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A Conjecture of Chinese Monetary Policy Rule: Evidence from Survey Data, Markov Regime Switching and Drifting Coefficients*

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This paper studies a modified Chinese Taylor's rule with money supply growth rate as the intermediate target. We develop survey data to measure the expected inflation and use the real marginal cost as a proxy for the output gap. The Markov regime switching model, the TVP model and the split sample OLS estimation are employed to estimate the changing coefficients of the monetary reaction function. We find that there were two structural changes in the Chinese monetary policy rule, which take the form of discrete jumps rather than continuous adjustments. Besides, the Chinese central bank is not purely forward-looking as most literature assumed.

Key Words: Taylor's rule; Survey data; Real marginal cost; TVP model; Markov regime switching model.

JEL Classification Numbers: E52, C32, C50.

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1. INTRODUCTION

Since Taylor proposed the simple discretionary monetary policy rule in 1993, it has influenced the macroeconomic research greatly[1]. According to the original Taylor's rule, the stabilization monetary policy should be conducted in the way that the response to the inflation gap is larger than 1 (1.5, typically) and the response to the output gap is between 0 and 1 (0.5, typically). The debates over the Taylor's rule never stop and the extension of policy form emerges quite often.

In recent years, the forward-looking Taylor's rule has been popular in empirical research. In this line of literature, three issues stand out: (1) the measurement of output gap. Traditionally, it has been necessary to estimate the potential output in order to measure the output gap. The general approach is to decompose the output into the trend part and the cyclical part, using Kalman filter, HP filter, linear time trend method or quadric time trend method. Another approach is to use unemployment as a proxy for the output gap, based on Okun's law (Orphanides, 2001, 2002). (2) The central bank's information problem. In Taylor's (1993) seminal paper, the Fed Fund interest rate reacts to the contemporary output gap and inflation gap, whereas in reality these data are not available to central bank at the time they determine the interest rate. Using ex post data provides distorted description of historical policy (Orphanides, 2001, 2002, 2004). To avoid this problem, a widely used approach is the Generalized Method of Moments estimation (GMM), following Clarida, Gali and Gerter 1998 (CGG). A better way to deal with the information problem is to employ the real time data, such as FOMC's "Greenbook Data" (Orphanides, 2001, 2004 and Boivin, 2006). Recently, Kim and Nelson (2006) developed a modified Heckman-type two-step procedure to estimate the policy rule with ex post data. They solve the endogenous problem by decomposing the disturbance term into two components, one is correlated with contemporary information and the other is not. (3) The estimates of structural change. During different periods (typically under different presidents of the Fed), the reaction functions of the central bank are usually thought different. The test for structural change could be a traditional Chow test, a usual split sample estimation (CGG, 2000; Boivin, 2005), the time-varying parameter (TVP) model (Boivin, 2006; Kim and Nelson, 2006; Cogley and Sargent, 2001, 2005) or the Markov regime switching model (Sims and Zha, 2006; Davig and Leeper, 2006).

Most empirical research on Chinese monetary policy rule follows CGG (1998, 2000), which estimates a forward-looking Taylor's rule using GMM. Xie and Luo (2002) estimated a forward-looking Taylor's rule with quarterly data, but concluded that the response to the inflation gap is smaller

[1] For a survey, see Asso, Kahn and Leeson (2007).

than 1, implying that the monetary policy cannot stabilize the economy. Zhang and Zhang (2007) added the money supply growth rate into the reaction function and estimated the forward-looking Taylor's rule with monthly data. Among the five types of interest rates they chose, only one of them has the response to inflation gap greater than 1, and the rest are below 0.5. Another strand of literature uses the cointegration test and error correction mechanism to estimate whether or not the People's Bank of China (the Chinese central bank, PBC hereafter) follows Taylor's rule (Lu and Zhong, 2004; Bian, 2006). A main concern is that they did not solve the information problem. Moreover, this approach has an obvious weakness; the purpose of cointegration is to test the long-term relationship among variables, but the monetary reaction function may change frequently according to various macroeconomic conditions. Other scholars attempt to apply different monetary policy rules in China, including inflation target rule, LWW rule and McCallum rule (Burdekin and Siklos, 2008).

Although there are numerous papers on the Chinese monetary policy rule, few take the following questions into consideration: whether or not it is suitable to estimate Taylor's rule in China, whether or not the Chinese monetary policy rule goes through structural change and whether or not the Chinese monetary policy rule is pure forward-looking as the recent research tends to assume. To answer the questions above, this paper extends the previous research along the following dimensions:

Firstly, we estimate the Taylor's rule with money supply growth rate as monetary intermediate target instead of the interest rate. The Taylor's rule was proposed when the U.S. Fed began to use the interest rate as intermediate target of monetary policy. The Fed adjusted the interest rate to affect consumption, investment, net exports and capital market to prevent inflation and stabilize output. Yet, the interest rate is under strict control and cannot vary flexibly with macroeconomic variables in China, which is quite different from the situation in the U.S. or other countries with highly developed capital market. Actually, the Chinese policy makers almost do not care about the interest rate when they plan to adjust the economy through monetary policy. Moreover, People's Bank of China claimed explicitly in 1998 that their only intermediate target is the money supply growth rate. Due to the above reasons, it is essentially problematic to estimate the Chinese monetary reaction function using interest rate as the dependent variable. In this paper, we conjecture that the PBC uses money supply growth rate to react to the inflation gap and output gap, and the other aspects are in line with the standard Taylor' rule. We will estimate this modified Taylor's rule, which may suit the Chinese policy-making methods better.

Secondly, we compare the results from Markov regime switching model and TVP model to investigate the structural changes in the policy. The robustness of the structural change is tested by the TVP-GARCH model, TVP-Markov model and split sample estimation. The Markov regime switching model is good at capturing the unobserved discrete jump, while the TVP model can find the continuous changes of the coefficients. Previous literature usually focuses on one of the two types of changing coefficient models. In this paper, we compare the results from both of them. As argued by Sims and Zha (2006), the changing variance might be more important than the changing coefficient when modeling the policy rule. We investigate the changing conditional variance resulted from the GARCH process and Markov regime switching process to verify our estimation of structural changes. We also use the split sample estimation after we find the break dates. This gives support to the policy change and helps us decide whether TVP model or Markov regime switching model fits Chinese monetary policy better.

Thirdly, different from others who adopted pure forward-looking or backward-looking model, we put forward a hybrid model in our study. Choosing only the forward-looking rule or back-looking rule may risk mis-specifying the monetary policy, because there is possibility that it cannot fully explain the PBC's behavior. In this paper, we study the PBC's response to lagged output gap, actual inflation rate and money supply growth rate.

Fourthly, we use the survey data to measure the inflation expectation. The forward-looking policy rule needs to estimate the expected inflation, but the methods of previous literature have serious weakness or can not be applied to Chinese monetary policy. First, the real time data (like "Greenbook data") is not available in China, even though it is the easiest way to estimate a forward-looking policy rule in empirical research. Second, the GMM is too sensitive to the choice of instrumental set and there are no clear criteria on whether to add a certain variable into the instrumental set or how many lags of the instruments should be included. Third, the method developed by Kim and Nelson (2006) solves the endogenous problem from an econometric viewpoint, but it still makes use of the ex post data, which is the same with GMM in nature. It is unrealistic to assume that the policy makers apply these methods to anticipate the development of future inflation. To obtain an effective measure of inflation expectation, we employ the data from micro survey conducted by People's Bank of China, and convert the qualitative data into quantitative time series (Chen, 2008; Xiao and Chen, 2004; Fluri and Spoerndli, 1987; Carlson and Parkin, 1975). If we assume that the consumers are rational (we verify the consumers' expectation are intermediate rational in section 3.2.3) and the central bank holds no information advantage, then the micro-survey data can be used as an efficient measure of future inflation for the central bank. The main benefit of using the survey data is that we can estimate the policy reaction function by OLS or changing coefficients model directly, without resorting to the instrumental variables which are far from reliable.

Fifthly, we use the real marginal cost as a proxy for the output gap. It is widely agreed that the output gap obtained from filters are questionable. It depends heavily on the choice of sample interval, and consequently is neither stable nor consistent. Using unemployment rate as proxy is also unsuitable in China because the published data only include the urban registered unemployment rate, which does not entirely reflect the employment fluctuation in small private sectors, or the large number of floating population. We follow Gali and Gertler (1999, 2001) and use the real marginal cost as a proxy for the output gap. The advantage of this method is that the volatility of real marginal cost does not change with the sample interval, suggesting that it is a consistent and stable measure for output gap.

Applying these methods, three results seem to be robust: (1) there are two structural changes of Chinese monetary policy rule. The first one is around 1998 and the second one is around 2002 to 2003. We have clear evolution paths of the responses to expected inflation, output gap (measured by real marginal cost), lagged inflation rate and lagged money supply growth rate. The coefficients of these variables change greatly from one period to the other. (2)The coefficients of the lagged variables are statistically significant, implying the PBC is partly backward-looking. The generally used pure forward-looking type monetary policy rule cannot fully explain the Chinese situation. (3) The structural change of the policy rule is more likely to be a discrete pattern. The Markov regime switching model fits Chinese monetary policy better.

This paper proceeds as follows: Section 2 specifies the Markov regime switching model and the TVP model. Then, we discuss our method in dealing with the heteroskedasticity. The estimation procedure will also be given in this section. Section 3 first describes the data we use. Then we present our approach to measure expected inflation and provides a short proof for the rational properties of the consumers' expectation. We will also show how to use real marginal cost as a proxy to measure the output gap. A comparison between the output gap obtained from the filters and real marginal cost gives us sufficient reason to forsake the former. Section 4 displays the empirical results of baseline models as well as the hybrid models with lagged inflation and money supply growth rate. Section 5 tests the robustness of the structural changes of Chinese policy rule. Split sample estimation helps us determine whether the Markov regime switching model or the TVP model is better. Section 6 concludes.

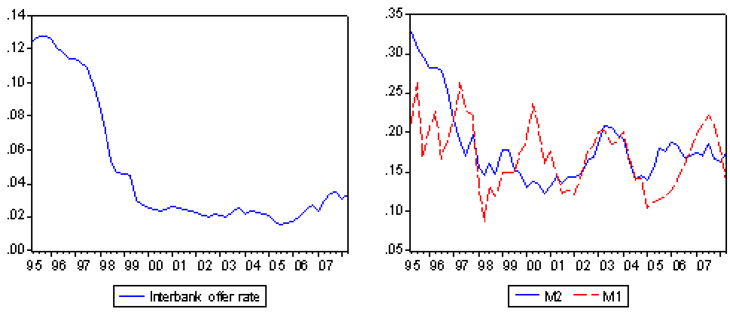
2. MODEL AND ESTIMATION PROCEDURE

2.1. Model Specification

The creation and extension of Taylor's rule were in the context that the U.S. Federal Reserve used interest rate as the intermediate target to stabilize the economy. The U.S. Fed increases or decreases the interest rate in response to the output gap and inflation gap. When we observe the movements of a U.S. interest rate (the Federal Fund's Interest Rate, for example), we can find that its time series varies relatively in sync with the changing macroeconomic conditions. At the same time, the U.S. Fed gave up controlling money supply and left it to the market.

However, the way the PBC implements its monetary policy is in sharp contrast with the U.S. Fed. First, the interest rate is under strict control. People's Bank of China sets the benchmark interest rate at the beginning of each year and then determines the deposit rates and loan rates for the commercial banks based on the benchmark interest rate. Consequently, these interest rates cannot change with the money demand and supply conditions. Even the interest rates with the highest level of marketization, such as the interbank offer rate or the repo rate, can only float around the benchmark rate. Applying Taylor's rule directly in China is problematic because the PBC does not use it as the tool to stabilize economy. Second, the PBC claimed explicitly that their monetary intermediate target is the money supply growth rate rather than interest rate. It is logical to conjecture that the monetary rule is to use money supply growth rate to react on the inflation gap and the output gap. As shown in Figure 1, the money supply growth rate, whether $M1$ or $M2$, has enough volatility, whereas the interbank offer rate is almost constant after 1998. This provides indirect evidence of the monetary policy change and motivates us to using the changing coefficient model.

FIG. 1. Interbank Offer Rate, Money Supply Growth Rate of $M1$ and $M2$



In this paper, we consider two kinds of monetary policy rule. The disparity between them is whether to include lagged variables. If the coefficients of the lagged variables are significantly different from 0, we tend to accept that the PBC is not purely forward-looking and a mixed-type policy rule is more suitable.

2.1.1. Baseline Monetary Policy Rule

The main difference between our structure of monetary policy conduct and the popular Taylor's rule is that we change the interest rate to the money supply growth rate. Usually, Taylor's rule is estimated with a forward-looking vision, but as discussed above, we want to avoid the GMM or other methods involving using ex post data as instrumental variables. We develop the measurement for expected inflation, but there is no proper proxy for expected output gap. We also have the lagged output gap enter the reaction function. Another reason is that even though the forward-looking Taylor's rule seems more appealing from a theoretical standpoint, there is no evidence that the central bank's behavior is purely forward-looking. It is difficult to believe that the central bank does not care about the recent history of inflation and output. In fact, our empirical results show that the PBC is partly backward-looking. Formally, we specify our baseline model as follows:

$$M2_t = \beta_{0,t} + \beta_{1,t} E_t \pi_{t+1} + \beta_{2,t} mc_{t-1} + \varepsilon_t \tag{1}$$

where $M2_t$ is the $M2$ money measure at time t, $E_t\pi_{t+1}$ is the expected inflation (it can be derived from the net balance statistics method or from the probability method with three different distributions, and we will give a detailed description in section 3.2), mc_{t-1} is the real marginal cost which is the proxy for output gap (we will discuss the real marginal cost in section 3.3). We allow the coefficient $\beta_{i,t}$ to change over time. In this paper, it can also follow a random walk process or a Markov switching process. In the traditional Taylor's rule, the disturbance term is usually interpreted as the monetary policy shock. Yet, we use $M2$ as the indicator of the amount of money supply, but the PBC cannot fully control this variable. It also depends on the behavior of the private sector, such as the consumers' expectation of the future economic condition, or the commercial banks' risk aversion degree. The disturbance term ε_t in equation (1) also captures the endogenous component of the money supply. Another disparity between the Taylor's rule and our monetary policy rule is the signs of the coefficients. The coefficients in Taylor's rule should be positive in order to stabilize the economy. In contrast, for the policy rule specified above, the coefficients ought to be negative except for the intercept term. This is because the increase of money supply will typically cause the rise of the inflation rate and stimulates the output.

2.1.2. Hybrid Monetary Policy Rule

As discussed in Sims and Zha (2006), there is no special reason to reject the backward-looking behavior of the monetary authority. If the monetary policy makers partly depend on the lagged variables when they determine the monetary growth rate, then the policy rule specified above will miss some variables that the central bank indeed reacts on. Even though we cannot add all the possibly relevant variables into the baseline model, we should still take those lagged variables, which are most likely to affect the policy, into consideration.

First, it is sensible that the central bank will observe the recent inflation history when they decide whether to increase or decrease the money supply growth rate. Most of the existing literature on the Taylor's rule includes the lagged inflation rate into the models. Second, for the typical Taylor's rule model, a large body of research has acknowledged the central bank's tendency to smooth the interest rate, in order to build the central bank's reputation and stabilize the capital market. To capture the smoothing behavior in empirical research, a generally used approach is to add the lagged interest rate to test whether it is statistically significant. Similarly, we expand the baseline monetary policy rule to incorporate lagged inflation and money supply growth rate, and name it the hybrid monetary policy rule:

$$M2_t = \beta_{0,t} + \beta_{1,t}E_t\pi_{t+1} + \beta_{2,t}mc_{t-1} + \beta_{3,t}\pi_{t-1} + \beta_{4,t}M2_{t-1} + \varepsilon_t \quad (2)$$

where π_{t-1} denotes the lagged actual inflation rate, $M2_{t-1}$ the lagged money supply growth rate, and the rest are the same with the baseline policy rule. Different from other variables, we expect the coefficient of the lagged money supply growth rate to be positive, for the reason that the central bank should decrease the policy volatility.

2.2. Modeling Changing Coefficients

An important task in this paper is to detect whether there is structural change of the policy rule. The regular OLS estimation or other fixed coefficient econometric methods cannot distinguish the policy changes in different periods. If we use split sample estimation to test whether there is a structural break, we have to accept the assumption that all the coefficients change at the same time, which is not necessarily the case (Boivin, 2006).

It is entirely possible that only some of the coefficients change and they change at different times. Besides, if we use the split sample estimation directly, we have to first know the break dates, but the test of break dates using traditional stability test involves uncertainty, which is unacceptable.

As a result, it is better to use the changing coefficients models to estimate the Chinese monetary policy rule considering the possibility of structural change. The recent research tends to use the Markov regime switching model or the TVP model. Both of them permit the breaks of different coefficients to happen during different periods. The difference between them is the Markov regime switching model assumes that the coefficient change is discrete, but the TVP model assumes the change is continuous. There is no prior knowledge about what the pattern of Chinese monetary policy rule transition is, even if it exists, so we compare the results of the two models.

2.2.1. Markov Regime Switching Model

A first conjecture is after the 1998's reform, the monetary policy rule may be under a different regime. For example, suppose that before 1998, the PBC focused on stabilizing the output and the response to output gap was large. After 1998, the PBC paid little attention to the output. In this case, a discrete jump of the coefficients could capture this change. The advantage of a Markov regime switching model is to detect the potential or unobserved changes, especially discrete ones, or else there would be no difference compared with the conventional dummy variable method. By applying the Markov regime switching model, we are able to know whether the reform was conducted exactly in 1998, or whether the reform was conducted, and whether there were some policy changes at some other time. Specifically, we set the model for the baseline monetary policy rule in the following way:

$$\begin{aligned} M2_t &= \beta_{0,S_t^c} + \beta_{1,S_t^c} E_t \pi_{t+1} + \beta_{2,S_t^c} mc_{t-1} + \varepsilon_t \\ \varepsilon_t &\sim N(0, \sigma_\varepsilon^2) \\ \beta_{i,S_t^c} &= \alpha_{0,\beta_i} + (\alpha_{1,\beta_i} - \alpha_{0,\beta_i}) S_t^c + \nu_{i,t} \\ \nu_{i,t} &\sim N(0, \sigma_{\nu i}^2), \quad i = 0, 1, 2 \end{aligned} \tag{3}$$

Similarly, the hybrid monetary policy rule with regime switching can be specified as:

$$\begin{aligned} M2_t &= \beta_{0,S_t^c} + \beta_{1,S_t^c} E_t \pi_{t+1} + \beta_{2,S_t^c} mc_{t-1} + \beta_{3,S_t^c} \pi_{t-1} + \beta_{4,S_t^c} M2_{t-1} \varepsilon_t \\ \varepsilon_t &\sim N(0, \sigma_\varepsilon^2) \\ \beta_{i,S_t^c} &= \alpha_{0,\beta_i} + (\alpha_{1,\beta_i} - \alpha_{0,\beta_i}) S_t^c + \nu_{i,t}, \\ \nu_{i,t} &\sim N(0, \sigma_{\nu i}^2), \quad i = 0, 1, 2, 3, 4 \end{aligned} \tag{4}$$

where equation (3) and (4) describe the evolution of coefficient β_{i,S_t^c}. The coefficient depends on the policy shock $\nu_{i,t}$ and an unobserved variable S_t^c ($S_t^c = 0\text{or}1$), which follows a first-order 2-state Markov switching process with transition probability given by:

$$p = \begin{pmatrix} p_{00} & p_{10} \\ p_{01} & p_{11} \end{pmatrix} \tag{5}$$

where $p_{i,j} = Pr[S_t^c = t|S_{t-1}^c = j]$ with $\sum_{j=0}^{1} p_{ij} = 1,\ i = 0, 1$. When $S_t^c = 0,\ E[\beta_{i,S_t^c}] = \alpha_{0,\beta_i}$, we interpret this as a small response; when $S_t^c = 1, E[\beta_{i,S_t^c}] = \alpha_1, \beta_i$, we interpret this as a large response. Each coefficient in our model has two regimes, and the number of possible regimes is 8 for the baseline monetary policy rule, 32 for the hybrid monetary policy rule. In the process of estimation, we can calculate the probability of a certain regime appearing at period t, $p_{i,t} = Pr[S_t^c = i|\varphi_t],\ i = 0, 1$, where φ is the information available up to time t (the method of calculating the probability is given in section 2.3.2). The coefficient estimate should be given as the expected value:

$$\beta_{i,t} = \alpha_{0,\beta_i} Pr[S_t^c = 0|\varphi_t] + \alpha_{1,\beta_i} Pr[S_t^c = 1|\varphi_t],\ i = 0, 1, 2 \tag{6}$$

The estimation procedure involves Kalman filter and Hamilton filter, and the details will be shown in section 2.3.2.

2.2.2. TVP Model

Another approach to estimate the changing coefficients is the TVP model. The Markov regime switching model assumes the changes of the coefficients are discrete (two states in this paper), but there exists the possibility that the policy parameters could change gradually. For instance, if the Chinese monetary reform is conducted in a step by step manner, rather than with a radical change, the coefficients incline to change little by little, as opposed to discrete jump. In this case, the TVP model is more suitable than the Markov regime switching model.

However, if the economy does jump at a certain time, the TVP model would only show a gradual adjustment, or alternatively, "produce a smooth estimate" (Boivin, 2006). In this case, the TVP model is misleading and a Markov regime switching model will function better. Even in this case, the TVP model can still provide us with some intuition about what happened and can be an approximation of the real situation. Besides, if the discrete jump happens very often, we have to specify more states and the parameters needing to be estimated with a Markov regime switching model will increase considerably. The worst situation would be if the Markov regime switching model were unable to specify these parameters when we did not add enough restriction to them. In that case, the TVP model is more practical and operable, and the more the states, the more precise approximation the TVP model will produce. We specify the baseline monetary policy rule to be estimated using the TVP model in the following way:

$$\begin{aligned} M2_t &= \beta_{0,t} + \beta_{1,t} E_t \pi_{t+1} + \beta_{2,t} mc_{t-1} + \varepsilon_t \\ \varepsilon_t &\sim N(0, \sigma_t^2) \\ \beta_{i,t} &= \beta_{i,t-1} + \nu_{i,t}, \quad \nu_{i,t} \sim N(0, \sigma_{\nu i}^2, \quad i = 0, 1, 2 \end{aligned} \tag{7}$$

At the same time, the hybrid monetary policy rule is specified in a similar way:

$$\begin{aligned} M2_t &= \beta_{0,t} + \beta_{1,t} E_t \pi_{t+1} + \beta_{2,t} mc_{t-1} + \beta_{3,t} \pi_{t-1} + \beta_4 M2_{t-1} + \varepsilon_t \\ \varepsilon_t &\sim N(0, \sigma_\varepsilon^2) \\ \beta_{i,t} &= \beta_{i,t-1} + \nu_{i,t}, \quad \nu_{i,t} \sim N(0, \sigma_{\nu i}^2), \quad i = 0, 1, 2, 3, 4 \end{aligned} \tag{8}$$

The coefficients follow a random walk process. The estimation procedure basically uses the Kalman filter, and it will be presented in section 2.3.1.

2.3. Modeling Heteroskedasticity

Sims and Zha (2006) emphasized the importance of the heteroskedasticity of the monetary policy rule, which is even more important than the changing coefficients. This aspect is also acknowledged by Kim and Nelson (2006) and Boivin (2006). Another motivation for us to consider the conditional variance is that the conditional variance contains information about the uncertainty of the regression model. The uncertainty of the policy rule, or alternatively, the conditional variance, is partly due to the changing coefficients, and another important source of the uncertainty is the heteroskedastic disturbance.

The arguments above can be applied to both Taylor's rule and our modified version. For our model, the disturbance term also includes information about endogenous money supply. Not taking the changing nature of the endogenous money supply into consideration may result in mis-specification of the monetary policy rule. Therefore, we follow the approach proposed by Kim and Nelson (1994, 1998), which allows us to decompose the conditional variance into two distinct parts[2].

[2]There are four types of models we can use to integrate the heteroskedastic disturbance term. Besides the models we introduce in section 2.3.1 and 2.3.2, we can use Markov

2.3.1. TVP Model with GARCH Disturbance Term

The heteroskedasticity of the baseline monetary policy rule may result from failing to add the lagged variables, so we only model the changing variance for the hybrid monetary policy rule. In this section, we consider the TVP-GARCH model which means that the disturbance term will follow a GARCH (1, 1) process:

$$\begin{aligned} M2_t &= \beta_{0,t} + \beta_{1,t} E_t \pi_{t+1} + \beta_{2,t} mc_{t-1} + \beta_{3,t} \pi_{t-1} + \beta_4 M2_{t-1} + \varepsilon_t \\ & \quad \varepsilon_t | \varphi_{t-1} \sim N(0, e_t^2), \quad e_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \alpha_2 e_{t-1}^2 \\ \beta_{i,t} &= \beta_{i,t-1} + \nu_{i,t}, \quad \nu_{i,t} \sim N(0, \sigma_{\nu i}^2), \quad i = 0,1,2,3,4 \end{aligned} \tag{9}$$

The standard Kalman filter cannot be used now, since the variance of the disturbance term changes over time. In order to apply an augmented Kalman filter, we rewrite the above equations in the following vector form:

$$\begin{aligned} M2_t &= [H_t \;\; 1] \begin{bmatrix} \beta_t \\ \varepsilon_t \end{bmatrix} = H_t^* \beta_t^* \\ \beta_t^* &= \begin{bmatrix} I_5 & 0 \\ 0 & 0 \end{bmatrix} \begin{bmatrix} \beta_{t-1} \\ \varepsilon_{t-1} \end{bmatrix} + \begin{bmatrix} \nu_t \\ \varepsilon_t \end{bmatrix} = F \beta_{t-1}^* + \nu_t^* \\ & \quad \nu_t \sim N(0, Q), \quad \varepsilon_t | \varphi_{t-1} \sim N(0, e_t^2) \\ Q_t^* &= E[\nu_t^* \nu_t^{*\prime} | \varepsilon_{t-1}] = \begin{bmatrix} Q & 0 \\ 0 & e_t^2 \end{bmatrix} \end{aligned} \tag{10}$$

where H_t denotes the exogenous variables. Now, we can go through the following basic procedure:

$$\begin{aligned} & \textit{Prediction}: \\ & \qquad \beta_{t|t-1}^* = F \beta_{t-1|t-1}^* \\ & \qquad P_{t|t-1}^* = F P_{t-1|t-1}^* F' + Q_t^* \\ & \qquad \eta_{t|t-1} = M2_t - H_t^* \beta_{t|t-1}^* \\ & \qquad f_{t|t-1} = H_t^* P_{t|t-1}^* H_t^{*\prime} \end{aligned} \tag{11}$$

regime switching model with GARCH disturbance term and Markov regime switching model with Markov switching disturbance term. Unfortunately, there are too many parameters in these two models that they cannot be specified. As a result, we only consider the TVP-type models with heteroskedastic disturbance term in this paper.

$$Updating:$$
$$\begin{aligned}\beta^*_{t|t} &= \beta^*_{t|t-1} + P^*_{t|t-1}H^{*\prime}_t f^{-1}_{t|t-1}\eta_{t|t-1} \\ P^*_{t|t} &= P^*_{t|t-1} - P^*_{t|t-1}H^{*\prime}_t f^{-1}_{t|t-1}H^*_t P^*_{t|t-1}\end{aligned} \tag{12}$$

where $\beta^*_{t|t-1}$ is the prediction of β^*_t based on the information up to time $t-1$, $P^*_{t|t-1}$ is the conditional covariance of β^*_t based on the information up to $t-1, \eta_{t|t-1}$ is the forecast error, $f_{t|t-1}$ is the conditional variance of the forecast error, $\beta{*}_{t|t}$ is the estimate of β^*_t based on the information up to t, and $P^*_{t|t}$ is the conditional covariance of β^*_t based on information up to t. However, the variance of the disturbance term e^2_t is not available at time $t-1$; instead, we use the expected value based on the information up to $t-1$ as a substitute:

$$E[e^2_t|\varphi_{t-1}] = \alpha_0 + \alpha_1 E[\varepsilon^2_{t-1}|\varphi_{t-1}] + \alpha_2 e^2_{t-1} \tag{13}$$

Notice that:

$$\begin{aligned}\varepsilon_{t-1} &= E[\varepsilon_t|\varphi_{t-1}] + (\varepsilon_{t-1} - E[\varepsilon_t|\varphi_{t-1}]) \\ E[\varepsilon^2_{t-1}|\varphi_{t-1}] &= E[\varepsilon_t|\varphi_{t-1}]^2 + E[\varepsilon_{t-1} - E[\varepsilon_t|\varphi_{t-1}]]^2\end{aligned} \tag{14}$$

where $E[\varepsilon_t|\varphi_{t-1}]$ is the last element of β^*_t, and $E[\varepsilon_{t-1} - E[\varepsilon_t|\varphi_{t-1}]]^2$ is the last element of $P^*_{t|t-1}$, we can calculate the expected e^2_t based on information up to $t-1$. The estimation procedure of the standard TVP model we discussed in the previous section is the same as the procedure given above except that the variance of the disturbance term is a constant.

The variance of forecast errors can be written in the following way:

$$f_{t|t-1} = H_t P_{t|t-1} H'_t + e^2_t \tag{15}$$

where $P_{t|t-1}$ is the conditional covariance of β_t. This expression shows that the uncertainty of the monetary policy can be decomposed into two components: the uncertainty arising because of the changing coefficients and the uncertainty arising because of endogenous money supply, the disturbance term.

2.3.2. TVP Model with Markov Switching Disturbance Term

The conditional heteroskedasticity may also result from a regime switching disturbance term, and it means that the uncertainty can come from the endogenous money supply with Markov regime switching, which is out of the central bank's control. The economic implication may be explained as when the agents expect there would be some changes in the policy, they adjust their behavior, which follow a Markov switching process here. To capture this effect we use the TVP-Markov model:

$$\begin{aligned}
M2_t &= \beta_{0,t} + \beta_{1,t} E_t \pi_{t+1} + \beta_{2,t} mc_{t-1} + \beta_{3,t} \pi_{t-1} + \beta_4 M2_{t-1} + \varepsilon_t \\
\beta_{i,t} &= \beta_{i,t-1} + \nu_{i,t}, \quad \nu_{i,t} \sim N(0, \sigma^2_{\nu i}), \quad i = 0,1,2,3,4 \\
& \varepsilon_t \sim N(0, \sigma^2_{\varepsilon, S^d_t}) \\
\sigma^2_{\varepsilon, S^d_t} &= \sigma^2_{\varepsilon,0} + (\sigma^2_{\varepsilon,1} - \sigma^2_{\varepsilon,0}) S^d_t
\end{aligned} \tag{16}$$

where S^d_t is a 2-state Markov switching variable ($S^d_t = 0, 1$), with transition probability:

$$p = \begin{pmatrix} p_{00} & p_{10} \\ p_{01} & p_{11} \end{pmatrix} \tag{17}$$

The definition of the probability in the parenthesis is given as before. We also rewrite the model in a vector form:

$$\begin{aligned}
M2_t &= H_t \beta_t + \varepsilon_t \\
\beta_t &= \beta_{t-1} + \nu_t, \quad \nu_t \sim N(0, Q)
\end{aligned} \tag{18}$$

The first step is to use the Kalman filter as TVP-GARCH model:

$$\begin{aligned}
& Prediction: \\
\beta^{(i,j)}_{t|t-1} &= \beta^i_{t-1|t-1} \\
P^{(i,j)}_{t|t-1} &= P^i_{t-1|t-1} + Q \\
\eta^{(i,j)}_{t|t-1} &= M2_t - H_t \beta^{(i,j)}_{t|t-1} \\
f^{(i,j)}_{t|t-1} &= H_t P^{(i,j)}_{t|t-1} H'_t + \sigma^2_{\varepsilon,j}
\end{aligned} \tag{19}$$

$$\begin{aligned}
& Updating: \\
\beta^{(i,j)}_{t|t} &= \beta^{(i,j)}_{t|t-1} + P^{(i,j)}_{t|t-1} H'_t [f^{(i,j)}_{t|t-1}]^{-1} \eta^{(i,j)}_{t|t-1} \\
P^{(i,j)}_{t|t} &= P^{(i,j)}_{t|t-1} - P^{(i,j)}_{t|t-1} H'_t [f^{(i,j)}_{t|t-1}]^{-1} H_t P^{(i,j)}_{t|t-1}
\end{aligned} \tag{20}$$

The subscript $t|t-1$ denotes the prediction of variables at time t based on the information up to $t-1$, and $t|t$ the estimate of variables at time t based on the information up to t. The superscript (i,j) $i,j = 0,1$, denotes that $S_t^d = j$, $S_{t-1}^d = i$. To make the above procedure possible, we need to make inferences about the probability of different regimes in order to estimate $\beta_{t|t}^i$ and $P_{t|t}^i$ in equation (19). It can be implemented in the following way: given that $Pr[S_{t-1}^d = i|\varphi_{t-1}]$, we are able to obtain the conditional probability of (i,j):

$$Pr[S_t^d = j, S_{t-1}^d = i|\varphi_{t-1}] = p_{j,i} Pr[S_{t-1}^d = i|\varphi_{t-1}] \tag{21}$$

Then we use the prediction errors and their variances to infer the conditional density of $M2_t$ based on the information up to $t-1$ and given $S_t^d = j$:

$$\begin{aligned} & f(M2_t|S_t^d = j, S_{t-1}^d = i, \varphi_{t-1}) \\ = \; & (2\pi|f_{t|t-1}^{(i,j)}|)^{-\frac{1}{2}} \exp(-\frac{1}{2}\eta_{t|t-1}^{(i,j)\,\prime} f_{t|t-1}^{(i,j)-1} \eta_{t|t-1}^{(i,j)}) \end{aligned} \tag{22}$$

It is also possible to calculate the conditional density of $M2_t$ based only on information up to $t-1$:

$$\begin{aligned} & f(M2_t|\varphi_{t-1}) \\ = \; & \sum_{i=0}^{1}\sum_{j=0}^{1} f(M2_t, S_t^d = j, S_{t-1}^d = i, \varphi_{t-1}) \qquad (23) \\ = \; & \sum_{i=0}^{1}\sum_{j=0}^{1} f(M2_t|S_t^d = j, S_{t-1}^d = i, \varphi_{t-1}) p_{j,i} Pr[S_{t-1}^d = i|\varphi_{t-1}] \end{aligned}$$

When we observe $M2_t$, we can update the probabilities in the following ways:

$$\begin{aligned} & Pr[S_t^d = j, S_{t-1}^d = i|\varphi_t] = Pr[S_t^d = j, S_{t-1}^d = i|\varphi_{t-1}, M2_t] \\ = \; & \frac{f(M2_t|S_t^d = j, S_{t-1}^d = i, \varphi_{t-1}) Pr[S_t^d = j, S_{t-1}^d = i|\varphi_{t-1}]}{f(M2_t|\varphi_{t-1})} \end{aligned} \tag{24}$$

$$Pr[S_t^d = j|\varphi_t] = \sum_{i=0}^{1} Pr[S_t^d = j, S_{t-1}^d = i|\varphi_t] \tag{25}$$

With the probabilities given above, we can approximate $\beta_{t|t}^j$ and $P_{t|t}^j$:

$$\beta_{t|t}^j = \frac{\sum_{i=0}^{1} Pr[S_t^d = j, S_{t-1}^d = i|\varphi_{t-1}]\beta_{t|t}^{(i,j)}}{Pr[S_t^d = j|\varphi_t]} \tag{26}$$

$$P^j_{t|t} = \frac{\sum_{i=0}^{1} Pr[S^d_t = j, S^d_{t-1} = i|\varphi_{t-1}][P^{(i,j)}_{t|t} + (\beta^j_{t|t} - \beta^{(i,j)}_{t|t})(\beta^j_{t|t} - \beta^{(i,j)}_{t|t})']}{Pr[S^d_t = j|\varphi_t]} \tag{27}$$

By now, the filter can be iterated following equations (19) and (20). The Markov regime switching model introduced in section 2.2.2 may be estimated using a similar procedure, but with a slightly different prediction approach in the Kalman filter. Specifically, the filter for the Markov regime switching model can be given in the following way:

$$\begin{array}{rcl} Prediction: & & \\ \beta^{(i,j)}_{t|t-1} & = & \mu^i_{t-1|t-1} \\ P^{(i,j)}_{t|t-1} & = & P^i_{t-1|t-1} + Q \\ \eta^{(i,j)}_{t|t-1} & = & M2_t - H_t\beta^{(i,j)}_{t|t-1} \\ f^{(i,j)}_{t|t-1} & = & H_tP^{(i,j)}_{t|t-1}H'_t + \sigma^2_\varepsilon \end{array} \tag{28}$$

$$\begin{array}{rcl} Updating: & & \\ \beta^{(i,j)}_{t|t} & = & \beta^{(i,j)}_{t|t-1} + P^{(i,j)}_{t|t-1}H'_t[f^{(i,j)}_{t|t-1}]^{-1}\eta^{(i,j)}_{t|t-1} \\ P^{(i,j)}_{t|t} & = & P^{(i,j)}_{t|t-1} - P^{(i,j)}_{t|t-1}H'_t[f^{(i,j)}_{t|t-1}]^{-1}H_tP^{(i,j)}_{t|t-1} \end{array} \tag{29}$$

We focus on the conditional variance of the forecast errors. Similar to the TVP-GARCH model, the conditional variance can be broken down into two parts:

$$f_{t|t-1} = H_tP_{t|t-1}H'_t + \sigma^2_{\varepsilon,S_t} \tag{30}$$

where the two parts of the variances are their expected value obtained from the Kalman filter using the probabilities calculated in equation (24) and equation (25). The first part is the uncertainty due to changing policy rule, and the second part is due to the Markov regime switching disturbance.

3. MEASUREMENT OF INFLATION EXPECTATION AND OUTPUT GAP

3.1. Data Description

The data we employ are quarterly data over the period 1995:II to 2008:II. The money supply growth rate is the average of three consecutive monthly growth rates of $M2$ in one quarter. $M2$ measure includes money in circulation, corporate demand, time deposit, rural deposit, individual time deposit, deposit of organizations and army, and deposit for infrastructure construction. The actual inflation rate is usually measured by the GDP deflator, but it is not published in China. Instead, we use the average of three consecutive monthly CPI in one quarter to measure actual inflation rate. The expected inflation rate is obtained by converting micro-survey data into time series data. We give a formal discussion in section 3.2. The real GDP is obtained by dividing the nominal GDP by the inflation rate. To give a comparison, we use the traditional HP filter and the quadric time trend method to calculate the output gap. An important attempt of this paper is to use real marginal cost as a proxy for output gap. The introduction to this method is given in section 3.3. All the data we use are from "China Monthly Economic Indicators", "People's Bank of China Quarterly Statistical Bulletin" and the website of the National Bureau of Statistics of China.

3.2. Inflation Expectation Measurement: A Micro Survey Approach

To estimate a forward-looking-type monetary policy rule, it is necessary to calculate the expectation term. However, the central bank's information problem makes this task difficult to deal with. The first concern lies in the fact that at the time the central bank determines the policy rule, they do not have access to the inflation or output data of the current period. When we use the ex post data to estimate the policy rule in history, it may cause serious distortion. The second concern is that we actually do not have the knowledge about how the policy makers form the inflation expectation. The widely used GMM analysis is too sensitive to the instrumental set. At the same time, neither economics nor mathematics provides a standard which can decide whether or not to add a new variable to the instrumental set or how many lags of the instruments should be included. The modified Heckman-type two-step procedure is smart from an econometric standpoint; however, it suffers the same problem as the GMM in the way that it also depends on the ex post data. Another famous method to calculate the expectation is to build a VAR system, but it is still difficult to encompass all the relevant variables. Besides, the central bank probably does not form their expectations using these methods.

In our study, we develop a measure of expected inflation. Our approach is to transform the qualitative data from the Consumers Saving Survey System (CSSS) conducted by the People's Bank of China into quantitative expected inflation data. The first advantage of our approach is that the survey data can be obtained at the time when the policy makers determine the money supply growth rate. Therefore, it is real time data like the data from the Greenbook. We can use the traditional econometric method or the time dependent coefficient model in this paper directly, without resorting to the GMM or modified Heckman two step procedure. Another benefit is that it is most likely that the PBC forms their inflation estimation based on the survey data. Using the survey data is apparently much more plausible than the pure "econometrics oriented guess", which involves using lagged data or prediction error decomposition from a Kalman filter. Xiao and Chen (2004) test the properties of the inflation expectation from CSSS from 1995 to 2004 both in the short term and the long term and conclude it is a good predictor. Chen (2008) used this expected inflation to estimate a New Keynesian Phillips Curve successfully. Their results prove that this approach is reliable.

Before introducing our method of calculating the expected inflation, it is necessary to discuss its suitableness for our purpose. The primary question we need to address is whether or not the consumers' expected inflation is consistent with that of the policy makers. On the one hand, the central bank may hold some information advantage over consumers, and they may have a better estimation of inflation; on the other hand, the policy maker may make use of economic agents' expectation to increase or decrease the money supply growth rate surprisingly, which is the famous time inconsistent policy. Here, we need to maintain three assumptions. Firstly, the consumers are rational; secondly, the central bank does not have an information advantage; thirdly, given the central bank's social loss function, the cost of making a surprising policy will exceed the benefit. The first assumption can be proved by testing the properties of the expected inflation, as discussed in section 4.1.3. The second and the third assumptions are difficult to test, particularly when we permit the response to the inflation gap and the output gap to change over time. However, to test whether the last two assumptions hold is beyond the scope of our current research. In this paper, we will mainly focus on the effectiveness of the expected inflation from micro-survey. Even though it might be attacked from a theoretical standpoint, it still amounts to a good measurement for the inflation forecast of the central bank if it is close to the actual inflation rate.

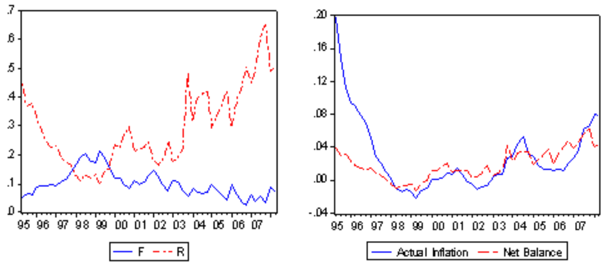
3.2.1. Qualitative Description

The CSSS designs three kinds of answers for the consumers to choose about the direction of inflation rate to change in the next period: Price Up, Price Down and No Change. Let R_t denote the percentage of population who choose "Price Up", F_t denote the percentage of the population who choose "Price Down", and then the net balance B_t is defined as $B_t = R_t - F_t$. The net balance itself is not the expected inflation series, but the magnitude of B_t can be viewed as a direct reflection of the extent to which the consumers expect the inflation to rise. The range of net balance is $[-1, 1]$. If all of the people expect the inflation to increase, then the net balance is 1; vice versa, the net balance is -1. If the net balance is greater than 0, it means more people expect inflation to rise; if less than 0, more people expect inflation to fall.

Figure 2 shows that the net balance evolves closely with the actual inflation, and this evolution fits the macroeconomic condition in China quite well. From 1992 to 1996, China went through high inflation and was in a prosperous period, and consequently more people expected the inflation to rise. Later, affected by the Asia financial crisis, the economic growth rate declined gradually and the number of people who expected inflation to fall exceeded that of people who thought it to rise during this period. Since 2000, the economy grew steadily and maintained mild inflation. As a result, more people expected the inflation to rise. However, as the domestic demand and the production cost rose remarkably in the recent two years, the inflation as well as the expected inflation rate also has a notable increase. More importantly, it is shown that the expected inflation leads the actual inflation, which is quite clear during 2000 and 2006, the times when China's overall economic condition changed from recession to boom, or boom to recession. Using ex post data to estimate the future usually fails when there is a transition of the economy, but the expectation made by the consumers has made use of all the available information and consequently the micro-survey data should be more accurate than the traditional methods. The expected inflation indeed has the power to forecast future inflation, and it is much better than using the historical data to estimate when the economy is in transition.

Although the net balance provides us with a baseline result of inflation expectation and some intuition of the future inflation, it is not the precise measure of the expected inflation. The original purpose to introduce the micro-survey data is to estimate the monetary policy rule, so it is essential to convert the qualitative data-net balance into certain quantitative mea-

FIG. 2. Percentage of Different Forecasts, Net Balance and Actual Inflation Rate



surement of expected inflation. In section 3.2.2, we present two types of methods to calculate the expected inflation.

3.2.2. Qualitative Expected Inflation Series

A. Net Balance Statistics

We first follow the method proposed by Fluri and Spoerndli (1987) to calculate the expected inflation as a linear transformation of net balance, using the following formula:

$$E_t(\pi_{t+1}) = k(R_t - F_t) \tag{31}$$

where $E_t(\pi_{t+1})$ denotes the expectation of inflation in period $t+1$ based on the information up to period t. The coefficient k can be calculated using the sample data as follows:

$$k = \frac{\sum_{t=1}^{T} \pi_t}{\sum_{t=1}^{T}(R_t - F_t)} \tag{32}$$

where π_t is the actual inflation in period t.

The net balance statistics method is easy to implement, but it overlooks the percentage of the population who choose “No Change”, which also contains the information about the expected inflation. For example, suppose the net balance is 30%. One extreme case is that the percentage of the people who choose “Price Up” is 65%, “Price Down” is 35%, and “No Change” is 0%. Another extreme case is the percentage of those who choose “Price Up” is 30%, “Price Down” is 0%, “No Change” is 70%. It is obvious that the inflation expectations in the two cases should be different even though the net balance is the same.

B. Probability Approach

Theil (1952) first proposed the probability method to analyze the survey data in early literature, then Carlson and Parkin (1975) and Taylor (1988) added several assumptions and improved the algorithm in accordance with the type of the survey. Xiao and Chen (2004) develop the method suitable for CSSS. Here, we follow the approach of Xiao and Chen (2004) and Carlson and Parkin (1975).

In order to transform the qualitative into expected inflation series, we need to make several assumptions. We first assume that people questioned follow a certain distribution, and this distribution will determine how people choose their answers. The next assumption is there exists a sensitive interval with a 0 central, $(-a_t, a_t)$. Those who expect the inflation rate to lie in this interval will choose "No Change".

Suppose that the expected inflation for period $t+1$ of population questioned is a random variable x^e_{t+1} with density function $f_{t+1}(x)$. We can write R_t, F_t, B_t as:

$$\begin{aligned} R_t &= P(x^e_{t+1} > a_t) \\ F_t &= P(x^e_{t+1} \leq -a_t) \\ B_t &= P(-a_t < x^e_{t+1} \leq a_t) \end{aligned} \tag{33}$$

If x^e_{t+1} follow the normal distribution, then the equation above could be rewritten in the following way:

$$\begin{aligned} P\left(\frac{x^e_{t+1} - E_t(\pi_{t+1})}{\sigma} > \frac{a_t - E_t(\pi_{t+1})}{\sigma}\right) &= R_t \\ P\left(\frac{x^e_{t+1} - E_t(\pi_{t+1})}{\sigma} \leq \frac{-a_t - E_t(\pi_{t+1})}{\sigma}\right) &= F_t \end{aligned} \tag{34}$$

where, σ is the standard error of the random variable x^e_{t+1}, $E_t(\pi_{t+1})$ denotes the expected value of x^e_{t+1}. Let $z_1(t) = \Phi^{-1}(F_t)$, $z_2(t) = \Phi^{-1}(1 - R_t)$, where $\Phi(\Box)$ is the cumulative distribution function. The expected inflation and its standard error can be derived as:

$$E_t(\pi_{t+1}) = \frac{z_1(t) + z_2(t)}{z_1(t) - z_2(t)} a_t, \quad std_t(\pi_{t+1}) = \left| \frac{2a_t}{z_1(t) - z_2(t)} \right| \tag{35}$$

Besides the normal distribution, we can also assume the random variable follows other distribution forms. Here, we consider uniform distribution and Logistic distribution, which have been widely used.

If x^e_{t+1} follows the uniform distribution instead, then $z_1(t) = (F_t - 0.5)\sqrt{12}$, $z_2(t) = (0.5 - R_t)\sqrt{12}$. The expected inflation and its standard error can be given in a similar way as:

$$E_t(\pi_{t+1})) = \frac{R_t - F_t}{1 - R_t - F_t} a_t, \quad std_t(\pi_{t+1}) = \left| \frac{a_t}{\sqrt{3}(1 - R_t - F_t)} \right| \tag{36}$$

If x^e_{t+1} follows the Logistic distribution, then $z_1(t) = -\log(1/F_t - 1)\sqrt{3}/\pi$, $z_2(t) = \log(1/R_t - 1)\sqrt{3}\pi$. The expected inflation and its standard error are:

$$E_t(\pi_{t+1}) = \frac{\log(1/F_t - 1) - \log(1/R_t - 1)}{\log(1/F_t - 1) + \log(1/R_t - 1)} a_t \tag{37}$$

$$std_t(\pi_{t+1}) = \left| \frac{2a_t}{\log(1/F_t - 1) + \log(1/R_t - 1)} \right| \frac{\pi}{\sqrt{3}} \tag{38}$$

Now, only the sensitive interval $(-a_t, a_t)$ remains undetermined. According to our assumption, the mean of the actual inflation rate should equal to that of the expected inflation, and we also assume that the interval does not change in the sample period. Under these assumptions, we can estimate the interval a:

$$\frac{\sum_{t=1}^{T} \pi_t}{T} = \frac{\sum_{t=1}^{T} [z_1(t) + z_2(t)]/[z_1(t) - z_2(t)]}{T} a \tag{39}$$

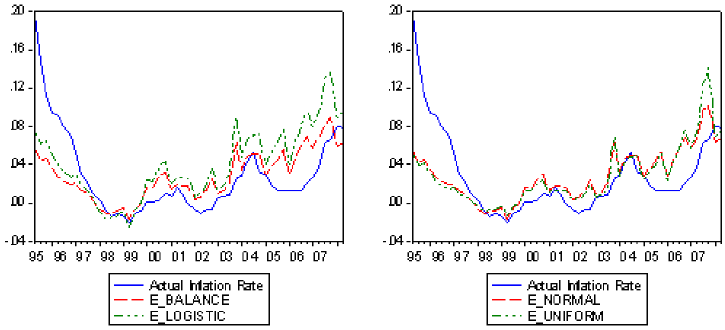
$$a = \frac{\sum_{t=1}^{T} \pi_t}{\sum_{t=1}^{T} [z_1(t) + z_2(t)]/[z_1(t) - z_2(t)]}$$

Applying the method above, we are able to convert the qualitative data into quantitative data. The comparisons between the four kinds of expected inflation with different distribution assumptions and the actual inflation rate are shown in Figure 3. From the figures we can see that the four types of expected inflation generally fit the actual inflation rate well. However, before 1999, the expected inflation tended to underestimate the inflation, partly because China's economy had not gone through serious inflation before and the actual inflation was somewhat beyond people's expectations. After 2000, people inclined to overestimate the inflation rate. This can be attributed to the fact that the economic growth rate in China maintained a high level and people tended to overestimate inflation based on the observation of output.

3.2.3. A Rational Expectation Test

As mentioned at the beginning of this section, it is important to test whether the expected inflation obtained from the CSSS is rational. If the inflation expectation is rational, the consumers will make use of all the available information and will not make systematic errors. Mathematically,

FIG. 3. Comparison between Inflation Expectations and Actual Inflation Rate



the rational expectation properties require that the predictor is unbiased, efficient, and there is no significant autocorrelation (Forsells and Kenny, 2002). Because the four types of inflation expectations are similar, we report the test results of only one of them.

We first test whether the expectation has bias. The existence of bias implies that the consumers systematically overestimate or underestimate the inflation rate on average. To carry out the bias test, we can test the following simple equation:

$$e_t = \pi_t - E_{t-1}(\pi_t) \tag{40}$$

where $E_{t-1}(\pi_t)$ denotes the expected inflation at time t formed at time $t-1$. If the null-hypothesis that the mean of e_t is 0 cannot be rejected, then we can conclude that the expected inflation is unbiased in the long run. The t-statistics and the p-value are 0.6923 and 0.4918, suggesting that the null-hypothesis cannot be rejected.

A more formal test for bias can be implemented in the following way:

$$\pi_t = \beta_0 + \beta_1 E_{t-1}(\pi_t) + \varepsilon_t \tag{41}$$

If the joint null-hypothesis ($\beta_0 = 0, \beta_1 = 1$) cannot be rejected, we conclude that the inflation expectation is unbiased. The Wald test reports the F-statistics is 0.0838 with p-value 0.9198, implying that the expected inflation is unbiased.

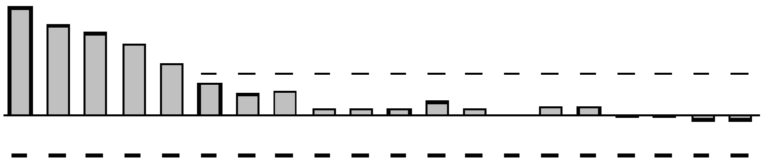
Secondly, we test whether the expectation inflation is autocorrelated. The Q-statistics of different orders are all significant even at 1% level. Consequently, we cannot reject that the expected inflation is autocrrelated. However, the consumers have the ability to adjust their expectations to weed out the systematic errors, reflected in the decreasing autocorrelation. It is reasonable to tolerate the autocorrelation because it takes some time for the consumers to distinguish whether the forecast error is transient or lasting, given that the macroeconomic conditions are quite complicated, and different kinds of information may conflict with one another. If consumers realize that the forecast error is persistent, then they will correct it in the future. Table 1 displays the value of autocorrelation up to 20 orders and Figure 4 shows a clearly decreasing autocorrelation.

TABLE 1.

Forecast Error: Autocorrelation

Lag	1	2	3	4	5	6	7	8	9	10
AC	0.7302	0.6099	0.5621	0.4922	0.3437	0.2075	0.1469	0.1490	0.0365	0.0371
Lag	11	12	13	14	15	16	17	18	19	20
AC	0.0306	0.0858	0.0345	−0.0053	0.0428	0.0518	−0.0089	−0.0104	−0.0313	−0.0360

FIG. 4. Forecast Error: Correlogram



Thirdly, we test whether the expected inflation is efficient in the sense that consumers can take advantage of all the information available. If the inflation expectation is efficient, the forecast error should be independent from other relevant macroeconomic variables. Alternatively, the forecast error should be orthogonal with respect to the information set including current and past macroeconomic variables, which is called strong-form efficiency. We estimate the following equation:

$$\pi_t - E_{t-1}(\pi_t) = \beta_0 + \beta_1 \Omega_{t-1} + \varepsilon_t \tag{42}$$

In our empirical analysis, the information sets encompass the nominal GDP, benchmark interest rate, interbank offer rate, foreign direct investment, net exports, and $M2$. The estimation results show that the coefficients of all the variables in the information set are not statistically significant except the net exports. This result illustrates that the expected inflation obtained from CSSS is generally efficient.

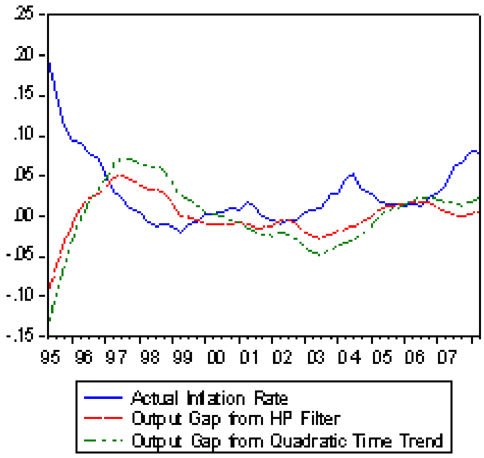
The null-hypothesis tests show that the inflation expectation from CSSS has no bias. Even though we cannot reject that the expected inflation is autocorrelated, the degree of autocorrelation is clearly decreasing, suggesting the consumers adjust their forecast error over time. The efficiency test verifies that the expected inflation is efficient in general. These results illustrate that the inflation expectation is intermediate rational, rather than fully rational.

3.3. Real Marginal Cost as the Proxy for Output Gap

It is widely agreed that the conventional measures of output gap involve considerable measurement error and are very fragile. The general approach is to find the potential output using a fitted deterministic trend, such as HP filter, Kalman filter and so on. This type of approaches heavily depends on the choice of sample interval. Specifically, as the starting point and the end point changes, the output gap for a fixed period can change dramatically. Another approach to look for the potential output is estimating the capital stock and employment rate and calculating the output according to a certain assumption of the production function. Unfortunately, it is extremely difficult to estimate the capital stock and employment rate in China due to lack of data and the imprecise measure of unemployment rate.

In this paper, we use the real marginal cost as the proxy for output gap, as discussed in the literature on New Keynesian Phillips Curve (Gali and Gertler, 1999, 2001; Walsh, 2003; Chen, 2008). From a theoretical point of view, when the labor market is efficient (without the wage make-up), the output gap is in proportion to the real marginal cost. Gali and Gertler (1999, 2001) pointed out that using real marginal cost instead of output gap generates sound econometric results, and they attribute this success partly to the existence of imperfect labor market and to the fact that the evolution of real marginal cost is sluggish compared to the output gap. Whether the labor market friction plays an important role in disconnecting the output gap and real marginal cost is open to challenge, but for our purpose of investigating the monetary policy rule, the real marginal cost will be a good proxy. In the first place, the output gap estimated from the filters or other detrending methods generates serious measurement error. Using these measurements will obviously produce distorted results. As shown in Figure 5A, the output series generated by the HP filter and the quadratic time trend method are very similar, but in sharp contrast to our basic intuition. First, from 1995 to 1997 and 2001 to 2003, China's economy was in prosperity, but the output gap was below 0. Second, from 1997 to 2000, the economy was in recession, but the output gap shows a boom in this period. Third, the Phillips curve tells us that the output gap should move with the inflation gap and it is indeed the fact in many countries. But, as shown in Figure 5A, the output gap and inflation often run in the opposite direction. In the second place, the output gap reflected by the real marginal cost does not depend on the sample interval. At different intervals, the output gap measured by real marginal cost only varies in the quantitative magnitude, but not the relative trend. The most significant advantage of using the real marginal cost is that this measurement is consistent and stable.

FIG. 5A. Comparison between Output Gap and Actual Inflation Rate



For simplicity, we assume a Cobb-Douglas technology. The output Y_t is given by:

$$Y_t = A_t K_t^{1-\alpha} N_t^{\alpha} \tag{43}$$

where A_t denote technology, K_t capital and N_t labor. Real marginal cost is defined as the real cost for producing one unit of extra product, mathematically:

$$MC_t = \frac{\partial C_t}{\partial Y_t} = \frac{W_t}{P_t} / \frac{\partial Y_t}{\partial N_t} = \alpha \frac{W_t N_t}{P_t Y_t} = \alpha S_t \tag{44}$$

where $W_t N_t$ is the nominal total wage, $P_t Y_t$ is the nominal GDP, and $S_t = W_t N_t / P_t Y_t$ is the labor income share. Taking logarithm and dividing the above equation by steady state value yields:

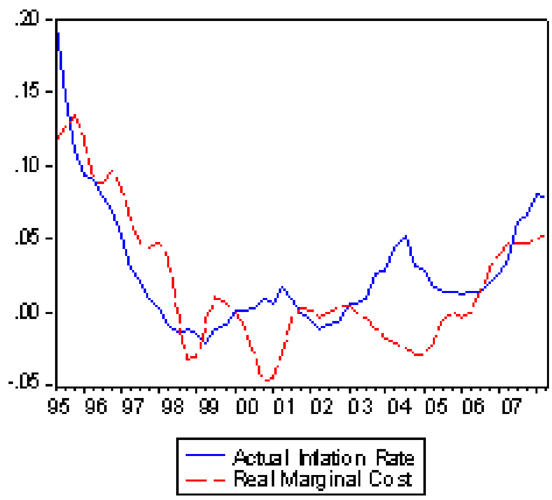
$$mc_t = s_t \tag{45}$$

where mc_t and s_t are the percentage deviations of real marginal cost and labor income share from their steady states respectively. Under the assumption of an efficient labor market, we can derive the result that the real marginal cost is in line with the output gap (see Walsh (2003) for details):

$$mc_t = \kappa(y_t - y_t^f) \tag{46}$$

where y_t denotes the real output and y_t^f potential output. Gali and Gertler (1999, 2001) use the non-farm business sector labor income to calculate the labor income share, but the non-farm business sector GDP is not published in China. Instead, we assume that the ratio of urban labor income to the nominal national income remains constant during the sample period, which is proportional to the ratio of non-farm business sector labor income to the non-farm business sector GDP. In this way, we can use the latter to obtain the percentage change from its steady state. It can be illustrated from Figure 5B that the real marginal cost (refer to the percentage change) co-moves with the inflation and accords with the economic reality. The real marginal cost and the inflation diversified during the period of 2000 and 2005, when the inflation rate began to rise and the economy was ready to boom. However, the real marginal cost generally reflects the evolution of output gap and is much better than the fitted deterministic trend method. We will use it as the proxy for output gap when we estimate the monetary reaction function.

FIG. 5B. Comparison between Real Marginal Cost and Actual Inflation Rate



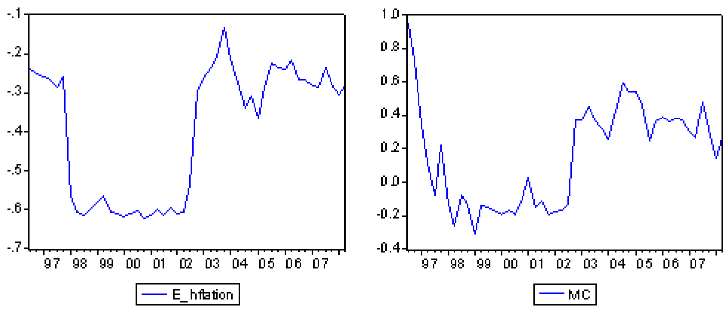
4. EMPIRICAL RESULTS

The estimation procedure requires the initial value of the parameters, so we use the data of the first several quarters to make an inference. The starting point of the sample may vary from model to model, but the estimation results are consistent. Because the four types of expected inflation evolve in a very similar manner, we only provide the estimation results using the inflation expectation based on the assumption of Logistic distribution.

4.1. Baseline Monetary Policy Rule Results[3]

The first column in Table A1 displays the parameter estimates of the Markov regime switching model for baseline monetary policy rule specified in equation (3). The first column in Table A2 presents the parameter estimates of the TVP model specified in equation (7). Figure 6A and Figure 6B depict the responses to the expected inflation and the output gap in the Markov regime switching model and the TVP model respectively.

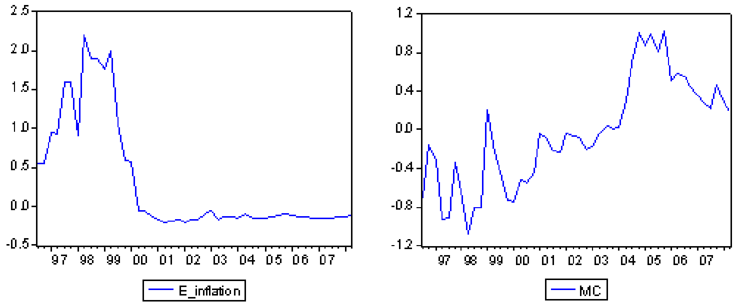
FIG. 6A. Markov Baseline Model: Coefficients Estimates



We begin with the response to the expected inflation. In the case of the Markov regime switching model, the coefficient is always negative, which is consistent with our expectation. The magnitude was larger from 1998 to 2002, about −0.6, than that before 1998, about −0.3. We could interpret this result in the way that the PBC paid more attention to preventing high inflation from 1998 to 2002. In the case of the TVP model, the coefficient of expected inflation is above 0 before 2000, conflicting with our expectation and suggesting an unscientific policy. Before 2002, the results of the two kinds of models are quite different. After 2002, the responses to

[3]The forecast errors of most models show no significant serial correlation, revealing no evidence of model misspecification. The Q-statistics are reported in Table A1 and A2.

FIG. 6B. TVP Baseline Model: Coefficients Estimates



the expected inflation are similar to each other, about -0.2, except that the evolution path in the TVP model is more stable.

We now turn to the response to the real marginal cost. The results of the Markov regime switching model and the TVP model share almost the same trend. The response to the real marginal cost is around -0.2 from 1998 to 2002, but is above 0 before 1998 and after 2002, which means that the response to output gap was not always stabilizing the economy. The disparity between the two models is that the TVP model provides a smoother estimate than the Markov regime switching model.

Finally, we consider the structural changes in the monetary policy rule. The Markov regime switching model reveals three states in the sample period: before 1998, 1998 to 2002 and after 2002. They are clearly shown from the evolution of the responses to expected inflation and real marginal cost. Figure 7A depicts the probability of regime 0, which also provides strong support for this finding. Our initial guess is that there should be some structural change of the monetary policy around 1998 due to the monetary and financial market reform, and the estimation results have ratified our conjecture. In addition, the Markov regime switching model detects another potential structural change around 2002, which is unexpected.

Yet, the structural changes are less obvious in the TVP model. During the first transition time 1998, the degree of response to the expected inflation increases but not as dramatically as the Markov regime switching model. The volatility of the response to the real marginal cost around 1998 is much higher than other periods, which might provide evidence for this transition. The second transition time of the response to the expected inflation is about 2000, two years sooner than the Markov switching regime model. The second transition time of the response to the real marginal cost is around 2002, during which the coefficient rises above 0, the same with the

FIG. 7A. Markov Baseline Model: Probability of Regime 0

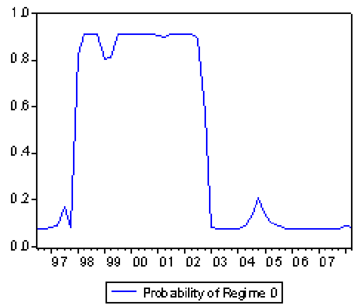
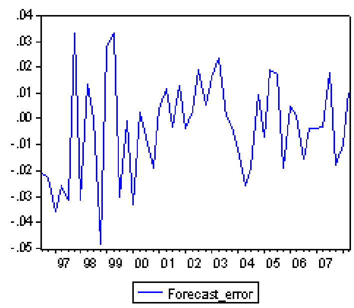


FIG. 7B. TVP Baseline Model: Forecast Errors



Markov regime switching model. When we apply the TVP model, there is no probability of a certain state, but the volatility of forecast errors shown in Figure 7B is obviously higher around 1998 than in any other period, suggesting that a structural change may have happened at this time.

4.2. Hybrid Monetary Policy Rule Results[4]

The baseline results provide us with some intuitions about how the monetary policy rule changes its reaction to expected inflation and real marginal cost during different macroeconomic conditions. In this section we report the estimation results of the hybrid models with the lagged inflation rate and money supply growth rate.

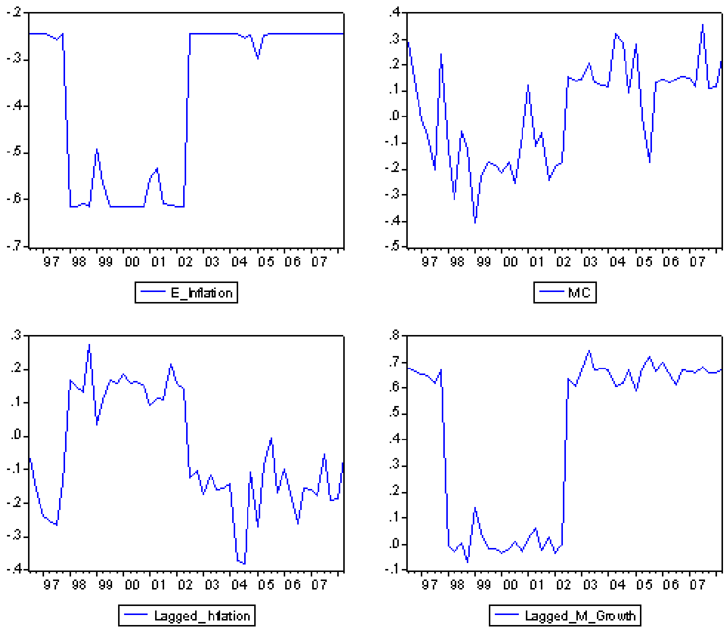
The forth column in Table A1 displays the parameter estimates of the Markov regime switching model for the hybrid monetary policy rule specified in equation (4). The forth column in Table A2 reports the parameter estimates of the TVP model specified in equation (8). Figure 8A and Figure 8B present the evolution of the responses to various macroeconomic variables in the Markov regime switching model and the TVP model respectively, where a sharp transition path can be traced.

The responses to the expected inflation and the real marginal cost show almost the same trend with the baseline results. For the Markov regime switching model, the coefficient of the expected inflation first changes from -0.25 to -0.6 at the end of 1998, implying an increasing trend of the importance attached to preventing inflation. Around 2002, the value of the coefficient changes from -0.6 to about -0.25. The response to the output

[4]Besides the hybrid monetary policy rule with both lagged inflation rate and lagged money supply growth rate, we also estimate the baseline monetary policy rule plus each one of them. The estimation result is similar to the hybrid monetary policy rule, and we report the results in the Appendix.

gap is also similar to the baseline result, but the coefficient of the hybrid policy rule is a littler larger than the baseline policy rule after 2002. For the TVP model, the responses to both the expected inflation and the real marginal cost share the same trend with the baseline monetary policy rule, except that the coefficients are generally slightly larger in the hybrid policy rule. Therefore, we can conclude that the results of the baseline monetary rule are robust when adding new independent variables.

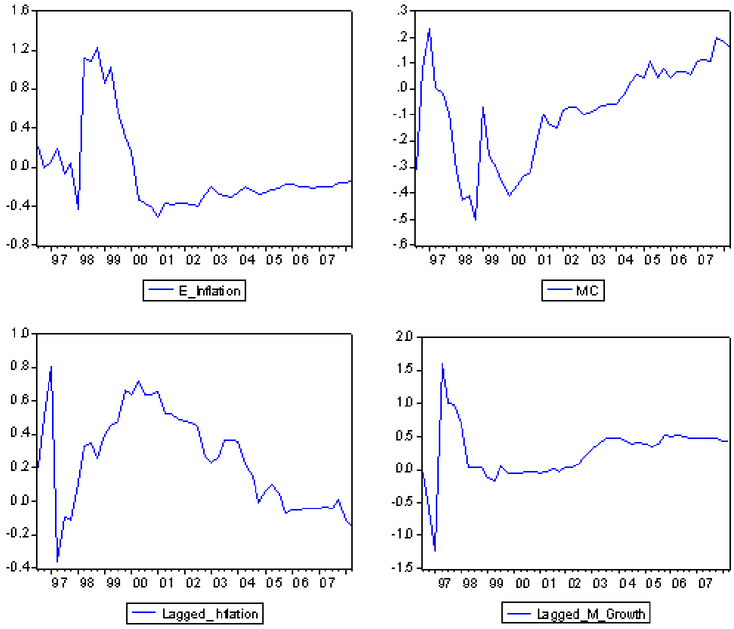
FIG. 8A. Markov Hybrid Model: Coefficients Estimates



According to the new added variables, the response to the lagged inflation rate in the Markov regime switching model is negative before 1998 but remains positive between 1998 and 2002 at around 0.15. It then turns negative after 2002, fluctuating around −1.5. The response to the lagged inflation in the TVP model evolves similarly to the Markov regime switching model, but generally larger than the former. Moreover, the TVP model provides a smoother estimate of the coefficient.

The responses to the lagged money supply growth rate are nearly the same between the two models, with clear transition points in 1998 and 2002. The coefficient is around 0.6 before 1998. It then declines to about 0 from 1998 to 2002. After 2002, it returns to 0.6. The coefficient of the lagged money supply growth rate is always above 0, which fits our

FIG. 8B. TVP Hybrid Model: Coefficients estimates



expectation that the PBC does not want to make a big surprise of the policy. These results also support our hypothesis that the behavior of the PBC is partly backward-looking.

The probability of regime 0 estimated from the Markov regime switching model is depicted in Figure 9A. It is consistent with the structural changes of the coefficients, suggesting two transitions happen in 1998 and 2002. The forecast errors of the TVP model shown in Figure 9B fluctuate heavily around 1998, which may be an evidence of structural change.

The empirical results of the Markov regime switching model and the TVP model for both baseline and hybrid monetary policy rules show that there are two structural changes, at 1998 and 2002 respectively. This finding is supported by the evolution of the coefficients of the four variables. The probability of regime 0 in the Markov regime switching model and the volatility of forecast errors in the TVP model lend further credentials. A comparison between the two kinds of changing coefficients models show that the TVP model gives smoother estimates, while the Markov regime switching model tends to generate discrete estimates. Even though in most cases the results obtained from the two approaches resemble each other, the estimations of response to expected inflation before 2002 are in a consider-

FIG. 9A. Markov Hybrid Model: Probability of Regime 0

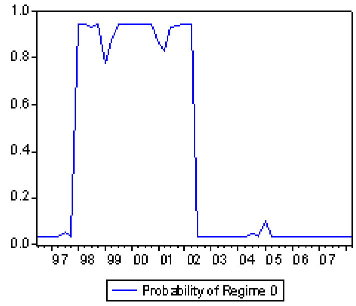
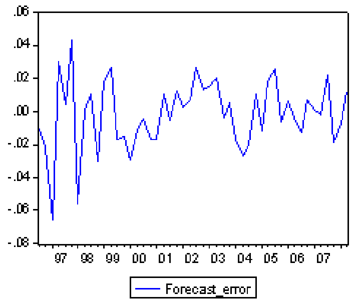


FIG. 9B. TVP Hybrid Model: Forecast Errors



ably different pattern. In the next section, we will determine whether the Markov regime switching model or the TVP model provides more accurate estimation.

5. ROBUSTNESS OF THE STRUCTURAL CHANGE

Section 5 displays the estimation results of the Markov regime switching model and the TVP model. These results have shown the evolution paths of the responses to variables under different macroeconomic conditions, which are of great interest and reflect how the Chinese monetary policy was conducted during the past ten years. It is clear that the policy rule is far from stable and the coefficients of these variables change considerably during several possible transition times, some of which were anticipated, and some of which were not expected but captured by our model. The probability of a certain regime and our rough observation of the coefficients have already provided strong support for the structural changes of the policy rule, and in this section we will provide more support to this finding. We finish this task by two approaches: the first is to investigate the conditional variance of the forecast errors. The larger the variance, the higher the probability of the existence of a transition will be. The second is to estimate the split sample. The break points we adopt are the periods that coefficients change dramatically in the changing coefficient models.

5.1. Heteroskedasticity

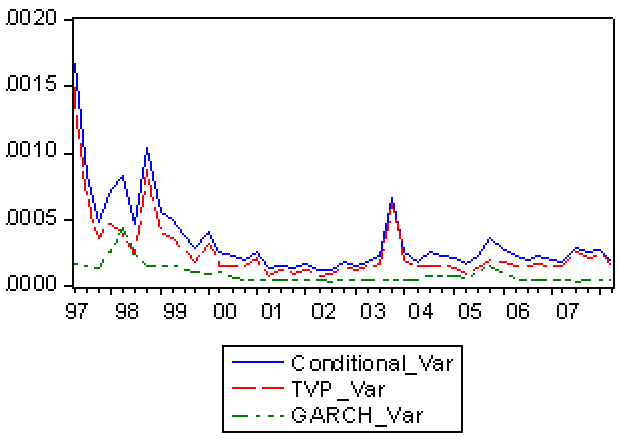
Turn to the TVP-GARCH model first, which is specified in equation (9). We focus on the conditional variance now. The variance of forecast errors can be modified in the following way:

$$f_{t|t-1} = H_t P_{t|t-1} H_t' + e_t^2$$

as we discussed in section 2.3.1. The uncertainty of the monetary policy is decomposed into two components: the uncertainty arising due to the changing coefficients and the uncertainty arising due to the endogenous money supply.

The parameter estimates are reported in the fifth column in Table A2. The general trends of these coefficients are in line with the standard TVP model. Figure 10A shows the conditional variance and its decomposition. The first finding is that the conditional variance is irregularly high around 1998 and 2003, which confirms our previous assumption that there are structural changes during the two periods. The second finding is that the variance is primarily from the changing coefficients rather than conditional heteroskedasticity. The intuition underlying the second finding is that the switches of the monetary policy target are the main factor to explain the policy uncertainty, and the endogenous money supply is not important relative to the monetary policy change. It also suggests that the PBC controls the money supply well and the market demand and supply cannot affect the money supply efficiently.

FIG. 10A. TVP-GARCH Model: Conditional Variance

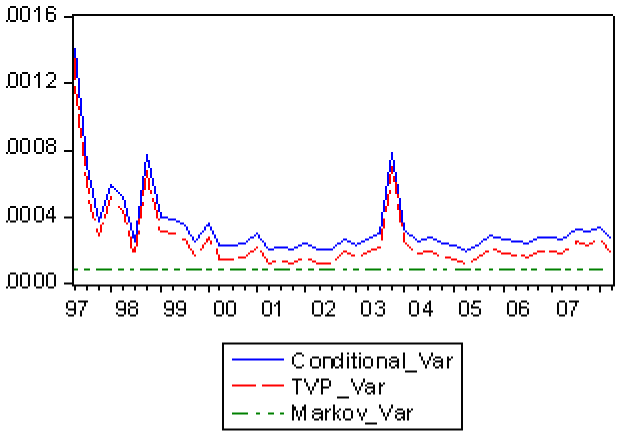


We consider the TVP-Markov model, as in equation (16). The conditional variance could be broken down into two parts, which we discussed in section 2.3.2:

$$f_{t|t-1} = H_t P_{t|t-1} H_t' + \sigma^2_{\varepsilon,S_t}$$

The parameter estimates are reported in the sixth column in Table A2. As shown in Figure 10B, the monetary policy uncertainty is generally higher around 1998 and 2003 than in any other periods, which is not surprising. The variance due to the disturbance term with endogenous regime switching is constant and of a small value, suggesting a low possibility of the existence of such disturbance. The variance of the changing coefficient amounts to a large part of the total variance, the same as the results from the TVP-GARCH model. Again, the PBC's policy change is a main factor for the overall uncertainty of the monetary policy rule.

FIG. 10B. TVP-Markov Model: Conditional Variance



The results from the changing variance models illustrate that around 1998 and 2002 to 2003, the conditional variance is irregularly high, suggesting a high probability of structural change. Also, the decomposition of the conditional variance shows that the main factor causing uncertainty is the changing coefficients of monetary policy rule.

5.2. Split Sample Estimates

An interesting phenomenon is that if we estimate the monetary policy rule by OLS over the full sample, the coefficient estimates are not statistically significant. It may be confusing and misleading in the sense that we tend to conclude that the money supply growth rate has nothing to do with those macroeconomic variables. However, when we already know how the policy rule was conducted in the past using the TVP model and the Markov regime switching model, the OLS results are easy to understand. Because the policy rule went through several transitions and the responses to some variables typically changed from negative to positive, or positive to negative, the OLS estimates calculate the average of the responses and the results will be biased toward 0, resulting in statistical insignificance. The advantage of the TVP model or the Markov regime switching model is that they can discern the differences of coefficients in different periods, and this makes it possible for us to use OLS to estimate the sub sample to confirm whether there are structural changes.

Another virtue of employing OLS is that it can help us to judge whether the type of structural change is discrete, as suggested by the Markov regime switching model, or is continuous, as suggested by the TVP model. If we estimate a sub sample and the result is similar to the Markov regime switching model, we tend to accept that the TVP model is mis-specified.

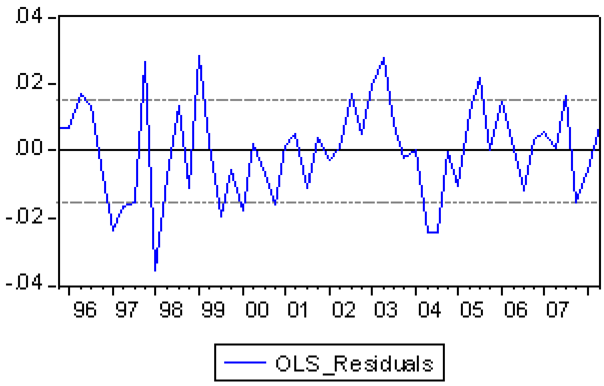
TABLE 2.

Parameter Estimates of the Split Sample OLS

Parameter	1995:III-1997:IV	1998:I-2002:II	2002:IV-2008:II
Expected Inflation	1.8059	−0.6878	−0.3373
	(0.1232)	(0.0000)	(0.0277)
Real Marginal Cost	0.2838	−0.1816	0.3961
	(0.7113)	(0.0026)	(0.0302)
Lagged Inflation	0.6448	0.0004	−0.4620
	(0.3080)	(0.9986)	(0.0138)
Lagged $M2$	0.8193	0.0470	0.6501
	(0.0561)	(0.7572)	(0.0001)

Table 2 presents the OLS estimates of the coefficients in three periods: 1995:III-1997:IV, 1998:I-2002:II, 2002:IV-2008:II. The split sample estimation results fit the changing coefficients model quite well, proving that our previous results are robust. The values of these coefficients change from one period to another period, showing the structural change of the monetary policy rule. The coefficients are generally not statistically significant in the first sub sample. We conclude that during this period, there was no clear monetary policy rule using money supply growth rate as a target. For the second sub sample, the responses to expected inflation and real marginal cost were significant even at the 1% level and with the correct sign, but the responses to lagged inflation rate and money supply growth rate were not clear. In this period, the PBC followed the baseline monetary policy rule quite well. After 2002, the responses are strong and sensible to all the variables, but the sign of real marginal cost is positive. We provide two possible explanations to this unsettling result. First, the real marginal cost may mismatch the output gap. As shown in Figure 5B, the real marginal cost did not move closely with the inflation rate after 2000, and this fact might result in the wrong sigh of the coefficient. Second, the nominal wages rise quickly after 2002, and a scientific policy rule should decrease the money supply growth rate to stabilize the output. However, during this period, there is strong pressure for Renminbi appreciation. Under a fixed exchange rate regime, the PBC may be unable to fully control the money supply growth and have to increase the money supply to maintain the fixed exchange rate. It is possible that the conflicting targets the PBC confronted made the policy maker neglect to react to the output gap correctly. The full sample OLS residual in Figure 11 presents higher volatility during 1998 and 2002, revealing that the policy went through structural changes.

FIG. 11. Full Sample OLS Residuals



In section 4, we have already observed that the coefficients of the lagged variables are different from 0 in most periods. The OLS estimates also show that the responses to most of the lagged variables are significant after 2002, suggesting that the PBC is indeed partly backward-looking. The purely forward-looking Taylor rule cannot fully explain the Chinese monetary policy rule.

If we compare the OLS result with the results of the Markov regime switching model, we will find they match surprisingly well. However, the

TVP model estimates somewhat mismatch the OLS result, especially the coefficient of expected inflation from 1998 to 2002. This fact, plus the obvious sudden increase of variance in all the models we have estimated, suggests that the structural change are more likely to be discrete jumps rather than gradual transitions, and the Markov regime switching model may fit the Chinese situation better than the TVP model.

6. CONCLUSION

In this paper, we have proposed a modified Taylor's rule in agreement with the Chinese macroeconomic conditions. We introduce our new approach to measure expected inflation and output gap. The survey data can forecast inflation quite well and the real marginal cost as a proxy for the output gap fits the reality much better than the traditional output gap from filters. With these data, we employ the Markov regime switching model and the TVP model to trace the responses to different variables and detect structural change. The TVP-GARCH and the TVP-Markov models are also used for robustness tests. The split sample estimation gives clear-cut results about how the monetary policy was conducted and the fitness of the two kinds of changing coefficient models. We summarize our findings as follows:

(1) There are two structural changes of the Chinese monetary policy rule from 1995 to 2008. The first is around 1998 and the second is around 2002 to 2003. These results are supported by all the methods we have used, both by observing the coefficients and the conditional variances of the forecast errors.

(2) There is no clear policy rule before 1998. The coefficients of all the variables are not statistically significant except the lagged money supply growth rate. From 1998 to 2002, the responses to the expected inflation and the real marginal cost are negative and significant. The lagged inflation rate and money supply growth rate are not significant. From 2003 to 2008, all the variables are significant and with the correct signs except that the response to real marginal cost is positive. The responses to the lagged variables are statistically significant after 2002, showing that the behavior of the PBC is not purely forward-looking.

(3) The split sample estimates results show that the monetary policy changes are more likely to be discrete. The TVP model smoothes the jump, more or less mis-specified. It may be better to use the Markov regime switching model to estimate the Chinese monetary policy rule.

APPENDIX

TABLE A. 1.

Parameter Estimates for Markov Regime Switching Models

Parameters	Baseline	Lagged π	Lagged M_2	Complete
σ_τ	0.0110	−0.0124	0.0000	0.0000
	(0.0172)	(0.0000)	(0.0048)	(0.0057)
$\sigma_{\nu 0}$	0.0008	0.0052	0.0000	0.0000
	(0.2415)	(0.0000)	(0.0048)	(0.0057)
$\sigma_{\nu 1}$	0.0948	0.0000	0.0834	0.0000
	(0.0882)	(0.0867)	(0.0604)	(0.0000)
$\sigma_{\nu 2}$	0.3042	0.1758	−0.2863	0.2243
	(0.0761)	(0.0979)	(0.0920)	(0.0903)
$\sigma_{\nu 3}$		0.0000	0.0322	−0.1845
		(0.0886)	(0.0118)	(0.1546)
$\sigma_{\nu 4}$				0.0366
				(0.0101)
α_{0,β_0}	0.1506	0.1586	0.0727	0.0735
	(0.0035)	(0.0037)	(0.0163)	(0.0161)
α_{0,β_1}	−0.6097	−1.3340	−0.2562	−0.2452
	(0.1830)	(0.3217)	(0.0742)	(0.1073)
α_{0,β_2}	−0.1659	−0.0468	0.1244	0.1355
	(0.1598)	(0.1371)	(0.1030)	(0.0998)
α_{0,β_3}		1.4002	0.6758	−0.1685
		(0.2673)	(0.0997)	(0.1152)
α_{0,β_4}				0.6615
				(0.0957)
α_{1,β_0}	0.1907	0.1895	0.1591	0.1537
	(0.0061)	(0.0079)	(0.0272)	(0.0065)
α_{1,β_1}	−0.2603	−0.1673	−0.5858	−0.9353
	(0.0960)	(0.1885)	(0.1667)	(0.1720)
α_{1,β_2}	0.3861	0.3771	−0.1739	−0.1778
	(0.1038)	(0.1416)	(0.1074)	(0.0968)
α_{1,β_3}		−0.4045	−0.0647	0.1575
		(0.1655)	(0.1764)	(0.2746)
α_{1,β_4}				−0.0165
				(0.0463)

TABLE A. 1—*Continued*

Parameters	Baseline	Lagged π	Lagged M_2	Complete
p_{00}	0.9653	1.0000	0.9430	0.9429
	(0.0342)	(0.0000)	(0.0555)	(0.0556)
p_{11}	0.9436	0.9617	0.9658	0.9658
	(0.0549)	(0.0376)	(0.0336)	(0.0336)
$Q(10)$	19.6462	23.8070	11.6077	10.3389
	(0.0328)	(0.0081)	(0.3122)	(0.4113)
Log Likelihood	−125.3427	−132.7340	−144.5327	−146.6070

Note: standard errors in parenthesis, except the Q-statistics with p-value in parenthesis. The sample interval varies from model to model.

FIG. A1. Baseline Markov Model with Lagged Inflation: Coefficients Estimates

FIG. A2. Baseline TVP Model with Lagged Inflation: Coefficients Estimates

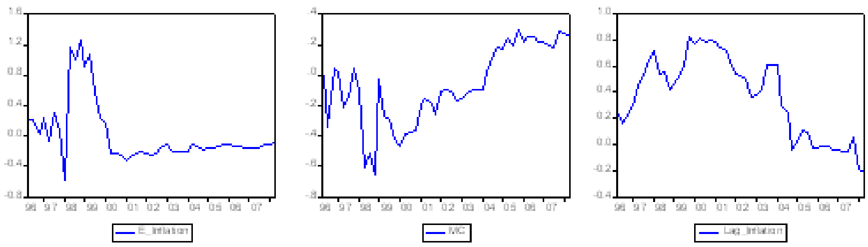


TABLE A. 2.

Parameter Estimates for TVP Models

Parameters	Baseline	Lagged π	Lagged M_2	Complete	GARCH	Markov
σ_ε	0.0000	0.0000	−0.0094	−0.0086		
	(0.0063)	(0.0000)	(0.0033)	(0.0038)		
$\sigma_{\nu 0}$	−0.0127	0.0141	0.0000	0.0000	0.0000	0.0090
	(0.0031)	(0.0025)	(0.0070)	(0.0092)	(0.0084)	(0.0021)
$\sigma_{\nu 1}$	0.0000	0.0000	0.0000	0.0000	−0.0496	0.0000
	(0.0447)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
$\sigma_{\nu 2}$	0.2352	−0.1054	−0.0533	0.0664	0.0543	0.0000
	(0.1312)	(0.1103)	(0.0682)	(0.0776)	(0.0000)	(0.0155)
$\sigma_{\nu 3}$		0.0901	−0.0509	0.0468	0.0411	0.0641
		(0.0594)	(0.0236)	(0.0534)	(0.0000)	(0.0627)
$\sigma_{\nu 4}$				−0.0548	0.0477	0.0401
				(0.0287)	(0.0188)	(0.0000)
α_0					0.0000	
					(0.0000)	
α_1					0.9915	
					(0.0245)	
α_2					0.0000	
					(0.0000)	
$\sigma_{\varepsilon,0}$					0.0534	
						(0.0170)
$\sigma_{\varepsilon,1}$						2.9829
						(0.0000)
p_{00}						0.9178
						(0.0000)
p_{11}						1.0000
						(0.0000)
$Q(10)$	14.6099	14.3092	11.9139	4.1517	13.5890	10.6141
	(0.1469)	(0.1593)	(0.2909)	(0.9402)	(0.1926)	(0.3884)
Log Likelihood	126.3851	127.8461	127.8168	123.9677	120.1803	−118.3207

Note: standard errors in parenthesis, except the Q-statistics with p-value in parenthesis. The sample interval varies from model to model.

FIG. A3. Baseline Markov Model with Lagged Money Supply Growth Rate: Coefficients Estimates

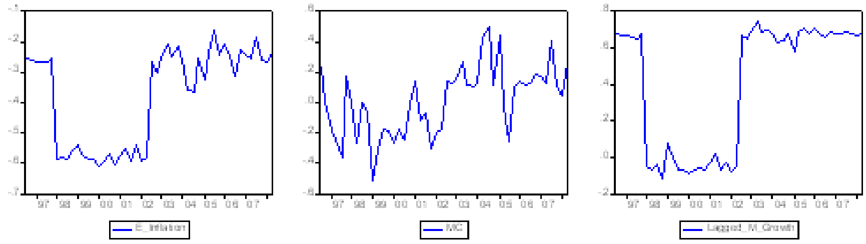
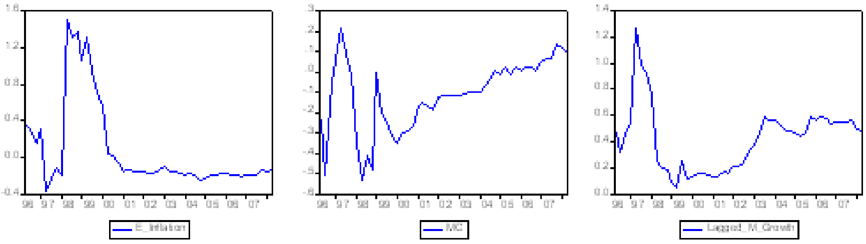


FIG. A4. Baseline TVP Model with Lagged Money Supply Growth Rate: Coefficients Estimates



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