed Funds futures contract, the next month's Fed Funds futures contract and the Fed Funds futures contract for the month of the next FOMC meeting (which typically occurs in one or two months). And we make use of the Eurodollar futures that expire in approximately two, three and four quarters.

The Federal Funds futures contract for a particular month (say April 2004) trades at price  $p$  and pays off  $100 - \bar{r}$  where  $\bar{r}$  is the average of the effective Federal Fund Rate over the month. The effective Federal Fund Rate is the rate that is quoted by the Federal Reserve Bank of New York on every business day. This rate is computed as a weighted average rate from trades that day. The price of the futures contract can, thus, be used to construct market based expectations of the average Fed Funds rate over the month in question.<sup>5</sup>

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<sup>5</sup>See the Chicago Board of Trade Reference guide [http://www.jamesgoulding.com/Research\\_II/FedFundFutures/FedFunds\(FuturesReferenceGuide\).pdf](http://www.jamesgoulding.com/Research_II/FedFundFutures/FedFunds(FuturesReferenceGuide).pdf) for a detailed description of Fed futures contracts. On a trading day in March (say), the April Federal Funds futures contract is labeled as 2nd expiration nearby and also as 1st beginning nearby, in reference to the month over which  $\bar{r}$  is computed.

A Eurodollar futures contract expiring in a particular quarter (say 2nd quarter 2004) is an agreement to exchange, on the second London business day before the third Wednesday of the last month of the quarter (typically a Monday near the 15th of the month), the price of the contract  $p$  for 100 minus the then current three-month US dollar BBA LIBOR interest rate. The contract thus provides market-based expectations of three month nominal interest rates on the expiration date.<sup>6</sup>

To measure movements in Treasuries at horizons of 1 year or more, we use daily data on zero-coupon nominal treasury yields and instantaneous forward rates constructed by Gurkaynak, Sack, and Swanson (2007). These data are available on the Fed’s website at <http://www.federalreserve.gov/pubs/feds/2006/200628/200628abs.html>. We also use the yields on 3M and 6M Treasury bills. We retrieve these from the Federal Reserve Board’s H.15 data release.

To measure movements in real interest rates, we use zero-coupon yields and instantaneous forward rates constructed by Gurkaynak, Sack, and Wright (2010) using data from the TIPS market. These data are available on the Fed’s website at <http://www.federalreserve.gov/pubs/feds/2008/200805/200805abs.html>. TIPS are “inflation protected” because the coupon and principal payments are multiplied by the ratio of the reference CPI on the date of maturity to the reference CPI on the date of issue.<sup>7</sup> The reference CPI for a given month is a moving average of the CPI two and three months prior to that month, to allow for the fact that the Bureau of Labor Statistics publishes these data with a lag.

TIPS were first issued in 1997 and were initially sold at maturities of 5, 10 and 30 years, but only the 10-year bonds have been issued systematically throughout the sample period. Other maturities have been issued more sporadically. While liquidity in the TIPS market was initially poor, TIPS now represent a substantial fraction of outstanding Treasury securities. We start our analysis in 2000 to avoid relying on data from the period when TIPS liquidity was limited.

We obtain the dates and times of FOMC meetings up to 2004 from the appendix to Gurkaynak, Sack, and Swanson (2005). We obtain the dates of the remaining FOMC meetings from the Federal Reserve Board website at <http://www.federalreserve.gov/monetarypolicy/fomccalendars.htm>. For the latter period, we verified the exact times of the FOMC announcements using the first news article about the FOMC announcement on *Bloomberg*.

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<sup>6</sup>See the CME Group Eurodollar futures reference guide <http://www.cmegroup.com/trading/interest-rates/files/eurodollar-futures-reference-guide.pdf> for more details about how Eurodollar futures are defined.

<sup>7</sup>This holds unless inflation is negative, in which case no adjustment is made for the principle payment.

We use data on inflation swaps from *Bloomberg*. An inflation swap is a financial instrument designed to help investors hedge inflation risk. As is standard for swaps, nothing is exchanged when an inflation swap is first executed. However, at the maturity date of the swap, the counterparties exchange  $R_t^x - \Pi_t$ , where  $R_t^x$  is the  $x$ -year inflation swap rate and  $\Pi_t$  is the reference inflation over that period. If agents were risk neutral, therefore,  $R_t^x$  would be expected inflation over the  $x$  year period.

Finally, we use data on expectations of future nominal interest rates and inflation from the *Blue Chip Economic Indicators*. *Blue Chip* carries out a survey during the first few days of every month soliciting forecasts of these variables for up to the next 8 quarters.

### 3 Empirical Analysis

Our goal in this section is to identify the effect of the monetary news contained in scheduled FOMC announcements on nominal and real interest rates and inflation. Our identification approach makes use of the discontinuous increase in the volatility of monetary shocks at the time of FOMC announcements. We therefore consider changes in interest rates in a narrow window around FOMC announcements. We consider two time intervals. The first is a 30 minute window from 10 minutes before the FOMC announcement to 20 minutes after it. The second is a 1 day window from the close of markets the day before the FOMC meeting to the close of markets the day of the FOMC meeting.

In their post-meeting announcements, the FOMC conveys information not only about immediate changes in the Federal Funds Rate but also about likely changes in monetary policy at later dates. In fact, over the last 15 years, changes in the Federal Funds Rate have often been largely anticipated by markets once they occur, while FOMC announcements have come to focus more and more on guiding expectations about future changes in the Federal Funds Rate (Gurkaynak, Sack, and Swanson, 2005). Motivated by these developments, we construct a measure of monetary policy news  $\Delta i_t$  by taking the first principle component of changes in five interest rates of maturity less than one year which can be inferred from futures data. We use Federal Funds futures and Eurodollar futures to infer changes in the market's expectations about the Federal Funds rate immediately following the FOMC meeting, the Federal Funds rate immediately following the next FOMC meeting, and the 3-month

Eurodollar interest rate at horizons of two, three and four quarters.<sup>8</sup> We refer to  $\Delta i_t$  as the “policy news shock.”<sup>9</sup> The scale of the policy news shock is arbitrary. For convenience, we rescale it such that an OLS regression of the 1-year Treasury yield on the policy news shock yields a coefficient of one. Appendix A provides details about the construction of the policy news shock.

### 3.1 Identification

If we were confident that movements in the policy news shock  $\Delta i_t$  over the windows of time we consider around FOMC announcements were due to monetary shocks and nothing else, then this variable would constitute a pure measure of monetary shocks. We could thus regress any other variable of interest on the policy news shock to assess the effect of monetary shocks on that variable. This is the approach taken by Cook and Hahn (1989), Kuttner (2001) and Cochrane and Piazzesi (2002) (all with a one-day window) and more recently by Hanson and Stein (2012) (with a two-day window). A potential concern with this approach is that other shocks may occur over the course of FOMC days. Interest rates fluctuate substantially on non-FOMC days. This suggests that other shocks than FOMC announcements affect interest rates on FOMC days. There is no way of knowing whether these other shocks are monetary shocks or non-monetary shocks.

We would, therefore, like to allow for “background noise” in interest rates on both FOMC and non-FOMC announcement days. To this end we adopt a heteroskedasticity-based estimator of monetary shocks developed by Rigobon (2003) and Rigobon and Sack (2004). Let  $\epsilon_t$  denote a pure monetary shock and suppose that movements in the policy news shock we measure in the data is governed both by monetary and non-monetary shocks:

$$\Delta i_t = \alpha_i + \epsilon_t + \beta_i \eta_t, \tag{1}$$

where  $\eta_t$  is a vector of all other shocks that affect  $\Delta i_t$ . Here  $\alpha_i$  and  $\beta_i$  are constants and we normalize the impact of  $\epsilon_t$  on  $\Delta i_t$  to one. We wish to estimate the effects of the monetary shock  $\epsilon_t$  on an outcome variable  $s_t$ . This variable is also affected by both the monetary and non-monetary shocks:

$$\Delta s_t = \alpha_s + \gamma \epsilon_t + \beta_s \eta_t. \tag{2}$$

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<sup>8</sup>More precisely, the expiration date of the “ $n$  quarter” Eurodollar future is between  $n - 1$  and  $n$  quarters in the future at any given point in time. See our discussion in section 2 on the exact expiration dates of Eurodollar futures.

<sup>9</sup>Our policy news shock variable is closely related to the path factor considered by Gurkaynak, Sack, and Swanson (2005). The five interest rate futures that we use to construct our policy new shock are the same five futures as Gurkaynak, Sack, and Swanson (2005) use. They motivate the choice of these particular futures by liquidity considerations.

Our objective is to estimate  $\gamma$ , which should be interpreted as the impact of the pure monetary shock  $\epsilon_t$  on  $s_t$  relative to its effect on  $i_t$ .

Our identifying assumption is that the variance of monetary shocks increases at the time of FOMC announcements, while the variance of other shocks is unchanged. Define  $R1$  as a sample of narrow time intervals around FOMC announcements, and define  $R2$  as a sample of equally narrow time intervals that do not contain FOMC announcements but are comparable on other dimensions (e.g., same time of day, same day of week, etc.). We refer to  $R1$  as our “treatment” sample and  $R2$  as our “control” sample. Our identifying assumption is that  $\sigma_{\epsilon,R1} > \sigma_{\epsilon,R2}$ , while  $\sigma_{\eta,R1} = \sigma_{\eta,R2}$ .

We show in Appendix B that given these assumptions  $\gamma$  is given by

$$\gamma = \frac{\text{cov}_{R1}(\Delta i_t, \Delta s_t) - \text{cov}_{R2}(\Delta i_t, \Delta s_t)}{\text{var}_{R1}(\Delta i_t) - \text{var}_{R2}(\Delta i_t)}. \quad (3)$$

Notice that if we set the variance of the “background noise”  $\eta_t$  to zero, then this estimator reduces to the coefficient from an OLS regression of  $\Delta s_t$  on  $\Delta i_t$ . Intuitively, the full heteroskedasticity-based estimator can be thought of as the simple OLS estimator, adjusted for the “normal” covariance between  $\Delta s_t$  and  $\Delta i_t$ .

As we discuss above, we present results where the policy news shock is constructed using 30-minute and 1-day time intervals surrounding FOMC announcements. Our control samples are then 30-minute or 1-day intervals that are chosen to be as comparable as possible except that they do not include FOMC announcements. Specifically, in the case of 30-minute windows, we choose the same 30-minute window (from 2:05pm to 2:35pm) on all non-FOMC Tuesdays and Wednesdays as our control sample (since scheduled FOMC meetings tend to occur on Tuesdays and Wednesdays), and in the case of 1-day windows, we choose all non-FOMC Tuesdays and Wednesdays as our control sample.<sup>10</sup> For our treatment sample, we focus on only scheduled FOMC meetings, since unscheduled meetings may occur in reaction to other shocks and thus be endogenous. In all cases, the outcome variables are measured over a 1-day window. Our sample period starts on January 1st 2000 and extends to January 25th 2012. We drop data before 2000 because of concerns about liquidity or TIPS and because very few TIPS securities were trading at the time. In our baseline analysis, we drop the second half of 2008 and the first half of 2009 to avoid the period when disruption of financial markets in the Great Recession was most severe.

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<sup>10</sup>For the case of 30-minute windows, we have also tried using a 30-minute window one hour before FOMC announcements on FOMC days as our control sample. This yield very similar results.

### 3.2 Main Estimates

Table 1 presents our baseline estimates of monetary shocks on nominal and real interest rates and inflation. The first column presents the effects of the policy news shock on nominal Treasury interest rates. By construction, the policy news shock has large effects on nominal yields. The effect of a 1% policy news shock on the zero-coupon 2-Year Treasury Yield is 105 basis points, and declines monotonically to 29 basis points at 10 years. Since longer-term yields reflect expectations about the average short-term interest rate over the life of the long bond, it is easier to interpret the time-path of the response of instantaneous forward rates. A 2-year instantaneous forward rate (say) is the short-term interest rate that the market expects to prevail in 2 years time. The impact of our policy news shock on forward rates is also monotonically declining in maturities. For maturities of 2, 3, 5, and 10 years, its effects on forward rates are 100, 60, 13 and -13 basis points, respectively. We show below that the negative effect on long-horizon nominal interest rates reflects a decline in long-horizon inflation expectations.

The second column of Table 1 presents the effects of the policy news shock on real interest rates measured using TIPS. While the effects on nominal rates are by construction, the impact of monetary shocks on real interest rates is not. In neoclassical models of the economy, the Fed controls the nominal interest rate but has no impact on real interest rates. Our estimate of the impact of a 1% the policy news shock on the 2-year real yield is 100 basis points, and the impact on the 3-year real yield is 94 basis points. Once again, the time-path of effects is easier to interpret using evidence on instantaneous forward rates. The effect of the shock on the 2-year real forward rate is 86 basis points. It falls monotonically at longer horizons to 72 basis points at 3 years, 39 basis points at 5 years, and 9 basis point at 10 years (which is not statistically significantly different from zero). Evidently, monetary policy shocks can affect real interest rates for substantial amounts of time. However, in the long-run, the effect of monetary policy shocks on real interest rates is zero as theory would predict.

The third column of Table 1 presents the effect of the policy news shock on break-even inflation, calculated as the difference between the nominal and real interest rate effects. The first several rows provide estimates based on bond yields, which indicate that the inflation response is small. The shorter horizon estimates are actually slightly positive but but then becomes negative at longer horizons. None of these estimates are statistically significantly different from zero. Again, it is

helpful to consider instantaneous forward inflation rates to get estimates of inflation at points in time in the future. The inflation response implied by the 2 year forwards is slightly positive, though statistically insignificant. The inflation response is negative at longer horizons: for maturities of 3, 5 and 10 years, the effect is -12, -27 and -22 basis points. It is only the responses at 5 and 10 years that are statistically significantly different from zero. Our evidence thus points to inflation responding quite gradually to monetary shocks that have a substantial effect on real interest rates. In section 4 below, we discuss what we can infer about the structure of the economy from these estimates.

Our policy news shock captures the effects of FOMC meetings on expectations about nominal interest rates over the next year. An alternative approach would be to focus on the impact of FOMC announcements on market expectations about the level of the Federal Funds Rate immediately following the announcement. This is the approach taken by much of the early literature. For example, Cochrane and Piazzesi (2002) consider changes in one-month Eurodollar rates at the time of FOMC announcements as a proxy for changes in expectations about the Federal Funds Rate. The disadvantage of this approach, however, is that it captures less of the variation in interest rates in response to monetary shocks than the policy news shock we construct. The remaining columns of Table 1, nevertheless, present estimates based on this approach. The conclusions are very similar. Nominal and real rates respond by roughly the same amount at horizons out to about 3 years. At longer horizons, the response of nominal rates is smaller than real rates, implying that inflation falls.<sup>11</sup>

### 3.3 Alternative Estimates

Table 2 compares our baseline methodology to alternative methods of identifying the monetary policy shock. The first two columns reproduce our baseline results for nominal and real yields. These results are based on Rigobon’s heteroskedasticity-based estimator, and use a 30-minute interval to measure the policy news shock. We first compare these results to results using a one-day window to compute the monetary policy shock. Columns 3 and 4 present estimates based on applying the Rigobon estimator with a 1-day window. The standard errors on these estimates are very large. Intuitively, there is too much “background noise” in the policy news shock variable over a 1-day window to be able to estimate its effect on the term structure with any precision.

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<sup>11</sup>Beechey and Wright (2009) analyze the effect of Federal Funds rate shocks at the time of FOMC announcements on nominal and real 5-year and 10-year yields and the five-to-ten year forward for the sample period February 17th 2004 to June 13th 2008. Their results are similar to ours for the 5-year and 10-year yields.

Columns 5 and 6 present results based on OLS and a 1-day window. These results implicitly make the (much stronger) identifying assumption that *only* monetary shocks occur on the day of an FOMC announcement. A comparison of columns 3 and 4 and columns 5 and 6 shows that OLS massively underestimates the standard errors on the estimated effects of monetary policy shocks relative to the Rigobon estimator. The much larger Rigobon standard errors reflect the large amount of “background” noise in interest rates over an entire day. These differences show that the OLS identifying assumption is too strong when a 1-day window is being used.

These concerns loom even larger when longer-term interest rates are used as proxies for monetary shocks. Columns 7 and 8 present the results of applying the Rigobon estimator with the monetary shock measure  $\Delta i_t$  constructed as one-day changes in the two-year nominal yield. The standard errors are even larger than in the case of the policy-news shock, and are in most cases many times larger than the coefficient of interest. These results arise because of the large amount of background noise in longer-term interest rates. The increase in volatility associated with FOMC announcements is not large enough over a one-day horizon to accurately assess its impact on the term structure.

Interestingly, the Rigobon and OLS estimation approaches yield quite similar results when applied to the case where the policy news shock is measured over a 30-minute window. Columns 9 and 10 present results for a 30-minute window based on OLS estimation. Even with this narrow window, OLS yields standard errors that are somewhat too tight. However, the difference is much smaller than with a one-day window. Intuitively, the difference in volatility of monetary shocks in the 30-minute window surrounding an FOMC announcement relative to other 30-minute windows is much larger than the difference over entire FOMC days relative to other days, implying that the “background noise” effect is much smaller when a 30-minute window is used.

The analysis in tables 1 and 2 is for the sample period from Jan 1st 2000 to Jan 25th 2012, except that we drop the period spanning the height of the financial crisis in the second half of 2008 and the first half of 2009. Numerous well-documented asset pricing anomalies arose during this crisis period, and we wish to avoid the concern that our results are driven by these anomalies. We have, however, also carried out our analysis on the full sample including the crisis, as well as a more restrictive data sample ending at the beginning of 2008. Table A.1 presents the results of our analysis for these two alternative sample periods. The pre-crisis sample yields very similar results to the baseline sample. For the full sample the response of both nominal and real rates is somewhat larger at longer horizons. In all three cases, the effect of the monetary shock on inflation is initially small and positive, but

becomes increasingly negative at longer horizons.

### 3.4 Survey Measures of Interest Rates and Inflation

An important question when it comes to interpreting our results is to what extent the movements in long-term interest rates we identify reflect movements in risk premia as opposed to changes in expected future short-term interest rates. In this regard, it is important to keep in mind that constant risk premia will not affect our results, since our identification is based on changes in bond yields at the time of FOMC announcements. However, if risk premia change at the time of FOMC announcements this could confound our results.

To study this issue directly, we analyze the impact of our policy news shock on direct measures of expectations from the *Blue Chip Economic Indicators*. *Blue Chip* surveys professional forecasters on their beliefs about macroeconomic variables over the next two years in the first few days of every month. We study the impact of our policy news shock on survey expectations about future short-term interest rates and inflation. By construction, these effects reflect expected movements in rates, as opposed to risk premium effects.

We measure the change in expected interest rates for a particular quarter in the future by the change in the *Blue Chip* forecast about that quarter from one month to the next. We regress this measure on the the sum of the policy news shocks that occur over the month except for those that occur in the first week (because we do not know whether these occurred before or after the survey response). We use *Blue Chip* forecasts of the 3-month T-Bill rate and the GDP deflator in our analysis. We construct a measure of expected short-term real interest rates by taking the difference between the expected 3-month T-bill rate and the expected GDP deflator for a given quarter. Unfortunately, *Blue Chip* asks respondents only about the current and subsequent calendar year, so fewer observations are available for longer-term expectations, leading to larger standard errors.<sup>12</sup> The sample period for this analysis is January 1995 to January 2012, except that we exclude the apex of the 2008-2009 financial crisis as we do in the rest of our analysis.

Table 3 presents the results of this analysis. The table shows that the policy news shock has a persistent impact on expected short-term interest rates, both nominal and real. The interest rate effects are somewhat larger than in our baseline analysis, but rather noisily estimated. As

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<sup>12</sup>For example, in the last quarter of the year, forecasters are only asked about their beliefs 1-year in advance; while in the first quarter they are asked about their beliefs for the next full 2-years.

in our analysis using financial variables, the effect on expected inflation is small and statistically insignificant at all horizons except that it is marginally significantly negative at 2 quarters. The much larger standard errors on our estimates in this analysis arise from the fact that, unlike in our analysis of financial variables, the changes in survey variables are available only at a monthly as opposed to a daily frequency.

### 3.5 Inflation Swaps

We also consider an alternative measure of inflation expectations based on inflation swap data. Fleckenstein, Longstaff, and Lustig (2013) point out that measures of breakeven inflation from the TIPS and inflation swap markets are not equal and that this difference increases during the crisis. Table 4 compares our estimates of the effects of the policy news shock on breakeven inflation from TIPS to that on inflation from inflation swaps. The sample period for this analysis is limited by the availability of swaps data to beginning in January 1st 2005. The results are quite similar for these two variables at longer horizons. At shorter horizons the “price puzzle”—i.e., the positive inflation effect at the shortest horizons—is larger for the inflation swap data than the TIPS data, though statistically insignificant in both cases.

### 3.6 Evidence of Mean Reversion

Finally, one additional question that merits attention is whether there is any evidence that the effects we identify on nominal and real yields tend to mean-revert over time, as some theories of liquidity premia might predict. Table 5 presents the effects of our policy news shock on nominal and real interest rates at horizons of 5, 10, 20, 60, 125 and 250 trading days. While the estimates are extremely noisy, there is little evidence that the effects on interest rates tend to dissipate over time. Indeed, in most cases, after a dip around 10 days the point estimates appear to grow over time (though, again, the standard errors are extremely large).

## 4 Evidence on Monetary Non-Neutrality

To be able to more clearly interpret our evidence on monetary non-neutrality, we follow in the tradition of work by Rotemberg and Woodford (1997), Christiano, Eichenbaum, and Evans (2005), and others who fit structural models of monetary policy to evidence on the response of real variables

to monetary shocks. Unlike this earlier work, we focus on fitting the response of the real interest rate and inflation to monetary shocks. The key advantage of looking at these variables is that we are able to obtain high frequency evidence on their dynamics—allowing us to identify monetary shocks with weaker assumption than those required by structural VAR’s.

We begin by developing the intuition for what parameters of the New Keynesian model can be identified using our evidence on the real interest rate and inflation responses. We do this in the context of a simple New Keynesian model: a three equation model consisting of a Euler equation, a Phillips curve, and a monetary policy rule. We show that our results on the response of real interest rates and inflation to monetary news provides evidence on two key features of this model: the responsiveness of prices to the real economy and the extent of inflation persistence. We then analyze the quantitative implications of our empirical results for monetary non-neutrality in the workhorse medium-scale business cycle model proposed by Christiano, Eichenbaum, and Evans (2005) and developed further by Altig et al. (2011).

## 4.1 Intuition in a Simple New Keynesian Model

### 4.1.1 Private Sector Behavior

Consider a setting in which private sector behavior can be described by the following Euler equation and Phillips curve:

$$\hat{x}_t = E_t \hat{x}_{t+1} - \sigma(\hat{i}_t - E_t \hat{\pi}_{t+1} - \hat{r}_t^n), \tag{4}$$

$$\hat{\pi}_t = \beta E_t \hat{\pi}_{t+1} + \kappa \zeta \hat{x}_t. \tag{5}$$

Hatted variables denote percentage deviations from steady state. The variable  $\hat{x} = \hat{y}_t - \hat{y}_t^n$  denotes the “output gap”—the difference between actual output  $\hat{y}_t$  and the “natural” level of output  $\hat{y}_t^n$  that would prevail if prices were flexible,  $\hat{\pi}_t$  denotes inflation,  $\hat{i}_t$  denotes the gross return on a one-period, risk-free, nominal bond, and  $\hat{r}_t^n$  denotes the “natural rate of interest.” Both the natural rate of output and the natural rate of interest are functions of exogenous shocks to tastes and technology. Appendix C presents a detailed derivation of these equations from primitive assumptions about tastes and technology. Woodford (2003) and Gali (2008) present textbook treatments.

The Euler equation (4) is common to both Real Business Cycle and New Keynesian models, and describes how households adjust their consumption behavior to movements in the real interest rate. The parameter  $\sigma$  in the Euler equation denotes the intertemporal elasticity of substitution. The

Phillips curve is fundamental to the New Keynesian paradigm, and describes the effects changes on output relative to the natural rate of output on inflation. Intuitively, the greater are price adjustment frictions, the smaller is the slope of the Phillips curve  $\kappa\zeta$ . Conversely, as the model approaches the frictionless limit of a Real Business Cycle model, the slope of the Phillips curve becomes infinite—in other words, prices are perfectly responsive to any deviation of the economy from the natural rate of output.

We have split the slope of the Phillips curve into two parameters  $\kappa$  and  $\zeta$ . The parameter  $\kappa$  governs the degree of nominal rigidity in the economy. Specifically,  $\kappa = (1 - \alpha)(1 - \alpha\beta)/\alpha$ , where  $(1 - \alpha)$  is the frequency of price change and  $\beta$  is the subjective discount factor of households in the model. The parameter  $\zeta$  governs the degree of “real rigidity” in the economy. Specifically,  $\zeta = (\omega + \sigma^{-1})/(1 + \omega\theta)$ , where  $\omega$  is the elasticity of a particular firm’s marginal costs with respect to that firm’s output holding other firms’ output fixed— $\omega$  is a function of the elasticity of labor supply and the curvature of the production function—and  $\theta$  is the elasticity of substitution between different goods in the economy. The numerator in  $\zeta$  reflects the curvature of labor demand and labor supply which imply that marginal costs rise when production rises. The denominator is due to the heterogeneous nature of the the labor markets in the model we lay out in appendix C. Intuitively, when firms in a particular industry raise their prices relative to the firms in other industries, this lowers demand which reduces the wage demands of workers in that industry implying that the firms don’t want to raise their prices as much as they otherwise would.

#### 4.1.2 Monetary Policy and Information Structure

In the simple model described above, good monetary policy varies the short-term interest rate in such a way that it tracks the natural rate of interest. If the monetary authority is able to vary the short-term interest rate so that it perfectly tracks the natural rate of interest, the monetary authority can achieve both a zero output gap and zero inflation (see Woodford, 2003, ch. 4).<sup>13</sup> With this in mind, we specify the following policy rule for the monetary authority:

$$\hat{i}_t - E_t \hat{\pi}_{t+1} = \bar{r}_t + \phi_\pi \hat{\pi}_t. \quad (6)$$

We have written this policy rule as a rule for the short term real interest rate. The first term in the rule is a time varying intercept term. We think of the monetary authority as using this term to track

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<sup>13</sup>Woodford (2003, ch. 4) shows that this constitutes optimal monetary policy in the model presented above.

variation in the natural rate of interest  $r_t^n$ . The second term is a conventional endogenous feedback term implying that the monetary authority raises the real interest rate as inflation increases. If the monetary authority is successful at varying  $\bar{r}_t$  so that it tracks  $r_t^n$ , inflation will be stable at zero and the endogenous feedback term will not come into play.

Let  $\epsilon_{t,t+j}$  denote the change in the private sector's expectations about the intercept term in the Fed's policy rule in period  $t+j$  that results from a Fed announcement at time  $t$ . In other words,  $\epsilon_{t,t+j} = E_t \bar{r}_{t+j} - E_{t-\delta} \bar{r}_{t+j}$ , where  $\delta$  denotes a small amount of time (30 minutes in our empirical work in section 3). To capture the term structure of changes in expectations about interest rates that we estimate occurring at the time of Fed announcements, we assume that

$$\epsilon_{t,t+j} = (\rho_1 + \rho_2)\epsilon_{t,t+j-1} - \rho_1\rho_2\epsilon_{t,t+j-2}. \quad (7)$$

This implies that the entire path of changes in private sector beliefs about Fed behavior at the time of an FOMC meeting can be characterized by three numbers:  $\epsilon_{t,t}$ , which gives the size and direction of the shock, and the parameters  $\rho_1$  and  $\rho_2$ , which govern the term structure of news about future interest rates. We have chosen to parameterize equation (7) in terms of the roots of its lag polynomial for ease of interpretation.

We assume that the Fed does not have an informational advantage over the private sector. However, the private sector and the central bank have different priors about how to interpret the information they receive (different models). Fed announcements lead the private sector to update what it thinks the Fed thinks the path for the natural rate of interest will be in the future. Since the private sector and the Fed have the same information set, the private sector is not using the announcements of the Fed to update its own views about future natural rates. The private sector has already seen all the information that the Fed is basing its announcements on and has incorporated this information into its forecast about the natural rate of interest. This means there is nothing the private sector can learn from the Fed's announcement about the natural rate. The difference in priors about how to interpret information implies that the private sector and the Fed agree to disagree about the future path of the natural rate of interest. Nonetheless, the Fed's views about the natural rate of interest affect the private sector through future monetary policy.

An alternative view would be that the public and the central bank share the same model of the world and therefore agree about how to interpret new information, but the central bank receives additional information about economic fundamentals that the public does not receive directly. In

this case, Fed announcements will lead the private sector to update its view about the future path of the natural rate of interest. We discuss this case in more detail in section 4.3.<sup>14</sup>

### 4.1.3 What Our Evidence Identifies

In this simple model, it is straightforward to show how our evidence on the response of the real interest rate and inflation to monetary shocks identifies key parameters relating to the extent of monetary non-neutrality. Taking as given that monetary shocks have no effect on output in the long run, we can solve the Euler equation (4) forward and get that the response of output to a monetary shock is,

$$\hat{y}_t = -\sigma \sum_{j=0}^{\infty} E_t \hat{r}_{t+j} = -\sigma \hat{r}_t^{\ell}. \quad (8)$$

where  $\hat{r}_{t+j}$  denotes the response of the short-term real interest rate at time  $t + j$ —i.e.,  $\hat{r}_{t+j} = \hat{r}_{t+j} - E_t \hat{r}_{t+j+1}$ —and  $\hat{r}_t^{\ell}$  denotes the response of the long-run real interest rate.<sup>15</sup> This shows that given the response of real interest rates and the assumption that the monetary policy shock has no effect on output in the long-run, the determination of output is a “partial equilibrium” exercise relying only on the Euler equation. The rest of the model does not affect the determination of output.

Similarly, we can solve forward the Phillips curve—equation (5) and get that the response of inflation to a monetary shock is

$$\hat{\pi}_t = \kappa \zeta \sum_{j=0}^{\infty} \beta^j E_t \hat{y}_{t+j}. \quad (9)$$

This shows that the response of inflation is fully determined by  $\kappa \zeta$ —the slope of the Phillips curve—and the sum of the response of output at different horizons. Combining equations (8) and (9), we get a relationship between the response of inflation and the response of real interest rates:

$$\hat{\pi}_t = -\kappa \zeta \sigma \sum_{j=0}^{\infty} \beta^j E_t \hat{r}_{t+j}^{\ell}. \quad (10)$$

If monetary shocks have long-run effects on inflation, equation (10) becomes

$$\hat{\pi}_t = -\kappa \zeta \sigma \sum_{j=0}^{\infty} \beta^j E_t \hat{r}_{t+j}^{\ell} + \hat{\pi}_{\infty}, \quad (11)$$

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<sup>14</sup>These two views of how the public interprets Fed statements are closely related to the notion of endogenous and exogenous monetary policy actions in Ellingsen and Soderstrom (2001).

<sup>15</sup>We will allow for long-run effects of monetary shocks on output below.

where  $\hat{\pi}_\infty$  denotes the long-run response of inflation to the monetary shock.<sup>16</sup> The monetary rule we introduce above implies that  $\pi_\infty = 0$ .

We wish to draw two main conclusions from equation (11). First, the relative size of the response of inflation and real interest rates pins down  $\kappa\zeta\sigma$ . This holds independent of the assumed monetary policy. A small response of inflation relative to the magnitude of the real interest rate response implies a small value of  $\kappa\zeta\sigma$ . In other words, such a pattern of responses implies a large amount of nominal and real rigidities, a small value of the intertemporal elasticity of substitution, or both.<sup>17</sup>

Second, the dynamics of the inflation response to a monetary shock are informative about the degree of inflation inertia in the economy. Equation (11) shows clearly that (almost) irrespective of the values of the parameters of the model, inflation should fall more in the short run than in the long run in response to a positive shock to real interest rates (since positive real interest rate terms “fall out” of the infinite sum on the right hand side of equation (11) as time passes).<sup>18</sup> This effect is illustrated in Figure 1 for particular values of the structural parameters.<sup>19</sup> However, the general shape of the inflation response—initial drop and then increase back to long-run response—is the same (almost) irrespective of the values of these parameters.

Figure 2 presents our estimated response of inflation and nominal and real interest rates in the form of a figure for ease of comparison with the results from the model. In sharp contrast with the predictions of equation (11) the inflation response we estimate in the data is small initially but builds over time. To be able to capture this inflation inertia, we augment the model discussed above by considering a hybrid Phillips curve that allows current inflation to be influenced by past inflation

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<sup>16</sup>The extra term arises because the long-run Phillips curve in our model is not completely vertical (because of discounting). For this reason, a monetary shock can have a (small) permanent effect on output if it has a permanent effect on inflation. Specifically, the Phillips curve implies that  $E_t y_\infty = (1 - \beta)E_t \pi_\infty / (\kappa\zeta)$ , where  $\hat{y}_\infty$  denotes the long-run response of output to a monetary shock. This implies that solving forward the Euler equation yields  $\hat{y}_t = -\sigma r_t^\ell + E_t \hat{y}_\infty$ . Plugging this into equation (9) yields equation (11).

<sup>17</sup>If we had data on the response of output to the monetary shock, we could distinguish between  $\sigma$  and  $\kappa\zeta$ . Distinguishing between  $\kappa$  and  $\zeta$  requires micro data.

<sup>18</sup>The exception to this is if the persistence of the monetary policy shock is sufficiently high (average persistence over time larger than  $\beta$ ). In this case, the fact that the terms further out in the sum are getting closer to the present as time passes will lead inflation to fall over time. Our estimated policy news shock is far less persistent than it would need to be to generate this effect.

<sup>19</sup>The autoregressive parameters  $\rho_1 = 0.93$  and  $\rho_2 = 0.61$  are chosen to roughly match the change in real interest rates at the time of FOMC announcements in the data and we calibrate the coefficient on the endogenous feedback term in the monetary policy rule to  $\phi_\pi = 0.5$ . The value of  $\kappa\zeta\sigma$  is illustrative and we allow for a non-zero value of  $\hat{\pi}_\infty$ , which is also chosen in an illustrative manner. To allow for a non-zero value of  $\hat{\pi}_\infty$  we add a second permanent component to the monetary policy rule.

in addition to future output gaps:

$$\hat{\pi}_t = \gamma \hat{\pi}_{t-1} + \kappa \zeta \sum_{j=0}^{\infty} \beta^j E_t \hat{y}_{t+j}. \quad (12)$$

Phillips curves of this form have been widely used in the recent literature (see, e.g., Woodford, 2003; Christiano et al., 2005).

By choosing a very high degree of inflation inertia ( $\gamma = 0.999$ ), we can match the empirical responses we estimate in section 3 reasonably well. Figure 3 shows that in this case, the model no longer counterfactually implies that inflation “jumps” down before reverting to its steady state. Rather, the response of inflation builds over time and then starts to dissipate slowly. Moreover, the nominal and real interest rate move together, as they do in the data.

To match the empirical responses we estimate in section 3, we must also choose a very small value for  $\kappa \zeta \sigma$ . The response plotted in Figure 3 sets  $\kappa \zeta \sigma = 0.0002$ . This implies a very large amount of nominal and real rigidities for conventional values of the intertemporal elasticity of substitution. For example, Rotemberg and Woodford (1997) estimate  $\kappa \zeta = 0.024$  and  $\sigma = 6.25$ , which implies  $\kappa \zeta \sigma = 0.15$ . Woodford (2003, ch. 3) discusses values of  $\zeta$  in the range  $[0.06, 2.25]$ . Combining this range with  $\sigma = 1$  and the assumption that prices change on average once a year implies a range for  $\kappa \zeta \sigma$  of  $[0.01, 0.59]$ . However, parameters in this range generate highly counterfactual patterns for the nominal and real interest rate patterns we study.

Figure 4 illustrates this by plotting the impulse response functions of nominal and real interest rates and inflation for a case where  $\kappa \zeta \sigma = 0.005$ . In this case, the inflation response is so large after the monetary shock that the nominal interest rate response becomes negative only a few periods after the initial shock and largely tracks inflation. Intuitively, in a model with small amounts of nominal and real rigidities, monetary policy shocks largely result in inflation as opposed to movements in real interest rates.<sup>20</sup>

One way of interpreting a low value of  $\kappa \zeta \sigma$  is as evidence of a low value of the intertemporal elasticity of substitution  $\sigma$  as opposed to evidence of a high degree of nominal and real rigidities.

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<sup>20</sup>One way to better reconcile our empirical evidence with the standard calibrations of simple New Keynesian models would be for us to adopt the hybrid Phillips curve specification in Gali and Gertler (1999) as opposed to the specification in equation (12). The difference is that in the Gali-Gertler specification, the coefficient on the output terms becomes  $(1 - \gamma)\kappa \zeta$  as opposed to  $\kappa \zeta$ . With a value of  $\gamma$  close to one (recall that we use 0.999 above), this would yield a much smaller response of inflation to variation in output for a given calibration of  $\kappa \zeta$ . In the Gali-Gertler model,  $\gamma$  represents the fraction of “rule-of-thumb” price setters. These price setters do not respond to fluctuations in output. With a high value of  $\gamma$  this implies that the Gali-Gertler specification yields a very small response of inflation to variations in output. We thank Eric Swanson for pointing out this interpretation to us.

In other words, a small response of inflation to large real interest rate movements could reflect a lack of responsiveness of inflation to output (nominal and real rigidities) or it could reflect a small response of output to real interest rates (the intertemporal elasticity of substitution). However, even for calibrations of our model with the intertemporal elasticity of substitution dramatically below conventional values—say 0.01—the implied value of  $\kappa\zeta$  needed to match our empirical evidence is 0.002—implying substantially more nominal and real rigidities than, e.g., Rotemberg and Woodford (1997) estimate. Thus, our evidence appears to point toward either much larger values of nominal and real rigidities, or much smaller values of the responsiveness of output to real interest rates than are conventionally assumed in simple business cycle models.

## 4.2 Estimating the CEE/ACEL Model with High Frequency Data

Our analysis above using a simple three-equation New Keynesian model provides intuition for how the evidence on the response of the real interest rate and inflation to a monetary shock that we document in section 3 sheds light on the extent of monetary non-neutrality. However, the simple three-equation New Keynesian model we consider above abstracts from many features that have been shown to be important in generating realistic business cycles. It is therefore not well suited for quantitative analysis.

We next investigate the quantitative implications of our empirical evidence by estimating the workhorse medium-scale business cycle model developed in Christiano, Eichenbaum, and Evans (2005, henceforth CEE) and Altig et al. (2011 henceforth ACEL). ACEL develop a version of this model in which capital is firm specific. They show that this version of the model is equivalent to the homogeneous capital version of the model up to a linear approximation (though with different parameter interpretations, as we discuss below). We therefore refer to this model as the CEE/ACEL model.

### 4.2.1 Selective Model Description

CEE and ACEL present detailed descriptions of the CEE/ACEL model. We refrain from repeating this material here. Rather, we discuss the elements of the model that are most relevant for our analysis. Consider first, the elasticity of output with respect to changes in real interest rates. In the simple New Keynesian model, this is governed by the intertemporal elasticity of substitution of households  $\sigma$ . In the CEE/ACEL model, this is governed by—on the one hand—the elasticity of

consumption to the real interest rate and—on the other hand—the elasticity of investment to the real interest rate.

For consumption, CEE/ACEL assume that the utility derived from consumption in each period is  $\log(C_t - bC_{t-1})$ , where  $b$  is a parameter that governs the degree of habit formation. The curvature of the log function pins down the elasticity of consumption with respect to real interest rates. (Without habit, this would imply an intertemporal elasticity of substitution of one.) Turning to investment, CEE show that the linearized first-order condition for investment in their model may be solved forward to yield

$$\hat{\lambda}_t = \hat{\lambda}_{t-1} + \frac{1}{k_I} \sum_{j=0}^{\infty} \beta^j E_{t-1} \hat{p}_{k,t+j}, \quad (13)$$

where  $\hat{\lambda}_t$  denotes investment and  $\hat{p}_{k,t}$  is the shadow value of a unit of installed capital. The key parameter is  $k_I$ . From equation (13), we see that  $1/k_I$  is the elasticity of investment with respect to a 1 percent temporary increase in the current price of installed capital. CEE estimate  $1/k_I = 0.40$ , while ACEL estimate  $1/k_I = 0.66$ . This is one of the parameters that we estimate.

Next, consider the response of inflation to variation in output. The two key parameters governing this response in the homogeneous capital version of the CEE/ACEL model are  $\xi_p$  and  $\xi_w$ . These parameters govern the frequency of price change and the frequency of wage change. Specifically, the frequency of price change is  $1 - \xi_p$  and the frequency of wage change is  $1 - \xi_w$ . We will estimate these two parameters using the homogeneous capital version of the CEE/ACEL model. ACEL show that the homogeneous capital version of the model with a particular value for  $\xi_p$  yields the same aggregate dynamics as the firm-specific capital version of the model with a much lower value of  $\xi_p$ . The reason for this is that firm-specific capital is a powerful source of real rigidity that dramatically lowers the slope of the price Phillips curve in the model for any given values of  $\xi_p$ .

CEE/ACEL assume that firms that do not have an opportunity to reoptimize their prices index their prices to past inflation. This is analogous to setting  $\gamma = 1$  in the hybrid Phillips curve model above. Likewise, CEE/ACEL assume that unions that do not have an opportunity to reoptimize their wages index their wages to past wage inflation. In other words, CEE/ACEL build the high degree of price and wage inflation inertia that we show above is essential in fitting our high frequency data into their model.

The only change we make to the CEE/ACEL model is that we replace the monetary policy rule in that model with the monetary policy rule we discuss above—equation (6)—and we consider

our policy news shock—equation (7). We fix  $\phi_\pi = 0.5$  but estimate  $\rho_1$  and  $\rho_2$ . We fix all other parameters equal to their estimated values in CEE. The primary reason that we do not estimate a larger set of parameters is that our empirical evidence provides us with information about certain aspects of the CEE/ACEL model, but not all aspects.

#### 4.2.2 Estimation Approach

We estimate the model by indirect inference. The moments we use in our estimation are the responses of 2, 3, 5, and 10-year nominal and real yields and the responses of 2, 3, 5, and 10-year instantaneous nominal and real forward rates to our policy news shock. We minimize the sum of the squared difference between the moments in the data and the model. So as not to have to estimate the size of the shock, we scale the responses from the model in such a way that they perfectly match the 3Y real forward rate.

We construct standard errors by bootstrapping. Our bootstrap procedure is to re-sample the data with replacement, estimate the empirical moments using the Rigobon method on the re-sampled data, and then estimate the structural parameters using a loss function based on the estimated empirical moments for the re-sampled data. We repeat this procedure 1000 times. Importantly, this procedure for constructing the standard errors captures the statistical uncertainty in our structural parameter estimates arising from the statistical uncertainty about our empirical estimates in Table 1.

#### 4.2.3 Estimates of Monetary Non-Neutrality

Our primary interest is the extent of monetary non-neutrality implied by our high frequency evidence. It is useful to define a unidimensional summary measure of monetary non-neutrality. The statistic we use for this purpose is the sum of the absolute value of the response of output over time divided by the sum of the absolute value of the response of inflation over time:

$$M = \frac{\sum_{j=0}^{\infty} |\hat{y}_{t+j}|}{\sum_{j=0}^{\infty} |\hat{\pi}_{t+j}|}. \quad (14)$$

A key advantage of focusing on this statistic as opposed to the primitive parameters of the model is that it takes account of the correlations between our estimated values for the primitive parameters. Since our estimates of the primitive parameters are correlated—e.g., a relatively high values of  $\xi_w$  tends to be associated with a relative low value for  $\xi_p$ —it is difficult to assess the strength of

our evidence regarding monetary non-neutrality from looking at the confidence intervals of these parameters.

Table 6 presents results on the value of  $M$  implied by our estimation of the CEE/ACEL model. Our estimation of the CEE/ACEL model yields a value of 16.2 for  $M$  with a 95% confidence interval of [6.8, 146.7].<sup>21</sup> For comparison, we also report the value of  $M$  implied by the estimates obtained by ACEL and CEE. The parameters that ACEL estimate imply a value of  $M$  of 10.7. Our estimation of the ACEL/CEE model thus implies about 50 percent more monetary non-neutrality than ACEL. But we cannot reject the degree of monetary non-neutrality estimated by ACEL. The parameters that CEE estimate imply a value of  $M$  of only 1.4. This is substantially below the lower bound of our 95% confidence interval. Our estimates thus clearly imply substantially more monetary non-neutrality than CEE's.

Figure 5 presents the response of nominal and real interest rates and inflation for our estimation of the CEE/ACEL model. Comparing these responses to those in Figure 2, we see that the model fits the data quite well. The response of inflation is very small initially and then gradually increases. The response of nominal and real interest rates is close to identical out to about 3 years. At longer horizons, the response of nominal interest rates falls below the response of real interest rates.

In contrast, Figure 6 presents the response of nominal and real interest rates and inflation for CEE's estimates of the CEE/ACEL model in response to our estimated policy news shock. These responses stand in stark contrast with the responses we estimate in the data. In particular, the response of inflation is much larger than in the data and as a consequence, the response of nominal interest rates largely track the response of inflation rather than largely tracking the response of real interest rates as they do in the data.

One difference between our evidence and the evidence in CEE is that the shock we analyze is much more persistent than the shock in CEE. To see whether this is an important contributing factor to our results, we also consider the value of  $M$  for our estimation of the CEE/ACEL model in response to the Taylor rule shock considered by CEE (which CEE show yields virtually identical results to their baseline shock to the money supply). For this shock our estimation of the CEE/ACEL model yields a value of 19.4 for  $M$  with a 95% confidence interval of [6.2, 40.7]. The parameters that ACEL estimate imply a value of  $M$  of 11.1 for this shock, while the parameters CEE estimate yield

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<sup>21</sup>We construct this confidence interval by calculating  $M$  for each iteration of the bootstrap we run using the estimates of  $(k_I, \xi_p, \xi_w)$  from that iteration but holding the shock process  $(\rho_1, \rho_2)$  fixed at their point estimates.

a value of  $M$  of 6.9. Again, our estimates imply substantially more monetary non-neutrality than the parameters estimated by ACEL and CEE. However, for this shock, we cannot reject the degree of monetary non-neutrality implied by the parameters estimated by ACEL or CEE. The persistence of the shock we consider in this paper thus contributes to our ability to narrow the set of parameters consistent with the data.

Table 7 presents our individual parameter estimates. We estimate  $\rho_1 = 0.93$  and  $\rho_2 = 0.61$ . This matches the hump-shaped response of interest rates to the policy news shock. We estimate  $\xi_p = 0.93$  with a 95% confidence interval of  $[0.79, 0.99]$  and we estimate  $\xi_w = 0.90$  with a 95% confidence interval of  $[0.00, 0.99]$ . The large confidence interval for the wage rigidity parameter reflects a substantial negative correlation between price and wage rigidity parameters,  $\xi_p$  and  $\xi_w$ , in our estimation. Figure 7 presents a scatterplot of the joint distribution of  $\xi_p$  and  $\xi_w$  that we estimate. The figure shows clearly that low values of  $\xi_w$  are accompanied by very high values of  $\xi_p$ . Our results, thus, provide strong evidence for a large amount of nominal and real rigidities, but are less sharp when it comes to determining whether the source of these rigidities is wage rigidity or price rigidity.

For comparison, CEE estimate  $\xi_p = 0.60$  and  $\xi_w = 0.64$ , while ACEL estimate  $\xi_p = 0.90$  and  $\xi_w = 0.78$ . Our point estimates of the extent of price and wage rigidity are, thus, much higher than CEE's but only slightly higher than ACEL's estimates. ACEL show that in the firm-specific capital model, their parameter estimates can be explained by a model in which prices change every 2-3 quarters. Intuitively, the presence of firm-specific capital implies that even when firms do adjust their prices, they adjust by only a fraction of the amount required to reach the flexible price equilibrium.

An additional interesting feature of our results is that our point estimate of the investment adjustment cost parameter  $k_I = 6877.4$  is much higher than the estimates obtained by CEE and ACEL. This is the analog to our finding in the simple three-equation model that the data appears to suggest a low responsiveness of the economy to movements in the real interest rate. However, the 95% confidence interval we estimate for  $k_I$  is very wide— from 1.1 to 7966.6—and the values estimates by CEE and ACEL are contained within our confidence interval. This wide confidence interval is a results of the fact that the loss function is very flat in  $k_I$  for large values of  $k_I$ . In particular, the loss function is close to constant in  $k_I$  for values of  $k_I$  larger than 20. ACEL's estimate of  $k_I = 1.5$  implies that a 1% permanent increase in the price of installed capital leads to a 66% change in investment, while a value of  $k_I = 25$  implies that such a change in the price of installed capital leads

to a 4% increase in investment.

### 4.3 Fed Information Case

In the analysis above, we assume that the Fed and the private sector receive the same information about future movements in the natural rate of interest but hold different views about the workings of the economy and therefore interpret these signals differently. This implies that announcements by the Fed lead the private sector to update its beliefs about what the Fed thinks about the future evolution of the natural rate but doesn't affect the private sector's own beliefs about future evolution of the natural rate. An alternative view is that the Fed and the private sector receive share the same model of the world but the Fed receives additional signals about economic fundamentals that the public does not receive directly. More specifically, consider a case where there are two types of signals; signals that are seen by both the public and the central bank and signals that only the central bank receives. These assumptions imply that movements in the term structure of interest rates at the time of FOMC announcements should be interpreted as being due to the private sector using what the Fed says to update its own beliefs about the future path of the natural rate of interest. If the public believes that the Fed is committed to vary short term interest rates in such a way as to track the natural rate of interest, FOMC announcements will in this case not change the public's views about future deviations between interest rates and the natural rate. In other words,  $E_t r_{t+j}^n - E_{t-1} r_{t+j}^n = E_t \bar{r}_{t+j} - E_{t-1} \bar{r}_{t+j}$  for all  $j$ . This implies that the Fed's announcement does not change the current or expected future "interest rate gap" ( $\hat{i}_{t+j} - E_t \hat{\pi}_{t+j+1} - r_{t+j}^n$ ) and therefore also leaves the current and expected future level of the output gap and inflation unaffected. As a consequence, the response of nominal rates and real rates should be the same at all horizons.

The response of nominal interest rates, real interest rates, and inflation we estimate in section 3 is consistent with this prediction at the short end of the term structure, but not at the long end. The response of inflation is close to zero and not statistically significantly different from zero at horizons out to 3 years. At the 5 and 10 year horizon, however, the nominal rate response is smaller than the real rate response implying that the response of inflation is significantly negative.

In addition, the plausibility of the Fed information case depends on how plausible it is to think that the Fed has a significant informational advantage over the private sector. To our knowledge, the Fed does not have access to a significant amount of information about the economy that is outside

the public domain. Any informational advantage by the Fed about the natural rate of interest must thus be due to an advantage in processing information. Romer and Romer (2000) argue that monetary policy actions by the Fed reveal information to the public that is useful for forecasting inflation and that this informational advantage is due to superior information processing.<sup>22</sup> However, an alternative reason why the Fed might have superior information about the future evolution of inflation is that it has superior information about its future monetary policy ( $\bar{r}_t$ ) rather than superior information about the natural rate of interest ( $r_t^n$ ). This alternative reason falls under our baseline case.

## 5 Conclusion

In this paper, we follow in the tradition of work by Christiano, Eichenbaum, and Evans (2005) and others who attempt to fit structural models of monetary policy to evidence on the response of real variables to monetary shocks. We focus on the effects of a “policy news shock” that we construct as a summary measure of the Fed’s impact on nominal interest rates over the year following an FOMC announcement. By construction, this variable has strong predictive power for movements in nominal interest rates. However, we document that it also has strong predictive power for movements in real interest rates. In fact, real interest rates move close to one-for-one with nominal rates in response to a policy news shock at horizons out to 3 years. Despite large movements in real interest rates, the response of inflation is small.

We show that the sluggish response of prices to movements in real interest rates associated with monetary shocks provides a great deal of information about the degree of monetary non-neutrality in business cycle models. The two key parameters in determining the response of inflation to movements in real interest rates associated with monetary shocks are: 1) the responsiveness of output to movements in the real interest rate, as determined by the intertemporal elasticity of substitution and in the elasticity of investment to real interest rate movements and 2) the responsiveness of inflation to output, as determined by the magnitude of nominal and real rigidities.

We develop a method-of-moments estimation approach to assess the implications of the empirical evidence we document for the structural parameters of a workhorse monetary model. Despite the

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<sup>22</sup>Faust, Swanson, and Wright (2004) argue that Romer and Romer’s results do not hold up for a more recent sample period and are sensitive to using the unexpected component of the change in the Federal Funds rate as the monetary surprise as opposed to the entire change in the Federal Funds rate.

short time-period over which real interest rate data are available, and the resulting large standard errors on the interest rate responses, our analysis yields strong conclusions about the parameters of our model. We find that matching our evidence on the response of inflation to real interest rate movements requires a small elasticity of output with respect to the real interest rate, a large amount of nominal and real rigidities, or both.

Our estimates thus provide strong support for the mechanisms generating large real rigidities that have been analyzed in the monetary economics literature. We explicitly investigate the ability of two such models of real rigidities—heterogeneous factor markets, and wage rigidities—to explain our empirical results. We find that these models can match the responses we observe in the data, albeit with a somewhat higher degree of real rigidity and a lower responsiveness of output to the real interest rate than is typically assumed in the existing literature. We also find strong support for mechanisms generating inflation inertial. In the data, we find that the inflation response to a monetary shock is initially small and grows over time. However, the bare bones New Keynesian Phillips curve predicts exactly the opposite: a large immediate inflation response to a monetary shock, which dies out over time.

Business cycle models with modest price adjustment frictions generate radically different predictions from our baseline estimates. In such models, the response of inflation to our monetary shocks is large. The response of nominal interest rates largely track the response of inflation. This implies that nominal rates fall after a deflationary monetary shock. Nominal and real interest rates, thus, move in opposite directions—in contrast to the nearly one-for-one movements that we observe in the data.

## A Construction of the Policy New Shock

The policy news shock is constructed as the first principle component of the change in five interest rates. The first of these is the change in market expectations of the Federal Funds Rate over the remainder of the month in which the FOMC meeting occurs. To construct this variable from the change in the price of the current month's Federal Funds Rate futures contract, we must adjust for the fact that a part of the month has already elapsed when the FOMC meeting occurs. Suppose the month in question has  $m_0$  days and the FOMC meeting occurs on day  $d_0$ . Let  $f_{t-\Delta t}^1$  denote the price of the current month's Federal Funds Rate futures contract immediately before the FOMC announcement and  $f_t^1$  the price of this contract immediately following the FOMC announcement. Let  $r_0$  denote the average Federal Funds Rate during the month up until the point of the FOMC announcement and  $r_1$  the average Federal Funds Rate for the remainder of the month. Then

$$\begin{aligned} f_{t-\Delta t}^1 &= \frac{d_0}{m_0} r_{-1} + \frac{m_0 - d_0}{m_0} E_{t-\Delta t} r_0, \\ f_t^1 &= \frac{d_0}{m_0} r_{-1} + \frac{m_0 - d_0}{m_0} E_t r_0. \end{aligned}$$

As a result

$$E_t r_0 - E_{t-\Delta t} r_0 = \frac{m_0}{m_0 - d_0} (f_t^1 - f_{t-\Delta t}^1).$$

When the FOMC meeting occurs on a day when there are 7 days or less remaining in a month, we instead use the change in the price of next month's Fed Funds Futures contract. This avoids multiplying  $f_t^1 - f_{t-\Delta t}^1$  by a very large factor.

The second variable used in constructing the policy news shock is the change in the expected Federal Funds Rate at the time of the next scheduled FOMC meeting. Similar issues arise in constructing this variable as with the variable described above. Let  $m_1$  denote the number of days in the month in which the next scheduled FOMC meeting occurs and let  $d_1$  denote the day of the meeting. The next scheduled FOMC meeting may occur in the next month or as late as 3 months after the current meeting. Let  $f_{t-\Delta t}^n$  denote the price of the Federal Funds Rate futures contract for the month of the next scheduled FOMC meeting immediately before the FOMC announcement and  $f_t^n$  the price of this contract immediately following the FOMC announcement. Let  $r_1$  denote the Federal Funds Rate after then next scheduled FOMC meeting. Analogous calculations to what we present above yield

$$E_t r_1 - E_{t-\Delta t} r_1 = \frac{m_1}{m_1 - d_1} \left[ (f_t^n - f_{t-\Delta t}^n) - \frac{d_1}{m_1} (E_t r_0 - E_{t-\Delta t} r_0) \right].$$

As with the first variable, if the next scheduled FOMC meeting occurs on a on a day when there are 7 days or less remaining in a month, we instead use the change in the price of next month's Fed Funds Futures contract.

The last three variables used are simply the change in the price of the Eurodollar futures at the time of the FOMC announcements.

We approximate the change in these variables over a 30-minute window around FOMC by taking the difference between the price in the last trade that occurred more than 10 minutes before the FOMC announcement and the first trade that occurred more than 20 minutes after the FOMC announcement. On control days, we take the last trade before 2:05pm and the first trade after 2:35pm (since FOMC announcements tend to occur at 2:15pm). On some days (most often control days), trading is quite sparse and there sometimes is no trade before 2:05 or after 2:35. To limit the size of the windows we consider, we only consider trades on the trading day in question and until noon the next day. If we do not find eligible trades to construct the price change we are interested in within this window, we set the price change to zero (i.e., we interpret no trading as no price change).

## B Derivation of Our Heteroskedasticity-Based Estimator

Let  $\Omega_{Ri}$  denote the variance-covariance matrix of  $[\Delta i_t, \Delta s_t]$  in regime  $Ri$ . Then  $\Omega_{Ri}$  is given by

$$\Omega_{Ri} = \begin{bmatrix} \sigma_{\epsilon, Ri}^2 + \sum_j \beta_{i,j}^2 \sigma_{\eta,j}^2 & \gamma \sigma_{\epsilon, Ri}^2 + \sum_j \beta_{i,j} \beta_{s,j} \sigma_{\eta,j}^2 \\ \gamma \sigma_{\epsilon, Ri}^2 + \sum_j \beta_{i,j} \beta_{s,j} \sigma_{\eta,j}^2 & \gamma^2 \sigma_{\epsilon, Ri}^2 + \sum_j \beta_{s,j}^2 \sigma_{\eta,j}^2 \end{bmatrix},$$

where  $j$  indexes the elements of  $\eta_t$ . Notice that

$$\Delta \Omega = \Omega_{R1} - \Omega_{R2} = (\sigma_{\epsilon, R1}^2 - \sigma_{\epsilon, R2}^2) \begin{bmatrix} 1 & \gamma \\ \gamma & \gamma^2 \end{bmatrix}.$$

Thus,

$$\gamma = \frac{\Delta \Omega_{12}}{\Delta \Omega_{11}} = \frac{\text{cov}_{R1}(\Delta i_t, \Delta s_t) - \text{cov}_{R2}(\Delta i_t, \Delta s_t)}{\text{var}_{R1}(\Delta i_t) - \text{var}_{R2}(\Delta i_t)}.$$

## C A Simple New Keynesian Model

This section lays out micro-foundations for the simple New Keynesian business cycle model discussed in section 4 in the main text. See Woodford (2003) and Gali (2008) for thorough expositions of New

Keynesian models.

## C.1 Households

The economy is populated by a continuum of household types indexed by  $x$ . A household's type indicates the type of labor supplied by that household. Households of type  $x$  seek to maximize their utility given by

$$E_0 \sum_{t=0}^{\infty} \beta^t [u(C_t, \xi_t) - v(L_t(x), \xi_t)], \quad (15)$$

where  $\beta$  denotes the household's subjective discount factor,  $C_t$  denotes household consumption of a composite consumption good,  $L_t(x)$  denotes household supply of differentiated labor input  $x$ , and  $\xi_t$  denotes a vector of preference shocks. There are an equal (large) number of households of each type. The composite consumption good in expression (15) is an index given by

$$C_t = \left[ \int_0^1 c_t(z)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta}{\theta-1}}, \quad (16)$$

where  $c_t(z)$  denotes consumption of products of variety  $z$ . The parameter  $\theta > 1$  denotes the elasticity of substitution between different varieties.

Households have access to complete financial markets. Households of type  $x$  face a flow budget constraint given by

$$P_t C_t + E_t[M_{t,t+1} B_{t+1}(x)] \leq B_t(x) + W_t(x) L_t(x) + \int_0^1 \Xi_t(z) dz - T_t, \quad (17)$$

where  $P_t$  is a price index that gives the minimum price of a unit of the consumption good  $C_t$ ,  $B_{t+1}(x)$  is a random variable that denotes the state contingent payoff of the portfolio of financial securities held by households of type  $x$  at the beginning of period  $t+1$ ,  $M_{t,t+1}$  is the stochastic discount factor that prices these payoffs in period  $t$ ,<sup>23</sup>  $W_t(x)$  denotes the wage rate received by households of type  $x$  in period  $t$ ,  $\Xi_t(z)$  denotes the profits of firm  $z$  in period  $t$ , and  $T_t$  is a lump-sum tax levied by the government. To rule out Ponzi schemes, household debt cannot exceed the present value of future income in any state of the world.

Households face a decision in each period about how much to spend on consumption, how many hours of labor to supply, how much to consume of each differentiated good produced in the economy

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<sup>23</sup>The stochastic discount factor  $M_{t,t+1}$  is a random variable over states in period  $t+1$ . For each such state it equals the price of the Arrow-Debreu asset that pays off in that state divided by the conditional probability of that state. See Cochrane (2005) for a detailed discussion.

and what portfolio of assets to purchase. Optimal choice regarding the trade-off between current consumption and consumption in different states in the future yields the following consumption Euler equation:

$$\frac{u_c(C_{t+j}, \xi_{t+j})}{u_c(C_t, \xi_t)} = \frac{M_{t,t+j} P_{t+j}}{\beta^j P_t} \quad (18)$$

as well as a standard transversality condition. Subscripts on the function  $u$  denote partial derivatives. Equation (18) holds state-by-state for all  $j > 0$ . Optimal choice regarding the intratemporal trade-off between current consumption and current labor supply yields a labor supply equation:

$$\frac{v_\ell(L_t(x), \xi_t)}{u_c(C_t, \xi_t)} = \frac{W_t(x)}{P_t}. \quad (19)$$

Households optimally choose to minimize the cost of attaining the level of consumption  $C_t$ . This implies the following demand curves for each of the differentiated products produced in the economy:

$$c_t(z) = C_t \left( \frac{p_t(z)}{P_t} \right)^{-\theta}, \quad (20)$$

where  $p_t(z)$  denotes the price of product  $z$  and

$$P_t = \left[ \int_0^1 p_t(z)^{1-\theta} dz \right]^{\frac{1}{1-\theta}}. \quad (21)$$

## C.2 Firms

There are a continuum of firms indexed by  $z$  in the economy. Firm  $z$  specializes in the production of differentiated good  $z$ , the output of which we denote  $y_t(z)$ . For simplicity, labor is the only variable factor of production used by firms. Each firm is endowed with a fixed, non-depreciating stock of capital. The production function of firm  $z$  is

$$y_t(z) = A_t f(L_t(z)), \quad (22)$$

where  $A_t$  denotes aggregate productivity. The function  $f$  is increasing and concave. It is concave because there are diminishing marginal return to labor given the fixed amount of other inputs employed at the firm. We follow Woodford (2003) in introducing heterogeneous labor markets. Firm belongs to an industry  $x$ . There are many firms in each industry. The goods in industry  $x$  are produced using labor of type  $x$  and all firms in industry  $x$  change prices at the same time. This heterogeneous labor market structure is a strong source of real rigidities in price setting.

Firm  $z$  acts to maximize its value,

$$E_t \sum_{j=0}^{\infty} M_{t,t+j} [p_{t+j}(z)y_{t+j}(z) - W_{t+j}(x)L_{t+j}(z)]. \quad (23)$$

Firm  $z$  must satisfy demand for its product given by equation (20). Firm  $z$  is therefore subject to the following constraint:

$$C_t \left( \frac{p_t(z)}{P_t} \right)^{-\theta} \leq A_t f(L_t(z)). \quad (24)$$

Firm  $z$  takes its industry wage  $W_t(x)$  as given. Optimal choice of labor demand by the firm is given by

$$W_t(x) = A_t f_\ell(L_t(z)) S_t(z), \quad (25)$$

where  $S_t(z)$  denotes the firm's nominal marginal cost (the Lagrange multiplier on equation (24) in the firm's constrained optimization problem).

Firm  $z$  can reoptimize its price with probability  $1 - \alpha$  as in Calvo (1983). With probability  $\alpha$  it must keep its price unchanged. Optimal price setting by firm  $z$  in periods when it can change its price implies

$$p_t(z) = \frac{\theta}{\theta - 1} E_t \sum_{j=0}^{\infty} \frac{\alpha^j M_{t,t+j} y_{t+j}(z)}{\sum_{k=0}^{\infty} \alpha^k M_{t,t+k} y_{t+k}(z)} S_{t+j}(z). \quad (26)$$

Intuitively, the firm sets its price equal to a constant markup over a weighted average of current and expected future marginal cost.

### C.3 A Linear Approximation of Private Sector Behavior

We seek a linear approximation of the equation describing private sector behavior around a zero-growth, zero-inflation steady state. We start by deriving a log-linear approximation for the consumption Euler equation that related consumption growth and a one-period, riskless, nominal bond. This equation takes the form  $E_t[M_{t,t+1}(1 + i_t)] = 1$ , where  $i_t$  denotes the yield on a one-period, riskless, nominal bond. Using equation (18) to plug in for  $M_{t,t+1}$  and rearranging terms yields

$$E_t \left[ \beta U_c(C_{t+1}, \xi_{t+1}) \frac{P_t}{P_{t+1}} \right] = \frac{U_c(C_t, \xi_t)}{1 + i_t}. \quad (27)$$

The zero-growth, zero-inflation steady state of this equation is  $\beta(1 + \bar{i})$ . A first order Taylor series approximation of equation (27) is

$$\hat{c}_t = E_t \hat{c}_{t+1} - \sigma(\hat{i}_t - E_t \hat{\pi}_{t+1}) - \sigma E_t \Delta \hat{\xi}_{ct+1}, \quad (28)$$

where  $\hat{c}_t = (C_t - C)/C$ ,  $\hat{\pi}_t = \pi_t - 1$ ,  $\hat{i}_t = (1 + i_t - 1 - \bar{i})/(1 + \bar{i})$ , and  $\hat{\xi}_{ct} = (U_{cc}/U_c)(\xi_t - 1)$ . The parameter  $\sigma = -U_c/(U_{cc}C)$  denotes the intertemporal elasticity of substitution of households.

We next linearize labor demand, labor supply, and the production function and combine these equations to get an expression for the marginal costs in period  $t + j$  of a firm that last changed its price in period  $t$ . Let  $\ell_{t,t+j}(x)$  denote the percent deviation from steady state in period  $t + j$  of hours worked for workers in industry  $x$  that last was able to change prices in period  $t$ . Let other industry level variables be defined analogously. We assume that  $f(L_t(x)) = L_t^a(x)$ .

A linear approximation of labor demand—equation (25)—in period  $t + j$  for industry  $x$  that was last able to change its prices in period  $t$  is then

$$\hat{w}_{t,t+j}(x) = \hat{a}_{t+j} - (1 - a)\hat{\ell}_{t,t+j}(x) + \hat{s}_{t,t+j}(x), \quad (29)$$

where  $\hat{w}_{t,t+j}(x)$  and  $\hat{s}_{t,t+j}(x)$  denote the percentage deviation of real wages and real marginal costs, respectively, from their steady state values.

A linear approximation of labor supply—equation (19)—in period  $t + j$  for industry  $x$  that was last able to change its prices in period  $t$  is

$$\hat{w}_{t,t+j}(x) = \eta^{-1}\hat{\ell}_{t,t+j}(x) + \sigma^{-1}\hat{c}_{t+j} + \hat{\xi}_{\ell,t+j} - \hat{\xi}_{c,t+j}, \quad (30)$$

where  $\hat{\xi}_{\ell,t+j} = (V_{\ell\xi}/V_\ell)(\xi_t - 1)$ . The parameter  $\eta = V_\ell/(V_{\ell\ell}L)$  is the Frisch elasticity of labor supply.

A linear approximation of the production function—equation (22)—in period  $t + j$  for industry  $x$  that was last able to change its prices in period  $t$  is

$$\hat{y}_{t,t+j}(x) = \hat{a}_{t+j} + a\hat{\ell}_{t,t+j}(x). \quad (31)$$

Combining labor demand and labor supply—equations (29) and (30)—to eliminate  $\hat{w}_{t,t+j}(x)$  yields

$$\hat{s}_{t,t+j}(x) = (\eta^{-1} + 1 - a)\hat{\ell}_{t,t+j}(x) + \sigma^{-1}\hat{c}_{t+j} - \hat{a}_{t+j} + \hat{\xi}_{\ell,t+j} - \hat{\xi}_{c,t+j}.$$

Using the production function—equation (31)—to eliminate  $\hat{\ell}_{t,t+j}(x)$  yields

$$\hat{s}_{t,t+j}(x) = \omega\hat{y}_{t,t+j}(x) + \sigma^{-1}\hat{c}_{t+j} - (\omega + 1)\hat{a}_{t+j} + \hat{\xi}_{\ell,t+j} - \hat{\xi}_{c,t+j}, \quad (32)$$

where  $\omega = (\eta^{-1} + 1 - a)/a$ .

Taking logs of consumer demand—equation (20)—in period  $t + j$  for industry  $x$  what was last able to change its prices in period  $t$  yields

$$\hat{y}_{t,t+j}(z) = -\theta\hat{p}_t(x) + \theta \sum_{k=1}^j \hat{\pi}_{t+k} + \hat{y}_{t+j}, \quad (33)$$

where we use the fact that  $Y_t = C_t$  and  $y_t(x) = c_t(x)$ . Plugging this equation into equation (32) and again using the fact that  $Y_t = C_t$  yields

$$\hat{s}_{t,t+j}(x) = -\omega\theta\hat{p}_t(x) + \omega\theta\sum_{k=1}^j\hat{\pi}_{t+k} + (\omega + \sigma^{-1})\hat{y}_{t+j} - (\omega + 1)\hat{a}_{t+j} + \hat{\xi}_{\ell,t+j} - \hat{\xi}_{c,t+j} \quad (34)$$

It is useful to derive the level of output that would prevail if all prices were flexible. Since our model does not have any industry specific shocks (other than the opportunity to change prices), marginal costs of all firms are the same when prices are flexible. Firm price setting in this case yields  $p_t(x) = \mu S_t$ , where  $\mu = \theta/(\theta - 1)$ . This implies that all prices are equal and that  $S_t/P_t = 1/\mu$ . Since real marginal cost is a constant, we have  $\hat{s}_t = 0$ . The flexible price version of equation (34) is then

$$(\omega + \sigma^{-1})\hat{y}_t^n = (\omega + 1)\hat{a}_t - \hat{\xi}_{\ell,t} + \hat{\xi}_{c,t}, \quad (35)$$

where we use the fact that output in all industries is the same under flexible prices and  $\hat{y}_t = \hat{c}_t$  and denote the rate of output under flexible prices as  $y_t^n$ . We will refer to  $y_t^n$  as the natural rate of output.

Combining equations (34) and (35) yields

$$\hat{s}_{t,t+j}(x) = -\omega\theta\hat{p}_t(x) + \omega\theta\sum_{k=1}^j\hat{\pi}_{t+k} + (\omega + \sigma^{-1})(\hat{y}_{t+j} - \hat{y}_{t+j}^n) \quad (36)$$

We next linearize the price setting equation—equation (26). This yields:

$$\sum_{j=0}^{\infty}(\alpha\beta)^j\hat{p}_t(x) - \sum_{j=0}^{\infty}(\alpha\beta)^j E_t\hat{s}_{t,t+j}(x) - \sum_{j=1}^{\infty}(\alpha\beta)^j\sum_{k=1}^j E_t\hat{\pi}_{t+k} = 0.$$

Manipulation of this equation yields

$$\hat{p}_t(x) = (1 - \alpha\beta)\sum_{j=0}^{\infty}(\alpha\beta)^j E_t\hat{s}_{t,t+j}(x) + \alpha\beta\sum_{j=1}^{\infty}(\alpha\beta)^j E_t\hat{\pi}_{t+j}. \quad (37)$$

Using equation (36) to eliminate  $\hat{s}_{t,t+j}(x)$  in equation (37) and manipulating the resulting equation yields

$$\hat{p}_t(x) = (1 - \alpha\beta)\zeta\sum_{j=0}^{\infty}(\alpha\beta)^j E_t(\hat{y}_{t+j} - \hat{y}_{t+j}^n) + \alpha\beta\sum_{j=1}^{\infty}(\alpha\beta)^j E_t\hat{\pi}_{t+j}, \quad (38)$$

where  $\zeta = (\omega + \sigma^{-1})/(1 + \omega\theta)$ . A linear approximation of the expression for the price index—equation (21)—yields

$$\hat{\pi}_t = \frac{1 - \alpha}{\alpha}\hat{p}_t(x). \quad (39)$$

Using this last equation to replace  $\hat{p}_t(x)$  in equation (38) yields

$$\hat{\pi}_t = \kappa\zeta \sum_{j=0}^{\infty} (\alpha\beta)^j E_t(\hat{y}_{t+j} - \hat{y}_{t+j}^n) + (1-\alpha)\beta \sum_{j=1}^{\infty} (\alpha\beta)^j E_t \hat{\pi}_{t+j},$$

where  $\kappa = (1-\alpha)(1-\alpha\beta)/\alpha$ . Quasi-differencing the resulting equation yields

$$\hat{\pi}_t - \alpha\beta E_t \hat{\pi}_{t+1} = \kappa\zeta(\hat{y}_t - \hat{y}_t^n) + (1-\alpha)\beta E_t \hat{\pi}_{t+1},$$

which implies

$$\hat{\pi}_t = \beta E_t \hat{\pi}_{t+1} + \kappa\zeta(\hat{y}_t - \hat{y}_t^n). \quad (40)$$

Finally, we rewrite the household's Euler equation—equation (28) in terms of the output gap:

$$y_t - y_t^n = E_t(y_{t+1} - y_{t+1}^n) - \sigma(\hat{y}_t - E_t \hat{\pi}_{t+1} - r_t^n), \quad (41)$$

where  $r_t^n$  denotes the “natural rate of interest” as is given by

$$r_t^n = E_t \Delta \xi_{c,t+1} + \frac{1}{\sigma} E_t \Delta y_{t+1}^n. \quad (42)$$

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TABLE 1  
Response of Interest Rates and Inflation to Monetary Shocks

	Policy News Shock			Fed Funds Shock		
	Nominal	Real	Inflation	Nominal	Real	Inflation
3M Treasury Yield	0.68 (0.16)			0.51 (0.19)		
6M Treasury Yield	0.84 (0.12)			0.59 (0.12)		
1Y Treasury Yield	0.98 (0.15)			0.41 (0.18)		
2Y Treasury Yield	1.05 (0.37)	1.00 (0.28)	0.05 (0.21)	0.49 (0.55)	0.53 (0.34)	-0.04 (0.25)
3Y Treasury Yield	0.97 (0.40)	0.94 (0.28)	0.03 (0.20)	0.37 (0.55)	0.42 (0.34)	-0.05 (0.29)
5Y Treasury Yield	0.63 (0.21)	0.58 (0.15)	0.05 (0.12)	0.10 (0.18)	0.21 (0.14)	-0.12 (0.10)
10Y Treasury Yield	0.29 (0.18)	0.38 (0.14)	-0.09 (0.09)	-0.02 (0.14)	0.10 (0.11)	-0.12 (0.08)
2Y Treasury Inst. Forward Rate	1.00 (0.49)	0.86 (0.31)	0.13 (0.26)	0.27 (0.64)	0.31 (0.40)	-0.04 (0.38)
3Y Treasury Inst. Forward Rate	0.60 (0.45)	0.72 (0.32)	-0.12 (0.17)	0.03 (0.55)	0.12 (0.34)	-0.10 (0.35)
5Y Treasury Inst. Forward Rate	0.13 (0.20)	0.39 (0.17)	-0.27 (0.09)	-0.11 (0.15)	0.07 (0.13)	-0.18 (0.08)
10Y Treasury Inst. Forward Rate	-0.13 (0.19)	0.09 (0.13)	-0.22 (0.10)	-0.13 (0.17)	-0.03 (0.12)	-0.09 (0.10)

Each estimate comes from a separate "regression." The dependent variable in each regression is the one day change in the variable stated in the left-most column. The independent variable is a change in the policy news shock (first three columns) or a change in the expected federal funds rate (last three columns) over a 30 minute window around the time of FOMC announcements. For the expected federal funds rate, this is the expected federal funds rate over the remainder of the current month unless the FOMC date in question occurs when there are 7 days or less remaining in the month, in which case it is the change in the expected federal funds rate over the next month. All results are based on Rigobon's (2003) method of identification by heteroskedasticity. The sample period is Jan 1st 2000 to Jan 25th 2012, except that we drop the second half of 2008, the first half of 2009 and a 10 day period after 9/11/2001. The "treatment" sample is a 30-minute window around all regularly scheduled FOMC announcements. The "control" sample is 2:05pm to 2:35pm on all Tuesdays and Wednesdays that are not FOMC meeting days. For 2Y and 3Y yields and real forwards, the sample starts in 2004. The sample size of the treatment sample for the 2Y and 3Y yields and forwards is 57. The sample size of the treatment sample for all other regressions is 89. Standard errors are calculated using a non-parametric bootstrap with 5000 iterations.

TABLE 2  
Comparison with Alternative Methodologies

	30-Minute Window Rigobon Policy News Shock		One-Day Window Rigobon Policy News Shock		One-Day Window OLS Policy News Shock		One-Day Window Rigobon 2Y Nominal Yield		30-Minute Window OLS Policy News Shock	
	Nominal	Real	Nominal	Real	Nominal	Real	Nominal	Real	Nominal	Real
2Y Treasury Yield	1.05 (0.37)	1.00 (0.28)	1.00 (1.90)	0.84 (0.95)	1.16 (0.16)	0.98 (0.15)	1.00 --	0.72 (6.34)	1.08 (0.33)	1.03 (0.23)
3Y Treasury Yield	0.97 (0.40)	0.94 (0.28)	0.91 (1.30)	0.79 (1.87)	1.12 (0.18)	0.95 (0.16)	0.97 (1.47)	0.73 (2.31)	1.00 (0.36)	0.96 (0.24)
5Y Treasury Yield	0.63 (0.21)	0.58 (0.15)	0.39 (0.89)	0.52 (0.28)	0.81 (0.11)	0.67 (0.10)	0.63 (5.69)	0.65 (2.10)	0.67 (0.20)	0.60 (0.14)
10Y Treasury Yield	0.29 (0.18)	0.38 (0.14)	0.01 (0.68)	0.28 (0.20)	0.48 (0.11)	0.45 (0.11)	0.12 (283.60)	0.37 (6.47)	0.33 (0.18)	0.39 (0.13)

Each estimate comes from a separate "regression." The dependent variable in each regression is the one day change in the variable stated in the left-most column. The independent variable is a change in the policy news shock over a 30-minute window around FOMC announcements (first two and last two columns) or a change in the policy news shock over a one-day window around FOMC announcements (columns three through six) or a change in the 2-Year nominal yield over the one-day window around FOMC announcements (columns seven and eight). Results in columns 1-4 and 7-8 are based on Rigobon's (2003) method of identification by heteroskedasticity, while results in columns 5-6 and 9-10 are based on OLS. The sample period is Jan 1st 2000 to Jan 25th 2012, except that we drop the second half of 2008, the first half of 2009 and a 10 day period after 9/11/2001. The sample of "treatment" days for the Rigobon method is all regularly scheduled FOMC meeting day. The sample of "control" days is all Tuesdays and Wednesdays that are not FOMC meeting days. For 2Y and 3Y yields and real forwards, the sample starts in 2004. Standard errors for the Rigobon method are calculated using a non-parametric bootstrap with 5000 iterations.

TABLE 3  
Effects of Monetary Shocks on Survey Expectations

	Nominal	Real	Inflation
1 quarter	1.05 (0.50)	1.17 (0.52)	-0.11 (0.27)
2 quarters	1.17 (0.52)	0.86 (0.52)	-0.48 (0.23)
3 quarters	0.97 (0.54)	1.65 (0.51)	-0.28 (0.22)
4 quarters	0.84 (0.52)	1.25 (0.51)	-0.31 (0.21)
5 quarters	0.63 (0.66)	0.45 (0.66)	0.18 (0.24)
6 quarters	1.79 (0.65)	1.56 (0.67)	0.23 (0.29)
7 quarters	3.88 (1.23)	3.75 (1.25)	0.13 (0.52)

This table presents the results of regressing changes in survey expectations from the *Blue Chip Economic Indicators* on the policy news shock. Since the *Blue Chip* survey expectations are available at a monthly frequency, we construct a corresponding monthly measure of our policy news shock. In particular, we calculate the sum of the policy news shocks that occur over the month except for those that occur in the first week (because we do not know whether these occurred before or after the survey response). The dependent variable is the change in the forecasted value of a variable N quarters ahead, between this month's survey and last month's survey. We consider the effects on expected future 3-month T-Bill rates, short-term real interest rates and inflation, where the inflation rate is the GDP deflator and the short-term real interest rate is calculated as the difference between the expected 3-month T-bill rate and the expected GDP deflator for a given quarter. The sample period is January 1995 to January 2012, except that we exclude the second half of 2008 and the first half of 2009.

TABLE 4  
Breakeven Inflation versus Inflation Swaps

	Breakeven	Swaps
Inflation Over Next 2 Years	0.05 (0.22)	0.33 (0.40)
Inflation Over Next 3 Years	0.03 (0.20)	0.32 (0.36)
Inflation Over Next 5 Years	0.05 (0.17)	-0.05 (0.18)
Inflation Over Next 10 Years	-0.09 (0.15)	-0.23 (0.19)

Each estimate comes from a separate "regression." The dependent variable in each regression is the one day change in expected inflation measured either by breakeven inflation from the difference between nominal Treasuries and TIPS (first column) or from inflation swaps (second column) for the period stated in the left-most column. The independent variable is a change in the policy new shock over a 30 minute window around the time of FOMC announcements. All results are based on Rigobon's (2003) method of identification by heteroskedasticity. The sample period is Jan 1st 2005 to Jan 25th 2012, except that we drop the second half of 2008 and the first half of 2009. The sample of "treatment" days is all regularly scheduled FOMC meeting days. The sample of "control" day is all Tuesdays and Wednesdays that are not FOMC meeting days. The sample size of the treatment sample is 49. Standard errors are calculated using a non-parametric bootstrap with 5000 iterations.

TABLE 5  
Mean Reversion

Horizon (Trading Days)	Nominal Yields			Real Yields		
	2-Year	3-Year	5-Year	2-Year	3-Year	5-Year
1	1.14 (0.21)	1.11 (0.23)	0.97 (0.25)	1.20 (0.37)	1.13 (0.36)	0.86 (0.23)
5	0.84 (0.63)	0.76 (0.58)	0.64 (0.48)	0.92 (0.67)	0.80 (0.61)	0.26 (0.35)
10	0.11 (0.81)	-0.02 (0.78)	-0.14 (0.70)	1.33 (0.94)	1.14 (0.83)	-0.05 (0.51)
20	0.28 (0.94)	0.16 (0.95)	0.14 (0.90)	1.49 (1.53)	0.93 (1.28)	0.13 (0.74)
60	0.76 (1.66)	0.21 (1.62)	-0.38 (1.48)	2.11 (2.84)	1.87 (2.33)	-0.13 (1.22)
125	4.53 (2.34)	3.86 (2.16)	2.82 (1.91)	7.41 (3.72)	6.22 (3.08)	2.38 (1.60)
250	6.20 (3.95)	5.52 (3.38)	4.18 (2.62)	9.95 (4.95)	8.42 (4.24)	3.81 (2.12)

This table presents the results of regressing the cumulative change in yields between the day before the FOMC announcement and 1, 5, 10, 20, 60, 125 and 250 trading days after the announcement on the policy news shock in the 30 minute interval surrounding the FOMC announcement. The first three columns present results for nominal zero coupon yields, and the next three columns present results for real zero coupon yields. Standard errors are in parentheses.

TABLE 6  
Monetary Non-Neutrality

<i>Panel A: Our Estimated Shock</i>	
Our Estimation of CEE/ACEL Model	16.2 [6.8, 146.7]
ACEL	10.7
CEE	1.4
<i>Panel B: CEE's Taylor Rule Shock</i>	
Our Estimation of CEE/ACEL Model	19.4 [6.2, 40.7]
ACEL	11.1
CEE	6.9

TABLE 7  
Estimates of Structural Parameters

$\xi_p$	0.93 [0.79, 0.99]
$\xi_w$	0.90 [0.00, 0.99]
$k_I$	6877.4 [1.1, 7966.6]
$\rho_1$	0.93 [0.85, 0.96]
$\rho_2$	0.61 [0.01, 0.88]

TABLE A1  
Response of Interest Rates to Monetary Shocks for Different Sample Periods

	Baseline Sample		Pre-Crisis (2000-2007)		Full Sample	
	Nominal	Real	Nominal	Real	Nominal	Real
3M Treasury Yield	0.68 (0.16)		0.77 (0.14)		0.61 (0.19)	
6M Treasury Yield	0.84 (0.12)		0.85 (0.13)		0.81 (0.15)	
1Y Treasury Yield	0.98 (0.15)		0.98 (0.15)		0.99 (0.16)	
2Y Treasury Yield	1.05 (0.37)	1.00 (0.28)	1.08 (0.43)	1.00 (0.31)	1.09 (0.33)	1.56 (0.39)
3Y Treasury Yield	0.97 (0.40)	0.94 (0.28)	1.00 (0.46)	0.93 (0.31)	1.09 (0.37)	1.35 (0.35)
5Y Treasury Yield	0.63 (0.21)	0.58 (0.15)	0.63 (0.21)	0.56 (0.16)	0.81 (0.22)	0.86 (0.20)
10Y Treasury Yield	0.29 (0.18)	0.38 (0.14)	0.33 (0.19)	0.42 (0.14)	0.52 (0.22)	0.63 (0.19)
2Y Treasury Inst. Forward Rate	1.00 (0.49)	0.86 (0.31)	1.04 (0.55)	0.87 (0.35)	1.17 (0.45)	0.90 (0.34)
3Y Treasury Inst. Forward Rate	0.60 (0.45)	0.72 (0.32)	0.62 (0.50)	0.73 (0.36)	0.98 (0.45)	0.97 (0.41)
5Y Treasury Inst. Forward Rate	0.13 (0.20)	0.39 (0.17)	0.17 (0.20)	0.46 (0.17)	0.44 (0.28)	0.70 (0.25)
10Y Treasury Inst. Forward Rate	-0.13 (0.19)	0.09 (0.13)	-0.03 (0.20)	0.20 (0.14)	0.06 (0.24)	0.19 (0.15)

Each estimate comes from a separate "regression." The dependent variable in each regression is the one day change in the variable stated in the left-most column. The independent variable is a change in the policy news shock over a 30 minute window around the time of FOMC announcements. All results are based on the Rigobon's (2003) method of identification by heteroskedasticity. The "treatment" sample is all regularly scheduled FOMC meeting days. The "control" sample is all Tuesdays and Wednesdays that are not FOMC meeting days and excluding a 10 day period after 9/11/2001. The baseline sample period is Jan 1st 2000 to Jan 25th 2012, except that we drop the second half of 2008 and the first half of 2009. The "Full Sample" is Jan 1st 2000 to Jan 25th 2012. The "Pre-Crisis" sample is 2000-2007. For 2Y and 3Y yields and real forwards, the sample starts in 2004. Standard errors are calculated using a non-parametric bootstrap with 5000 iterations.

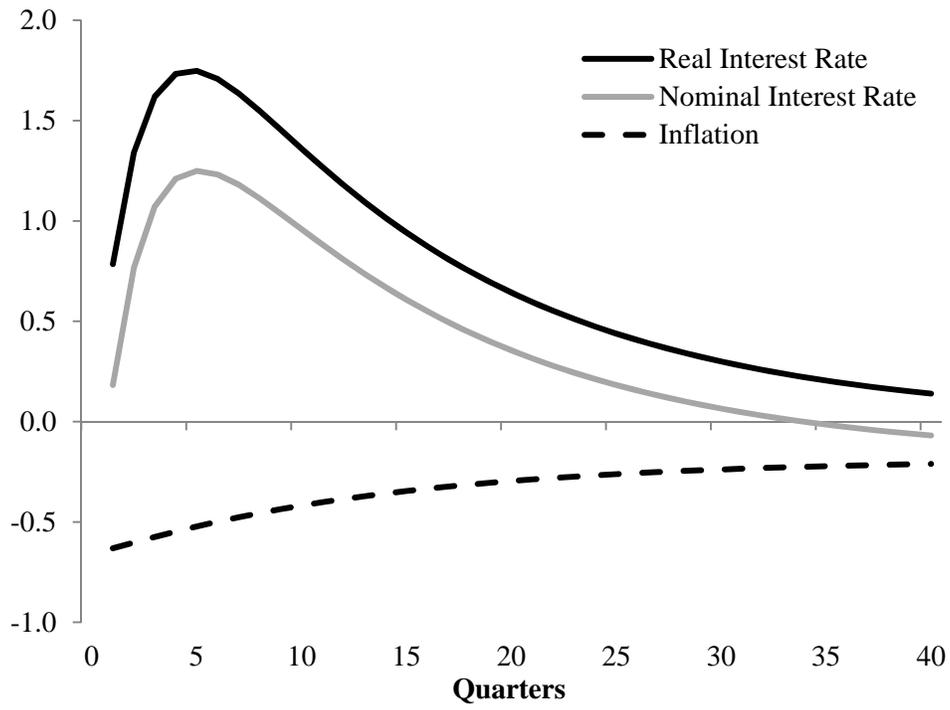


Figure 1: Interest Rate and Inflation in the Simple New Keynesian Model

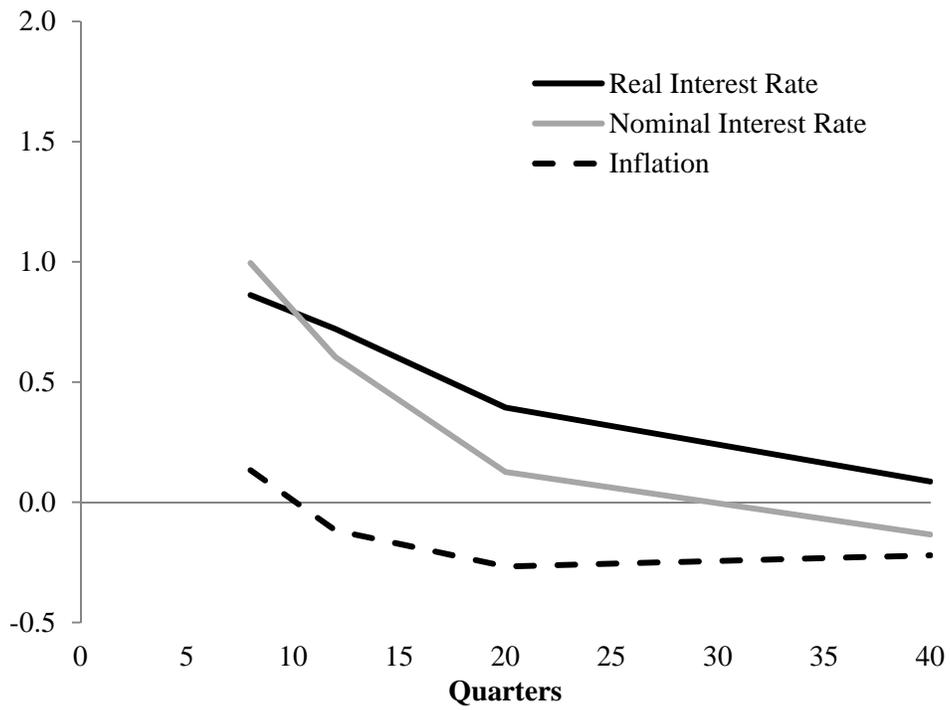


Figure 2: Interest Rates and Inflation in the Data

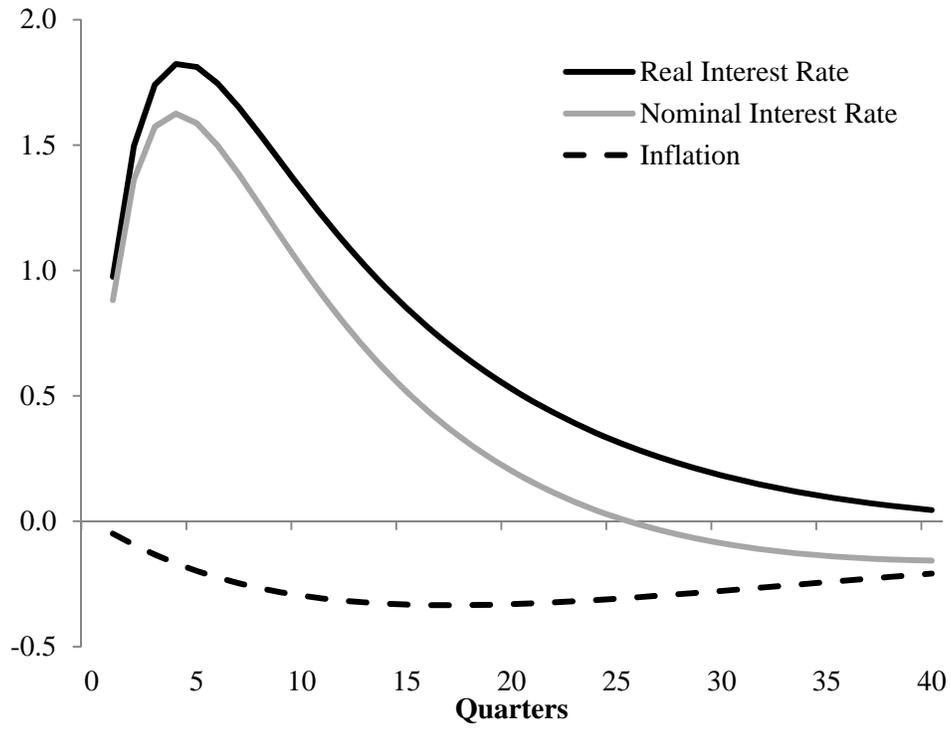


Figure 3: Response of Inflation and Interest Rates in Model with Hybrid Phillips Curve

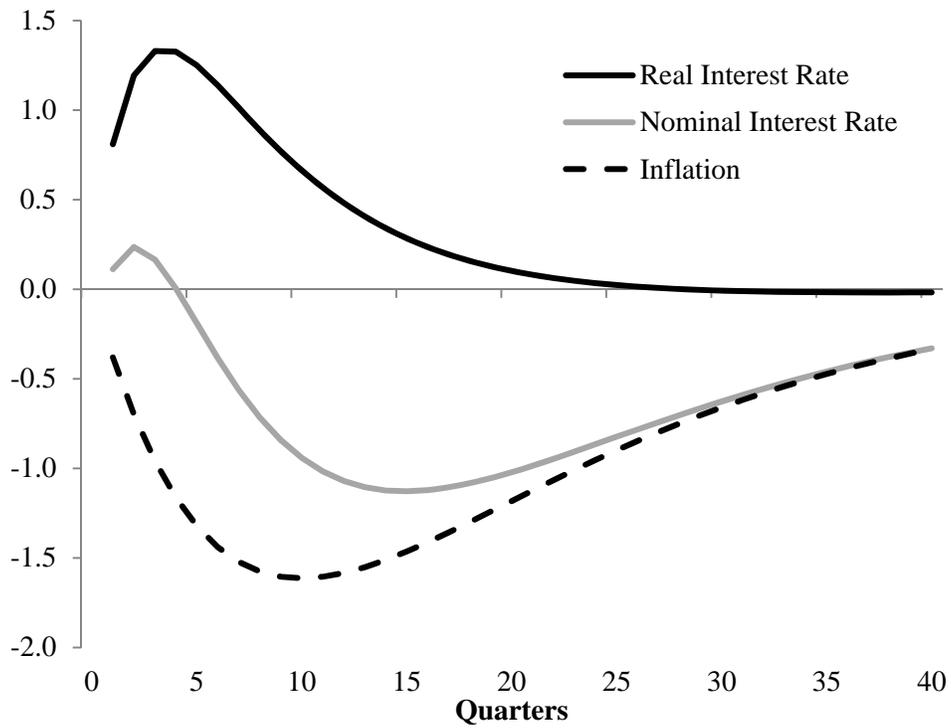


Figure 4: Response of Inflation and Interest Rates in Model with Hybrid Phillips Curve with Counter-Factually Large  $\kappa\zeta\sigma$

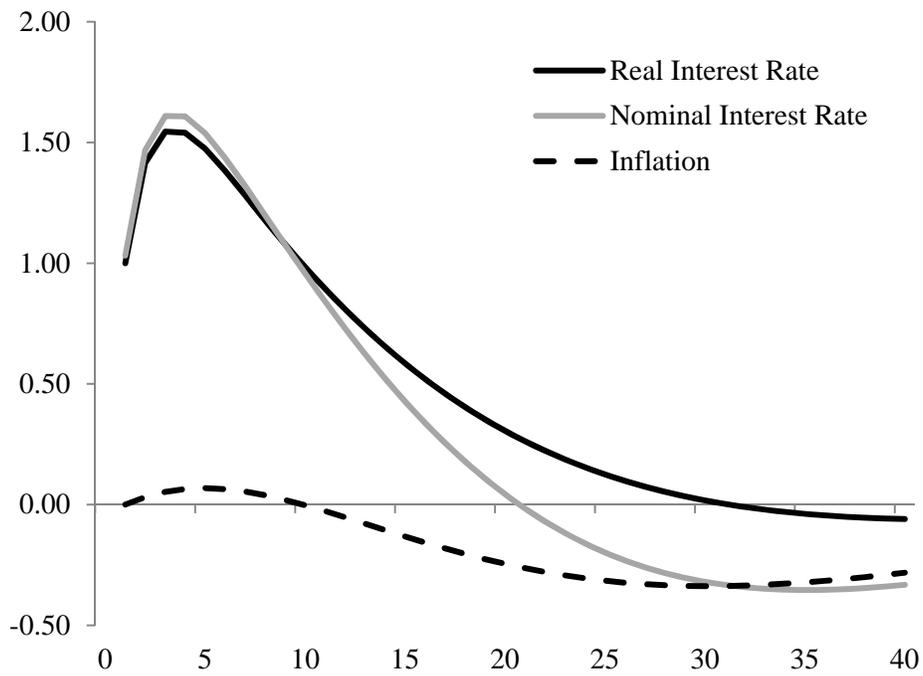


Figure 5: Response of Inflation and Interest Rates to Policy News Shock in Our Estimation of CEE/ACEL Model

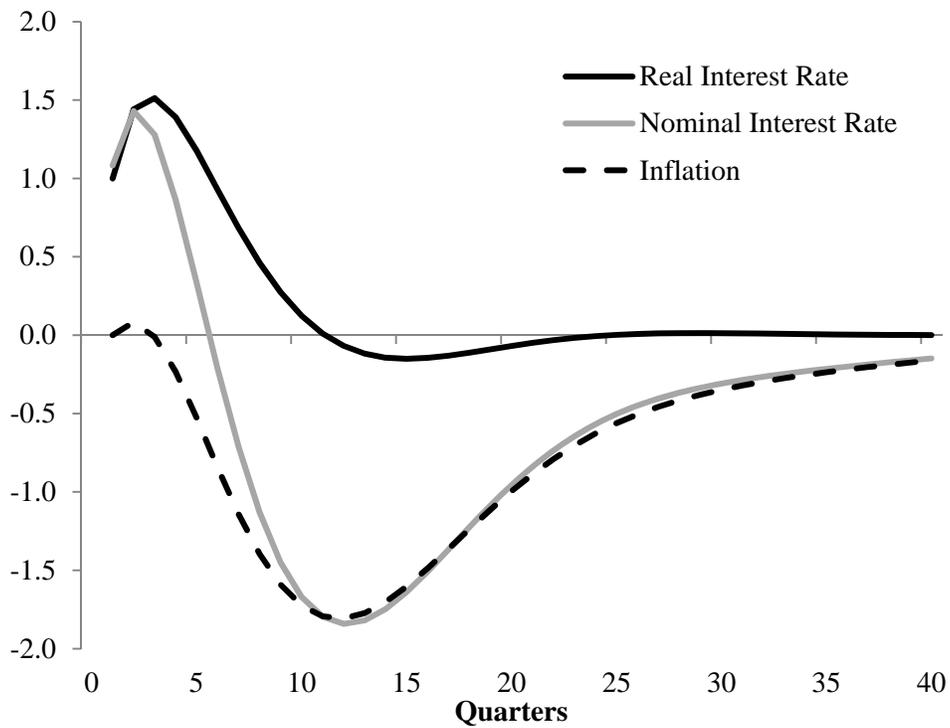


Figure 6: Response of Inflation and Interest Rates to Policy News Shock in CEE/ACEL Model with CEE Parameters

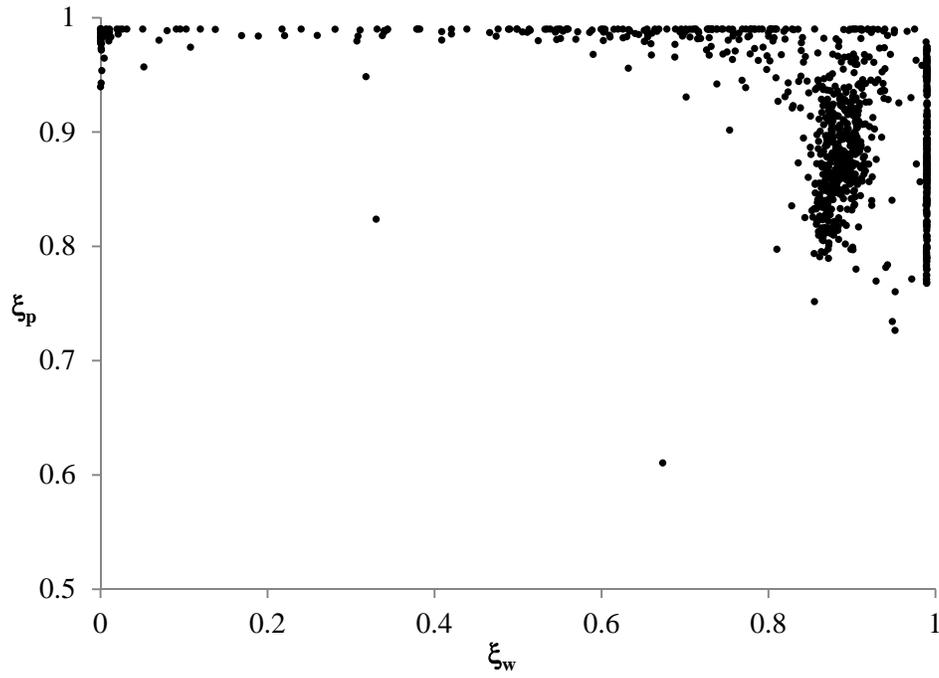


Figure 7: Scatterplot of Estimated Joint Distribution of  $\xi_w$  and  $\xi_p$

Note: The figure plots the values of  $\xi_w$  and  $\xi_p$  from the 1000 bootstrap draws we calculate.