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Review

Estimating forward-looking rules for China's Monetary Policy: A regime-switching perspective

Tingguo ZHENG *, Xia WANG, Huiming GUO

Wang Yanan Institute for Studies in Economics, Xiamen University, Xiamen, Fujian, 361005, China

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A B S T R A C T

This paper introduces a regime-switching forward-looking Taylor rule to describe the monetary policy behavior and considers its estimation using a two-step MLE procedure due to Kim and Nelson (2006), Kim (2009) and Zheng and Wang (2010). By doing an empirical analysis on quarterly data for China over the period 1992–2010, our results show that the actual reactions of China's monetary policy can be well characterized by a two-regime forward-looking Taylor rule. Furthermore, it is also suggested that the interest rate policy in response to inflation and output gap is asymmetric, behaving a significant characteristic of regime-switching nonlinearity. Specifically, in the first regime the People's Bank of China targets inflation, but not focuses on the output gap; while in the second regime the central bank targets the output gap and the policy rule is not a stable framework.

1. Introduction

Since 1990s, interest rate, by replacing the money supply, has gradually become the intermediate target of monetary policy in the western countries, and it has played a more and more important role in the conduct of monetary policy. Taylor rule, firstly proposed by Taylor (1993), characterizes central banks' behavior through a linear function of interest rate to inflation gap (the deviation of inflation rate from its target) and output gap (the deviation of real output from its potential value). Now it is an important reference for central banks when implementing monetary policy to stabilize price level and smooth output fluctuations. In fact, Taylor rule not only performs well in empirical study, but also has fundamental economic theory. Given a quadratic central bank loss function and linear aggregate demand and supply curves in the dynamic structure of economy, we can obtain Taylor rule by minimizing the loss function.

After Taylor (1993), a stream of empirical literature studied and tested Taylor rule and its extensions. Clarida, Galí, and Gertler (1998) used the forward-looking reaction function to test Taylor rule. They reported estimates of monetary policy functions for two sets of countries: the G3 (Germany, Japan, and U.S.) and the E3 (UK, France, and Italy), and the results lend support to the view that some form of inflation targeting may be superior to fixing exchange rates, and then took it as a mean to gain a nominal anchor for monetary policy. McCallum (2000) adopted historical analysis to test Taylor rule using the economic data of U.S. and U.K. during the period from 1962 to 1999, and Japan from 1972 to 1998. He suggested that rules' messages are more dependent upon which instrument rather than which target variable is used. Clarida, Galí, and Gertler (2000) estimated a forward-looking U.S. monetary policy reaction function. They found that there are substantial differences in the estimated rule across periods before

* Corresponding author at: A202, Economics Building, Xiamen University, Xiamen, 361005, Fujian, China. Tel.: +86 592 2180695; fax: +86 592 2187708.
E-mail addresses: zhengtg@gmail.com (T. ZHENG), wxia820@163.com (X. WANG), guohuiming1203@163.com (H. GUO).

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and after 1979, and used these differences to show the price stability since 1980s. Judd and Rudebusch (1998) and Nelson (2000) combined the historical analysis with the reaction function method. Based on the analysis of monetary historical data, they successfully estimated the central bank's reaction function of U.S. during the period from 1970 to 1997 and U.K. during the period from 1992 to 1997, respectively. Ball (1999) established the policy rule under the open economy, and added the exchange rate in Taylor rule to decide the interest rate. In his paper, interest rate or monetary condition index are chosen by central bank as policy tool. Giannoni and Woodford (2002) introduced sticky price and/or wages into optimal monetary policy rules, and studied the robustness of these new rules. Semmler and Zhang (2007) and Taylor (2007) further applied Taylor rules into the analysis of asset price and estate price.

However, some recent research argued that the reaction of the interest rate to inflation and output gap could be nonlinear rather than linear. These studies can be classified into two categories. The first is about econometric modeling and testing for nonlinearity. For example, Kim and Nelson (2006) and Boivin (2006) adopted time-varying parameter models to examine the U.S. data during the latest 50 years, and found that the reaction of federal fund rate to inflation is obviously unstable, and what's more, the U.S. monetary policy began to step into a more active state since 1980s. Rabanal (2004) estimated a two-state Taylor rule to explain the behavior of the Federal Reserve over business cycle phases, and found that during expansions the Fed follows a rule that can be characterized as inflation targeting, whereas the Fed targets output growth during recessions. Bruggemann and Riedel (2008) used logistic smooth transition regression (LSTR) models with time varying parameters to estimate the U.K. monetary policy reaction function, and found that in times of recessions the Bank of England put more weight on the output gap and less on inflation, and a reverse pattern is observed in non-recession periods. The second is about theoretical analysis and explanation in terms of three different ways, e.g. the aggregate supply curve being nonlinear, the central bank's preferences being asymmetric around the targeted rate, and the central banker facing uncertainty about the model describing the economy. For example, Dolado, Maria-Dolores, and Naveria (2005) relaxed the linear assumption of Phillips curve, and assumed that inflation is a concave function of output gap, which means a nonlinear Phillips curve. Nobay and Peel (2003), Ruge-Murcia (2003), Surico (2007a, b)[1] set nonlinear welfare loss functions of central bank, which can represent central bank's asymmetric preference to positive or negative deviations of inflation and output gap from its target. Tillmann (2010) discussed the optimal monetary policy under parameter uncertainty, and showed that if the central bank is uncertain about the slope of Phillips curve and applied monetary policy according to a min–max strategy, then the response of interest rate to inflation becomes stronger when inflation is away from its target. In addition, Davig and Leeper (2007) generalized the Taylor principle to an environment in which reaction coefficients in the monetary policy rule evolve according to a Markov process.

From the practical situation of China's economic development, there are still a lot of controversies about whether the interest rate or the money supply is more appropriate as an intermediate goal for monetary policy. If Taylor rule may track China's monetary policy well, then taking the interest rate as an intermediate target will offer the central bank a wider selection of policy targets. So far, the Chinese scholars have done a lot of explorations and discussions. Xia and Liao (2001), after having examined the conduct of China's monetary policy in the past, have concluded that the quantity control of money supply is not suited any more for the Chinese economy. Xie and Luo (2002) firstly employed the historical analysis and reaction function method to conduct an empirical analysis of China's monetary policy in the framework of Taylor rule and draw the conclusion that this rule can accurately measure the operation level of China's monetary policy. Zhao and Gao (2004) constructed a more suitable interest rate rule in China by allowing the exchange rate to influence the long-run inflation target. Bian (2006) used both general method of moments (GMM) and cointegration test method to examine the applicability of Taylor rule in China. Considering standard Taylor rule and the optimal monetary policy raised by Clarida, Gali, and Gertler (2002) in a comprehensive way, Wang and Zou (2006) proposed an extensive Taylor rule in the opening-up economic condition and made empirical test on China's monetary policy. In short, these studies have largely promoted and enriched the understanding of monetary policy rules for China. However, the estimated results of policy reaction functions showed unstable behavior of China's monetary policy, which is not consistent with the traditional Taylor rule. A more likely explanation is that these studies specified the policy reaction functions improperly.

Recently, a few studies attempt to employ nonlinear methods to investigate China's monetary policy reaction function. Zhang and Zhang (2008) applied a threshold regression model and found that all reaction coefficients in the regime of high money growth are larger than those in the regime of low money growth. Ouyang and Wang (2009), based on the situation of asymmetric reaction of monetary policy to inflation and real GDP, constructed a nonlinear monetary policy reaction function using economic growth rate and inflation rate as threshold variables. The results showed that the reaction of monetary policy to inflation and GDP gap is nonlinear and asymmetric. Zheng and Liu (2010) developed a regime switching Taylor rule with time-varying inflation target and showed that China's monetary policy has significant regime switching feature, that is, different regimes reflect different monetary policy reactions. However, possible two drawbacks of their study are worthy of consideration: the first is that the estimated model is not based on the forward-looking Taylor rule, so it might be inconsistent with the People's Bank of China (PBC)'s actions; the second is that they do not deal with the endogeneity problem in the regime switching model which may lead to inconsistent parameter estimates for Taylor rule.

In this study, we consider a regime-switching forward-looking Taylor rule that nests the simple rule with regime switching in Zheng and Liu (2010) as a special case. For this regime switching model, the traditional generalized method of moments (GMM) is

[1] Surico (2007b) showed that the first six years of ECB monetary policy can be characterized by a nonlinear, state-dependent policy rule. The empirical analysis on synthetic euro-area data suggests that the objective of price stability is symmetric, whereas the objectives of real activity and interest-rate stabilizations are not. Output contractions imply larger policy responses than output expansions of the same size, while no asymmetric reaction for inflation.

no longer suitable for estimating this type of nonlinear reaction function, since the forward-looking rule has a dynamic structure with endogenous regressors. Therefore, we realize its appropriate estimation of this nonlinear rule via the two-step MLE procedure due to Kim and Nelson (2006), Kim (2009) and Zheng and Wang (2010) to deal with the potential endogeneity problem.

According to formal econometric methods, we would select two-regime Markov-switching forward-looking Taylor rule to investigate China's monetary policy reaction function from 1992 to 2010. As far as we know, the forward-looking rule is more reasonable than the simple rule due to at least the following three reasons. Firstly, it's helpful to increase the degree of transparency and accountability and thus maintain stable economic growth (Svensson, 1999) and it's also useful to stabilize public expectation and improve social welfare (Liu, 2004). Secondly, the forward-looking rule is more consistent with the PBC's actions. The fact that the central government determines the coming year's monetary policy at the end of each year illuminates the important role of expected component in the PBC's monetary policy decision. Thirdly, the state council executive meeting on 21st Oct. 2009 formally proposed “management of inflation expectations”, which confirms that the PBC targeted on expected inflation rather than past or actual inflation.

The remainder of the paper is organized as follows. Section 2 briefly reviews traditional and forward-looking Taylor rules. Section 3 introduces a regime-switching forward-looking Taylor rule and presents its estimation procedure. Section 4 describes the data selection and processing. In Section 5, we implement formal econometric testing and model selection, and report the empirical results for China's monetary policy. Finally, the last section draws some conclusions.

2. Linear Taylor rules

2.1. Traditional Taylor rule

Taylor rule, proposed by Taylor (1993) for U.S. real data during the period from 1987 to 1992, is one of the simplest monetary policy rules in common use. Given the inflation target and potential output, it provides an adjustment criteria of short-term interest rate to the changes of inflation and real output.

Taylor (1993) assumes that the central bank follows the following general behavior criteria:

$$i_t^* = \bar{r} + \pi^* + \beta(\pi_t - \pi^*) + \gamma(y_t - y_t^*), \tag{1}$$

where i_t^* is the desired value of the short-term nominal interest rate. It is assumed to rise if inflation rate (π) rises above its target (π^*) or if real output (y) increases above its trend (y^*). Therefore, in Eq. (1), the parameter β indicates the sensitivity of interest rate policy to the deviation of inflation rate from its target, and the parameter γ indicates the sensitivity of interest rate policy to the output gap. However, once the economy reaches the long-run equilibrium, the deviation of inflation rate from its target value is zero and the output gap is zero. Consequently, the desired nominal interest rate (i^*) is simply the sum of the equilibrium real interest rate ($\bar{r}$) and the target value of inflation (π^*), reducing to the standard Fisher Equation.

An important empirical question is related to the estimated weight on inflation. Since it is the real interest rate that actually drives private decisions, the size of β needs to assure that the nominal interest rate is raised enough to actually increase the real interest rate. This so-called “Taylor principle” implies that the coefficient on the inflation gap (β) should exceed 1.[2] A coefficient larger than one on inflation means that the central bank increases the real interest rate in response to higher inflation, which exerts a stabilizing effect on inflation; on the other hand, $\beta<1$ indicates an accommodative behavior of interest rate to inflation, which may generate self-fulfilling bursts of inflation and output.[3] In addition, the coefficient on the output gap (γ) should be positive. A positive coefficient on output gap means that if real output is below its potential level, a decrease in the interest rate will have a stabilizing influence on the economy.

In his study on U.S. data, Taylor (1993) suggested that both the equilibrium real interest rate and the target inflation rate be 2%. He also assumed that the weight of inflation gap and output gap relative to Federal Funds rate might be 1.5 and 0.5, respectively, i.e. $\beta = 1.5, \gamma = 0.5$.

2.2. Forward-looking Taylor rule

In the traditional Taylor rule, the reaction of interest rate to inflation and output gap is contemporary or backward-looking. For example, Taylor (1993) considered the deviation of inflation from target over the last four quarters. However, in practice, central banks do not tend to take the past or actual inflation as the target but the expected inflation. Therefore, Clarida et al. (1998) suggested introducing the expectation to construct a forward-looking version of Taylor rule, which allows the central bank to consider a broad array of information to form beliefs about the future condition of the economy. According to Clarida et al. (1998, 2000), the central bank's desired target interest rate (i_t^*) depends not only on the deviation of k periods ahead expected inflation

[2] See, e.g. Taylor (1999) and Clarida et al. (1998).

[3] Benassy (2006) characterized optimal interest rate rules in the framework of a dynamic stochastic general equilibrium model, and found that the elasticity of response for interest rate depends on the degree of price rigidity, the autocorrelation of the underlying shocks, or which measure of inflation is used. It is also suggested that in general the optimal elasticity of the interest rate with respect to inflation needs not be great than one.

(in annual rates) from its target value but also the p periods ahead expected output gap. Then we have the following forward-looking Taylor rule[4]:

$$i_t^* = \bar{r} + \pi^* + \beta\left[E_t\left(\pi_{t,k}\right) - \pi^*\right] + \gamma E_t\left(y_{t,p} - y_{t,p}^*\right), \tag{2}$$

where E_t is the conditional expectation operator, conditional on information available to the monetary authority at time t, $E_t(\pi_{t,k})$ is the k periods ahead forecast for inflation at time t, and $E_t(y_{t,p} - y_{t,p}^*)$ is the p periods ahead forecast for output gap at time t.

Moreover, we also assume that the central bank adjusts interest rates in a cautious way through smoothing in the form of partial adjustment as follows[5]:

$$i_t = (1-\theta)i_t^* + \theta i_{t-1} + m_t \tag{3}$$

where $\theta \in (0,1)$ is the smoothing parameter, m_t is a random disturbance term caused by the central bank's control of interest rate. According to this partial adjustment behavior, the central bank at each period adjusts its instrument in order to eliminate only a fraction $(1-\theta)$ of the gap between its current target level and some linear combination of its past values. Therefore, the parameter θ can be considered as an index which captures the degree of interest rate smoothing (Clarida et al., 2000).

Then, let $\tilde{y}_{t,p} = y_{t,p} - y_{t,p}^*$ denote the output gap, $c = \bar{r} - (\beta - 1)\pi^*$, and substitute Taylor rule Eq. (2) into Eq. (3). Rearranging, we can get the following simplified forward-looking monetary policy function:

$$i_t = (1-\theta)\left[c + \beta\pi_{t,k} + \gamma\,\tilde{y}_{t,p}\right] + \theta i_{t-1} + e_t \tag{4}$$

where $e_t = -(1-\theta)\left\{\beta\left[\pi_{t,k} - E_t(\pi_{t,k})\right] + \gamma\left[\tilde{y}_{t,p} - E_t\left(\tilde{y}_{t,p}\right)\right]\right\} + m_t$.[6] Eq. (4) is regarded as the forward-looking monetary reaction function considering interest rate smoothing. Kim and Nelson (2006) showed that there may be some dependence between the error term (e_t) and two explanatory variables ($\pi_{t,k}$ and $\tilde{y}_{t,p}$), which produces the problem of endogeneity.

In the existing literature, Eq. (4) is usually estimated by the generalized method of moments (GMM). According to Clarida et al. (1998, 2000), this estimation method is well suited for the econometric analysis of interest rate rules when the regressions are made on that are not known by the central bank at the decision-making moment. To implement this estimation procedure, the following orthogonal condition is imposed on:

$$E_t\left\{i_t - (1-\theta)\left[c + \beta\pi_{t,k} + \gamma\,\tilde{y}_{t,p}\right] - \theta i_{t-1} \,|\, z_t\right\} = 0, \tag{5}$$

where z_t is a vector of instrument variables that are orthogonal to e_t. These variables can be a set of lagged variables and additional exogenous variables that help to predict inflation and output gap. In addition, in order to account for the possible heteroscedasticity and serial correlation in e_t, an optimal weighting matrix is often used in estimation. Therefore, the GMM estimator is robust to heteroscedasticity and serial correlation.

In Kim and Nelson's (2006) study, they applied Kim's (2006) time-varying parameter model with endogenous regressors to model the forward-looking Taylor rule. Their work made two important progresses. First, it can be extended to deal with nonlinearities, which allow all parameters in the model to be time-varying. Second, it can deal with heteroscedasticity in the disturbance terms for the monetary policy rule, which allows us to consider the degree of uncertainty associated with economic conditions. In order to obtain the consistent estimation in the time-varying parameter model, they suggested a two-step MLE procedure to deal with endogeneity in regressors.

3. Regime-switching Taylor rule and its estimation

In this section, the linear forward-looking Taylor rule is generalized to the regime switching form. We first introduce the regime switching forward-looking Taylor rule, and then set up its econometric model. After that, we apply the two-step MLE procedure suggested by Kim and Nelson (2006) and Kim (2004, 2006, 2009) and Zheng and Wang (2010) to estimate this nonlinear model.

[4] Although empirically motivated, the forward-looking Taylor rule can also be derived theoretically. Assuming that the evolution of economy described by the New-Keynesian model is linear, that is the linear IS curve and the linear Phillips curve, and the central bank uses a policy rule to implement monetary policy, then the central bank can get the first-order necessary conditions for the policy reaction function (i.e. the forward-looking Taylor rule) by minimizing the inter-temporal loss function.

[5] Traditional explanations for the interest rate smoothing hypothesis include: fear of disrupting capital markets, loss of credibility from a sudden large policy reversal, the need for consensus building for a policy change, etc. In addition, interest rate smoothing may be thought of as a learning device by the central bank, which may not have a full knowledge of economy due to imperfect information.

[6] When expected value equals actual value, i.e. expected inflation equals to actual inflation and expected output gap equals to actual output gap, this forward-looking Taylor rule degenerates into the contemporary Taylor rule.

3.1. Model specification

For the linear Taylor rule, policy reaction coefficients and interest rate smoothing parameter are generally assumed to be constant. But in the long run, these parameters may be time-varying or state dependent. As mentioned earlier, we consider the following extended forward-looking Taylor rule with regime switching nonlinearity:

$$i_t = \left(1-\theta_{S_t}\right)\left[c_{S_t} + \beta_{S_t}\pi_{t,k} + \gamma_{S_t}\tilde{y}_{t,p}\right] + \theta_{S_t}i_{t-1} + e_t \tag{6}$$

where $e_t = -(1-\theta_{S_t})\left\{\beta_{S_t}\left[\pi_{t,k}-E_t(\pi_{t,k})\right] + \gamma_{S_t}\left[\tilde{y}_{t,p}-E_t\left(\tilde{y}_{t,p}\right)\right]\right\} + m_t$ is the disturbance term. S_t is a discrete random variable, whose realization can be interpreted as the state of economy at period t. In this study, we assume that the state variable S_t can be 1, 2…, N, and it follows a first-order Markov chain with the constant transition probability matrix:

$$\mathbf{P} = \begin{bmatrix} p_{11} & p_{21} & \cdots & p_{N1} \\ p_{12} & p_{22} & \cdots & p_{N2} \\ \vdots & \vdots & \ddots & \vdots \\ p_{1N} & p_{2N} & \cdots & p_{NN} \end{bmatrix} \tag{7}$$

where $p_{ij} = \Pr(S_t = j|S_{t-1} = i)$ is the transition probability of moving from the state $S_{t-1} = i$ at time $t-1$ to the state $S_t = j$ at time t, $i,j = 1,2,\ldots,N$.

Eq. (6) relaxes the linearity assumption of the traditional Taylor rule, allowing all parameters to be possibly state dependent, i.e. the parameters c, β, γ and θ can switch across states. By taking the policy reaction coefficients as the regime switching form, we can capture possible regime shifts in monetary policy operations.

In addition, as argued by Sims (2001) and Sims and Zha (2006), we note the importance of time-varying variance of the shocks in the monetary policy rule. Thus, the disturbance term e_t is assumed to follow a Gaussian distribution with state-dependent variance, i.e.[7]

$$e_t \sim N\left(0, \sigma^2_{S_t}\right), \tag{8}$$

and without loss of generality, we assume that $\sigma_1^2 > \sigma_2^2 > \cdots > \sigma_N^2$.

In order to estimate above Eqs. (6)–(9), we need several instrumental variables (IV). Similar to Kim and Nelson (2006), we assume that the relationships between the endogenous regressors ($\pi_{t,k}$ and $\tilde{y}_{t,p}$) in Eq. (6) and the vector of IV z_t are given by:

$$\pi_{t,k} = z_t'\delta_{1t} + v_{1t}, \quad v_{1t} \sim N\left(0, \sigma^2_{v1}\right), \tag{9}$$

$$\tilde{y}_{t,p} = z_t'\delta_{2t} + v_{2t}, \quad v_{2t} \sim N\left(0, \sigma^2_{v2}\right), \tag{10}$$

with

$$\delta_{it} = \delta_{it-1} + u_{it}, \quad u_{it} \sim iidN\left(0, \Sigma_{u,i}\right), i = 1, 2, \tag{11}$$

where the error terms v_{1t} and v_{2t} are assumed to be correlated with the correlation coefficient ρ, i.e. $\mathrm{corr}(v_{1t}, v_{2t}) = \rho$.

As specified above, the relationships between the endogenous regressors ($\pi_{t,k}$ and $\tilde{y}_{t,p}$) and the vector of IV z_t in Eq. (6) could be time-varying. So, this time-varying parameter model could catch some gradual changes over time.

3.2. Two step estimation procedure

Now let us take in account the estimation of the regime switching model given by Eqs. (6)–(8) mentioned above. Obviously, as an unobservable discrete variable is added, it is not possible to apply the GMM estimation procedure to estimate this regime switching model. Therefore, this study estimates this nonlinear model with endogenous repressors based on the two-step MLE procedure suggested by Kim and Nelson (2006), Kim (2009) and Zheng and Wang (2010).[8]

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[7] An alternative possible setting for the time varying variance is using the GARCH (generalized autoregressive conditional heteroscedasticity) models. See for example Kim and Nelson (2006) and Kim, Kishor, and Nelson (2006).

[8] The two-step estimation procedure has at least two advantages. One is that it provides us a consistent estimator rather than an approximate estimator due to Kim's (1994) approximation (Zheng & Wang, 2010). Another important advantage is that as the number of states for the latent Markov-switching variables and endogenous regressors, the joint estimation procedure is infeasible due to the "curse of dimensionality" (Kim, 2009). In this case, we have no choice but to employ the two-step procedure introduced in this paper, at the cost of a loss of asymptotic efficiency.

For the model given by Eqs. (9)–(11), we can represent it as the following state space form:

$$\begin{bmatrix}\pi_{t,k}\\ \tilde{y}_{t,p}\end{bmatrix} = \left(I_2 \otimes z_t'\right)\delta_t + v_t = \left(I_2 \otimes z_t' \Sigma_v^{1/2}\right)\begin{bmatrix}\delta_t\\ v_t^*\end{bmatrix}, \quad \Sigma_v^{1/2} = \begin{bmatrix}\sigma_1 & 0\\ \rho\sigma_2 & \sigma_2\end{bmatrix}, \tag{12}$$

$$\begin{pmatrix}\delta_t\\ v_t^*\end{pmatrix} = \begin{pmatrix}I_k & 0\\ 0 & 0\end{pmatrix}\cdot\begin{pmatrix}\delta_{t-1}\\ v_{t-1}^*\end{pmatrix} + \begin{pmatrix}u_t\\ v_t^*\end{pmatrix}, \begin{pmatrix}u_t\\ v_t^*\end{pmatrix} \sim N\left(\begin{pmatrix}0\\ 0\end{pmatrix}, \begin{pmatrix}\Sigma_u & 0\\ 0 & I_2\end{pmatrix}\right), \tag{13}$$

where the time varying parameter vector $\delta_t = [\delta_{1t} \quad \delta_{2t}]'$, the mutually independent error vector $v_t^* = [v_{1t}^* \quad v_{2t}^*]'$ are iid standardized normal variables. It is obvious that the model given by Eqs. (12)–(13) can be estimated with the maximum likelihood estimation procedure by running the conventional Kalman filter. Furthermore, smoothed estimates of $\tilde{\delta}_{t|T}$ and its corresponding variance $\tilde{P}_{t|T}$ can also be obtained from the conventional Kalman filter. We denote the last $K\times 1$ block of $\tilde{\delta}_{t|T}$ as $v_{t|T}^*$, and the last $K\times K$ block of $\tilde{P}_{t|T}$ as $P_{t|T}^* = E[(v_t^* - v_{t|T}^*)\cdot(v_t^* - v_{t|T}^*)']$, which will be used in the second step.

Since there exists correlated structure between the shock term e_t and two regressors $\pi_{t,k}$ and $\tilde{y}_{t,p}$, we then assume the following covariance structure between $v_t^* = [v_{1t}^* \quad v_{2t}^*]'$ and e_t:

$$\begin{bmatrix}v_t^*\\ e_t\end{bmatrix} \sim N\left(\begin{bmatrix}0\\ 0\end{bmatrix}, \begin{bmatrix}I_2 & \rho_{S_t}\sigma_{S_t}\\ \rho_{S_t}'\sigma_{S_t} & \sigma_{S_t}^2\end{bmatrix}\right) \tag{14}$$

where $\rho_{S_t} = [\rho_1 S_t \quad \rho_2 S_t]'$ is a state-dependent 2×1 vector of correlation coefficients. As shown in Kim (2004, 2009), the Cholesky decomposition of the covariance matrix of $[v_t^{*\prime} \quad e_t]'$ in Eq. (14) results in the following representation:

$$\begin{bmatrix}v_t^*\\ e_t\end{bmatrix} = \begin{bmatrix}I_2 & 0\\ \rho_{S_t}'\sigma_{S_t} & \sqrt{1-\rho_{S_t}'\rho_{S_t}\sigma_{S_t}}\end{bmatrix}\begin{bmatrix}\varepsilon_t\\ \eta_t\end{bmatrix}, \begin{bmatrix}\varepsilon_t\\ \eta_t\end{bmatrix} \sim i.i.d.N\left(\begin{bmatrix}0\\ 0\end{bmatrix}, \begin{bmatrix}I_2 & 0\\ 0 & 1\end{bmatrix}\right) \tag{15}$$

where I_2 is a 2×2 identity matrix.

Then, e_t can be rewritten as a linear form of v_{1t}^* and v_{2t}^*, i.e.

$$e_t = \gamma_{S_t} v_t^* + \omega_t, \quad \omega_t \sim N\left(0, \sigma_{\omega,S_t}^2\right), \tag{16}$$

where $\gamma_{S_t} = \rho'_{S_t}\sigma_{S_t}$, $\sigma_{\omega,S_t}{}^2 = (1-\rho'_{S_t}\rho_{S_t})\sigma_{S_t}{}^2$, and the disturbance term ω_t is uncorrelated with either v_{1t}^* or v_{2t}^*. Thus, substituting Eq. (16) into Eq. (6), we get the transformed model as follows:

$$i_t = \left(1-\theta_{S_t}\right)\left[c_{S_t} + \beta_{S_t}\pi_{t,1} + \gamma_{S_t}\tilde{y}_{t,1}\right] + \theta_{S_t} i_{t-1} + \gamma_{S_t}' v_t^* + \omega_t. \tag{6′}$$

As the disturbance term ω_t is uncorrelated with $\pi_{t,1}$, $\tilde{y}_{t,1}$, v_{1t}^* and v_{2t}^*, the following two-step MLE would be valid:

Step 1: Estimate Eqs. (9)–(11) via the MLE procedure based on the conventional Kalman filter, and obtain standardized errors $\hat{v}_{1,t|T}^*$ and $\hat{v}_{2,t|T}^*$, and their corresponding covariance matrix $\hat{P}_{t|T}^* = E\left[\left(v_t^* - v_{t|T}^*\right)\cdot\left(v_t^* - v_{t|T}^*\right)'\right]$.
Step 2: Using the maximum likelihood method via the Hamilton's (1989) filter to estimate the following equation:

$$i_t = \left(1-\theta_{S_t}\right)\left[c_{S_t} + \beta_{S_t}\pi_{t,1} + \gamma_{S_t}\tilde{y}_{t,1}\right] + \theta_{S_t} i_{t-1} + \gamma_{S_t}\hat{v}_{t|T}^* + \omega_t^*, \omega_t^* \sim i.i.d.N\left(0, \sigma_{\omega t}^2\right), \tag{6″}$$

where $\omega_t^* = \omega_t + \gamma_{S_t}\left(v_t^* - \hat{v}_{t|T}^*\right)$ and $\sigma_{\omega t}^2 = \left[1 - \rho_{S_t}'\left(I_2 - \hat{P}_{t|T}^*\right)\rho_{S_t}\right]\sigma_{S_t}^2$ which is suggested by Zheng and Wang (2010). Given $\hat{v}_t^*$ and the corresponding covariance matrix $\hat{P}_{t|T}^*$, the model parameters can be estimated by the maximum likelihood procedure, and then the unobservable discrete variable S_t can be exactly inferred given estimated parameters to get the filtered and smoothed probabilities.

Remark 1. In the maximum likelihood estimation of Eq. (6″) based on the derived Hamilton filter, the parameter estimates are consistent, and the standard errors have been corrected by introducing the term $\hat{P}_{t|T}^*$ into the variance of ω_t^* for Pagan's (1984) generated regressors' problem. For further details, the readers are referred to a series of papers by Kim (2004, 2006, 2009), Kim and Nelson (2006), Kim and Kim (2007) and Zheng and Wang (2010).

Remark 2. The aforementioned two step estimation procedure can also be used to estimate the model of linear Taylor rule. If one takes account of one regime or the state variable only takes one number, then the regime switching Taylor rule would reduce to a linear rule. In the case of the two step estimation for the linear rule, the sole difference from that of the regime switching rule is that all parameters are not state dependent, leading to a very simple MLE procedure in the second step.

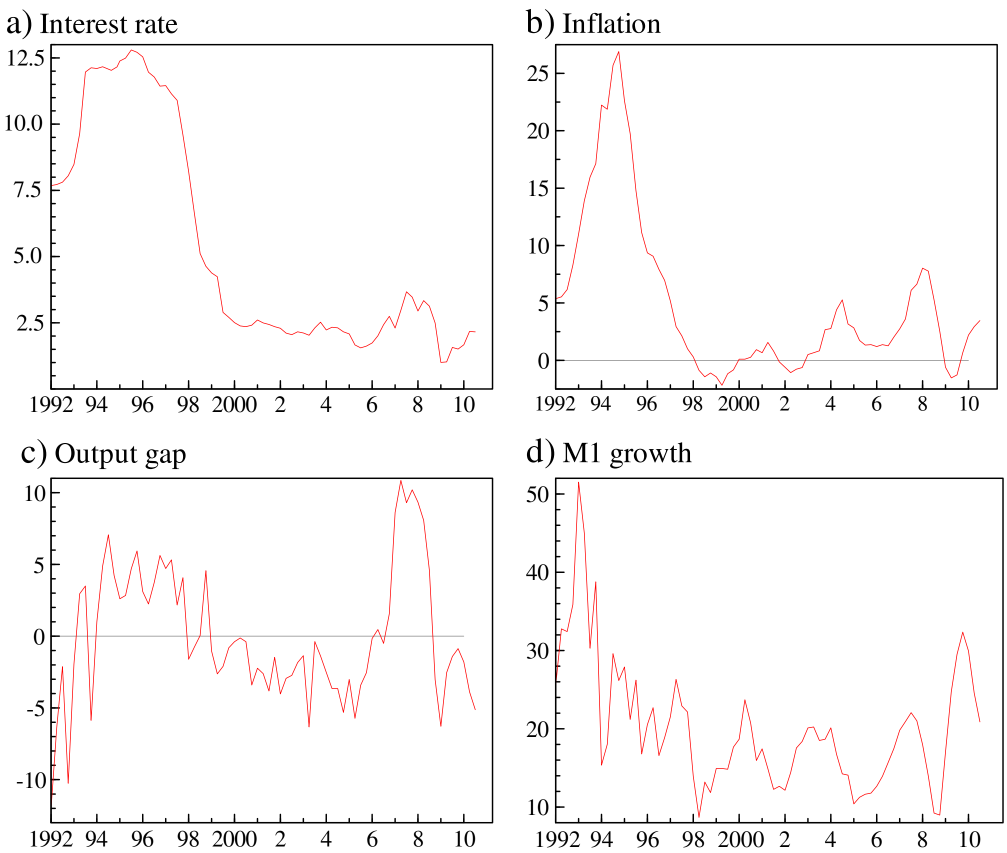

Fig. 1. Data series during the period 1992Q1 to 2010Q3.

4. Data

In this study, the data series we employed are quarterly data for China covering the period 1992:Q1-2010:Q3 with 75 observations in total. The detail of data selection, processing, and description is given as follows:

(1) Short-term nominal interest rate. According to most of Chinese scholar's studies such as Xie and Luo (2002), Lu and Zhong (2003), Zhao and Gao (2004), and recently Zheng and Liu (2010), we choose the China inter-bank offered rate (CHIBOR) as a proxy variable of interest rate. This index meets with the requirement of market interest rate, and can quickly reflect the demand and supply in money market. The data of quarterly interest rate during the period from 1992 to 1995 is the inter-bank offered rate in Shanghai Financial Center. Although the lending behavior of China's financial institutions is disordered before and after 1993, the inter-bank offered rate in Shanghai can still reflect the situation of nationwide inter-bank offered market before networking in 1996 (Xie & Luo, 2002). For the interest rate data during the period 1996–2010, we select 7-day China inter-bank offered rate,[9] which can essentially reflect the variability of recent market position. We get quarterly CHIBOR by quarterly averaging of monthly data, which are published by the PBC's (People's Bank of China) Quarterly Statistical Bulletin and the PBC's website (http://www.pbc.gov.cn).[10] The final data of quarterly short-term interest rate are depicted in Fig. 1(a).

(2) Inflation rate. Taking into account the availability and reliability of consumer price index (CPI) and GDP deflator, we choose the CPI as a measure of inflation rate. In our study, the annualized quarterly inflation rate is calculated according to the data of monthly year-on-year growth rate of CPI, which is obtained from CEI database (available from: http://db.cei.gov.cn) and China Monthly Economic Indicators. Since the official CPI is monthly year-on-year growth rate, three term moving average

[9] Due to the limited data sources, the weighting interest rates during the period 1992–1995 are calculated by all term interest rate weighting of Shanghai Financial Center, whereas for the period from 1996 to 2009 the interest rate is seven day inter-bank offered rate. Although the term in these two periods does not match, it has little effect on modeling in that there is not large in interest differential of different terms for Shanghai Financing Center. (Xie & Luo, 2002).

[10] Based on the data of the monthly inter-bank offered rate and its corresponding trading volume, we can calculate the quarterly weighted average inter-bank offered rate as follows:

$$\bar{i} = i_1 \frac{f_1}{\sum f} + i_2 \frac{f_2}{\sum f} + \cdots + i_n \frac{f_n}{\sum f} = \frac{\sum if}{\sum f},$$

where i_k is the monthly interest rate, f_k is its corresponding trading volume.

Table 1
AICs and BICs under different choice of k and p.

	AIC					BIC				
	$p=0$	$p=1$	$p=2$	$p=3$	$p=4$	$p=0$	$p=1$	$p=2$	$p=3$	$p=4$
$k=0$	111.86	102.05	111.07	106.89	105.76	127.70	117.79	126.71	122.42	121.19
$k=1$	106.70	108.19	105.10	103.04	103.00	122.44	123.93	120.73	118.58	118.43
$k=2$	100.09	99.56	**90.16**	98.62	99.16	115.73	115.20	**105.80**	114.15	114.59
$k=3$	109.01	100.32	99.41	99.90	98.93	112.06	115.86	114.95	115.44	114.37
$k=4$	105.76	103.00	99.16	98.93	95.41	121.19	118.43	114.59	114.37	110.84

Note: The figures are the corresponding Akaike information criterions (AIC) and Bayesian information criterions (BIC) computed by the linear model based on different values of k and p.

is adopted to get quarterly year-on-year growth rate of CPI, and then the quarterly data of annual inflation rate, π_t, is calculated as follows: $\pi_t=(\text{quarterly CPI}-1)\times 100\%$. The quarterly inflation rate based on CPI is shown in Fig. 1(b).

(3) Output gap. To get the real output gap, we should firstly measure potential output based on real output. The gross domestic product (GDP) is selected as a measure of total output. Official statistical data provides quarterly GDP at current prices and cumulative GDP growth rate over the same period last year since 1992. The data is available from CEI database (http://db.cei.gov.cn) and China Monthly Economic Indicators. In order to get real GDP, denoted by Y_t, we first re-calculate real value at comparable price (setting year 2000 = 100) based on cumulative GDP growth rate, and then compute the seasonally adjusted GDP series via X-12 seasonal adjustment method executed by the software Eviews. After applying the Hodrick–Prescott filter (HP filter) to obtain potential output, denoted by Y_t^*, the annualized output gap is computed as percentualized log-deviation of real output with respect to potential output, i.e. $y_t\equiv 400\times\log(Y_t/Y_t^*)$. Fig. 1(c) draws the final data of real output gap.

In addition, to estimate the forward-looking monetary policy reaction function, we need to construct instrumental variables (IV) to calculate standardized prediction errors $\{v_{1t}^*\}$ and $\{v_{2t}^*\}$. The IV include four lags of each of the following variables: the short-term inter-bank offered rate, inflation rate, output gap, and the M1 growth rate which is shown in Fig. 1(d).[11]

5. Empirical results

In this section, we first consider the selection of the model, then estimate the regime switching forward-looking Taylor rule, and at last analyze the asymmetric behavior of China's monetary policy.

5.1. Model selection

5.1.1. Determination of k and p

The first task of model selection is to determine the appropriate numbers of k and $p$for the linear forward-looking rule. We estimate the linear model with different numbers of k and p (i.e. taking the values 0, 1, 2, 3 or 4) using the two step estimation procedure, and then choose the model which minimizes the information criteria such as Akaike's information criterion (AIC) and Bayesian information criteria (BIC). The smaller is the AIC or the BIC, the better the model is.

Table 1 presents the AICs and BICs of the linear Taylor rule with different numbers of k and p. According to the results, both the AIC and BIC values arrive at the minimum value when $k=p=2$. Therefore, in the following context, we consider this appropriate setting $k=p=2$ and fix them.

The parameter estimates of linear forward-looking Taylor rule are reported in Table 2. As shown in the table, the interest rate smoothing parameter is very significant, indicating that the interest rate adjustment mechanism strongly depends on the previous interest rate. For the response of short-term interest rate to inflation, the policy parameter β is significant but less than 1, indicating that the interest rate policy rule is unstable; moreover, for the response of short-term interest rate to the output gap, the policy parameter (γ) is insignificant. The above results are consistent with a majority of Chinese scholars' researches, such as Xie and Luo (2002) and Bian (2006). In addition, the estimation results of a linear model without considering endogenous regressors are also compared with the two step estimation results. The results show that the correlation coefficient with respect to output gap is significant, showing its importance of considering the problem of endogeneity.

5.1.2. Determination of the number of regimes

For the specification of the Markov-switching model, it is important to know the number of states which is often not known a priori. Although Akaike's information criterion (AIC) has been used in various model selection contexts, Smith, Naik, and Tsai (2006) found that in Markov-switching models it misleads the users into selecting too many states. Using the Kullback–Leibler

[11] We do not adopt the M2 growth rate, mainly because the official provided M2 growth rate since 1996.

Table 2
Parameter estimates of linear Taylor rule.

Parameter	Estimation without endogeneity			Two-step estimation		
	Estimate	Std error	p-value	Estimate	Std error	p-value
Interest rate smoothing θ	0.9390	(0.0160)	[0.0000]	0.9178	(0.0165)	[0.0000]
Constant c	−0.5099	(1.6482)	[0.7581]	0.3913	(1.0114)	[0.7002]
Inflation reaction β	0.9710	(0.2356)	[0.0001]	0.8537	(0.1434)	[0.0000]
Output gap reaction γ	0.1405	(0.2349)	[0.5518]	0.2743	(0.1807)	[0.1340]
Correlation coefficient ρ_1				−0.1743	(0.1338)	[0.1975]
Correlation coefficient ρ_2				−0.5381	(0.1186)	[0.0000]
Standard deviation σ	0.4603	(0.0395)	[0.0000]	0.4684	(0.0407)	[0.0000]
Log-likelihood	−43.7275			−38.0813		

Note: Figures in parentheses and brackets are corresponding standard errors and p-values of parameter estimates.

divergence between the true and candidate models, Smith et al. (2006) derive the following formal procedure called Markov switching criterion (MSC) to select the number of states and variables simultaneously:

$$MSC = -2\log L + \sum_{i=1}^{N} \frac{T_i(T_i + \lambda_i K)}{\delta_i T_i - \lambda_i K - 2} \tag{17}$$

where N is the number of states and K is the number of the regressors in the model. Also, T_i is the sum of smoothed probabilities of being in the regime i. $\sum_{i=1}^{N} T_i = T$, where T is the number of the effective observations used in estimation. The second term in Eq. (17) is a penalty for the complexity of a model. Smith et al. (2006) recommend setting $\delta_i = 1$ and $\lambda_i = N$ (also $\lambda_i = 1$ or $\lambda_i = N^2$ is another possible choice but they show that $\text{MSC}_{\lambda=1}$ and $\text{MSC}_{\lambda=N^2}$ perform worse than $\text{MSC}_{\lambda=N}$ when both the sample size is small and the signal is weak).

We estimate one-state, two-state and three-state Markov-switching models, respectively, and compute two estimates of KL divergence: AIC and $\text{MSC}_{\lambda=N}$. Based on the minimum $\text{MSC}_{\lambda=N}$ value, we would select a model with $N^* = 2$ while the minimum value of AIC yields $N^* = 3$. This finding is consistent with the simulation evidence given by Smith et al. (2006) which reveals AIC's tendency to select more states than necessary. Therefore, this paper takes $N = 2$ as the number of Markov-switching states.

5.2. Results of the regime switching model

Table 3 reports the estimation results of the regime-switching forward-looking Taylor rule with two states. As shown in the table, all parameter estimates are significantly different from zero under 5% significant level, except for c_1, γ_1 and ρ_{12} (the correlation coefficient ρ_1 in the second regime or Regime 2). It seems that the regime switching model can well describe the dynamic behavior of interest rate policy. Moreover, for the regime switching model, its log likelihood value increases by 39.89 relative to the linear model, showing better performance in modeling the short term rate. In particular, the results also show important implication of considering the problem of endogeneity. As we can see from the table, the correlation coefficients ρ_2 in both regimes are significant and very high, while the correlation coefficient ρ_1 in the second regime is not significant, but significant in the first regime. Therefore according to the two-step MLE procedure, we can also capture the existence of significant correlation in endogenous structure using the regime switching model.

Table 3
Parameter estimates of regime-switching Taylor rule.

Parameter	Regime 1			Regime 2		
	Estimate	Std error	p-value	Estimate	Std error	p-value
Interest rate smoothing θ	0.8894	(0.0223)	[0.0000]	0.9474	(0.0082)	[0.0000]
Constant c	−1.6846	(1.6420)	[0.3094]	2.3796	(0.5075)	[0.0000]
Inflation reaction β	1.3312	(0.2362)	[0.0000]	0.3683	(0.0687)	[0.0000]
Output gap reaction γ	0.2499	(0.2049)	[0.2280]	0.7333	(0.1439)	[0.0000]
Correlation coefficient ρ_1	−0.4610	(0.1427)	[0.0021]	0.1138	(0.0938)	[0.2307]
Correlation coefficient ρ_2	−0.7364	(0.1268)	[0.0000]	−0.9460	(0.0587)	[0.0000]
Standard deviation σ	0.4398	(0.0622)	[0.0000]	0.1891	(0.0240)	[0.0000]
Transition probability matrix						
$\Pr(S_t = 1 \mid S_{t-1})$	0.7512	(0.1124)	[0.0000]	0.1449	(0.0675)	[0.0363]
$\Pr(S_t = 2 \mid S_{t-1})$	0.2488	(0.1124)	[0.0312]	0.8551	(0.0675)	[0.0000]
Log-likelihood	1.8082					
$\text{MSC}_{\lambda=N}$	161.67					

Note: Figures in parentheses and brackets are corresponding standard errors and p-values of parameter estimates.

Table 4
Wald test results.

$W_1\left(\hat{\theta}_1 = \hat{\theta}_2\right)$	$W_2\left(\hat{c}_1 = \hat{c}_2\right)$	$W_3\left(\hat{\beta}_1 = \hat{\beta}_2\right)$	$W_4\left(\hat{\gamma}_1 = \hat{\gamma}_2\right)$	$W_5\left(\hat{\rho}_{11} = \hat{\rho}_{12}\right)$	$W_6\left(\hat{\rho}_{21} = \hat{\rho}_{22}\right)$	$W_7\left(\hat{\sigma}_1 = \hat{\sigma}_2\right)$
6.5435	5.9510	15.562	4.2095	11.974	2.3825	14.405
[0.0105]	[0.0147]	[0.0001]	[0.0402]	[0.0005]	[0.1227]	[0.0001]

Note: The Wald statistic, W, is used to test the parametric stability of the regime switching model, i.e. whether a parameter in first regime (μ_1) is equal to another parameter in the second regime (μ_2), where the null hypothesis is $\mu_1 = \mu_2$, the alternative hypothesis is $\mu_1 \neq \mu_2$. Then Wald-test statistic is given by

$$W(\mu_1 = \mu_2) = \frac{(\mu_1 - \mu_2)^2}{\mathrm{Var}(\mu_1) + \mathrm{Var}(\mu_2) - 2\mathrm{Cov}(\mu_1, \mu_2)} \rightarrow \chi^2(1),$$

where $\mathrm{Var}(\mu)$ is the variance of μ, $\mathrm{Cov}(\mu_1,\mu_2)$ is the covariance of μ_1 and μ_2, $\chi^2(1)$ is the chi-square distribution with freedom of degree 1.

However, two important issues arise, i.e. whether the interest rate policy takes on regime switching nonlinearity, and whether interest rate smoothing parameter, policy reaction coefficients and standard deviation are asymmetric in different regimes. In order to check these problems, we carry out the following two formal tests, respectively.

5.2.1. Testing for linearity

Including regime switching nonlinearity appears to represent an improvement over the linear model, comparing Table 2 and Table 3. Testing the restriction of the linear model, i.e. that all parameters are not state dependent, the likelihood ratio (LR) test statistic is 79.78. This test statistic, however, is nonstandard.[12] In order to establish the statistical significance of this result, a parametric bootstrap test was performed in terms of Di Sanzo (2009).[13]

The bootstrap algorithm runs as follows: The data were simulated under the null of no regime switching nonlinearity, i.e. using the parameter estimates of the linear model from Table 2. The model was re-estimated for each sample under both the null and the alternative to obtain a likelihood ratio test statistic. The bootstrapped p-value, based on 10,000 bootstrap samples, is 0.0001, which is the fraction of simulated LR values that are greater than the observed LR value. This suggests that regime switching nonlinearity is indeed appropriate for explaining the movements in China's monetary policy reaction.

5.2.2. Testing parameter stability

After estimating the regime switching model, we proceed to test the possible asymmetries in the policy rule. By imposing restrictions on the parameters being not state dependent, we test the following null hypotheses of parameter stability: $\mu_1 = \mu_2$, where μ could be the interest rate smoothing parameter, policy reaction coefficients, correlation coefficients or standard deviation. The statistical significance of the restrictions imposed on the model is assessed by the usual Wald test. Under null hypothesis, the Wald test has a chi-square distribution with 1 degree of freedom.[14]

Table 4 reports the Wald test results of parameter stability across regimes. For the interest rate smoothing parameter (θ), policy reaction coefficients (β and γ) and standard deviation (σ), each Wald test for parameter stability shows that the null hypothesis can be strongly rejected under 5% significant level, indicating that these parameters are statistically different between two regimes. For the correlation coefficients (ρ_1 and ρ_2) capturing endogenous structure, the Wald test results show that the correlation coefficient of interest rate to the standardized prediction error of expected inflation (v_1^*) is significantly state dependent, while the correlation coefficient of interest rate to the standardized prediction error of expected output gap (v_2^*) is not state dependent.

5.2.3. The policy reaction function

In this part, let us turn to evaluate the nonlinear policy reaction function according to the parameter estimates according to Table 3 and Table 4. From the two parameters θ and σ, the result reveals that the interest rate smoothing mechanism and the interest rate shock are variant, depending on the state of economy; furthermore, the greater the degree of interest rate smoothing is, the smaller the variance of the interest rate shock is, and vice versa. Similarly, the standard deviation σ also depends on the discrete state, implying that the variance of the interest rate shock is time-varying, but here it is represented in a regime switching form. From policy reaction coefficients β and γ, the reactions of interest rate policy to inflation and output gap do not follow the traditional Taylor rule. This is because in the second regime, the inflation coefficient is less than 1 ($\beta<1$) and the output gap coefficient is larger than 0 ($\gamma>0$); however, in the first regime, the inflation coefficient β is larger than 1 and the output gap coefficient is not statistically different from zero.

Consequently, we can draw a conclusion that the monetary policy reaction function that relates the nominal interest rate to inflation and output fluctuation is nonlinear and asymmetric. It is revealed in this paper that the PBC targets inflation in the first

[12] Testing for linearity in the context of Markov-switching models is complicated because standard regularity conditions for likelihood based inference are not maintained under the null hypothesis due to the presence of nuisance parameters and singularity of the information matrix (Hansen, 1992). Thus, the asymptotic distribution of the relevant likelihood ratio (LR) statistic does not possess the standard chi-square-distribution.

[13] Although several alternative test methods such as Hansen (1992, 1996), Garcia (1998) and Carrasco, Liang, and Ploberger (2004) are available, this study adopts the bootstrap-based LR test introduced by Di Sanzo (2009) for testing linearity in Markov-switching models mainly due to its relative simplicity. Di Sanzo (2009) showed that a bootstrap of a likelihood ratio test statistic performs well for testing linearity in Markov-switching models.

[14] In more detail, this test for the regime switching model is referred to Hamilton (1996).

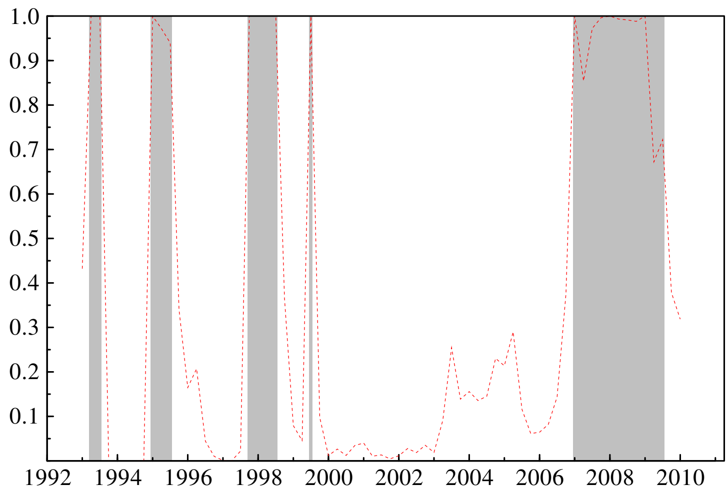

Fig. 2. Smoothed probabilities in the first regime.

regime while focuses on the output gap in the second regime. That is, in the first regime, if the inflation rate is 1% higher than its target value, the real interest rate would raise by 0.33%, and the nominal interest rate would raise by 1.33%; if the inflation rate is 1% lower than its target value, the real interest rate and the nominal interest rate would fall by 0.33% and 1.33%, respectively. But it is not obvious for the output gap. In the second regime, if the real output is 1% higher than its potential level, the nominal interest rate and real interest rate would raise by 0.73% simultaneously; and if the real output is 1% lower than its potential level, the nominal interest rate and real interest rate would fall by 0.73% simultaneously. Under the output gap unchanged, if the inflation rate is 1% higher (lower) than its target value, the nominal interest rate would raise (fall) by 0.37%, while the real interest rate would fall (raise) by 0.63%.

5.3. Regime switching behavior of China's monetary policy

To learn the regime switching characteristic of China's monetary policy, we further estimate the regime probabilities of the state variable S_t using parameter estimates of the regime switching model in Table 3. Fig. 2 depicts the time paths of the smoothed probabilities in the first regime, i.e. $\Pr(S_t = 1|\psi_T)$, $t = 1, \ldots, T$, where ψ_T denotes all available information in the data sample. According to the decision rule whether the probability is larger than 0.5, we also draw the corresponding state by the shading area.

As shown in Fig. 2, China's monetary policy is obviously in the first regime during the periods 1993:Q2-1993:Q3, 1995:Q1-1995:Q3, 1997:Q4-1998:Q3, 1999Q3 and 2007:Q1-2009:Q3, while it is in the second regime during other periods. This result implies that China's monetary policy experienced more periods in the second regime than those in the first regime. This can also be reflected by the probabilities of self-maintenance, showing that the expected duration ($D_i = 1/(1 - p_{ii})$) of the second regime is longer than that of the first regime, about 4 and 7 quarters, respectively.

To show the relationship between economic situations and monetary policy intuitively, Fig. 3 demonstrates the time paths of data series and the periods in the first regime. Obviously, the occurrences of the first regime mostly correspond to several major adjustments in the benchmark of lending and deposit interest rates. It suggests that the central bank prefers to adjust interest rate

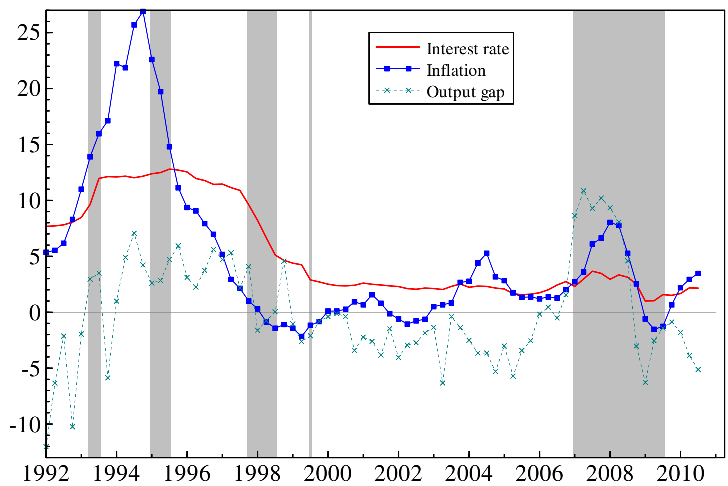

Fig. 3. Data series and regime shifts.

to manage public inflation expectations and thus curb deflation or high inflation effectively in the first regime. Moreover, the interest rate policy is always evident during the periods in the first regime, taking on larger interest rate changes. For instance, in 1993 China began to appear the most serious inflation since reform and opening up, the central bank increased the interest rate in May and July in 1993 and January and July in 1995 respectively, and then controlled the inflation effectively. When faced with Asian financial crisis and the clear signs of deflation in the domestic economy in the end of 1997, the central bank decreased the interest rate four times to curb the deepening deflation trend from October 1997 to December 1998. Similarly, the central bank increased interest rate six times continuously to prevent the evident inflation in 2007, which was the result of excessive credit growth and continuous price increase. However, as the “once in a century” international financial crisis deeply influenced China's economy and caused deflation, the central bank decreased the interest rate five times to achieve the goals of “maintain growth” and “anti-deflation” during the second half of 2008 and 2009.

On the contrary, the interest rate adjustment is relatively smooth in the second regime, and its changes are relatively small. During these periods, the central bank tends to carry out prudent monetary policy to promote the sustained, rapid and sound development of the national economy. In fact, as shown in the estimation results, the central bank focuses on economic fluctuation in the second regime, and the interest rate policy is sensitive to the output gap. From Fig. 3, we can observe that the interest rate and the output gap move in the same direction during the periods 1996–1997 and 2000–2006, reflecting the positive reaction of interest rate to output fluctuation in these periods.

6. Conclusions

In this study, we consider a regime-switching forward-looking Taylor rule that nests the simple rule with Markov switching in Zheng and Liu (2010) as a special case and realize its appropriate estimation via the two-step MLE procedure due to Kim and Nelson (2006), Kim (2009) and Zheng and Wang (2010). Based on this regime-switching rule, we do an empirical study on the dynamic reaction of interest rate policy to inflation and output gap for China during the period from 1992Q1 to 2010Q3. Our results show that a two-regime forward-looking rule performs very well in modeling actual reactions of China's monetary policy. It is also suggested that by employing this regime switching Taylor rule, we can capture significant asymmetry in the monetary policy reaction of the short-term interest rate to inflation and output gap.

Specifically, we can account for this asymmetry in the following three aspects. Firstly, this study reveals the regime switching characteristic of monetary policy. As can be seen earlier, all empirical estimates of interest rate smoothing, policy reaction coefficients and the variances of the interest rate shock are state-dependent. Moreover, the greater the interest rate smoothing is, the smaller variance of the interest rate shock is, and vice visa. Secondly, the PBC behaves quite different policy reactions in different regimes. For example, the central bank targets inflation in the first regime, i.e. the reaction of interest rate policy is mainly focused on inflation and deflation, but is not so obvious on output gap. However, in the second regime, the central bank mainly targets output gap. In this regime, the monetary policy reaction of the interest rate to inflation does not follow the traditional Taylor rule, but has an unstable framework. Finally, according to the monetary policy behavior of the central bank, the first regime is coincident with several major adjustments of the benchmark interest rates of lending and deposit, which plays an important role in curbing inflation and deflation. On the contrary, interest rate adjustment is relative smooth in the second regime, which plays an active role in stabilizing real output growth.

In summary, this paper, from perspective of regime switching, has investigated the monetary policy reaction of the interest rate to inflation and output gap, and verified the asymmetric reaction of the Chinese monetary policy. We strongly believe that it provides a new method and evidence to study China's monetary policy behavior. However, whether the monetary policy reaction by taking interest rate as its tool is linear or nonlinear remains a lot of support from empirical studies. In particular, its dynamic adjustment mechanism still requires some theoretical considerations. We leave this as a future research.

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