Estimating Monetary Policy Rules with Serially Correlated Monetary Policy Shocks

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Abstract

The general equilibrium of a typical new-Keynesian model indicates that inflation and output gaps are determined by contemporaneous and historical monetary policy shocks. Thus, if monetary policy shocks are serially correlated, as argued in the literature, the conventional estimation of Taylor rules using lags of inflation and output gaps as instruments, has the endogeneity problem. This paper investigates the magnitude of the bias. We estimate Taylor rules using two methods that allow for the presence of serial correlation in monetary policy shocks. The first method "purifies" inflation and output gaps by removing from them the components attributable to monetary shocks. The second method uses as instruments a set of strictly exogenous variables. Results from both methods show that the endogeneity problem does not cause much bias in the conventional estimation of Taylor rules.

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1 Introduction

Monetary shocks are probably serially correlated. Rudebusch (2002) argues that it is the serial correlation in monetary shocks that accounts for the strong persistence in federal funds rates. Taylor (2007) blames the Fed's prolonged deviation from the Taylor rule during the period from 2002 to 2006 for the housing bubble that led to the 2007-2008 financial crisis. The deviation that lasts for years implies persistent monetary shocks in that period. As one piece of empirical evidence, unanticipated changes in federal funds rates, as a measure of monetary shocks, indeed show significant serial dependence.[1] Therefore, it is very restrictive to assume that monetary shocks are serially uncorrelated.

Serial correlation in monetary shocks complicates the identification of monetary policy rules. The general equilibrium of a typical new-Keynesian model indicates that inflation and output gaps are determined by, and thus endogenous to, contemporaneous and historical monetary shocks. If monetary shocks are serially correlated, lags of inflation and output gaps are also endogenous to monetary shocks. As pointed out by Cochrane (2011), the conventional estimation of Taylor rules that treats inflation, output gaps and their lags as exogenous may be biased.

The literature on monetary policy rules largely ignores or eschews the endogeneity problem. For instance, Clarida, Gali and Gertler's (2000) Taylor rule has no monetary shocks, and the Fed's target rate is fully justified by the response to inflation and output gaps. More typically, Taylor rules are specified with serially uncorrelated monetary shocks (e.g., Mavroeidis, 2010). The literature studying the origin of the strong persistence in the federal funds rate (e.g., Rudebusch, 2002 and Coibion and Gorodnichenko, 2011) also ignores the endogeneity problem. They admit serial correlation in monetary shocks associated with Taylor rules, but simply treats inflation and output gaps as exogenous to monetary shocks.

This paper investigates how the estimation of Taylor rules is afflicted by the endogeneity problem when monetary shocks have serial correlation. To this end, we develop two estimation methods that allow for the presence of serial correlation in monetary shocks. In the first

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[1]See Section 3 for details on the measure of monetary shocks.

method, we start with estimating "model-free" monetary shocks as unanticipated changes in the federal funds rate implied by federal funds futures rates as in Rudebusch (1998) and Kuttner (2001), and unanticipated changes in the 3-month Treasury bill rate derived from the Survey of Professional Forecasters. We then use our monetary shock estimates to "purify" inflation and output gaps by removing from them the components attributable to monetary shocks. Finally, the purified variables are used as instruments to estimate Taylor rules.

The second method uses as instruments a set of strictly exogenous variables. Specifically, the instruments are exogenous oil shocks identified by Hamilton (2003), the Congressional Budget Office's potential output, and technology shocks constructed by Basu, Fernald and Kimball (2006).

Our main findings are as follows. First, our monetary shock estimates are significantly serially correlated. Moreover, exogeneity of some conventional instruments to monetary shocks is rejected. Second, the endogeneity problem does not cause substantial bias in the conventional estimation of Taylor rules. In particular, the endogeneity problem does not lead to wrong conclusions on whether the Taylor principle is satisfied. We apply the purification procedure to the sample 1990-2007 and estimate Taylor rules using real-time data. Both the conventional estimation and the purification find that the inflation response parameter is significantly greater than unity. Taylor rules are estimated with very large uncertainty using revised data in this period. Neither the conventional estimation nor the purification finds the inflation response parameter significantly greater than unity. We compare our strictly exogenous estimation with Clarida, Gali and Gertler (2000) for the period from 1960 to 1996. The strictly exogenous estimation confirms Clarida, Gali and Gertler's (2000) main conclusion that the Taylor principle was violated in the pre-Volker era (1960Q1-1979Q2) but not in the Volker-Greenspan era (1979Q3-1996Q4). Due to the possible weak identification of Taylor rules using GMM, we also considered our conclusion based on weak identification robust inference.

The next section explains, in two simple general equilibrium models, the identification difficulty of Taylor rules with serially correlated monetary shocks. The two identification

strategies proposed by this paper are illustrated using the two models. Section 3 discusses the construction and properties of our measures of monetary shocks. The first identification method, which uses purified explanatory variables as instruments, is implemented in Section 4. Section 5 estimates Taylor rules using strictly exogenous variables as instruments. Finally, section 6 concludes.

2　Identification with Serially Correlated Monetary Shocks

This section discusses the endogeneity and identification issues in estimating Taylor rules when monetary shocks are serially correlated. We first consider the endogeneity problem and two identification strategies in a simple general equilibrium model borrowed from Cochrane (2011) with a slight, yet important modification. We also show that the endogeneity problem prevails in, and the identification strategies that we propose apply to, a canonical new-Keynesian model.

2.1　The Simple Model

The simple model consists of a Fisher equation, a Taylor rule and an equation defining the natural rate of interest,

$$i_t = r_t + \mathbb{E}_t \pi_{t+1}, \tag{1}$$

$$i_t = r + \phi \pi_t + x_t, \tag{2}$$

$$r_t = r + a_t, \tag{3}$$

where i_t is the nominal interest rate, r_t is the natural rate of interest, $\mathbb{E}_t \pi_{t+1}$ denotes the expectation of inflation conditional on information at time t, x_t are possibly *serially correlated* monetary shocks, following $x_t = \rho x_{t-1} + \varepsilon_t$, and a_t are technology changes, following $a_t = \rho_a a_{t-1} + \varepsilon_{at}$. ε_t and ε_{at} are *i.i.d.* and not cross-correlated.

A unique stable solution of this model exists if $\phi > 1$. Solving the model forward, the

unique stable solution for π_t is[2]

$$\pi_t = \frac{a_t}{\phi - \rho_a} - \frac{x_t}{\phi - \rho}. \qquad (4)$$

The general equilibrium constraint given by (4) indicates that π_t is correlated with x_t. More-over, lags of π_t are also correlated with x_t, because of the serial correlation in x_t. Taking (4) into consideration, OLS estimator for ϕ in the Taylor rule (2) is inconsistent. Moreover, lags of π_t are not valid instruments for estimating (2). This is Cochrane's (2011) o the identification of Taylor rules.

Despite the identification difficulty, the Taylor rule (2) can still be consistently estimated. There are two ways. First, suppose we have an estimate for the monetary shocks, denoted by $\hat{x}_t$,

$$\hat{x}_t = x_t + \xi_{xt},$$

where ξ_{xt} are measurement errors. Assume that $\hat{x}_t$ are well estimated, i.e., ξ_{xt} are not correlated with x_t and π_t, and variance of ξ_{xt} is much smaller than that of x_t. Regressing π_t on $\hat{x}_t$, the residual from the regression, denoted by $\hat{\pi}_t$, is given by

$$\hat{\pi}_t = \pi_t - \hat{\beta}\hat{x}_t,$$

where $\hat{\beta} = \frac{\mathbb{Cov}(\pi_t, x_t + \xi_{xt})}{\mathbb{Var}(x_t + \xi_{xt})}$. Using the property that ξ_{xt} are not correlated with x_t and π_t, $\hat{\beta}$ simplifies to $-\frac{1}{\phi - \rho}\frac{\mathbb{Var}(x_t)}{\mathbb{Var}(x_t) + \mathbb{Var}(\xi_{xt})}$. Since $\mathbb{Var}(\xi_{xt})$ is much smaller than $\mathbb{Var}(x_t)$, $\hat{\pi}_t$ is approximated by

$$\hat{\pi}_t \approx \pi_t + \frac{1}{\phi - \rho}\hat{x}_t = \frac{a_t}{\phi - \rho_a} + \frac{\xi_{xt}}{\phi - \rho}.$$

$\hat{\pi}_t$ are correlated with π_t but not with x_t, and thus are valid instruments for estimating (2).

The second method to estimate the Taylor rule is more straightforward. It relies on estimates of fundamental shocks, a_t. Suppose we have an estimate for a_t, denoted by $\hat{a}_t$,

$$\hat{a}_t = a_t + \xi_{at},$$

[2]Combining (1),(2) and (3) yields $\pi_t = \frac{1}{\phi}(\mathbb{E}_t\pi_{t+1} + a_t - x_t)$. Using the fact that $\pi_{t+1} = \frac{1}{\phi}(\mathbb{E}_{t+1}\pi_{t+2} + a_{t+1} - x_{t+1})$, π_t is expressed as $\pi_t = \frac{1}{\phi}(\mathbb{E}_t(\frac{1}{\phi}(\mathbb{E}_{t+1}\pi_{t+2} + a_{t+1} - x_{t+1})) + a_t - x_t)$. Do the substitution recursively and apply the law of iterative expectation. It follows that $\pi_t = \frac{a_t}{\phi}(1 + \frac{\rho_a}{\phi} + \frac{\rho_a^2}{\phi^2} + ...) - \frac{x_t}{\phi}(1 + \frac{\rho}{\phi} + \frac{\rho^2}{\phi^2} + ...)$. It is easy to see (4) follows.

where ξ_{at} are measurement errors, uncorrelated with a_t and π_t. For instance, a_t can be estimated as the Solow residual. It is clear that $\hat{a}_t$ can be used to instrument π_t in estimating (2).[3]

2.2   A New-Keynesian Model

The endogeneity problem and the proposed identification strategies in the simple model have their analogies in a typical new-Keynesian model. The new-Keynesian model that we consider consists of an aggregate demand equation, a Phillips curve and a Taylor rule,

$$y_t = \mathbb{E}_t y_{t+1} - \sigma(i_t - \mathbb{E}_t \pi_{t+1}) + x_{dt}, \tag{5}$$

$$\pi_t = \beta \mathbb{E}_t \pi_{t+1} + \gamma(y_t - \bar{y}_t) + x_{\pi t}, \tag{6}$$

$$i_t = \alpha + \rho_1 i_{t-1} + \rho_2 i_{t-2} + (1 - \rho_1 - \rho_2)(\phi_\pi \mathbb{E}_t \pi_{t+1} + \phi_y(y_t - \bar{y}_t)) + x_{it}, \tag{7}$$

where y_t is the percentage deviation of output from its steady state, $\bar{y}_t$ is the flexible price counterpart of y_t, and $x_{dt}, x_{\pi t}, x_{it}$ are demand shocks, inflation shocks and monetary shocks, respectively. The inclusion of i_{t-1} and i_{t-2} in the Taylor rule implies that the central bank smoothes the interest rate. Monetary shocks are serially correlated. Assume $x_{it} = \rho_i x_{it-1} + \varepsilon_{it}$. ε_{it} are *i.i.d.* disturbances.

Define output gaps g_t as $g_t = y_t - \bar{y}_t$. Given $\phi_\pi > 1$, for a wide range of other parameters, the new-Keynesian model consisting of (5), (6) and (7) has a unique stable solution (Appendix A gives details on the solution and parameter restrictions for determinacy). In particular, solutions to i_t and $\mathbb{E}_t \pi_{t+1}$ take the form,

$$\mathbb{E}_t \pi_{t+1} = \kappa_1 i_{t-1} + \kappa_2 i_{t-2} + \kappa_i x_{it} + \kappa_{\bar{y}} \bar{y}_t + \kappa_d x_{dt} + \kappa_\pi x_{\pi t}, \tag{8}$$

$$i_t = \alpha + \lambda_1 i_{t-1} + \lambda_2 i_{t-2} + \lambda_i x_{it} + \lambda_{\bar{y}} \bar{y}_t + \lambda_d x_{dt} + \lambda_\pi x_{\pi t}, \tag{9}$$

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[3]If $a_t \equiv 0$ or the Taylor rule has a "stochastic intercept", i.e., $i_t = r_t + \phi \pi_t + x_t$, the solution for inflation (4) reduces to $\pi_t = -\frac{x_t}{\phi - \rho}$. Since π_t is completely determined by x_t, it is not possible to identify the Taylor rule. This is Cochrane's (2011) critique on the identification of Taylor rules. However, the "stochastic intercept" Taylor rule is is not a practical rule, because it requires the Fed to respond contemporaneously to unobservable technology changes. Therefore, Cochrane's critique is not a practical concern.

where κ's and λ's are parameters. Solutions to y_t takes a similar structure. The general equilibrium puts constraints on the estimation of the Taylor rule (7). First of all, g_t and $\mathbb{E}_t \pi_{t+1}$ are in general correlated with x_{it}. Moreover, because x_{it} are serially correlated, lags of inflation and output gaps are also correlated with x_{it}. Therefore, lags of g_t and π_t, which are widely used in the literature, are not valid instruments for estimating (7).

Suppose we have a good estimate for monetary shocks, $\hat{x}_{it} = x_{it} + \xi_{it}$, where ξ_{it} are measurement errors, not correlated with x_{it} and other structural shocks. Assume that variance of ξ_{it} is much smaller than that of x_{it}. Residuals from the regression of $\mathbb{E}_t \pi_{t+1}$ on $(1, \hat{x}_{it})$, denoted by $\widehat{\mathbb{E}}_t \pi_{t+1}$, are exogenous to x_{it}. To see this point, substituting recursively solutions for i_{t-1} and i_{t-2} into (8), $\mathbb{E}_t \pi_{t+1}$ can be expressed as a linear function of $x_{it}, \bar{y}_t, x_{dt}, x_{\pi t}$ and their infinite order lags,

$$\mathbb{E}_t \pi_{t+1} = \alpha_\pi + \sum_{j=0}^{\infty} \theta_{ij} x_{i,t-j} + \sum_{j=0}^{\infty} \theta_{\bar{y}j} \bar{y}_{t-j} + \sum_{j=0}^{\infty} \theta_{dj} x_{d,t-j} + \sum_{j=0}^{\infty} \theta_{\pi j} x_{\pi,t-j},$$

where α_π is a constant, and θ's are parameters. Projection of $\sum_{j=0}^{\infty} \theta_{ij} x_{i,t-j}$ on $(1, \hat{x}_{it})$ makes the residual orthogonal to x_{it}.[4] Apply the same procedure to g_t, i_{t-1}, i_{t-2} and denote the residuals by $\hat{g}_t, \hat{i}_{t-1}$ and $\hat{i}_{t-2}$, respectively. Purified variables, $\widehat{\mathbb{E}}_t \pi_{t+1}, \hat{g}_t, \hat{i}_{t-1}$ and $\hat{i}_{t-2}$ are exogenous to monetary shocks, and thus valid instruments for estimating Taylor rules.

As explained in the simple model, the quality of the purification relies on the quality of the estimate of monetary shocks. We assume our monetary shocks are precisely estimated in the sense that measurement errors are true errors, not correlated with structural shocks, and variance of measurement errors is much less than that of monetary shocks.

The estimation strategy using strictly exogenous variables is straightforward. Since $x_{dt}, x_{\pi t}$ and $\bar{y}_t$ are uncorrelated with monetary shocks, they are valid instruments for estimating (7). We now turn to applying these methods to estimate Taylor rules.

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[4]As an illustration, consider the regression of $\theta_1 x_{i,t-1} + x_{it}$ on $\hat{x}_{it}$. The estimated coefficient is $\rho_i \theta_1 + 1$, where ρ_i is the first order autocorrelation of x_{it}. Here we assumed that $\mathbb{V}\text{ar}(\xi_{it})$ is much smaller than $\mathbb{V}\text{ar}(x_{it})$. Residuals from the regression are given by $\theta_1(1 - \rho_i^2) x_{i,t-1} - \theta_1 \rho_i \varepsilon_{it}$. It is easy to verify that the residuals are orthogonal to x_{it}.

3 Monetary Shocks

We use two estimates for monetary shocks. The first estimate is market-based, derived from the unanticipated changes in the federal funds rate. To gauge expectations on federal funds rates, we use the 30-day federal funds futures, which provide a market-based measure for expected federal funds rates. Krueger and Kuttner (1996) show that federal funds futures rate is an "efficient" forecast for the actual funds rate in the sense that forecast errors are not significantly correlated with information available when the forecast is made.

Let $f_{t,t-1}$ denote the futures rate in quarter t implied by funds futures contracts trading at the end of quarter $t-1$.[5] Following Kuttner (2001), the futures rate, $f_{t,t-1}$, is interpreted as the sum of conditional expectation of the average funds rate in quarter t plus a term of risk premium, $\mu_{t,t-1}$, accruing to investors long in the futures contract,

$$f_{t,t-1} = \mathbb{E}_{t-1} \frac{1}{S} \sum_{s \in S} i_s + \mu_{t,t-1}, \tag{10}$$

where i_s is the overnight funds rate on day s of quarter t, $\mathbb{E}_{t-1} i_s$ is the expectation of i_s at the end of quarter $t-1$, and S is the number of days in quarter t. Rearranging (10), we get a measure of expected federal funds rates as

$$\mathbb{E}_{t-1} \frac{1}{S} \sum_{s \in S} i_s = f_{t,t-1} - \mu_{t,t-1}. \tag{11}$$

Monetary shocks as the average unanticipated changes in funds rates in quarter t are given by,

$$\hat{x}_{it} = \frac{1}{S} \sum_{s \in S} (i_s - \mathbb{E}_{t-1} i_s). \tag{12}$$

Substituting (11) into (12), yields

$$\hat{x}_{it}^{ff} = \frac{1}{S} \sum_{s \in S} i_s - f_{t,t-1} + \mu_{t,t-1}. \tag{13}$$

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[5]The federal funds futures contract on the average federal funds overnight rate for the delivery month. The quarterly futures rate is calculated as the average rate of contracts maturing in the quarter.

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(13) indicates that monetary shocks are measured as the difference between the actual average funds rate and the futures rate plus a risk premium term.[6] However, risk premia, $\mu_{t,t-1}$, are unknown. Following Guidolin and Thornton (2008), we assume $\hat{x}_{it}$ averages to zero in a rolling window with width τ.[7] It implies that $\mu_{t,t-1}$ can be expressed as

$$\mu_{t,t-1} = -\frac{1}{\tau}\sum_{j=0}^{\tau-1}(\frac{1}{S}\sum_{s\in S}i_s - f_{t-j,t-1-j}). \tag{14}$$

Monetary shocks estimates $\hat{x}_{it}^{ff}$ are available from 1988Q4, because federal funds futures contracts began to trade on the Chicago Board of Trade in October 1988.

Piazzesi and Swanson (2008) show that federal funds futures rates contain substantial risk premia, which are strongly countercyclical. Obviously, (14) is just a rough adjustment, and $\hat{x}_{it}^{ff}$ can still be contaminated by risk premia. As a robustness check, we also consider Kuttner's (2001) intraday monetary shocks, which measures changes of federal funds futures rates in a narrow time window around FOMC announcements.

Suppose day d of month t has an FOMC announcement, in a short time interval around the announcement, futures rate changes from f_d^- before the announcement to f_d^+ after. For the futures contract that matures in month t, $S-d$ days are affected by the announcement. If no further FOMC announcements occur in the month, the difference between f_d^+ and f_d^-, re-scaled to reflect the number of days affected measures the unanticipated changes in federal funds rates. Thus, Kuttner's measure of intraday monetary shocks, denoted by $\hat{x}_{it}^K$ is given by

$$\hat{x}_{it}^K = \frac{S}{S-d}(f_d^+ - f_d^-). \tag{15}$$

This measure is free of risk premia, provided that risk premia does not change in the short time window.[8]

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[6]Rudebusch (1998) first constructs monetary shocks using this approach and criticizes measures of monetary shocks from VARs by contrasting VAR residuals and $\hat{x}_{it}$.

[7]I set τ to eight quarters. Note that when τ equals to the sample size, it is equivalent to assuming constant risk premia. Changing τ does not change the main results of this paper.

[8]This is not an unreasonable assumption given the short time window as one day in Kuttner (2001), and thirty minutes or one hour in Gürkaynak, Sack and Swanson (2005).

Ignoring inter-meeting FOMC announcements, Kuttner's (2001) intraday measure of monetary shocks has usually eight observations in one year, as the FOMC normally meets eight times every year. To match the frequency of $\hat{x}_{it}^{ff}$, we use only announcements in the middle of a quarter for $\hat{x}_{it}^{K}$.[9] We instead denote observations at all announcement dates by $\hat{x}_{it}^{M}$, which has eight observations every year. Intraday monetary shocks used in this paper are constructed by Gürkaynak, Sack and Swanson (2005) with a time window of thirty minutes. They are available from the announcement on February 8, 1990.

The second estimate of monetary shocks is survey-based. Monetary shocks in quarter t is denoted by $\hat{x}_{it}^{spf}$. It is constructed as the difference between average 3-month Treasury bill rates in quarter t and the median forecast made in the middle of quarter t from the Survey of professional forecasters (SPF). Though SPF forecast errors do not correspond directly to the unanticipated changes in the federal funds rate (which is the Fed's policy instrument and the left hand side variable of Taylor rules), as shown in Appendix B, SPF forecast errors and the unanticipated changes in the federal funds rate are closely related through a term structure model. SPF forecasts for the 3-month Treasury bill rate since 1981Q3 are available from Philadelphia Fed's real-time data sets.

3.1 Serial Correlations in Monetary Shocks

Figure 1 shows, in percentage points, monetary shocks measured from quarterly unexpected changes in federal funds futures rates, $\hat{x}_{it}^{ff}$, intraday unexpected changes in funds futures rates at FOMC announcement dates, $\hat{x}_{it}^{K}$, and unexpected changes in 3-month Treasury bill rates from the Survey of Professional Forecasters, $\hat{x}_{it}^{spf}$. These shocks differ considerably. First of all, $\hat{x}_{it}^{K}$ are less volatile than the other two measures. Second, $\hat{x}_{it}^{ff}$ have moderate correlation with $\hat{x}_{it}^{spf}$, whereas $\hat{x}_{it}^{K}$ do not seem to move together with them at all, as shown in Table 1. Risk premia can be the underlying reason for the difference. The contamination of risk premia can make $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$ more volatile and correlated with each other.

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[9]These announcement are usually in February, May, August and December, which match as close as possible the timing of the Survey of Professional Forecasters, as we will discuss later.

However, they agree on the serial correlation. Table 2 summarizes serial autocorrelation in estimated monetary shocks. Sample autocorrelation functions for the first four lags are reported.[10] Q-test and F-test report p-values associated with Ljung-Box test with four lags and joint significance test for regressing monetary shocks on their one-to-four lags, respectively. It is shown that autocorrelations are considerable, and very big occasionally. For the full sample from 1990Q1 to 2007Q4, serial correlations in all estimates are found significant at a 5% significance level. More interestingly, even the risk-premium-free measures of monetary shocks, $\hat{x}_{it}^{K}$, and $\hat{x}_{it}^{M}$ are significantly serially correlated. Since they are pure monetary surprises, not contaminated by risk premia, we have very strong evidence against the idea that monetary shocks have no serial correlation.

Ljung-Box test with two lags and F-test for regressing monetary shock on their first two lags (results not reported) give similar results. All series except $\hat{x}_{it}^{ff}$ are found to be serially correlated by the Ljung-Box test. The F-test finds $\hat{x}_{it}^{spf}$ significantly serially correlated, though it does not reject the hypothesis that the other three series have no serial correlation.

To check whether the serial correlation is more pronounced in the years before the sub-prime mortgage crisis as suggested by Taylor (2007), Table 2 reports test results for the subsample 2001Q1-2007Q4. All series except $\hat{x}_{it}^{M}$ are found to be significantly serially correlated at a 5% significance level.

To capture the persistence of monetary shocks, Table 2 also reports the sum of coefficients of regressing monetary shocks on their first four lags.[11] The associated t-statistics are presented in parentheses following the persistence parameters. While persistence parameters for $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{K}$ are small and insignificant, they are substantial and significant for $\hat{x}_{it}^{K}$ and $\hat{x}_{it}^{spf}$. In a nutshell, monetary shocks seem to respond substantially to previous realizations.

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[10]Since FOMC announcements are not evenly distributed in one year, the time interval between two consecutive observations of $\hat{x}_{it}^{M}$ is not constant. However, we still treat observations prior to $\hat{x}_{it}^{M}$ as its lags.

[11]In an autocorrelation process, a big persistence parameter implies a large accumulative impulse response functions. Therefore, the persistence parameter measures the influence of a monetary shock on shocks in the following periods.

3.2 Exogeneity of Conventional Instruments

Since the evidence of serial correlation in monetary shocks is strong, we need to be skeptical on the validity of conventional instruments for estimating (7). We test exogeneity of these instruments by regressing estimated monetary shocks on these instruments and report in Table 3 p-values associated with F-statistics for testing joint significance of the instruments. Instruments that we consider are those used in Clarida, Gali and Gerlter (2000): *one-to-four* lags of inflation, π, output gaps, g, federal funds rates, i, interest rate spreads between 10-year Treasury bonds and 3-month Treasury bills, s, commodity inflation, π_{com}, and M2 growth rates, M2.[12]

In the full sample, exogeneity of all instrumental sets put together can not be rejected at a 10% significance level. For individual instrumental sets, however, results are mixed. Exogeneity of s is rejected at a 5% significance level for all three monetary shocks. Funds rates, i, are rejected to be exogenous at a 10% significance level by $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$, but not by $\hat{x}_{it}^{K}$. Exogeneity of the other four instrumental sets can not be rejected at a 10% significance level, except that of M2 to $\hat{x}_{it}^{K}$. Results from subsamples are also mixed. In the subsample 1990Q1-2000Q4, exogeneity of instrumental sets except i can not be rejected at conventional significance levels. In the second subample 2001Q1-2007Q4, exogeneity of i to $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$ is rejected.

The high chance for i and s to be rejected as exogenous to $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$ is probably caused by risk premia contained in $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$. Following Piazzesi and Swanson (2008), we use 3-moth Treasury bill rates as proxies for risk premia and study their relation with estimated monetary shocks. The last column of Table 4 shows clearly that $\hat{x}_{it}^{K}$ is exogenous to risk premia, but $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$ are not.

Instead of imposing rational expectation and using lagged explanatory variables as instruments, a large amount of research estimates (7) using nonlinear least squares and real-time forecasts for inflation and output gaps (e.g., Orphanides, 2001; Boivin, 2006). As pointed out

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[12]These instrumental sets are standard in estimating Taylor rules of the kind of (7). See also Consolo and Favero (2009) and Mavroeidis (2010).

by Boivin (2006), a potential problem for the real-time estimation is the contemporaneous correlation between forecasts and monetary shocks. We test exogeneity of real-time forecasts and summarize the results in Table 4, which reports p-values associated with F-tests of regressing estimated monetary shocks on federal funds target rates, i_{t-1}, i_{t-2}, average one-to-four quarter ahead inflation forecasts, $\mathbb{E}_t\pi_{t,4}$, and current-quarter forecasts for output gaps, $\mathbb{E}_t g_t$, respectively. The data that we use for $\mathbb{E}_t\pi_{t,4}$ and $\mathbb{E}_t g_t$ are Greenbook forecasts for GDP inflation and output gaps made by the staff economists of the Fed's Board of Governors.[13] Note that we replace $\mathbb{E}_t\pi_{t+1}$ in (7) by $\mathbb{E}_t\pi_{t,4}$, because it is argued that in the Greenspan era (1987Q3-2005Q4), which our sample roughly overlaps, the Fed targets medium term inflation.[14]

We can not reject, at conventional significance levels, exogeneity of real-time regressors to $\hat{x}_{it}^K$ and $\hat{x}_{it}^M$, in both the full sample and subsamples. However, exogeneity to $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$ is rejected. In a linear Taylor rule (i_{t-1} and i_{t-2} are excluded), exogeneity of explanatory variables $\mathbb{E}_t\pi_{t,4}$ and $\mathbb{E}_t g_t$ can not be rejected. Again, we suspect that the endogeneity is due to the fact that $\hat{x}_{it}^{ff}$ and $\hat{x}_{it}^{spf}$ contain risk premia, to which i_{t-1} and i_{t-2} are correlated.

4 Estimation Using Purification

As discussed in Section 2, serially correlated monetary shocks imply that neither lagged explanatory variables nor real-time forecasts are valid instruments. Moreover, our monetary shock estimates demonstrate strong serial correlation, but fail to reach a clear-cut conclusion about exogeneity of lagged explanatory variables and real-time forecasts. It is therefore

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[13]Before 1996Q1, $\mathbb{E}_t\pi_{t,4}$ are measured by GNP/GDP implicit deflator. After 1996Q2, $\mathbb{E}_t\pi_{t,4}$ are instead measured by chain-weight price index for GDP. The data are obtained from Philadelphia Fed's Greenbook Data Sets.

[14]In a testimony (quoted in Orphanides, 2001), the Fed's Chairman then Alan Greenspan said: "Because monetary policy works with a lag, it not the conditions prevailing today that are critical but rather those likely to prevail six to twelve months, or even longer, from now" (January 21, 1997 testimony by Chairman Greenspan before the Senate Committee on the Budget). Using $\mathbb{E}_t\pi_{t+1}$ does not change the main conclusion of Table 4.

worthwhile to estimate (7) using the purification method developed in Section 2.

4.1   Real-Time Data

We estimate (7) using real-time forecasting data to measure expectations on inflation and output gaps. The data that we use for $\mathbb{E}_t\pi_{t,4}$ and $\mathbb{E}_t g_t$ are Greenbook forecasts for GDP inflation and output gaps made by the staff economists of the Fed's Board of Governors. Due to the availability of our monetary shocks estimates, the sample is confined to 1990Q1-2007Q4. Table 5 reports estimation results. In the table, (1),(2) and (3) denote Taylor rules excluding lagged interest rates, including only i_{t-1} and including both i_{t-1} and i_{t-2}, respectively. In model (4), inflation expectations are measured by $\mathbb{E}_t\pi_{t+1}$, instead of $\mathbb{E}_t\pi_{t,4}$ as in the other three models. Panel A presents the conventional estimates using nonlinear least squares. Panel B,C and D present estimates using instruments purified by our three monetary estimates, respectively.

Panel A shows that estimates for ϕ_π is sensitive to the measure of inflation expectations, but robust to different specifications of interest rate smoothing. The estimate for ϕ_π is significantly greater than unity, when inflation expectations are measured by $\mathbb{E}_t\pi_{t,4}$. However, when inflation expectations are measured by $\mathbb{E}_t\pi_{t+1}$, the estimate for ϕ_π is smaller and not significantly different from unity.

Comparing estimates in Panel B,C and D with that in Panel A, two interesting results stand out. First, purifications using the three estimates of monetary shocks do not differ much from each other. Second, and more importantly, no matter which monetary shock estimate is used, point estimates and standard errors of Taylor rule parameters from the purification are very close to that from the ordinary estimation. In particular, estimates for ϕ_π are significantly greater than unity when inflation expectations are measured by $\mathbb{E}_t\pi_{t,4}$, but not significantly different from unity when inflation is measured by $\mathbb{E}_t\pi_{t+1}$.

To assess to which extent our monetary shocks estimates capture the true monetary shocks, the bottom rows of Table 5 report correlations of our monetary estimates and Taylor rule residuals, which, if correctly estimated, are true monetary shocks. First, it is shown

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that the correlation for model (3) and (4), which are featured with second-order interest rate smoothing, is stronger than that for model (1) and (2). Second, the intraday monetary shocks, $\hat{x}_{it}^K$, have the strongest correlations with Taylor rule residuals, and $\hat{x}_{it}^{spf}$ have the weakest correlation. Third, the correlation is not substantially affected by the purification. Since $\hat{x}_{it}^K$ is the most reliable estimate for monetary shocks, we may conclude that our monetary estimates are reasonably good.[15]

Table 6 presents estimates using real-time data at FOMC meeting dates. Different from results using quarterly data shown in Table 6, data at FOMC announcement dates estimate ϕ_π greater than unity at a 10% significance level (for the estimation using purification, at a 5% significance level) in model (4), where inflation is measured by $\mathbb{E}_t \pi_{t+1}$. Despite this difference, point estimates and standard errors using NLS (Panel A) do not differ much from that using $\hat{x}_{it}^M$ purified instruments (Panel B). In summary, purifying explanatory variables with monetary shocks does not change much estimates of Taylor rules. The ordinary estimation of Taylor rules in the literature does not seem to be seriously biased due to the endogeneity problem.

4.2   Revised Data

We estimate (7) using ex post, revised data. Specifically, i_t are quarterly-average effective federal funds rates, π_t GDP deflator inflation of the 2013 vintage, g_t percentage difference between actual GDP and CBO's potential GDP (2013 vintage). Assuming rational expectation and replacing $\mathbb{E}_t \pi_{t+1}$ and $\mathbb{E}_t g_t$ by π_{t+1} and g_t, respectively, (7) can be written as

$$i_t = \alpha + \rho_1 i_{t-1} + \rho_2 i_{t-2} + (1 - \rho_1 - \rho_2)(\phi_\pi \pi_{t+1} + \phi_y g_t) + \eta_t, \tag{16}$$

where $\eta_t = x_{it} - \phi_\pi(1 - \rho_1 - \rho_2)(\pi_{t+1} - \mathbb{E}_t \pi_{t+1})$. Note that η_t are serially correlated due to the serial correlation in x_{it}.

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[15]It also provides evidence that second-order interest rate smoothing is a more proper specification for Taylor rules, compared with specifications without interest rate smoothing or with first-order interest rate smoothing.

The sample that we consider is 1990Q1-2007Q4, due to the availability of our monetary shock estimates. Table 7 presents GMM estimates using both conventional instruments (0-4 lags of funds rates, inflation and output gaps) and the corresponding, $\hat{x}_{it}^K$ purified instruments. The first panel reports two-step GMM estimates and the second panel reports continuously updating GMM estimates. It is shown that the inflation response parameter ϕ_π is poorly estimated in this sample using revised data. Both GMM methods estimate ϕ_π with large uncertainty. ϕ_π is not significantly different from unity. The purification slightly increases point estimates and associated standard errors, but does not change the conclusion that estimates for ϕ_π is not significantly different from unity.

There are two possible reasons for the poor estimation results. First, as pointed out by Orphanides (2001), real-time data and revised data can lead to different Taylor rules. The revised data can have greater uncertainty about Taylor rues compared to real-time data. Second, the identification of GMM on Taylor rules in our short sample can be very weak (Stock, Wright and Yogo, 2002; Mavroeidis, 2010). The uncertainty about ϕ_π can be even larger than what Table 7 suggests. To check whether ϕ_π is well identified by GMM and the revised data, we construct a weak identification robust confidence interval for ϕ_π using Stock and Wright's (2000) S-set.

The 90% S-set for ϕ_π is constructed as follows. Let ϕ_π vary from a given lower bound to a upper bound with a certain step size. For each value of ϕ_π, the other three parameters of (16) are estimated using continuously updating GMM. The p-value associated with the J-test is calculated. The S-set collects all values of ϕ_π for which the overidentification restrictions are not rejected at a 10% significance level (i.e., p-values are greater than 0.10). The S-set consists of parameter values at which the joint hypothesis of $\phi_\pi = \phi_{\pi 0}$ and the overidentification restrictions can not be rejected. Figure 2 plots S-sets for the conventional instrument set and the purified instrument set. The 90% confidence interval seems to be infinite. It is a sign that both the conventional instruments and the purified instruments are very weak or irrelevant for estimating (16). Compared to conventional instruments, the purification does not substantially changes the poor identification.

5   Strictly Exogenous Estimation

In construction of monetary shocks in Section 3, we assume the Fed and the financial market have the same information set and use the same monetary policy rule to determine or forecast federal funds rates. If either the Fed has information advantage over the market as shown in Romer and Romer (2000) or the Fed uses a monetary policy rule different from that perceived by the market, interest rate changes unexpected by the market are not purely monetary shocks, but rather contaminated by information difference and model difference between the Fed and the financial market.

As discussed in Section 2, the validity of the purification procedure depends on one key assumption that measurement errors of our monetary shock estimates are relatively small, and not correlated with structural shocks. The potential information difference and model difference between the Fed and the financial market can make measurement errors of our monetary shock estimates very large, and even worse, correlated structural shocks. The moderate correlations between our monetary shock estimates and Taylor rule residuals, as shown in Table 5, indicate that our monetary shocks used for purification indeed capture partially monetary shocks. Nevertheless, the lack of strong correlation among our monetary shock estimates indicates that at least the other two are poor estimates if one of them is a sound measure. This casts doubts on the quality of the purification procedure. This section therefore uses the second method discussed in Section 2 to estimate Taylor rules.

5.1   The Data

Our strictly exogenous instruments are demand shocks, x_{dt}, inflation shocks, $x_{\pi t}$, and potential output, $\bar{y}_t$, and their lags. We proxy demand shocks and inflation shocks by oil shocks and technology shocks.[16] We use oil shocks constructed by Hamilton (2003), who measures oil shocks as the percentage drops of world oil production caused by five major wars in the

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[16]We also considered Ramey's (2011) military spending shocks as a proxy for demand shocks. However, in our sample, the military spending shocks do not have significant explanatory power for inflation and output gaps.

Middle East. Since it is not easy to doubt the exogeneity of those wars to monetary policy, Hamilton's (2003) measure avoids the problem of endogeneity suffered by measures of oil shocks based on oil prices. Hamilton's oil shocks have four observations in our sample range, corresponding respectively to Arab-Israel war (1973Q4), Iranian revolution (1978Q4), Iran-Iraq war (1980Q4) and Persian Gulf war (1990Q3).

Technology shocks that we use are constructed by Basu, Fernald and Kimball (2006). The shocks are essentially purified Solow residuals. The purified technology changes are defined as a weighted sum of sectoral technology change. Basu, Fernald and Kimball (2006) claim that their measures of technology changes control for the aggregation effect cross sectors, time varying utilization of capital and labor, and arbitrary returns to scale.

Finally, since potential output measures the long-run production capacity of the economy, it is independent from monetary policy shocks, which are believed to have only short-term effects. We use CBO's measure of potential output, which is derived from a neoclassical production accounting.

5.2   Results

Table 8 summarizes the first stage fit of strictly exogenous variables to Taylor rule variables, i_{t-1}, i_{t-2}, π_{t+1} and g_t. We regress Taylor rule variables on 0-4 lags of oil shocks (OIL), linearly detrended logarithm of potential GDP (POTGDP) and technology changes (TECH). Corresponding to each Taylor rule variable, the first row reports adjusted R^2 and the second row reports p-values associated F-statistics for the joint significance of 0-4 lags of the instrument. It is shown that potential GDP has significant explanatory power for all Taylor rule variables. In the pre-Volker sample (1960Q1-1979Q2), oil shocks and technology changes explain a substantial part of Taylor rule variables, except g_t. In the Volker-Greenspan sample (1979Q3-1996q4), however, only potential GDP has good fit to Taylor rule variables. This is a sign of weak instruments, to which we return in the next subsection.

We estimate (7) using two-step GMM. Instruments include 0-4 lags of oil shocks, technology changes, and detrended potential output. The weighting matrix is Newey-West type

with a bandwidth of 4. Table 9 presents the strictly exogenous estimation together with Clarida, Gali and Gertler's (2000) results for comparison purpose.

Though the strictly exogenous estimation gives less precise estimates, it still shows that in the pre-Volker era the inflation response parameter, ϕ_π, is significantly less than one ($\phi_\pi = 0.72$, $s.e.= 0.07$), but significantly greater than one ($\phi_\pi = 2.20$, $s.e.= 0.39$) in the Volker-Greenspan era. This confirms Clarida, Gali and Gertler's (2000) conclusion: the Taylor principle was violated in the pre-Volker era but was satisfied in the Volker-Greenspan era. It seems that the bias caused by serially correlated monetary shocks in Clarida, Gali and Gertler (2000) is small, at least not big enough to give misleading conclusion on the Taylor principle.

We do two robustness checks. First, as shown in Table 8, oil shocks and technology changes have very weak explanatory power to g_t in the pre-Volker sample, and to i_{t-1} and i_{t-2} in Volker-Greenspan sample. We thus drop oil shocks and technology changes in our instrument set, and use only potential GDP to estimate (7). Second, we consider alternative specifications of Taylor rules by excluding from the baseline Taylor rule i_{t-2}, i_{t-1} and i_{t-2}, and $\mathbb{E}_t g_t$, respectively. Table 4 reports results from the robustness check. Using 0-7 lags of potential GDP, the Taylor rule is less precisely estimated, compared to the baseline instrument set. ϕ_π is estimated significantly greater than unity in the Volker-Greenspan era, but not significantly different from unity in the pre-Volker era.[17] Alternative specifications also largely confirm that the inflation response parameter was less than one in the pre-Volker ear, but greater than one in the Volker-Greenspan era.

5.3 Weak Identification Robust Inference

Results in Table 9 and 10 should be taken with a grain of salt. Due to the low correlation between instruments and explanatory variables, inference based on the reported standard errors with the assumption of normal asymptotic distribution can be misleading (Stock,

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[17]Though ϕ_π is not significantly different from unity in the pre-Volker era, the one-side hypothesis that ϕ_π is greater than unity is rejected at a 10% significance level.

Wright and Yogo, 2002).[*to be added*]

6 Conclusions

Monetary shocks are indeed serially correlated. The serial correlation makes the conventional estimation of Taylor rules subject to the endogeneity problem, as implied by the general equilibrium of a new-Keynesian model. We propose two methods to estimate Taylor rules with serially correlated monetary shocks. Both methods show that the endogeneity problem does not cause much bias in the conventional estimations. In particular, the endogeneity problem does not affect the conclusion on the Taylor principle. Clarida, Gali and Gertler's (2000) main conclusion that the US economy in the pre-Volker era was subject to self-fulfilling inflation but not in the Volker-Greenspan era seems to hold in the presence of persistent monetary shocks. One explanation for the innocuous endogeneity problem in estimating Taylor rules is that monetary shocks account for only a very limited fraction of variations in inflation and output, as emphasized by the VAR literature.

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A   Solutions to the New-Keynesian Model

To solve the model consisting of (5), (6) and (7), we have to make some assumptions about the exogenous disturbances. Assume x_{dt} and $x_{\pi t}$ are *i.i.d.* Assume that monetary policy shocks, x_{it}, and potential output, $\bar{y}_t$, follow AR(1) process,

$$x_{it} = \rho_i x_{i,t-1} + \varepsilon_{it} \text{ and } \bar{y}_t = \rho_{\bar{y}} \bar{y}_{t-1} + \varepsilon_{\bar{y}t}.$$

Define $\omega_t = i_{t-1}$. The model can be written in a first-order vector autoregression form,

$$A\mathbb{E}_t X_{t+1} = BX_t + Dv_t + C \tag{17}$$

where

$$\mathbb{E}_t X_{t+1} = \left(\mathbb{E}_t y_{t+1}, \mathbb{E}_t \pi_{t+1}, i_t, \omega_t, x_{it}, \bar{y}_t\right)'$$
$$X_t = \left(y_t, \pi_t, i_{t-1}, \omega_{t-1}, x_{i,t-1}, \bar{y}_{t-1}\right)'$$
$$v_t = \left(x_{dt}, x_{\pi t}, \varepsilon_{it}, \varepsilon_{\bar{y}t}\right)'$$

and C is a vector of constants and A, B, D are conforming matrices.

To check the model's determinacy, following Mavroeidis (2010) I calibrate $\beta = 0.99$, $\gamma = 0.3$ and $\sigma = 1$. Set $\rho_i = \rho_{\bar{y}} = 0.95$. A numerical experiment shows that determinacy of the model is *not* sensitive to ρ_1 and ρ_2, given $\rho_1 + \rho_2 < 1$. The minimum ϕ_π above which the model attains determinant solution depends on ϕ_y. Given $\phi_y = 0, 0.5, 1, 2$, the minimum ϕ_π are $1.0, 0.98, 0.97, 0.93$, respectively. Therefore, in a model with interest rate smoothing, $\phi_\pi = 1$ is still an threshold for model's determinacy as in the canonical new-Keynesian model.

Partition X_t into two blocks with $X_{1t} = (y_t, \pi_t)'$ and $X_{2t} = (i_{t-1}, \omega_{t-1}, x_{i,t-1}, \bar{y}_{t-1})'$. $\mathbb{E}_t X_{t+1}$ is partitioned accordingly. Use Blanchard-Kahn method to solve (17), yielding

$$X_{1t} = MX_{2t} + Nv_t + Q \tag{18}$$

where Q is a vector of constants, M and N are conforming matrices. The solution shows that inflation and output gaps are linear functions of state variables, X_{2t}, and structural shocks, v_t. In particular,

$$y_t = \psi_1 i_{t-1} + \psi_2 i_{t-2} + \psi_i x_{it} + \psi_{\bar{y}} \bar{y}_t + \psi_d x_{dt} + \psi_\pi x_{\pi t},$$

where ψ's are parameters determined by M, N and Q. The solution to π_t takes a similar form.

Since the model is determinant, all forward looking variables, $\mathbb{E}_t X_{t+1}$, is completely determined by state variables and structural shocks. To see this point, substituting (18) into (17), $\mathbb{E}_t X_{t+1}$ is expressed as a linear function of state variables and structural shocks. Therefore, $\mathbb{E}_t \pi_{t+1}$ has a solution in the form of (8).

B   A Term Structure Model

Let r_t denote the average 3-month Treasury bill rate in quarter t, and $r_{t,d}$ the median forecast for r_t from the Survey of Professional Forecasters at day d in quarter t. It follows that

$$r_t = \frac{1}{S} \sum_{s=1}^{S} r_s \text{ and } r_{t,d} = \frac{1}{d} \sum_{s=1}^{d} r_s + \frac{1}{S-d} \sum_{s=d+1}^{S} \mathbb{E}_{t,d} r_s,$$

where r_s is the Treasury bill rate at day s of quarter t, S is the number of days in the quarter, and $\mathbb{E}_{t,d}r_s$ is the conditional expectation for r_s on day d. The expectation hypothesis for the term structure states that the 3-month Treasury bill rate, r_s, equals to the average expected funds rates over the life of the bill (including the funds rate on day s), plus a term of risk premium,

$$r_s = \frac{1}{m}(\sum_{j=0}^{m} \mathbb{E}_{t,s}i_{s+j}) + \mu_{t,s},$$

where $\mathbb{E}_{t,s}i_{s+j}$ is the expected overnight funds rate on the jth day after s, conditional on information available at day s in quarter t; $\mu_{t,s}$ denotes the risk premium and m is the bill's number of days to maturity.[18]

The unanticipated change of the bill rate is given by

$$r_t - r_{t,d} = \frac{1}{S-d} \sum_{s=d+1}^{S} (r_s - \mathbb{E}_{t,d}r_s), \tag{19}$$

$$= \frac{1}{S-d} \sum_{s=d+1}^{S} \left(\frac{1}{m}\sum_{j=0}^{m}(\mathbb{E}_{t,s}i_{s+j} - \mathbb{E}_{t,d}i_{s+j})\right) + \frac{1}{S-d} \sum_{s=d+1}^{S} (\mu_{t,s} - \mathbb{E}_{t,d}\mu_{t,s}),, \tag{20}$$

where $\mathbb{E}_{t,s}i_{s+j} - \mathbb{E}_{t,d}i_{s+j}$ are revisions on the expectation for the funds rate j days away from s. The revisions are made based on information arriving in the period between day d and day s, they can thus be used to proxy surprises in the federal funds rate realized during the period. Therefore, our monetary shocks from SPF are defined as

$$\hat{x}_{it}^{spf} = \frac{1}{S-d} \sum_{s=d+1}^{S} \left(\frac{1}{m}\sum_{j=0}^{m}(\mathbb{E}_{t,s}i_{s+j} - \mathbb{E}_{t,d}i_{s+j})\right),$$

a weighted average of funds rate shocks realized in quarter t.

To illustrate, assume $i_s, d < s \leq S$ follows a random walk process,

$$i_s = i_d + \epsilon_{d+1} + ... + \epsilon_s,$$

where i_d is the funds rate at day d, and ϵ_s denotes monetary shocks at day s. It follows that

$$\frac{1}{m}\sum_{j=0}^{m}(\mathbb{E}_{t,s}i_{s+j} - \mathbb{E}_{t,d}i_{s+j}) = \epsilon_{d+1} + ... + \epsilon_s$$

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[18]Note that m is not a constant. It equals to ninety on each Thursday and for the days from Friday to the next Wednesday, m reduces one by one day.

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and

$$\hat{x}_{it}^{spf} = \frac{1}{S-d} \sum_{s=d+1}^{S} (\epsilon_{d+1} + ... + \epsilon_s).$$

In this particular case, $\hat{x}_{it}^{spf}$ weights more on shocks occur days close to day d.

The risk premium surprise, $\frac{1}{S} \sum_{s=1}^{S} (\mu_{t,s} - \mathbb{E}_{t-1}\mu_{t,s})$ in (20), is unknown. I assume the risk premium is either constant for time-varying yet completely predictable, i.e., risk premium surprise equals to zero. Consequently, our measure of monetary shocks from SPF equals to forecast errors for the 3-month Treasury bill rate, $\hat{x}_{it}^{spf} = r_t - r_{t,d}$.

Table 1: Correlations of Monetary Shocks

	90:1-07:4			90:1-00:4			01:1-07:4		
	$\hat{x}_{it}^{ff}$	$\hat{x}_{it}^{K}$	$\hat{x}_{it}^{spf}$	$\hat{x}_{it}^{ff}$	$\hat{x}_{it}^{K}$	$\hat{x}_{it}^{spf}$	$\hat{x}_{it}^{ff}$	$\hat{x}_{it}^{K}$	$\hat{x}_{it}^{spf}$
$\hat{x}_{it}^{ff}$	1.00	**0.11**	**0.37**	1.00			1.00		
$\hat{x}_{it}^{K}$	0.11	1.00	**0.13**	0.08	1.00		0.12	1.00	
$\hat{x}_{it}^{spf}$	0.38	0.13	1.00	0.39	0.26	1.00	0.40	-0.05	1.00

Notes: This table reports cross-correlations of monetary shocks measured from quarterly unexpected changes in federal funds futures rates, $\hat{x}_{it}^{ff}$, intraday unexpected changes in funds futures rates, $\hat{x}_{it}^{K}$, and unexpected changes in 3-month Treasury bill rates from the Survey of professional forecasters. Bold numbers at the upper triangle are pairwise correlations for $\hat{x}_{it}^{ff}$ ranging from 1988Q3 to 2007Q4.

Table 2: Serial Correlations in Monetary Shocks

		Autocorrelation				Q-test	F-test	Persistence
		1	2	3	4			
$\hat{x}_{it}^{ff}$								
	90:1-07:4	0.18	0.13	0.14	-0.28	0.02	0.01	0.12 (0.66)
	90:1-00:4	-0.05	0.01	0.24	-0.32	0.09	0.13	-0.01 (-0.05)
	01:1-07:4	0.47	0.28	0.14	-0.13	0.03	0.02	0.43 (2.12)
$\hat{x}_{it}^{K}$								
	90:1-07:4	0.10	0.30	0.08	0.22	0.02	0.10	0.47 (2.44)
	90:1-00:4	0.29	0.40	0.21	0.12	0.01	0.09	0.53 (2.41)
	01:1-07:4	-0.34	0.04	-0.28	0.39	0.02	0.05	0.44 (-0.94)
$\hat{x}_{it}^{M}$								
	90:1-07:8	-0.21	0.11	-0.18	0.22	0.00	0.01	-0.07 (-0.37)
	90:1-00:8	-0.32	0.22	-0.22	0.30	0.00	0.00	-0.22 (-0.09)
	01:1-07:8	-0.02	-0.13	-0.06	0.00	0.88	0.65	-0.23 (-0.73)
$\hat{x}_{it}^{spf}$								
	90:1-07:4	0.41	0.10	0.07	0.00	0.01	0.00	0.60 (2.38)
	90:1-00:4	0.00	-0.08	0.07	-0.15	0.79	0.71	-0.15 (-0.44)
	01:1-07:4	0.56	0.19	0.02	0.09	0.02	0.00	1.12 (2.92)

Notes: This table summarizes serial correlations in monetary shocks measured from quarterly unexpected changes in federal funds futures rates, $\hat{x}_{it}^{ff}$, intraday unexpected changes in funds futures rates, $\hat{x}_{it}^{K}$, and unexpected changes in 3-month Treasury bill rates from the Survey of professional forecasters. $\hat{x}_{it}^{M}$ denotes intraday unexpected changes in funds futures rates at FOMC meeting dates. The first to the fourth order autocorrelations are reported. Q-test reports p-values for Ljung-Box test with four lags. F-test reports p-values for the joint significance of regressing monetary shocks on their *one-to-four* lags. The last column reports the sum of the regression coefficients from the F-test (t-statistics are in parentheses), a measure of persistence in monetary shocks.

Table 3: Exogeneity of Conventional Instruments to Monetary Shocks

		Clarida, Gali and Gertler (2000)							T-Bill
		i	π	g	s	π_{com}	M2	All	
$\hat{x}_{it}^{ff}$									
	90:1-07:4	0.00	0.27	0.22	0.05	0.98	0.93	0.17	0.01
	90:1-00:4	0.01	0.34	0.53	0.28	0.94	0.36	0.12	0.00
	01:1-07:4	0.02	0.00	0.00	0.11	0.25	0.19	–	0.02
$\hat{x}_{it}^{K}$									
	90:1-07:4	0.53	0.85	0.21	0.04	0.57	0.09	0.10	0.60
	90:1-00:4	0.09	0.90	0.82	0.11	0.75	0.39	0.18	0.22
	01:1-07:4	0.91	0.97	0.95	0.35	0.19	0.47	–	0.91
$\hat{x}_{it}^{spf}$									
	90:1-07:4	0.09	0.41	0.29	0.00	0.25	0.62	0.48	0.09
	90:1-00:4	0.15	0.36	0.20	0.87	0.12	0.62	0.02	0.05
	01:1-07:4	0.02	0.87	0.80	0.01	0.66	0.97	–	0.02

Notes: This table lists p-values associated with F-tests for regressing monetary shocks on Clarida, Gali and Gertler's (2000) instrument sets: *one-to-four* lags of federal funds rates, i, inflation, π, output gaps, g, interest rate spreads between 10-year Treasury bonds and 3-month Treasury bills, s, commodity inflation, π_{com}, and M2 growth rates, M2. The last column reports p-values associated with t-statistics from regressing monetary shock estimates on 3-month Treasury bill rates. See also notes to Table 2.

Table 4: Exogeneity of Real-Time Taylor Rule Explanatory Variables to Monetary Shocks

		i_{t-1}	i_{t-2}	$\mathbb{E}_t\pi_{t,4}$	$\mathbb{E}_t g_t$	π and g	All
$\hat{x}_{it}^{ff}$							
	90:1-07:4	0.00	0.00	0.23	0.34	0.27	0.00
	90:1-00:4	0.01	0.05	0.20	0.93	0.35	0.04
	01:1-07:4	0.00	0.02	0.21	0.00	0.00	0.02
$\hat{x}_{it}^{K}$							
	90:1-07:4	0.96	0.71	0.27	0.91	0.55	0.20
	90:1-00:4	0.17	0.09	0.80	0.61	0.88	0.15
	01:1-07:4	0.47	0.78	0.90	0.86	0.98	0.71
$\hat{x}_{it}^{M}$							
	90:1-07:8	0.18	0.21	0.30	0.38	0.35	0.65
	90:1-00:8	0.13	0.17	0.47	0.53	0.44	0.44
	01:1-07:8	0.41	0.21	0.32	0.42	0.61	0.76
$\hat{x}_{it}^{spf}$							
	90:1-07:4	0.02	0.01	0.78	0.36	0.62	0.05
	90:1-00:4	0.06	0.07	0.84	0.99	0.98	0.17
	01:1-07:4	0.00	0.00	0.02	0.00	0.01	0.01

Notes: This table lists p-values associated with F-tests for regressing monetary shocks on real-time explanatory variables of Taylor rules: federal funds target rates, i_{t-1} and i_{t-2}, one-to-four quarter ahead inflation forecasts, $\mathbb{E}_t\pi_{t,4}$, current-quarter forecasts for output gaps, $\mathbb{E}_t g_t$. The last two columns report p-values of regressing monetary shock estimates on $\mathbb{E}_t\pi_{t,4}$ together with $\mathbb{E}_t g_t$, and on all variables, respectively. See also notes to Table 2.

Table 5: Estimating Taylor Rules Using Real-Time Data

	A: Real-Time Data				B: Purified by $\hat{x}_{it}^{ff}$				C: Purified by $\hat{x}_{it}^{K}$				D: Purified by $\hat{x}_{it}^{spf}$			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
i_{t-1}		0.67	1.25	0.47		0.69	1.33	1.46		0.70	1.24	1.34		0.67	1.23	1.33
		(0.05)	(0.07)	(0.07)		(0.05)	(0.06)	(0.75)		(0.05)	(0.07)	(0.08)		(0.05)	(0.07)	(0.08)
i_{t-2}			-0.49	-0.54			-0.54	-0.59			-0.46	-0.49			-0.48	-0.51
			(0.07)	(0.07)			(0.07)	(0.07)			(0.07)	(0.08)			(0.07)	(0.08)
$\mathbb{E}_t \pi_{t,4}$	1.86*	1.70*	1.73*		1.84*	1.72*	1.80*		1.87*	1.60*	1.62*		1.86*	1.69*	1.72*	
	(0.13)	(0.20)	(0.16)		(0.14)	(0.22)	(0.20)		(0.14)	(0.19)	(0.15)		(0.13)	(0.19)	(0.15)	
$\mathbb{E}_t \pi_{t+1}$				1.24				1.29				1.10				1.26
				(0.24)				(0.33)				(0.28)				(0.22)
$\mathbb{E}_t g_t$	0.58	0.74	0.63	0.69	0.57	0.76	0.66	0.76	0.58	0.76	0.66	0.73	0.58	0.74	0.64	0.71
	(0.07)	(0.07)	(0.05)	(0.08)	(0.07)	(0.07)	(0.05)	(0.10)	(0.07)	(0.08)	(0.06)	(0.10)	(0.07)	(0.07)	(0.05)	(0.08)
Corr($\hat{x}_{it}^{ff}$, Resid)	-0.27	0.12	0.33	0.38	-0.27	0.15	0.40	0.44	-0.27	0.15	0.35	0.38	-0.27	0.12	0.33	0.36
Corr($\hat{x}_{it}^{K}$, Resid)	-0.03	0.39	0.41	0.43	-0.03	0.40	0.41	0.43	-0.03	0.42	0.44	0.46	-0.03	0.39	0.41	0.43
Corr($\hat{x}_{it}^{spf}$, Resid)	-0.27	0.07	0.13	0.20	-0.27	0.09	0.18	0.24	-0.27	0.10	0.16	0.21	-0.26	0.07	0.13	0.19

Notes: This table reports results of estimating Taylor rules using purified real-time explanatory variables. Results using unpurified real-time explanatory variables are presented in Panel A. Panel B, C and D report estimates using instruments purified by $\hat{x}_{it}^{ff}$, $\hat{x}_{it}^{K}$ and $\hat{x}_{it}^{spf}$, respectively. (1),(2) and (3) denote Taylor rules without lags of interest rates, with only i_{t-1}, with both i_{t-1} and i_{t-2}, respectively. The three rows at the bottom report correlations of Taylor rule residuals and our "model-free" estimates of monetary shocks. Newey-West standard errors with a bandwidth of 4 are presented in parentheses. * indicates that the estimate is significantly different from one at the 1% significance level. The sample is 1990Q1-2007Q4.

Table 6: Estimating Taylor Rules Real-Time Data at FOMC Announcement Dates

	A: Real-Time Data				B: Purified by x_{it}^M			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
i_{t-1}		0.81	1.14	1.21		0.81	1.16	1.22
		(0.04)	(0.10)	(0.09)		(0.04)	(0.09)	(0.09)
i_{t-2}			-0.30	-0.34			-0.31	-0.35
			(0.08)	(0.08)			(0.08)	(0.08)
$\mathbb{E}_t \pi_{t,4}$	1.85*	1.70*	1.70*		1.86*	1.77*	1.79*	
	(0.14)	(0.24)	(0.22)		(0.14)	(0.26)	(0.25)	
$\mathbb{E}_t \pi_{t+1}$				1.36				1.50
				(0.24)				(0.26)
$\mathbb{E}_t g_t$	0.60	0.74	0.70	0.73	0.61	0.76	0.73	0.77
	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)	(0.07)	(0.06)	(0.07)
$\mathrm{Corr}(\hat{x}_{it}^K, \mathrm{Resid})$	0.13	0.42	0.46	0.48	0.13	0.43	0.47	0.49

Notes: This table reports results of estimating Taylor rules using real-time explanatory variables at FOMC announcement dates. Results using NLS are presented in Panel A. Panel B shows estimates using instruments purified by $\hat{x}_{it}^M$. (1),(2) and (3) denote Taylor rules without lags of interest rates, with only i_{t-1}, with both i_{t-1} and i_{t-2}, respectively. The bottom row reports correlations of Taylor rule residuals and $\hat{x}_{it}^M$. Newey-West standard errors with a bandwidth of 4 are presented in parentheses. * indicates the estimate is significantly different from one at the 1% significance level. The sample is from the first meeting of 1990 to the last meeting of 2007.

Table 7: Estimating Taylor Rules Revised Data

	2SGMM		CUGMM	
	Conven.	Purified	Conven.	Purified
i_{t-1}	1.53	1.49	1.46	1.40
	(0.08)	(0.08)	(0.09)	(0.09)
i_{t-2}	-0.57	-0.52	-0.51	-0.45
	(0.08)	(0.08)	(0.08)	(0.08)
$\mathbb{E}_t\pi_{t+1}$	1.10	1.34	1.52	1.72
	(1.06)	(1.16)	(0.93)	(1.02)
$\mathbb{E}_t g_t$	1.19	1.57	0.69	1.05
	(0.63)	(0.63)	(0.53)	(0.54)
p	0.16	0.16	0.25	0.21

Notes: This table reports results of estimating (16) using revised data. The first panel reports two-step GMM estimates and the second panel reports continuously updating GMM estimates. The weight matrix is Newey-West type with a bandwidth of 4. Conven. stands for conventional instruments: 0-4 lags of funds rates, inflation and output gaps. Purified stands for $\hat{x}_{it}^K$ purified 0-4 lags of funds rates, inflation and output gaps. p's are p-values associated with J-tests of overidentification restrictions. Newey-West standard errors are reported in parentheses. The sample is 1990Q1-2007Q4.

Table 8: The First Stage Fit of Strictly Exogenous Instruments

| | 1960Q1-1979Q2 | | | | 1979Q3-1996Q4 | | |
	OIL	POTGDP	TECH		OIL	POTGDP	TECH
i_{t-1}	0.43	0.61	0.71		**0.06**	0.77	**0.07**
	0.00	0.00	0.00		**0.11**	0.00	**0.10**
i_{t-2}	0.30	0.59	0.66		**0.03**	0.72	**0.06**
	0.00	0.00	0.00		**0.20**	0.00	**0.11**
π_{t+1}	0.35	0.73	0.58		0.12	0.75	0.11
	0.00	0.00	0.00		0.02	0.00	0.03
g_t	**-0.03**	0.58	**0.05**		**-0.06**	0.22	0.10
	0.72	0.00	**0.14**		**0.98**	0.00	0.04

Notes: This table summarizes the first stage fit of strictly exogenous variables to Taylor rule variables, i_{t-1}, i_{t-2}, π_{t+1} and g_t. We regress Taylor rule variables on 0-4 lags of oil shocks (OIL), detrended logarithm of potential GDP (POTGDP) and technology changes (TECH). Corresponding to each Taylor rule variable, the first row reports adjusted R^2 and the second row reports p-values associated F-statistics for the joint significance of 0-4 lags of the instrument.

Table 9: Strictly Exogenous Instruments Estimation

| | 1960Q1-1979Q2 | | 1979Q3-1996Q4 | |
	CGG	Strict Exo.	CGG	Strict Exo.
i_{t-1}	1.13	1.30	0.88	0.57
	(0.05)	(0.08)	(0.05)	(0.11)
i_{t-2}	-0.37	-0.49	-0.07	0.24
	(0.04)	(0.05)	(0.04)	(0.08)
$\mathbb{E}_t \pi_{t+1}$	0.82	0.72	2.16	2.20
	(0.05)	(0.07)	(0.20)	(0.39)
$\mathbb{E}_t g_t$	0.18	0.27	0.22	1.05
	(0.05)	(0.13)	(0.12)	(0.66)
p	0.20	0.16	0.05	0.42

Notes: This table summarizes estimation results of (7) using strictly exogenous instruments. The Taylor rule is estimated with two-step GMM, Newey-West type weight matrix with a bandwidth of 4. CGG stands for the replication of Clarida, Gali and Gertler (2000), which uses the conventional instrument set: 1-4 lags of the funds rate, inflation, output gaps, commodity inflation, interest rate spreads and M2 growth rates. Strict Exo. stands for the strictly exogenous estimation, which uses 0-4 lags of potential output, oil shocks and technology changes. p's are p-values associated with J-tests of overidentification restrictions. Newey-West standard errors are reported in parentheses.

Table 10: Strictly Exogenous Instruments Estimation: Alternative Specifications

	1960Q1-1979Q2				1979Q3-1996Q4			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
i_{t-1}	0.55		1.78	1.75	0.89		1.13	0.42
	(0.18)		(0.20)	(0.32)	(0.07)		(0.15)	(0.25)
i_{t-2}			-0.87	-0.85			-0.27	0.43
			(0.17)	(0.24)			(0.09)	(0.23)
$\mathbb{E}_t \pi_{t+1}$	0.74***	0.78***	0.73	0.71	2.07**	1.49***	2.10***	2.30**
	(0.05)	(0.04)	(0.21)	(0.19)	(0.54)	(0.14)	(0.36)	(0.64)
$\mathbb{E}_t g_t$	0.31	0.21		0.03	1.37	-0.03		1.56
	(0.12)	(0.05)		(0.35)	(1.15)	(0.25)		(0.97)
p	0.39	0.01	0.85	0.74	0.54	0.05	0.20	0.47

Notes: This table summarizes estimation results of alternative specifications of Taylor rules using strictly exogenous instruments. Taylor rules are estimated with two-step GMM and the Newey-West type weight matrix with a bandwidth of 4. Instruments are 0-7 lags of potential output. Specifications (1), (2), (3) are more parsimonious than (4), which corresponds to (7). p's are p-values associated with J-tests of overidentification restrictions. Newey-West standard errors are reported in parentheses. *, ** and *** denote the estimate is significantly different from 1 at the 10%, 5% and 1% significance level, respectively.

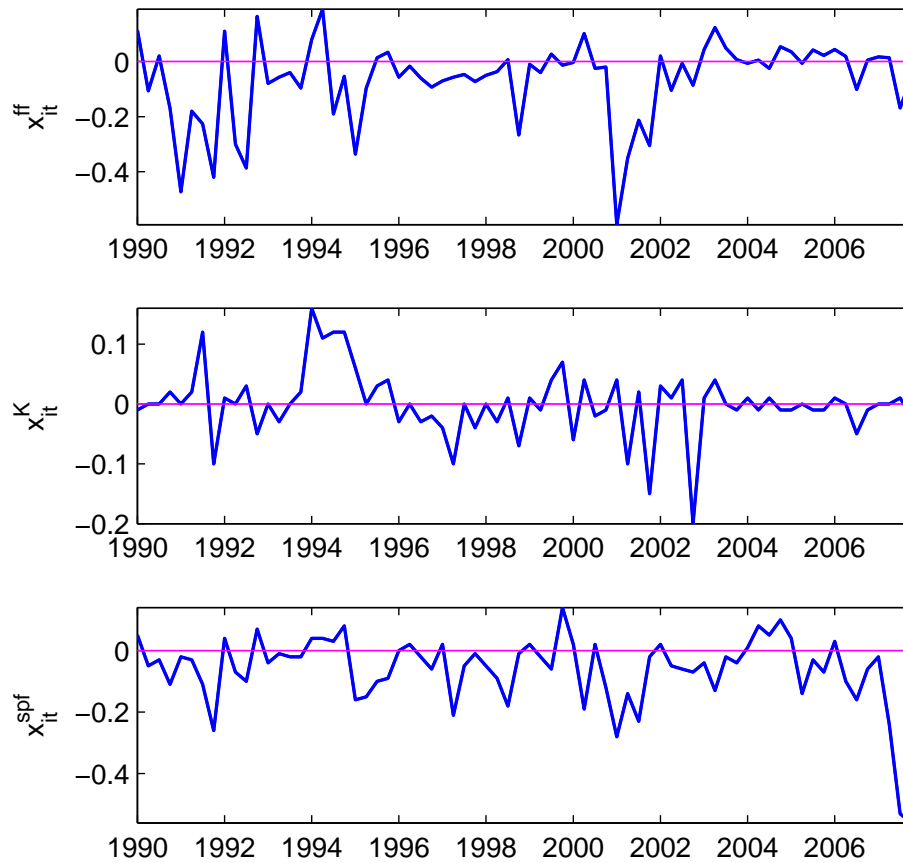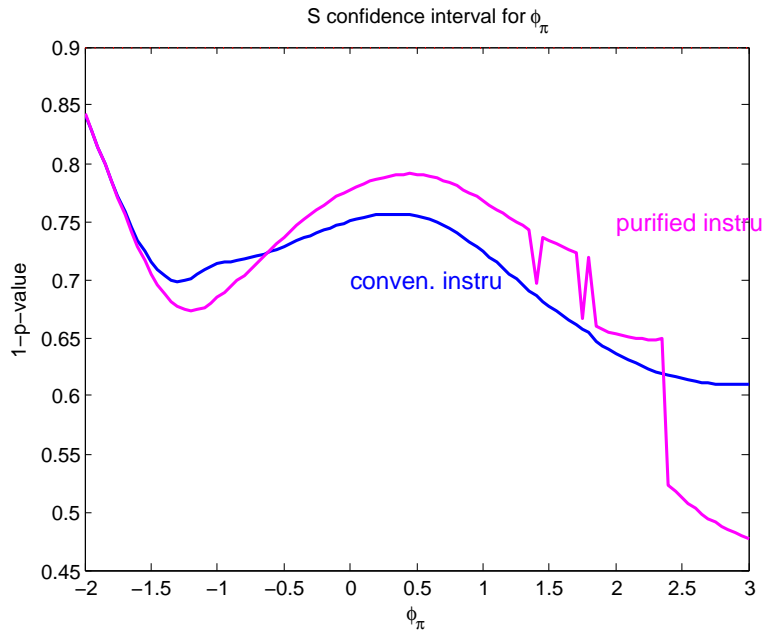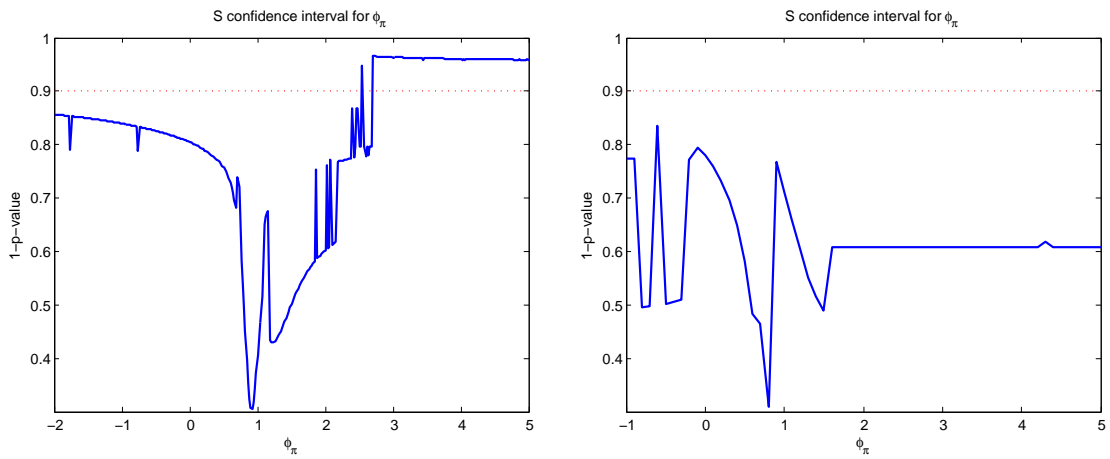Figure 1: Monetary Shocks from Federal Funds Futures and the Survey of Professional Forecasters

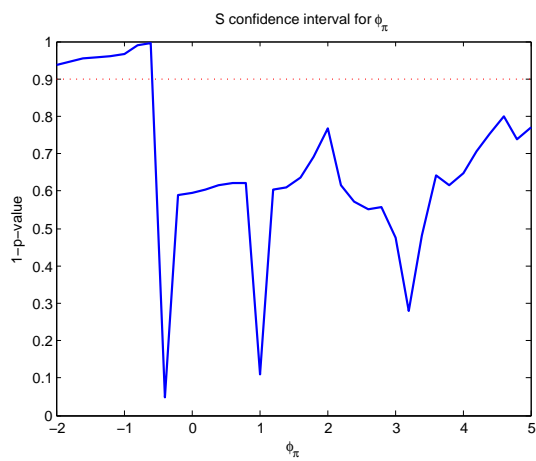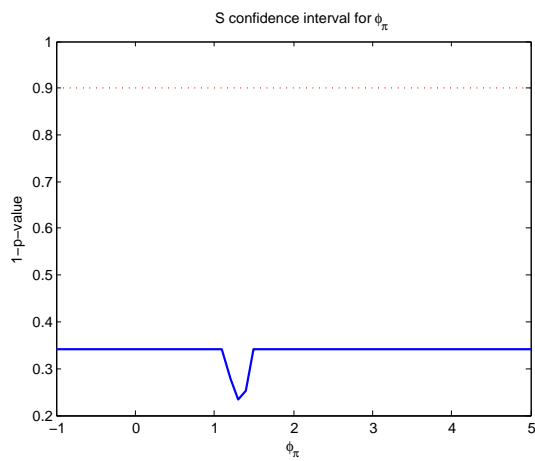Figure 2: S-Set for ϕ_π: 1990Q1-2007Q4

(a) Conventional Instruments

(b) Strict Exogeneity

Figure 3: S-Set of ϕ_π: 1960Q1-1979Q2

(a) Conventional Instruments   (b) Strict Exogeneity

Figure 4: S-Set of ϕ_π: 1979Q3-1996Q4

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