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An examination of macroeconomic fluctuations in Korea exploiting a Markov-switching DSGE approach☆

Jinho Choi [a,*], Joonyoung Hur [b]

[a] Monetary Policy Department, The Bank of Korea, 39, Namdaemun-ro, Jung-gu, Seoul, Republic of Korea
[b] Economic Research Institute, The Bank of Korea, 39, Namdaemun-ro, Jung-gu, Seoul, Republic of Korea

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ABSTRACT

This paper estimates a Markov-switching dynamic stochastic general equilibrium (MS-DSGE) model that allows shifts in the monetary policy rule coefficients as well as the shock volatilities with Korean data ranged from 1976 to 2013. We find that allowing for the regime-switching aspect both in monetary policy rules and shock volatilities is a crucial setup in improving the model's fit with Korean data. The regime estimates indicate that monetary policy more aggressively reacts to inflation, but less strongly to output, after launching the Inflation Targeting (IT) policy in the late 1990s. The identified regimes have three implications on macroeconomic performance in Korea. First, the introduction of the IT monetary policy has contributed to a sharp reduction in the level as well as the volatility of inflation in the 2000s. Second, technology shocks are the most important drivers of output fluctuations in Korea as the major economic crises in Korea are mainly explained by adverse shocks on technology. Finally, it would have been possible to achieve higher output and lower inflation simultaneously if the IT monetary policy regime was maintained over the entire sample period.

1. Introduction

This study seeks to empirically assess how monetary policy behavior in Korea has changed in response to structural shocks across different policy regimes, which are identified by estimating a Markov-switching dynamic stochastic general equilibrium (MS-DSGE) model. We posit that Korean economy may work as a natural laboratory for regime switching approaches. In order to rationalize this argument, Fig. 1 plots key macroeconomic time series of Korea from 1976:Q3 to 2013:Q3, in which large swings in output, inflation, and exchange rate are readily observed after outbreaks of major macroeconomic episodes for the Korean economy.

Over the last four decades, Korean economy had suffered from three significant economic crises—the second oil crisis in the late 1970s, the Asian currency crisis in 1997–98, and the global financial crisis in 2008–09. All these episodes have caused the severe recessions as well as the surges in price level and exchange rate.[1] Despite their similar patterns in outcomes, the aftermath of each episode entails quite differentiated characteristics, as well documented in Cho (2007) and Huh and Nam (2010). In particular, the Asian currency crisis might be the most crucial economic episode for the Korean economy. This economic turmoil has resulted in a sequence of institutional changes in Korea's economic stabilization policy. For instance, in December 1997 the Korean government has allowed foreign exchange rates to freely float, and in 1998, the Bank of Korea, the nation's central bank, has introduced an Inflation Targeting (IT) system to pursue its goal of price stability while having shifted to the interest rate based policy from the money aggregate based approach. In contrast, the recent global financial crisis has damaged Korea's export growth via the shrinking global demand for imported goods. Relatively minor institutional changes are involved in response to this economic event.

Grounded on these historical episodes, we formulate a small open economy MS-DSGE model for Korea, following Galí and Monacelli (2005) and Justiniano and Preston (2010), in order to capture changes in the magnitude of fundamental shocks as well as monetary policy responses.[2] Consequently, our MS-DSGE model allows for regime shifts in the monetary policy rule coefficients as well as the shock volatilities. We employ the algorithm in Farmer et al. (2011) to solve the MS-DSGE

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* Corresponding author. Tel.: + 82 2 759 4488.
E-mail addresses: irobot@bok.or.kr (J. Choi), joonyhur@gmail.com (J. Hur).

[1] The consequences of these economic upheavals are detailed in Kim (2009) and Yoon (2011).

[2] MS-DSGE models associated with monetary policy switches have been extensively studied for the US (e.g., Sims and Zha, 2006; Davig and Doh, 2008; Bianchi, 2013) and UK (Liu and Mumtaz, 2011) economies. To the best of our knowledge, however, the macroeconomic effects of introducing inflation targeting in small open economies have been explicitly examined mostly with fixed-coefficient DSGE models, e.g., for Chile (Del Negro and Schorfheide, 2009) and for Brazil (Palma and Portugal, 2014). As an exception, Cúrdia and Finocchiaro (2013) exploit a regime switching DSGE model for Sweden to measure the potential effects due to the transition from exchange rate targeting to inflation targeting.

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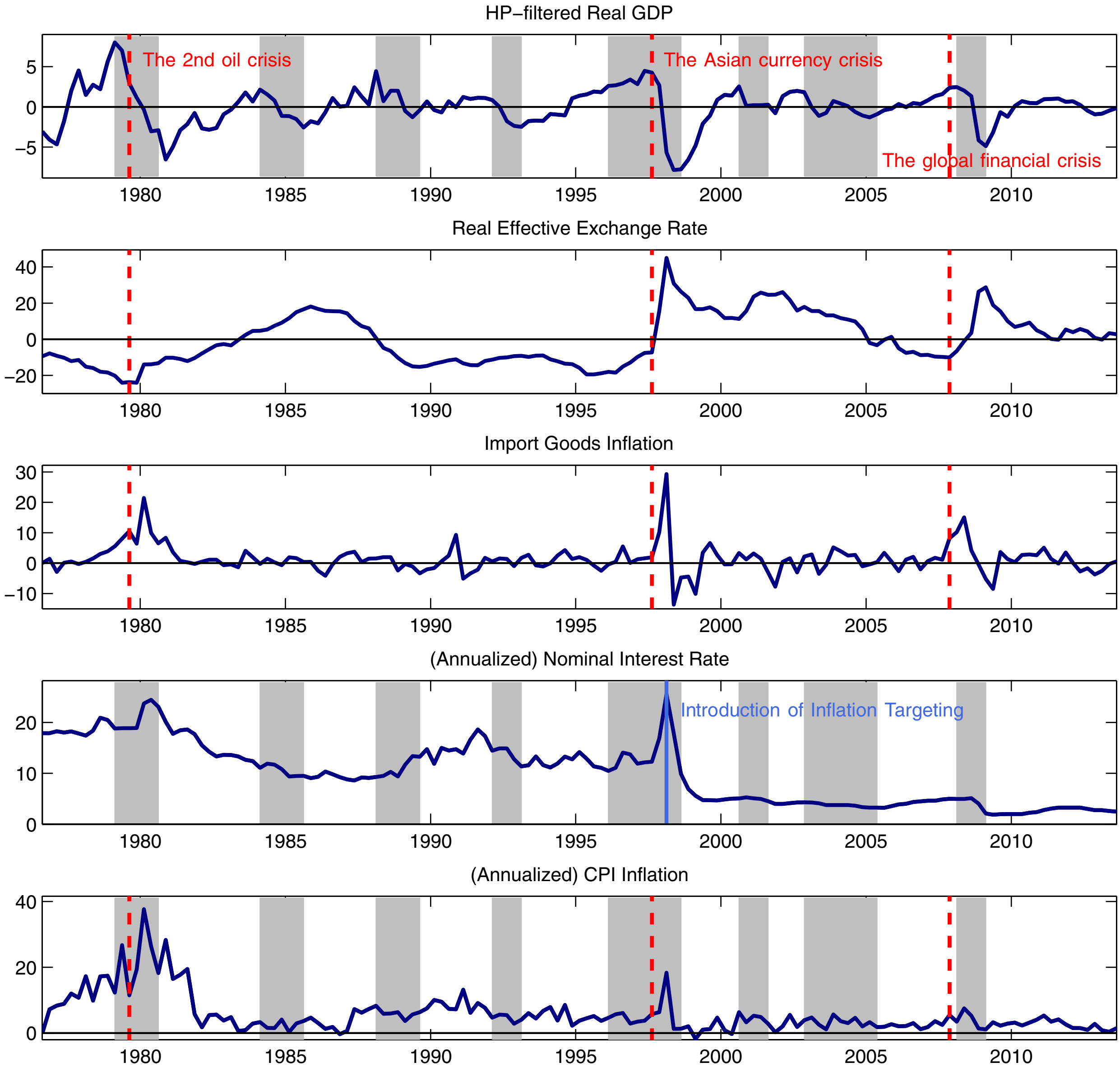

Fig. 1. Time series of Korean data. The shaded areas indicate the recession dates identified by the Korea National Statistical Office (KOSTAT). The vertical lines denote important macroeconomic events for Korean economy.

model, and estimate the model using Bayesian methods, based on Korean data ranged from 1976:Q3 to 2013:Q3. We estimate two versions of the model allowing for regime shifts in (i) the shock volatilities only, and (ii) both the shock volatilities and the monetary policy rule coefficients.[3] By doing so, we aim to judge which specification is the most preferred by data. To this end, we compute marginal data densities of the specifications, and make an explicit comparison of model fit, including a conventional no regime switching DSGE counterpart.

According to the estimation results, our MS-DSGE models' fit with Korean data outperforms the fixed parameter DSGE counterpart. This finding suggests that the MS-DSGE models are more suitable to investigate the structural changes in Korean economy. Among the MS-DSGE specifications, data strongly prefers the specification that allows for regime shifts in both the shock volatilities and the monetary policy rule coefficients. This finding reveals that the regime-switching aspect both in monetary policy rules and shock volatilities plays a crucial role in improving the model's fit with the data.

The regime probability estimates for the best-fitting model characterize four different regimes for Korean economy that are formulated as a combination of "*high*"/"*low*" volatility regimes and "*IT*"/"*Non-IT*" monetary regimes. Regarding the shock volatilities, our regime estimates perform quite well in capturing the major high volatile episodes in domestic output and CPI inflation, which coincide with the 1979 Oil crisis, the 1990–91 Gulf War, the 1997–98 Asian currency crisis, and the 2008–09 Global financial crisis. This finding suggests that a large fluctuation in the Korean economy tends to heavily depend upon external shocks. Tuning to the identified policy regimes, there is strong evidence that the monetary authority responds more aggressively to inflation, but less strongly to output, after launching the IT in the late 1990s. Meanwhile, its responsiveness to exchange rate only moderately chances across the two regimes.

Three main findings stand out based on the identified regimes. First, our impulse response analysis allowing for policy regime shifts lends

[3] Another possible specification is one that allows regime changes in the monetary policy coefficients only, while the heteroskedasticity of shocks is excluded. However, we do not pursue this avenue in this paper based on the following reasoning. As Sims and Zha (2006) illustrate, an absence of the heteroskedasticity may cause statistic biases in the policy coefficients. In addition, we find that incorporating regime-dependent shocks substantially enhances the data fit.

some support to the view that Korea's monetary policy might have contributed to a sharp reduction in the level as well as the volatility of inflation in the 2000s. This finding reconfirms the conclusion of the existing literature such as Kim and Park (2006) and Sánchez (2009). Second, our shock decomposition analyses suggest that technology shocks are the most important driver of output fluctuations in Korea. Indeed, we observe that both the Asian currency crisis and the global financial crisis are mainly explained by substantial drops in technology shocks. Regarding inflation dynamics, monetary policy and preference shocks are the main drivers over the sample period. Finally, a counterfactual exercise shows that it would have been possible to achieve higher output and lower inflation simultaneously if the "*IT*" monetary policy regime were maintained over the entire sample period.

The remainder of this article is organized as follows: Section 2 provides the overview of our DSGE model structure. Section 3 explains the solution method of the MS-DSGE model. Section 4 presents our Bayesian estimation procedure, followed by Section 5, providing a variety of policy discussions based on our MS-DSGE estimation results. Section 6 concludes.

2. The estimated model

The model follows Justiniano and Preston (2010), which is a small open-economy (SOE) new Keynesian (NK) model which extends the models in Monacelli (2005) and Galí and Monacelli (2005) by augmenting incomplete asset markets, habit formation and indexation of prices to past inflation.

2.1. Households

A representative household maximizes its utility function given by

$$\mathbb{E}_0\sum_{t=0}^{\infty}\beta^t\varepsilon_{g,t}\left[\frac{(C_t-hC_{t-1})^{1-\sigma}}{1-\sigma}-\frac{N_t^{1+\varphi}}{1+\varphi}\right],$$

where β is the discount factor, h is the external habit parameter, C_t is a composite consumption index, N_t is the labor input, $\sigma > 0$ and $\varphi > 0$ denote the inverses of intertemporal elasticity of substitution and Frisch labor supply elasticity respectively. The variable $\varepsilon_{g,t}$ is a general preference shock that follows

$$\varepsilon_{g,t}=\bar{\varepsilon}_g(\varepsilon_{g,t-1}/\bar{\varepsilon}_g)^{\rho_g}\exp\left[\sigma_g\left(\xi_t^Q\right)\epsilon_{g,t}\right],\epsilon_{g,t}\sim\mathbb{N}(0,1),$$

where ξ_t^Q is an unobservable state variable which governs the volatility regime at time t, and $\bar{\varepsilon}_g$ is the steady-state preference.

The consumption index C_t is given by

$$C_t=\left[(1-\alpha)^{\frac{1}{\eta}}C_{H,t}^{\frac{\eta-1}{\eta}}+\alpha^{\frac{1}{\eta}}C_{F,t}^{\frac{\eta-1}{\eta}}\right]^{\frac{\eta}{\eta-1}},$$

where $C_{H,t}$ and $C_{F,t}$ are the domestic and foreign produced goods which can be written by using a Dixit and Stiglitz (1977) aggregator as

$$C_{H,t}=\left[\int_0^1 C_{H,t}(i)^{\frac{\theta-1}{\theta}}di\right]^{\frac{\theta}{\theta-1}}\quad\text{and}\quad C_{F,t}=\left[\int_0^1 C_{F,t}(i)^{\frac{\theta-1}{\theta}}di\right]^{\frac{\theta}{\theta-1}},$$

where α is the degree of openness parameter, $\eta > 0$ is the elasticity of substitution between domestic and foreign goods, and $\theta > 1$ is the elasticity of substitution between domestic and foreign goods i.

The household's choices are constrained by

$$P_tC_t+D_t+\tilde{e}_tB_t=D_{t-1}\left(1+\tilde{i}_{t-1}\right)+\tilde{e}_tB_{t-1}\left(1+\tilde{i}^*_{t-1}\right)\phi_t(A_t)+W_tN_t+\Pi_{H,t}+\Pi_{F,t}+T_t,\tag{1}$$

where D_t and B_t denote the household's holding of one-period domestic and foreign bonds, respectively, with corresponding interest rates $\tilde{i}_t$ and $\tilde{i}^*_t$, $\tilde{e}_t$ is the nominal exchange rate, P_t is the domestic CPI, and W_t is the nominal wage of labor supply. $\Pi_{H,t}$ and $\Pi_{F,t}$ denote profits from holding shares in domestic and imported goods firms respectively, and T_t denotes lump-sum transfers. The function $\phi_t(\cdot)$ is interest rate premium with respect to debt given as

$$\phi_t=\exp[-\chi(A_t+\varepsilon_{rp,t})],$$

where

$$A_t\equiv\frac{\tilde{e}_{t-1}B_{t-1}}{\overline{Y}P_{t-1}}$$

is the real foreign debt outstanding converted in domestic currency as a fraction of domestic steady-state output, and χ is the debt elasticity with respect to the interest rate premium. $\varepsilon_{rp,t}$ is a risk premium shock that follows

$$\varepsilon_{rp,t}=\bar{\varepsilon}_{rp}(\varepsilon_{rp,t-1}/\bar{\varepsilon}_{rp})^{\rho_{rp}}\exp\left[\sigma_{rp}\left(\xi_t^Q\right)\epsilon_{rp,t}\right],\quad\epsilon_{rp,t}\sim\mathbb{N}(0,1),$$

where $\bar{\varepsilon}_{rp}$ is the steady-state risk premium.

The flow budget constraint (1) implicitly assumes that, for all households, nominal income in each period is $W_tN_t+\Pi_{H,t}+\Pi_{F,t}$. This in turn equals $P_{H,t}Y_{H,t}+(P_{F,t}-\tilde{e}P_t^*)C_{F,t}$ in the symmetric equilibrium, where $P_{H,t}$, $P_{F,t}$, and P_t^* denote the domestic goods prices, the domestic currency price of imported goods, and the foreign price, respectively.

The representative household's optimality conditions imply

$$C_{H,t}=(1-\alpha)(P_{H,t}/P_t)^{-\eta}C_t\ \text{and}\ C_{F,t}=\alpha(P_{F,t}/P_t)^{-\eta}C_t\tag{2}$$

$$\lambda_t=\varepsilon_{g,t}(C_t-hC_{t-1})^{-1/\sigma}\tag{3}$$

$$\lambda_t=\varepsilon_{g,t}P_tN_t^{\varphi}/W_t\tag{4}$$

$$\lambda_t\tilde{e}_tP_t=\mathbb{E}\left[\left(1+\tilde{i}_t^*\right)\beta\phi_{t+1}\lambda_{t+1}\tilde{e}_{t+1}P_{t+1}\right]\tag{5}$$

$$\lambda_tP_t=\mathbb{E}\left[\left(1+\tilde{i}_t\right)\beta\lambda_{t+1}P_{t+1}\right]\tag{6}$$

where $P_t=[(1-\alpha)P_{H,t}^{1-\eta}+\alpha P_{F,t}^{1-\eta}]^{\frac{1}{1-\eta}}$ and λ_t is the Lagrange multiplier on the budget constraint.

2.2. Domestic producers

The domestic production sector consists of monopolistically competitive intermediate goods producing firms who produce a continuum of differentiated inputs and a representative final goods producing firm. Each $i\in[0,1]$ in the domestic intermediate goods sector produces a differentiated good, $y_{H,t}(i)$ according to $y_{H,t}(i)=\varepsilon_{a,t}N_t(i)$, where $N_t(i)$ is the firm i's labor input and $\varepsilon_{a,t}$ is a technology shock that follows

$$\varepsilon_{a,t}=\bar{\varepsilon}_a(\varepsilon_{a,t-1}/\bar{\varepsilon}_a)^{\rho_a}\exp\left[\sigma_a\left(\xi_t^Q\right)\epsilon_{a,t}\right],\quad\epsilon_{a,t}\sim\mathbb{N}(0,1),$$

where $\bar{\varepsilon}_a$ is the steady-state technology. Each intermediate firm chooses its labor input to minimize its costs, $W_tN_t(i)$, subject to its production function.

Following Calvo (1983), a randomly chosen fraction $1-\theta_H$ of the domestic intermediate goods firms is allowed to reoptimize their prices every period. Firms that are unable to optimally reset their price partially index their price to past inflation according to

$$\log P_{H,t}(i)=\log P_{H,t-1}(i)+\delta_H\pi_{H,t-1},$$

where $\delta_H \in [0, 1]$ is the degree of indexation to past inflation and $\pi_{H,t} = \log(P_{H,t}/P_{H-1,t})$. Firms that are able to reset their price at t choose their optimal price, $P_{H,t}(i)$, to maximize the expected discounted present value of real profits:

$$\mathbb{E}_t \sum_{k=t}^{\infty} \theta_H^{k-t} Q_{t,k} y_{H,k}(i) \left[P_{H,t}(i) \left(\frac{P_{H,k-1}}{P_{H,t-1}} \right)^{\delta_H} - P_{H,k} MC_k \right],$$

subject to the demand curve

$$y_{H,k}(i) = \left(\frac{P_{H,t}(i)}{P_{H,k}} \cdot \left(\frac{P_{H,k-1}(i)}{P_{H,t-1}} \right)^{\delta_H} \right)^{-\epsilon} \left(C_{H,k} + C^*_{H,k} \right),$$

where Q is the discount factor between periods t and $k > t$, $MC_k = W_k/(P_{H,k}\varepsilon_{a,k})$ is the real marginal cost function for each firm, and $\epsilon > 1$ is the elasticity of substitution between types of differentiated domestic or foreign goods.

The optimality condition of the firm is given by

$$\mathbb{E}_t \sum_{k=t}^{\infty} \theta_H^{k-t} Q_{t,k} y_{H,k}(i) \left[P_{H,t}(i) \left(\frac{P_{H,k-1}}{P_{H,t-1}} \right)^{\delta_H} - \frac{\theta_H}{\theta_{H-1}} P_{H,k} MC_k \right] = 0. \tag{7}$$

2.3. Retail firms

Retail firms import differentiated foreign goods under which the law of one price holds. Retail firms are assumed to follow a Calvo-style price setting behavior augmented with indexation to past inflation. Accordingly, a random fraction $1 - \theta_F$ of firms reoptimize their prices every period, while those who are unable to optimally reset their price partially index their price to past inflation. Firms that are able to reset their price at t choose their optimal price, $P_{F,t}(i)$, to maximize the expected discounted present value of real profits:

$$\mathbb{E}_t \sum_{k=t}^{\infty} \theta_F^{k-t} Q_{t,k} C_{F,k}(i) \left[P_{F,t}(i) \left(\frac{P_{F,k-1}}{P_{F,t-1}} \right)^{\delta_F} - \tilde{e}_k P^*_{F,k}(i) \right],$$

subject to the demand curve

$$C_{F,k}(i) = \left(\frac{P_{F,t}(i)}{P_{F,k}} \cdot \left(\frac{P_{F,k-1}(i)}{P_{F,t-1}} \right)^{\delta_F} \right)^{-\epsilon} C_{F,k}.$$

The optimality condition of the firm is given by

$$\mathbb{E}_t \sum_{k=t}^{\infty} \theta_F^{k-t} Q_{t,k} \left[P_{F,t}(i) \left(\frac{P_{F,k-1}}{P_{F,t-1}} \right)^{\delta_F} - \frac{\theta_H}{\theta_H - 1} \tilde{e}_k P^*_{H,k}(i) \right] = 0. \tag{8}$$

2.4. International risk sharing

Incomplete asset substitution between domestic and foreign bonds yield the uncovered interest rate parity (UIP) condition given by

$$\mathbb{E}_t \lambda_{t+1} P_{t+1} \left[\left(1 + \tilde{i}_t\right) - \left(1 + \tilde{i}^*_t\right) (\tilde{e}_{t+1}/\tilde{e}_t) \phi_{t+1} \right] = 0, \tag{9}$$

and the real exchange rate is defined as $\tilde{q}_t \equiv \tilde{e}_t P^*_t / P_t$.

2.5. Monetary policy

The monetary authority sets policy according to

$$\frac{i_t}{\bar{i}} = \left(\frac{i_{t-1}}{\bar{i}} \right)^{\rho_i(\xi_t^P)} \left[\left(\frac{\pi_t}{\bar{\pi}} \right)^{\lambda_\pi(\xi_t^P)} \left(\frac{Y_t}{\bar{Y}} \right)^{\lambda_y(\xi_t^P)} \left(\frac{\tilde{e}_t}{\tilde{e}_{t-1}} \right)^{\lambda_{de}(\xi_t^P)} \right]^{1-\rho_i(\xi_t^P)} \exp\left[\sigma_i\left(\xi_t^Q\right) \epsilon_{i,t} \right], \tag{10}$$

where λ_π, λ_y, and λ_{de} measure the policy responses to inflation, output, and exchange rate gap respectively, and $\epsilon_{i,t} \sim \mathbb{N}(0, 1)$.

2.6. General equilibrium

The goods market clearing condition is

$$Y_{H,t} = C_{H,t} + C^*_{H,t}, \tag{11}$$

in the domestic economy. To close the model, the foreign demand for the domestically produced goods is assumed as

$$C^*_{H,t} = \left(\frac{P^*_{H,t}}{P^*_t} \right)^{-\lambda} Y^*_t, \tag{12}$$

where $\lambda > 0$.

3. Solution of the MS-DSGE model

Once the deterministic steady-state is calculated, the model can be log-linearized conditional on a particular monetary policy rule regime.[4] Note that the solution of the MS-DSGE model hinges only upon the monetary policy regime, but not upon the stochastic volatility regime. This is due to the usage of the first-order approximation in deriving the equilibrium conditions of households and firms. In addition, the steady-state is independent of regime-shifts in the monetary policy rule (Liu, et al., 2011; Bianchi, 2013).

Let S_t to be the DSGE state vector which contains all the model endogenous variables. Then the log-linearized system can be expressed as

$$\Gamma_0\left(\xi_t^P, \Theta^P\right) S_t = \Gamma_1\left(\xi_t^P, \Theta^P\right) S_{t-1} + \Psi M\left(\xi_t^Q, \Theta^Q\right) \epsilon_t + \Pi \eta_t, \tag{13}$$

where Θ^P and Θ^Q denote the regime-dependent structural parameters and shock standard deviations, respectively. The vector ϵ_t contains all the exogenous shocks defined in the previous section, and η_t is the vector of the expectations errors. Under the presence of potential regime switching, the model solution for the form (13) cannot be obtained via the standard solution methods for linearized rational expectations models, including the algorithm in Sims (2002). Instead, we solve the system using the minimum state variable (MSV) solution algorithm proposed by Farmer et al. (2011).[5] If a solution exists, the output of the solution algorithm is expressed in a regime-switching vector autoregression form:

$$S_t = T\left(\xi_t^P, \Theta^P, H^P\right) S_{t-1} + R\left(\xi_t^P, \Theta^P, H^P\right) M\left(\xi_t^Q, \Theta^Q\right) \epsilon_t, \tag{14}$$

where H^P is the probability of moving across difference structural parameter regimes. We impose a structure on H^P (and H^Q) by assuming that there are two unobserved regimes associated with the structural parameters and shock volatilities, respectively. We further assume

[4] The description of the entire log-linearized system is provided in Appendix A.

[5] Alternative solution algorithms are demonstrated in Davig and Leeper (2007) and Cho (2010).

that the two state variables ξ_t^P and ξ_t^Q follow a first-order Markov chain with the following transition probability matrices:

$$H^P = \begin{bmatrix} P_{11} & P_{12} \\ P_{21} & P_{22} \end{bmatrix} \text{ and } H^Q = \begin{bmatrix} Q_{11} & Q_{12} \\ Q_{21} & Q_{22} \end{bmatrix},$$

where $P_{ij} = \text{Prob}\ (\xi_t^P = j | \xi_{t-1}^P = i)$ and $Q_{ij} = \text{Prob}\ (\xi_t^Q = j | \xi_{t-1}^Q = i)$.

Let X_t to be the observable data used for the estimation. Then the measurement equation is given by

$$X_t = ZS_t \tag{15}$$

where Z is a matrix that maps the MS-DSGE law of motion in Eq. (14) into the observable variables.

4. Estimation of the MS-DSGE model

For the estimation of the model described above, we use Bayesian inference methods to construct the parameters' posterior distribution, which is a combination of the likelihood function and prior information.

4.1. Data

We use 8 observable variables including domestic real GDP, domestic real effective exchange rate (REER), domestic import goods inflation, domestic quarterly nominal interest rates, domestic CPI inflation, foreign quarterly nominal interest rate, foreign CPI inflation, and foreign real GDP ranged from 1976:Q3 to 2013:Q3 for the estimation. Foreign data use U.S. time series. Appendix B provides a detailed description of the data. We demean each time series, except for domestic and foreign GDP variables which are detrended by the Hodrick–Prescott filter.[6]

It is worthwhile mentioning that the starting date of the sample period is chosen to utilize the longest span of Korean data available. One prominent benefit of MS-DSGE models is to let the data speak about the timing of regime shifts, which requires no judgment calls for splitting the sample. In contrast, the existing DSGE literature on the Korean economy tends to restrict the sample period in order to avoid the potential regime-switching issue. For example, Elekdag et al. (2006) consider the post-1990 sample which corresponds to the abolishment of pegged exchange rates. To be robust, we also estimate the model by using the sample from 1990 and find that the main empirical results which will be presented below – posterior estimates for the parameters and regime probabilities, and model fit – are only mildly altered by a selection of the sample span.[7]

Before proceeding, we attempt to seek for evidence on how the volatility of Korean macroeconomic time series evolve over time. To this end, we begin by a CUSUM of squares test on the residuals from an AR(1) specification for the series. Fig. 2 plots the test statistics. The CUSUM of squares test illustrates that the Asian currency crisis is the most evident event causing instability in the variance of GDP, exchange rate, and import goods inflation. Table 1 demonstrates the volatility and normality of the key macroeconomic times series of Korea, prior to and following the introduction of the Inflation Targeting policy followed by the Asian currency crisis. Regarding volatility, the post-IT sample is associated with dramatic reduction in the variance of the key macroeconomic variables, which is more pronounced for GDP, nominal interest rate, and CPI inflation. As the p-values for the Kolmogorov–Smirnov normality test show, exchange rate, import goods inflation and CPI inflation in the pre-IT sample do not tend to follow a normal distribution. On a contrary, the normality cannot be rejected for all the variables in the post-IT sample.

4.2. Prior distributions

We calibrate several parameters that are difficult to identify from the data. The subjective discount factor, β, is set to 0.99, which implies an annual steady-state real interest rate of 4%. Following Justiniano and Preston (2010) and Liu and Mumtaz (2011), the degree of openness, α, is approximated by the average share of imports and exports to GDP in Korea over the sample span considered. Accordingly, we set α to be 0.35. Finally, the debt elasticity with respect to the interest rate premium, χ, is set to be 0.01 as in Justiniano and Preston (2010).

The second and third columns of Tables 3 and 4 display the prior distribution for all estimated parameters.[8] The priors for the parameters are largely drawn from Justiniano and Preston (2010) and Liu and Mumtaz (2011). The prior for the inverse intertemporal elasticity of substitution, σ, and the inverse of Frisch elasticity of labor supply, ω, parameters are drawn from Justiniano and Preston (2010) so that they follow Gamma distribution of means 1.2 and 1.5, and standard deviations 0.40 and 0.75, respectively. Priors for the Calvo price parameters, θ_H and θ_F, assume to follow Beta distribution centered at 0.5 with standard deviation of 0.1. A fairly diffuse prior is imposed on the elasticity of substitution between domestic and foreign goods, which follows Gamma distribution with mean of 1.5 and standard deviation of 0.75. Habit, indexation, and monetary policy rule autoregressive parameters also have quite diffuse priors to reflect the lack of *a priori* knowledge of these parameters.

Priors for monetary policy responses to inflation, output, and exchange rate are drawn from Liu and Mumtaz (2011). The prior for the nominal interest rate reaction to inflation, λ_π, is assumed to be independent of the monetary policy regimes. For both regimes, it follows Gamma distribution of mean 1.5 and standard deviation 0.25. The nominal interest rate reactions to output and exchange rate, λ_y and λ_{de}, are assumed to follow Gamma distribution of mean 0.25 and standard deviation 0.13. The exogenous shock autocorrelation parameters, ρ's, follow Beta distribution of mean 0.5 and standard deviation 0.15. We assume inverse Gamma distribution with mean of 0.5 and standard deviation of 10 for the standard deviation of the exogenous shock process.

Finally, priors for the regime switching probability parameters impose two conditions: non-negativity and sum-to-one constraints. To satisfy these constraints, we assign Dirichlet prior distributions. The parameters that govern the probability of switches in structural parameter regimes, P_{11} and P_{22}, and shock volatility regimes, Q_{11} and Q_{22}, are assumed to follow a Dirichlet prior of mean 0.96 and standard deviation 0.04.

4.3. Estimation procedure

Unlike fixed coefficient models, the standard Kalman filter is not applicable for a likelihood evaluation of MS-DSGE models whose transition dynamics is given as the form in (14). Instead, we use the modified Kalman filter proposed in Kim and Nelson (1999). This algorithm tracks only a limited number of states based upon their weight, which is given by the probability assigned to each path from the filter described in Hamilton (1989).

Together with the model with time-varying monetary policy rule and volatilities, we estimate two additional specifications for comparison,

[6] We use the HP filter to estimate *output gap* with the business cycle interpretation as the deviation of actual output from its potential level. Although no single solution would exist for extracting business cycle fluctuations as pointed out in Canova (1998), the HP filter is one of the most popular detrending methods in the exiting literature (e.g., Taylor, 1999; Orphanides and Van Norden, 2002). Hence, we have selected the HP filter to follow influential works in the DSGE literature as in Clarida et al. (2000), Smets and Wouters (2003) and Lubik and Schorfheide (2004), to name a few. In addition, our robustness checks have shown that using alternatives detrending methods (e.g., removing a quadratic trend) would not draw different conclusions.

[7] The supplementary online appendix C provides the estimation results based on the post-1990 sample.

[8] We additionally impose the lower and upper bounds of [0,15] for the shock standard deviation parameters. A similar approach is employed in Liu and Mumtaz (2011).

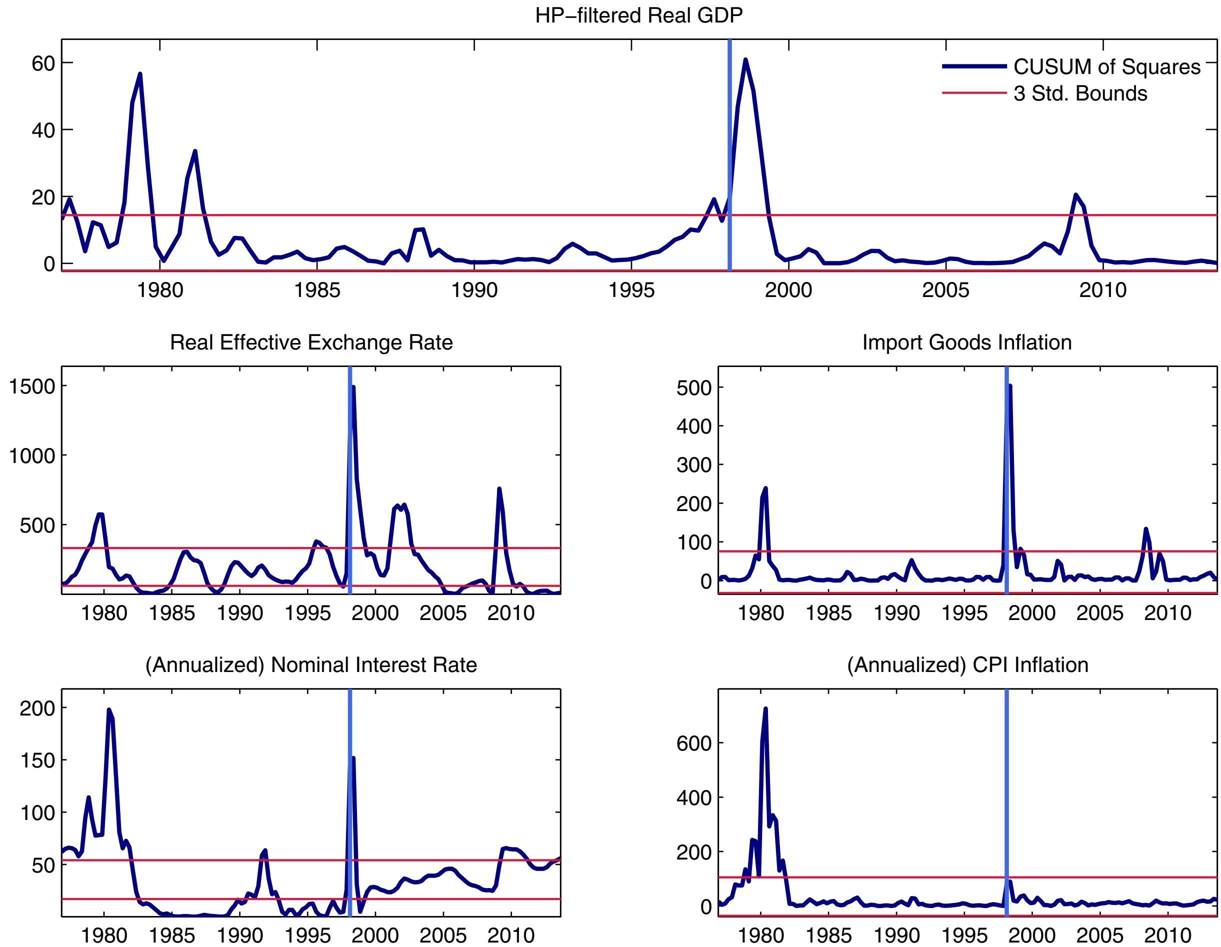

Fig. 2. CUSUM of squares test statistics for time series of Korean data. The thick and thin solid lines denote the CUSUM of squares and its ±3 standard error bounds. The vertical lines denote the introduction of the Inflation Targeting policy.

the conventional model with no regime switching and the model allowing for regime shifts in volatility only. By doing so, we let the data speak about the best SOE NK-DSGE specification in analyzing Korean data.

Table 1
Normality and volatility of time series of Korean data, before and after the introduction of the Inflation Targeting policy.

	Real GDP	Real exchange rate	Import goods inflation	Nominal interest rate	CPI inflation
p-Values for the Kolmogorov–Smirnov normality test					
Pre-IT	0.79	0.00	0.00	0.38	0.00
Post-IT	0.44	0.83	0.35	0.80	0.88
Standard deviations					
Pre-IT	2.94	13.45	5.08	4.03	6.84
Post-IT	1.57	10.60	4.07	1.02	1.76

Table 2
Average log marginal density for each estimated model.

Model	Average log marginal density
Fixed coefficient	−1923.13
Regime switch in volatility only	−1840.91
Regime switch in Taylor rule coefficients and volatility	−1803.48

We use the Sims's optimization routine *csminwel* to maximize the log posterior function, which combines the priors and the likelihood of the data. We then implement the random walk Metropolis–Hastings (MH) algorithm and simulate 50000 draws. We compute medians and the covariance matrix of the initial 50000 draws and update the initial value of the MH chain and covariance matrix of the proposal density. Then, we simulate additional 200000 draws, with the first 100000 used as a burn-in period and every 20th thinned, leaving a sample size of 5000.

5. Empirical results

5.1. Model fit

Table 2 summarizes model fit measured by the posterior log-likelihood.[9] In general, the models allowing for regime shifts fit the data significantly better than the fixed coefficient counterpart. The fixed coefficient model is the least preferred specification with a

[9] The posterior log-likelihood uses the average of Geweke's (1999) modified harmonic mean estimator. Sims et al. (2008) argue that the estimator proposed by Geweke (1999) performs poorly for the models with time-varying parameters whose posterior density tend to be non-Gaussian. However, Davig and Doh (2008) use the estimator, illustrating that a drawback associated with the estimator is its sensitivity to the scaling parameter of the covariance matrix of the proposal density. Consequently, we recalculate the posterior density using various values for the scaling parameter, and find that the ranking of model fit remains unaltered by the choice of a scaling parameter. More details on this issue are provided in the supplementary online appendix C.

Table 3
Prior and posterior distributions of each parameter. This table reports the median and associated [2.5%, 97.5%] percentile intervals (in brackets).

Parameter	Prior		Posterior				
	Dist.	Mean (std)	Fixed coeff.	Vol. only		Talyor rule & vol.	
				Regime 1	Regime 2	Regime 1	Regime 2
σ	G	1.2 (0.40)	0.45	0.08		0.04	
(Risk aversion)		[0.55, 2.10]	[0.24, 0.69]	[0.04, 0.13]		[0.02, 0.06]	
φ	G	1.5 (0.75)	1.26	4.29		3.92	
(Inverse Frisch elasticity)		[0.41, 3.29]	[0.72, 1.82]	[3.04, 6.34]		[2.93, 5.46]	
θ_H	B	0.5 (0.10)	0.86	0.25		0.21	
(Calvo domestic prices)		[0.31, 0.69]	[0.84, 0.88]	[0.14, 0.36]		[0.13, 0.31]	
θ_F	B	0.5 (0.10)	0.61	0.66		0.64	
(Calvo import prices)		[0.31, 0.69]	[0.53, 0.71]	[0.60, 0.73]		[0.55, 0.72]	
η	G	1.5 (0.75)	0.21	0.65		0.68	
(Elasticity home–foreign goods)		[0.41, 3.29]	[0.18, 0.23]	[0.51, 1.05]		[0.51, 1.12]	
h	B	0.5 (0.25)	0.91	0.13		0.07	
(Consumption habit)		[0.06, 0.94]	[0.86, 0.95]	[0.02, 0.30]		[0.01, 0.20]	
δ_H	B	0.5 (0.25)	0.02	0.13		0.16	
(Backward indexation, domestic)		[0.06, 0.94]	[0.00, 0.06]	[0.02, 0.52]		[0.02, 0.53]	
δ_F	B	0.5 (0.25)	0.03	0.02		0.01	
(Backward indexation, foreign)		[0.06, 0.94]	[0.00, 0.10]	[0.00, 0.06]		[0.00, 0.03]	
ρ_i	B	0.5 (0.25)	0.81	0.49		0.35	0.43
(Taylor rule AR(1))		[0.06, 0.94]	[0.77, 0.85]	[0.37, 0.59]		[0.14, 0.58]	[0.31, 0.55]
λ_π	G	1.5 (0.25)	0.50	1.64		2.19	0.91
(Taylor rule inflation)		[1.05, 2.03]	[0.40, 0.59]	[1.42, 1.95]		[1.91, 2.53]	[0.82, 0.97]
λ_y	G	0.25 (0.13)	0.13	0.03		0.02	0.05
(Taylor rule output)		[0.06, 0.56]	[0.08, 0.20]	[0.01, 0.06]		[0.01, 0.05]	[0.02, 0.12]
λ_{de}	G	0.25 (0.13)	0.18	0.06		0.04	0.03
(Taylor rule exchange rate)		[0.06, 0.56]	[0.12, 0.24]	[0.02, 0.11]		[0.01, 0.09]	[0.01, 0.07]
ρ_a	B	0.5 (0.15)	0.96	0.74		0.71	
(Technology shock AR(1))		[0.21, 0.79]	[0.94, 0.97]	[0.65, 0.83]		[0.63, 0.78]	
ρ_g	B	0.5 (0.15)	0.29	0.91		0.98	
(Preference shock AR(1))		[0.21, 0.79]	[0.18, 0.39]	[0.87, 0.93]		[0.96, 0.99]	
ρ_{rp}	B	0.5 (0.15)	0.92	0.88		0.89	
(Risk premium shock AR(1))		[0.21, 0.79]	[0.90, 0.94]	[0.84, 0.93]		[0.86, 0.94]	
ρ_{cp}	B	0.5 (0.15)	1.00	0.67		0.64	
(Import cost push shock AR(1))		[0.21, 0.79]	[0.99, 1.00]	[0.53, 0.78]		[0.49, 0.77]	

Table 4
Prior and posterior distributions of each parameter (continued). This table reports the median and associated [2.5%, 97.5%] percentile intervals (in brackets).

Parameter	Prior		Posterior				
	Dist.	Mean (std)	Fixed coeff.	Vol. only		Talyor rule & vol.	
				Regime 1	Regime 2	Regime 1	Regime 2
ρ_{π^*}	B	0.5 (0.15)	0.76	0.77		0.76	
(Foreign inflation shock AR(1))		[0.21, 0.79]	[0.65, 0.84]	[0.71, 0.82]		[0.70, 0.82]	
ρ_{y^*}	B	0.5 (0.15)	0.80	0.49		0.49	
(Foreign output shock AR(1))		[0.21, 0.79]	[0.74, 0.85]	[0.37, 0.63]		[0.37, 0.62]	
ρ_{i^*}	B	0.5 (0.15)	0.89	0.85		0.85	
(Foreign IntRate shock AR(1))		[0.21, 0.79]	[0.85, 0.91]	[0.79, 0.91]		[0.80, 0.90]	
σ_{π^*}	IG	0.5 (10)	0.55	0.36	1.77	0.41	0.76
(Foreign inflation shock std.)		[0.09, 2.07]	[0.50, 0.62]	[0.28, 0.51]	[1.07, 3.52]	[0.31, 0.59]	[0.51, 1.20]
σ_{y^*}	IG	0.5 (10)	0.69	1.01	8.01	0.41	1.34
(Foreign output shock std.)		[0.09, 2.07]	[0.62, 0.77]	[0.63, 1.63]	[4.45, 13.34]	[0.30, 0.60]	[0.60, 2.74]
σ_{i^*}	IG	0.5 (10)	0.27	0.64	2.20	0.52	1.58
(Foreign IntRate shock std.)		[0.09, 2.07]	[0.24, 0.31]	[0.48, 0.86]	[1.30, 3.87]	[0.39, 0.72]	[0.99, 2.68]
σ_a	IG	0.5 (10)	1.67	2.88	14.18	3.06	14.36
(Technology shock std.)		[0.09, 2.07]	[1.23, 2.24]	[1.89, 4.50]	[11.55, 14.95]	[2.06, 4.63]	[12.30, 14.96]
σ_{mp}	IG	0.5 (10)	0.35	0.33	1.32	0.33	1.33
(Monetary policy shock std.)		[0.09, 2.07]	[0.31, 0.40]	[0.26, 0.45]	[0.80, 2.26]	[0.26, 0.44]	[0.81, 2.23]
σ_g	IG	0.5 (10)	14.73	0.21	0.35	0.21	0.35
(Preference shock std.)		[0.09, 2.07]	[13.62, 14.99]	[0.18, 0.26]	[0.27, 0.49]	[0.18, 0.26]	[0.28, 0.48]
σ_{rp}	IG	0.5 (10)	0.60	0.28	0.69	0.28	0.69
(Risk premium shock std.)		[0.09, 2.07]	[0.44, 0.80]	[0.23, 0.35]	[0.49, 1.03]	[0.23, 0.34]	[0.50, 1.03]
σ_{cp}	IG	0.5 (10)	2.46	0.34	0.77	0.32	0.78
(Import cost push shock std.)		[0.09, 2.07]	[1.71, 3.42]	[0.27, 0.43]	[0.54, 1.15]	[0.26, 0.43]	[0.54, 1.13]
P_{11}	D	0.96 (0.04)				0.94	
(Prob. of coeff. regime 1)		[0.86, 1.00]				[0.91, 0.96]	
P_{22}	D	0.96 (0.04)					0.95
(Prob. of coeff. regime 2)		[0.86, 1.00]					[0.89, 0.99]
Q_{11}	D	0.96 (0.04)		0.96		0.96	
(Prob. of vol. regime 1)		[0.86, 1.00]		[0.91. 0.99]		[0.91, 0.99]	
Q_{22}	D	0.96 (0.04)			0.92		0.91
(Prob. of vol. regime 2)		[0.86, 1.00]			[0.81, 0.97]		[0.81, 0.97]

substantial margin in the posterior log-likelihood. The best-fitting specification is one that allows regime shifts in the monetary policy rule and shock volatilities. Next in the ordering comes the model with regime shifts in the volatilities only.

This finding advocates that the monetary policy regime switching aspect is a crucial setup in conducting policy analyses associated with the Korean economy. Indeed, the Bank of Korea has adopted Inflation Targeting on April 1998 as an alternative to the monetary aggregate targeting regime, seeking to unravel the decline in the effectiveness of monetary indicators since the mid-1990s (The Bank of Korea, 2012). Furthermore, our model selection allowing for regime switching lends some support to the presence of potential breaks in Korea's economic structures around the Asian currency crisis as argued in previous studies. In particular, Cho (2007) suggests for explicitly considering the crisis to assess the true economic properties underlying the Korean economy.

5.2. Parameter estimates and regime probabilities

Posterior median and 95% interval estimates of the model parameters are summarized in Tables 3 and 4.[10] Overall, the posterior estimates are substantially inconsistent across the fixed coefficient and Markov-switching models. As illustrated above, however, the fixed coefficient model is the least favored specification by the data. Rather, we focus on the posterior estimates for the best-fitting model, the model allowing for regime shifts in monetary policy rules as well as the shock volatilities.

Our estimate of the inverse of intertemporal elasticity of substitution parameter is somewhat smaller than the values reported in previous studies. For example, Elekdag et al. (2006) obtain the median parameter value of 0.74 by estimating a SOE NK-DSGE model using the Korean data spanned from 1990 to 2003. Meanwhile, our estimate of the inverse Frisch elasticity has the median of 3.92, which is relatively larger than the median of 1.89 estimated in Elekdag et al. (2006). The domestic Calvo parameter estimate is 0.21 in terms of the median value, which implies that firms reoptimize prices approximately every 1.3 quarters. This estimate is lower than the values estimated or calibrated for the advanced economies such as the United States and the Euro area (e.g., Smets and Wouters, 2005; Smets and Wouters, 2007; Justiniano and Preston, 2010). Instead, our estimate is relatively close to (Elekdag et al., 2006), who report 0.39 for the parameter estimate. Prices in the imported goods sector are adjusted less frequently than home goods prices, being reoptimized on average every 2.8 quarters.

The median estimate for the elasticity of substitution between domestic and foreign goods is 0.68, and the 2.5th–97.5th percentile interval is [0.51, 1.12]. It turns out that the data suggests a quite weak evidence of consumption habit formation as it characterizes the parameter estimates of [0.01, 0.20] in terms of the 95% interval. Also, the backward indexation parameters for both domestic and imported goods sectors appear to play a limited role in generating endogenous persistence in the dynamics of inflation in Korea, by having median values of 0.16 and 0.01 respectively.

Regarding the historical behavior of monetary policy in Korea, several observations can be made when comparing the estimates across the two regimes. And it turns out that the primary differences between these regimes emerge from the degree of inflation and output responsiveness, while they differ only marginally with respect to exchange rate. More specifically, under Regime 1, the degree of interest rate smoothing is slightly higher than that of Regime 2, with a more dispersed posterior interval. The median responsiveness to inflation is 2.19, which is substantially larger than the value of 0.91 in the second regime. The response to output is somewhat lower under Regime 1. Regarding the responsiveness to exchange rate, the median estimates illustrate that Regime 1 reacts slightly more actively to exchange rate than Regime 2. In sum, Regime 1 is characterized as one that targets inflation and exchange rate relatively more strongly, but output less actively than Regime 2. Note that the parameter estimates associated with the monetary regime 1 are consistently lower than the results in Elekdag et al. (2006), whose median estimates for the inflation, output, and exchange rate coefficients of 6.46, 0.15, and 0.11, respectively. The probabilities of monetary policy regimes 1 and 2, P_{11} and P_{22}, suggest that both regimes are quite persistent, as they have the median values of 0.94 and 0.95 respectively.

Throughout this work, we refer to Regime 1 and 2 as the "*IT*" and "*Non-IT*" monetary policy regimes, respectively. As will be illustrated formally below, the estimation results reveal that the monetary regime 1 prevails over the sample after the Asian currency crisis, in which the period largely overlaps with the monetary authority's Inflation Targeting behavior. Regime 2, on the contrary, dominates most of the pre-crisis sample up to 1998, with a few exceptions of only short deviations from it.[11]

Tuning to the shock processes, the preference and risk premium disturbances are highly persistent, having the median autoregressive coefficients of 0.98 and 0.89, respectively. There are two regimes in the shock volatilities characterized by the data. The estimated standard deviations of exogenous shocks are consistently higher under the volatility regime 2 than the first one. We refer to the first and second volatility regimes as the *low* and *high* volatility regimes, respectively. For both volatility regimes, it is worth mentioning that a technology shock is the most volatile component among the exogenous shocks. Finally, the probabilities of the volatility regimes, Q_{11} and Q_{22}, reveal that the low volatility regime is a lot more persistent than the high volatility regime.

The top panel of Fig. 3 shows the median and 95% interval estimates for the probability of the *IT* monetary regime. We observe that the parameter regime estimates for the post-Asian currency crisis are in sharp contrast to those for the pre-crisis period in both the magnitude and the precision. More specifically, our probability estimates suggest that until the late 1990s the monetary policy rule parameters tend to frequently switch across the two regimes with relatively low precisions. In contrast, throughout the 2000s the policy regime 1 appears fairly dominant with high precisions. We might attribute such a discrepancy in our probability estimates to the introduction of Inflation Targeting by the Bank of Korea in the late 1990s.

To assess our regime probability estimates, we present the estimated monetary policy regimes in the bottom panel of Fig. 3. Note that we identify the *IT* regime as the periods when the estimated probability of policy regime 1 exceeds 50%. Our identification gives rise to the *IT* regime dominance in the 2000s. From a broad perspective, this is consistent with the assessment of Cargill (2010), which provides a historical overview of Korea's monetary policy. One of the main findings of the paper is that until the mid-1990s Korea's monetary policy was most likely to be constrained by government-directed industrial policy. However, the study attributed improved monetary policy outcomes, especially in terms of price stability in the 2000s, to the institutional changes in December 1997 and August 2003, including the launch of Inflation Targeting.

From a narrow perspective, on the contrary, it seems that our identification fails to capture two remarkable regime change episodes. First, the period 1983–85 could be labeled as the *IT* monetary regime as the Bank of Korea had consecutively lowered its money supply target

[10] We conduct convergence diagnostic tests for the posterior distributions of the estimated parameters for both specifications. The convergence diagnostics use Geweke's chi-squared test for two sets of MCMC sample draws of the posterior distributions. The statistics for Geweke's chi-squared test show that most of the parameter draws for the MS-DSGE models converge to the stationary distribution under a significant level of 0.1. The supplementary online appendix C provides the convergence statistics.

[11] Regime 1 is also observed in the pre-1998 sample, and this finding is somewhat inconsistent with the monetary policy behavior by the Bank of Korea, whose main target is to control monetary aggregates. One potential source of this result ascribes to allowing for only two different monetary policy regimes, even though there would have been more regimes with distinct stances of monetary policy.

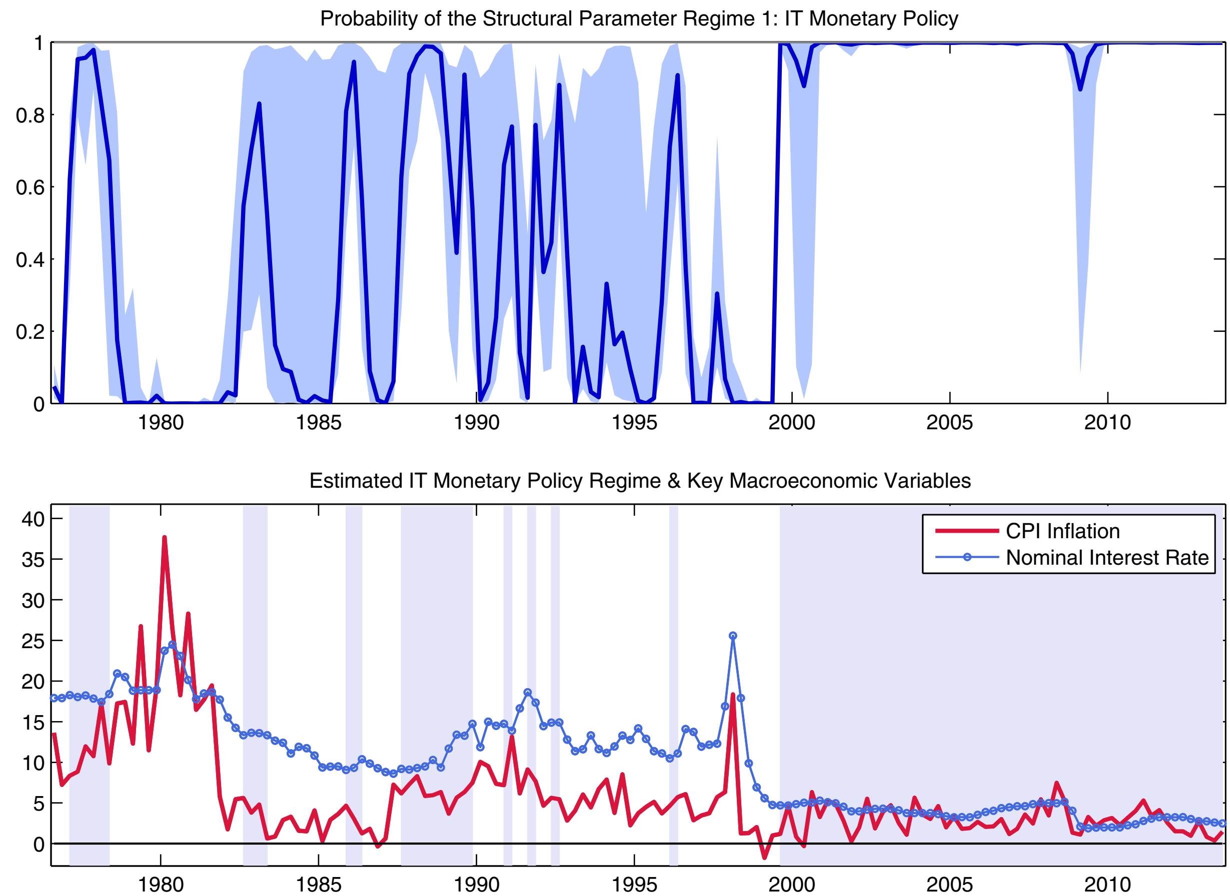

Fig. 3. [Top panel] Probabilities of the structural parameter regime 1 (*IT* monetary policy). The posterior median and corresponding 95% interval estimates are reported. [Bottom panel] Estimated *IT* monetary policy regime together with the key macroeconomic variables. The shaded areas correspond to the periods in which the probability of the structural parameter regime 1 exceeds 50% in terms of the median estimate.

for M2 growth to contain high inflations (The Bank of Korea, 2010). Note that our policy regime probability for this period is poorly estimated with fairly low precisions. Second, our identification using the median threshold is not consistent with previous studies, such as Oh (2014), regarding the three easing cycles of 2001, 2003–04, and 2008–09. We argue that our policy regime estimates mildly respond to the easing cycles of 2001 and 2008–09 in the sense that the precisions for such periods become exceptionally lower while failing to capture the 2003–04 easing cycle.

The top panel of Fig. 4 displays the median and 95% interval estimates for the probability of high volatility regimes. As with the policy regime, we also present the estimated volatility regimes in the bottom panel of Fig. 4 in which the high volatility regime is depicted as shaded area. Based on both panels, we argue that the volatility regime estimates perform quite well in capturing high volatile episodes in domestic output and CPI inflation, which coincide with the 1979 Oil crisis, the 1990–91 Gulf War, the 1997–98 Asian currency crisis, and the 2008–09 global financial crisis. This suggests that a large fluctuation in the Korean economy tends to heavily depend upon external shocks.

5.3. Impulse responses

Figs. 5 through 8 display impulse responses of key macroeconomic variables to various model shocks. Fig. 5 compares the estimated median impulse responses to monetary policy shocks across the two policy regimes. At the impact of a positive monetary policy shock, the regime 2 (*Non-IT*) responses of consumption and output are much larger than those in the regime 1 (*IT*), making domestic inflation lower. Moreover, the real exchange rate appreciation is larger under the regime 2, yielding a sharper reduction in import goods inflation. It follows from these observations that for a given monetary policy shock, output and inflation under the regime 2 tend to overreact, compared to the estimated responses under the regime 1.

Fig. 6 reports some nontrivial discrepancies in the two regime responses to a cost-push shock. For a positive cost-push shock, import goods inflation sharply rises under both the regimes. In contrast, nominal interest rate mildly rises in the regime 1 whereas persistently falling in the regime 2. Furthermore, we observe that given the cost-push shock, the dynamic response of headline inflation is more stable in the regime 1 than in the regime 2. It is worthwhile to note that the inflation stabilization in the regime 1 is achieved without amplifying output fluctuations in sharp contrast to the case of the UK investigated in Liu and Mumtaz (2011).

Some evidence of Korea's inflation stabilization with no larger output volatilities is also found in impulse responses to risk premium and technology shocks. As presented in Fig. 7, a positive risk premium shock raises nominal interest rate, entailing a sharp depreciation in real exchange rates which are governed by the UIP relation. The regime 2 response of nominal interest rate is approximately three times larger than the estimate in the regime 1, resulting in a large swing in CPI inflation. This suggests that inflation dynamics in the regime 1 might be less vulnerable to the risk premium as well as the cost-push disturbances. In addition, Fig. 8 exhibits that for a positive technology shock the regime 1 response of CPI inflation is less volatile than the estimate in the regime 2. Similarly, we observe that

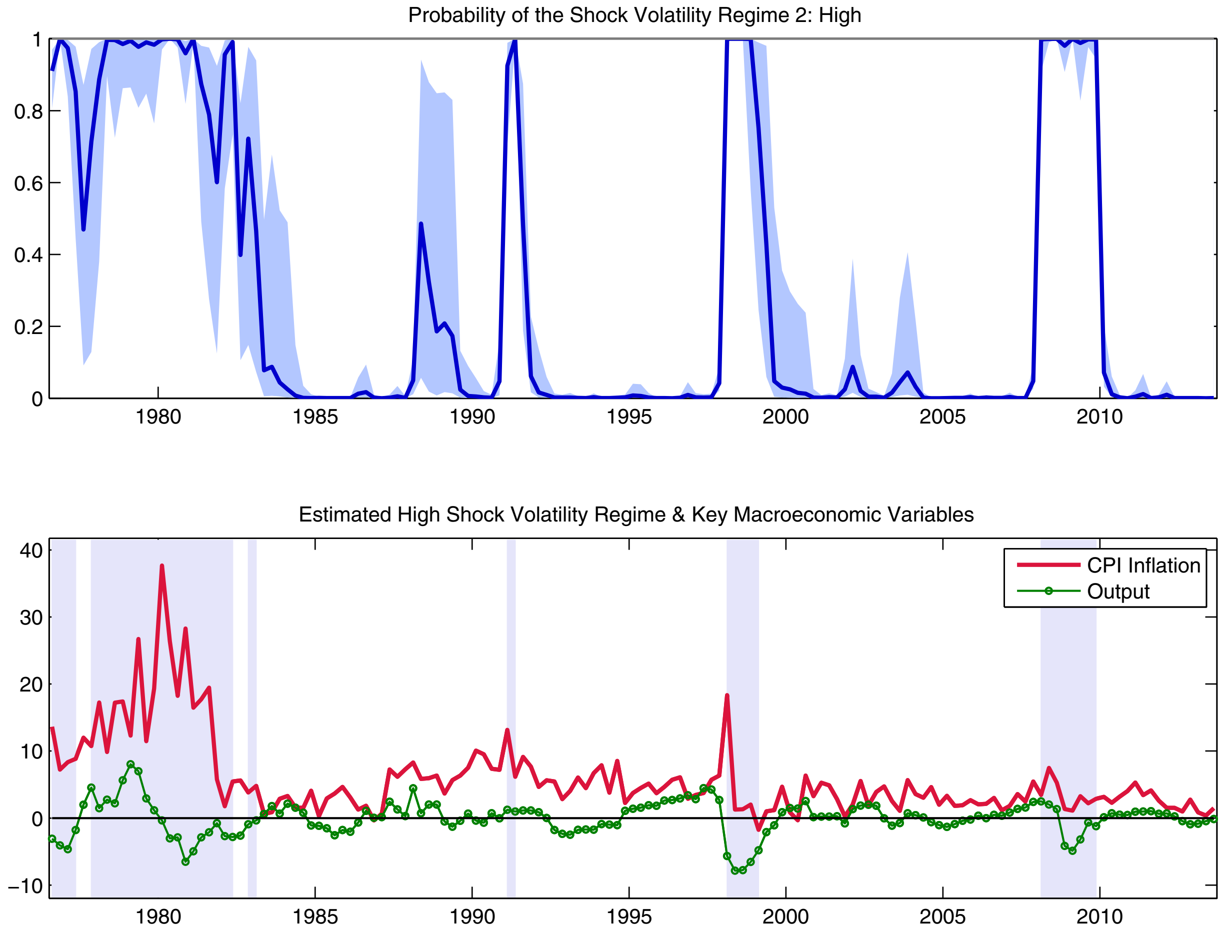

Fig. 4. [Top panel] Probabilities of the shock volatility regime 2 (high volatility). The posterior median and corresponding 95% interval estimates are reported. [Bottom panel] Estimated high shock volatility regime together with the key macroeconomic variables. The shaded areas correspond to the periods in which the probability of the shock volatility regime 2 exceeds 50% in terms of the median estimate.

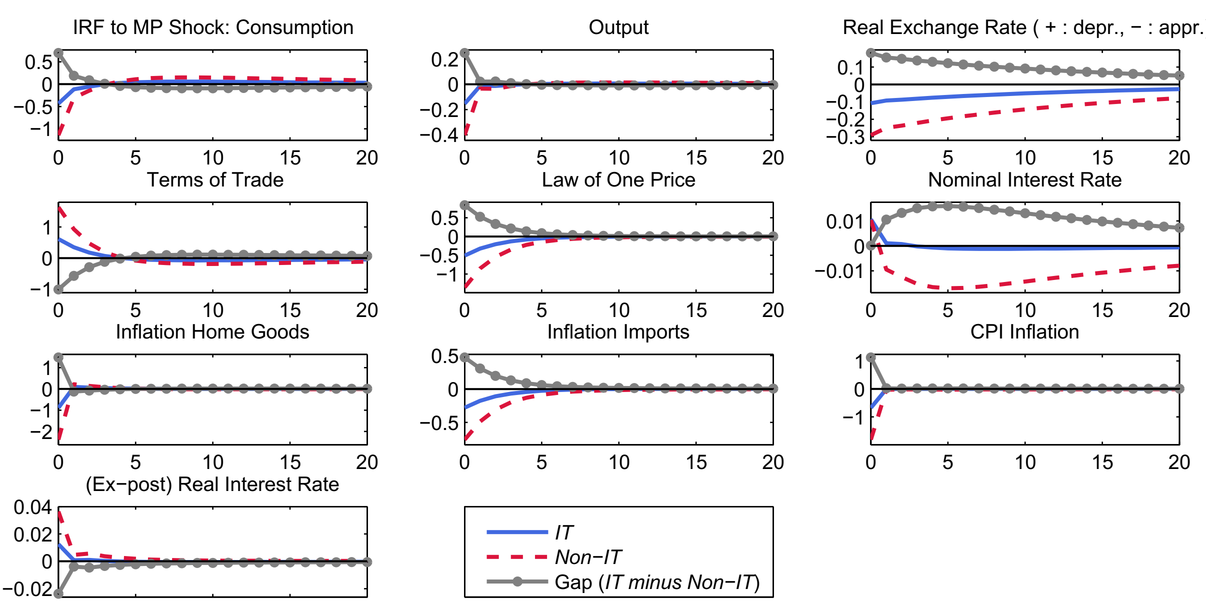

Fig. 5. Median impulse responses to monetary policy shocks under the monetary policy regime 1 (*IT*, solid lines) and regime 2 (*Non-IT*, dashed lines), and the median differences across the two impulse responses (solid lines with circles). The x-axis measures quarters.

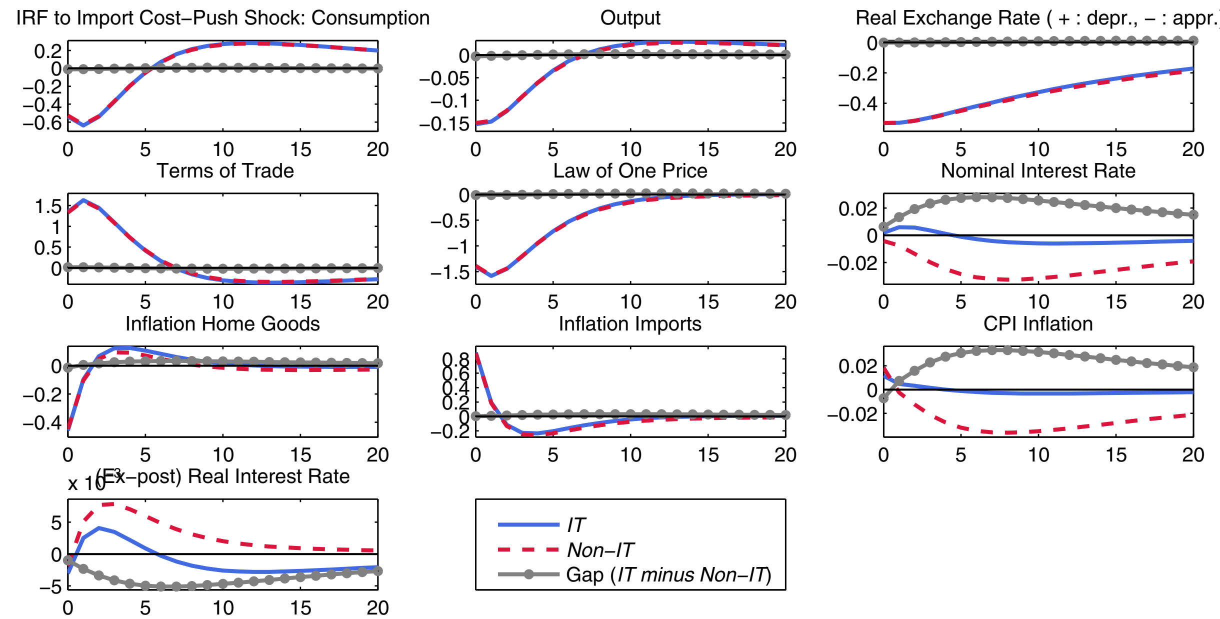

Fig. 6. Median impulse responses to import cost-push shocks under the monetary policy regime 1 (*IT*, solid lines) and regime 2 (*Non-IT*, dashed lines), and the median differences across the two impulse responses (solid lines with circles). The x-axis measures quarters.

the lower inflation volatility in the regime 1 is unaccompanied by a larger output fluctuation.

In sum, we find that under the *IT* monetary regime, to which the 2000s belongs, output and inflation tends to respond less to a monetary policy shock. Moreover, the inflation dynamics in response to non-monetary disturbances become more stable with no costs of larger output volatilities under the *IT* regime. In this sense, our impulse response analysis allowing for policy regime shifts lends some support to the view that Korea's monetary policy might have contributed to a sharp reduction in the level as well as the volatility of inflation in the 2000s. Note that this finding is consistent to the existing literature such as Kim and Park (2006) and Sánchez (2009).

5.4. Forecast error variance decompositions

In order to examine the sources of fluctuations in the model endogenous variables, we calculate the variance decomposition of the forecast errors of output and inflation. Since forecast error variance decompositions are determined by the model's structural parameters as well as shock standard deviations, there are four different outcomes conditional on the monetary policy and shock volatility regimes.

Table 5 summarizes the median forecast error variances of each regime, up to 20 quarters ahead. Overall, forecast error variance decompositions of output are quite insensitive to which regime is in place. Technology and import cost-push shocks account for most of the

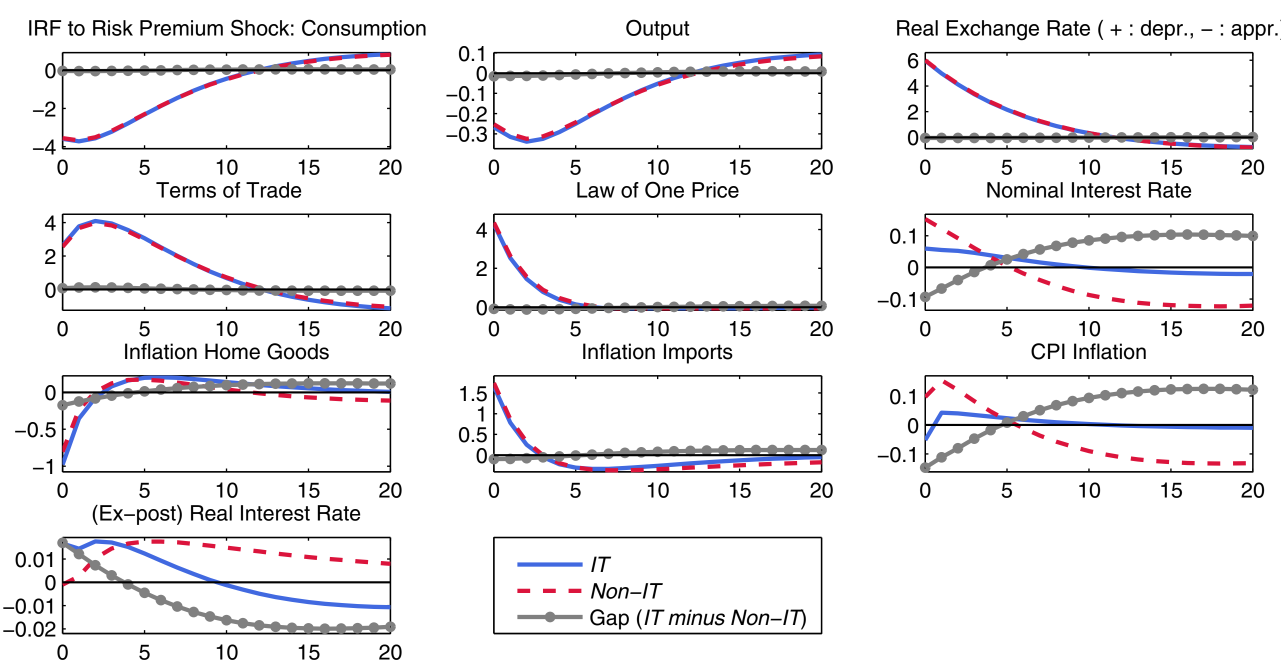

Fig. 7. Median impulse responses to risk premium shocks under the monetary policy regime 1 (*IT*, solid lines) and regime 2 (*Non-IT*, dashed lines), and the median differences across the two impulse responses (solid lines with circles). The x-axis measures quarters.

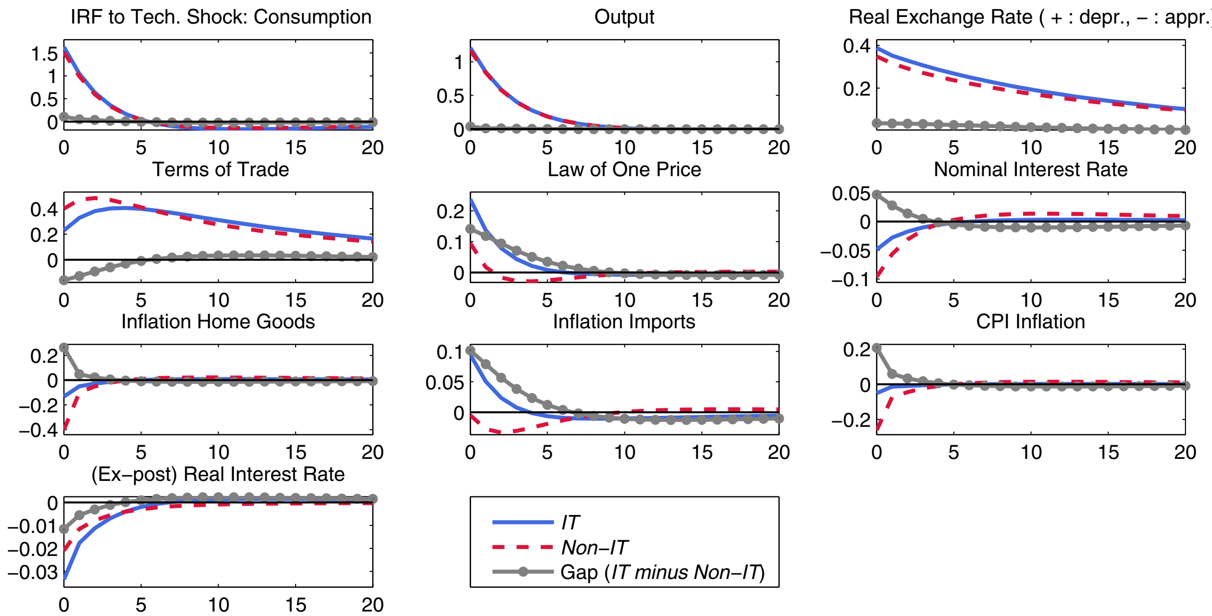

Fig. 8. Median impulse responses to technology shocks under the monetary policy regime 1 (*IT*, solid lines) and regime 2 (*Non-IT*, dashed lines), and the median differences across the two impulse responses (solid lines with circles). The x-axis measures quarters.

variation in output in the short- to medium-run. Longer-run output is heavily driven by import cost-push shocks alone.

Turning to forecast error variance decompositions of inflation, a universal finding across the regimes is that medium- to longer-run variances are driven mainly by preference shocks. Nevertheless, the source of short-run fluctuations in inflation varies substantially across the monetary policy and shock volatility regimes in place. More than 90% of inflation variations at the 1-quarter horizon is accounted for by monetary policy shocks, under the combination of *IT* monetary and low volatility regime. The contribution of preference shocks on short-run fluctuations in inflation, however, decreases under the other combinations of regimes. A notable difference is found in forecast error variances of inflation under the combination of *Non-IT* monetary and high volatility regime, as preference shocks are relatively more important that monetary policy shocks in explaining short-run inflation fluctuations.

Table 5
Median of forecast error variance share of output and inflation explained by each exogenous shock, under the four combinations of monetary and shock volatility regimes. The last column displays the summation of forecast error variances accounted for by foreign shocks (on foreign output, inflation, and nominal interest rate). Medians need not add up to one. Forecast horizon is measured in quarters.

	Horizon	Monetary policy	Preference	Technology	Import CP	Risk premium	Foreign shocks
IT monetary policy & low volatility regime							
Output	1	0.01	0.00	0.63	0.34	0.01	0.00
Inflation	1	0.95	0.02	0.01	0.01	0.00	0.00
Output	4	0.00	0.00	0.31	0.57	0.08	0.02
Inflation	4	0.00	0.70	0.01	0.03	0.17	0.05
Output	20	0.00	0.00	0.01	0.77	0.16	0.05
Inflation	20	0.00	0.80	0.00	0.14	0.03	0.01
Non-IT monetary policy & low volatility regime							
Output	1	0.04	0.01	0.59	0.33	0.01	0.00
Inflation	1	0.70	0.26	0.02	0.00	0.00	0.00
Output	4	0.00	0.00	0.31	0.58	0.08	0.02
Inflation	4	0.00	0.91	0.00	0.06	0.01	0.00
Output	20	0.00	0.00	0.00	0.79	0.15	0.04
Inflation	20	0.00	0.82	0.00	0.11	0.05	0.01
IT monetary policy & high volatility regime							
Output	1	0.00	0.00	0.43	0.55	0.01	0.00
Inflation	1	0.81	0.05	0.02	0.08	0.01	0.00
Output	4	0.00	0.00	0.17	0.72	0.08	0.00
Inflation	4	0.00	0.64	0.01	0.06	0.23	0.01
Output	20	0.00	0.00	0.00	0.85	0.13	0.01
Inflation	20	0.00	0.69	0.00	0.24	0.04	0.00
Non-IT monetary policy & high volatility regime							
Output	1	0.01	0.00	0.42	0.54	0.01	0.00
Inflation	1	0.41	0.50	0.03	0.02	0.00	0.00
Output	4	0.00	0.00	0.17	0.73	0.07	0.00
Inflation	4	0.00	0.86	0.00	0.11	0.01	0.00
Output	20	0.00	0.00	0.00	0.87	0.12	0.01
Inflation	20	0.00	0.72	0.00	0.18	0.07	0.00

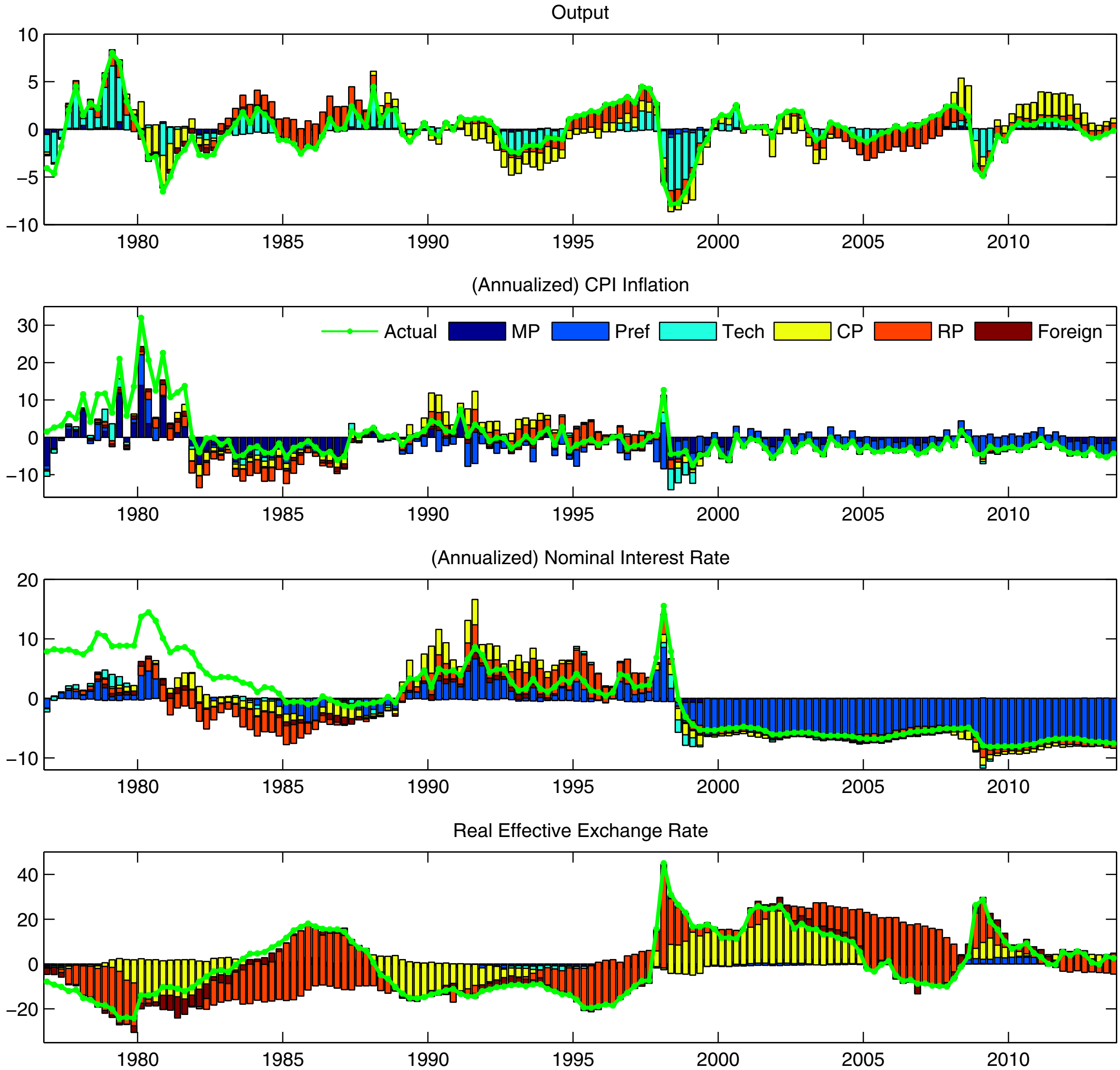

Fig. 9. Median historical decomposition estimates of output, inflation, interest rate and exchange rate for the best-fitting specification, the MS-DSGE model with regime shifts in the monetary policy rule coefficients and shock volatilities. Foreign shocks include output, inflation, and nominal interest rate shocks.

5.5. Historical decompositions

We evaluate the historical contribution of each exogenous shock in accounting for the macroeconomic fluctuations in Korea. To this end, we calculate the historical decomposition of output, inflation, nominal interest rate and exchange rate, and display the median estimates in Fig. 9.[12]

The first panel of Fig. 9 illustrates that technology shocks are the most important driver of output fluctuations in Korea. Most of post-1976 output booms and recessions are accounted for by changes in the sign and magnitude of technology shocks. More interestingly, a substantial portion of the two subsequent economic downturns, followed by the Asian currency crisis in 1997–98 and the global financial crisis in 2008–09, is characterized by adverse shocks on technology. Aside from technology shocks, it turns out that import cost-push and risk premium shocks follow next to technology shocks in terms of the contribution to output fluctuations.

Regarding the behavior of inflation, it is clear from the second panel of Fig. 9 that there seems to be a structural change occurring in the last 1990s, which coincides approximately with the wake of the Asian currency crisis. For the pre-Asian currency crisis period, four shocks – monetary policy, preference, risk premium, and import cost-push – are the main sources of the inflation dynamics. In particular, monetary policy shocks are an important driver of inflation up to the late 1980s, but their role is quite limited between the late 1980s and the late 1990s. Rather, preference shocks are given relatively more weight in accounting for the inflation dynamics over this sample period. Together with the four shocks, changes in technology shocks play a significant role in explaining the inflation variability during the Asian currency crisis period. On the contrary, the post-Asian currency crisis dynamics of inflation are mainly accounted for by the variability of preference shocks. Also monetary policy shocks have a positive, though much more limited contribution to the variation in inflation over the sample period.

As in the third panel of Fig. 9, the driving forces of the nominal interest rate change substantially around the Asian currency crisis. Prior to the event, major portion of the interest rate variability is generated

[12] In this figure, the actual series denote the model-implied filtered series of the corresponding variable.

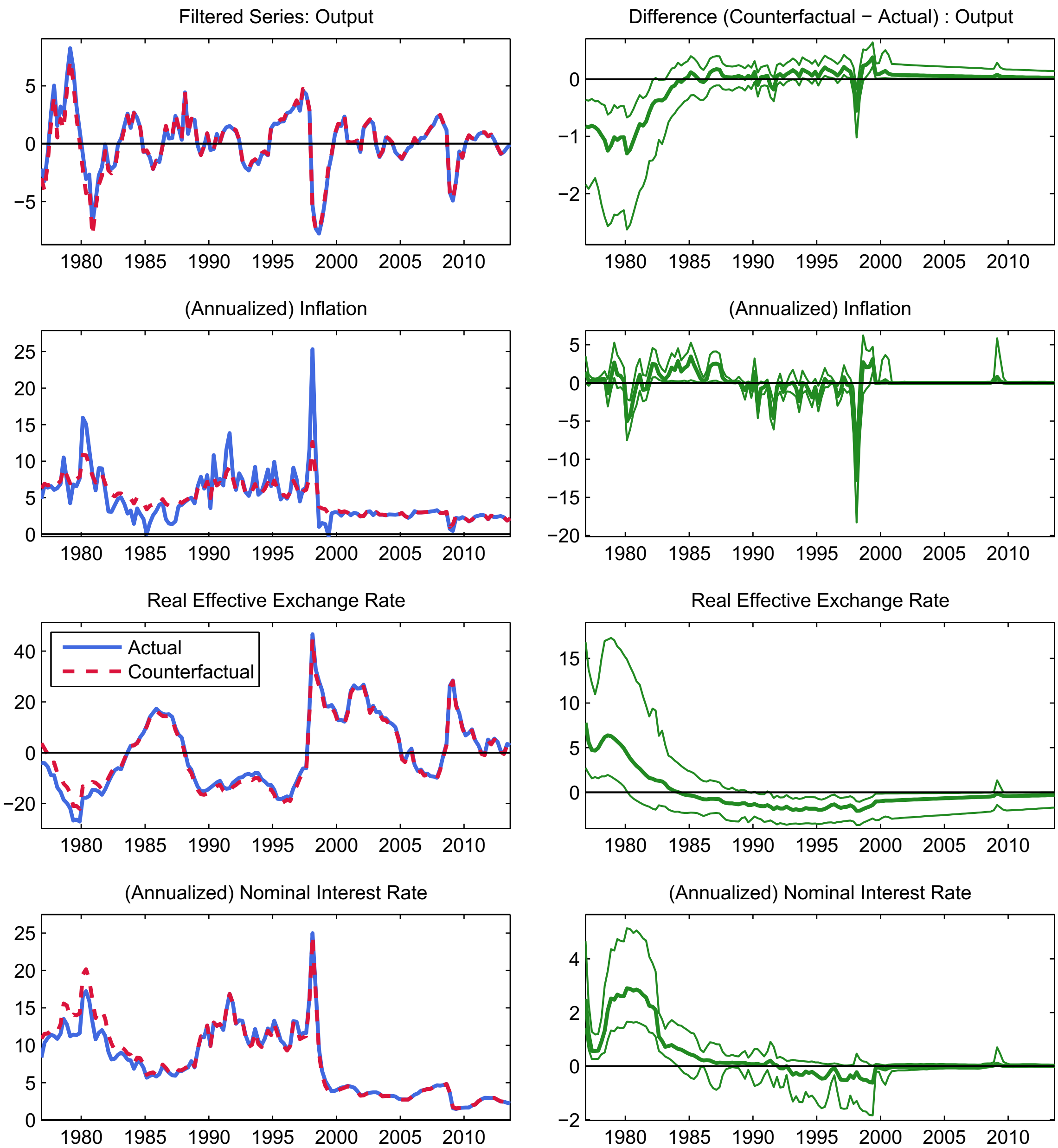

Fig. 10. [Left panels] Median estimates of actual (solid lines) and counterfactual (dashed lines) series implied by the MS-DSGE model with regime shifts in the monetary policy rule coefficients and shock volatilities. The counterfactual series are obtained under the assumption that the *IT* monetary policy regime is maintained over the entire sample period. [Right panels] Median and 95% interval difference in the actual and counterfactual series implied by the MS-DSGE model with regime shifts in the monetary policy rule coefficients and shock volatilities.

jointly by preference, risk premium, and import cost-push shocks. However, preference shocks dominate in explaining the dynamics of the interest rate after the Asian currency crisis.

Finally, the last panel of Fig. 9 demonstrates that the exchange rate dynamics over the sample period is largely attributed to risk premium and import cost-push shocks. In particular, the dramatic depreciations of the Korean currency around the Asian currency and global financial crisis periods are associated with dominant contributions from these two shocks.

5.6. Counterfactual exercises

We now turn to the implications of regime shifts in monetary policy regarding the macroeconomic performance in Korea. In particular, we conduct counterfactual analyses for the MS-DSGE model allowing for regime shifts in monetary policy and shock volatilities. The scenario that we consider herein is how macroeconomic outcomes are altered if the *IT* monetary policy regime is maintained over the entire sample period. In order to achieve this goal, we re-solve the MS-DSGE model with the monetary policy coefficients fixed at the values drawn from the *IT* regime, and generate the filtered series of key macroeconomic variables. In this exercise, the monetary policy shocks are set to be zero, in order to isolate the impacts of changes in the systematic component of monetary policy on the model economy.

The left panels of Fig. 10 plot the median filtered series of output, inflation, REER, and nominal interest rate, together with the median estimates under the counterfactual scenario. In addition, the differences between the actual and counterfactual series are displayed in the right panels of Fig. 10. For the four variables considered, the counterfactual series deviate from the actual ones mostly during the period of the *Non-IT* monetary policy, which takes place from the mid-1970s until

[13] This statistic is analogous to the counterfactual sacrifice ratio (CFSR) measure in Bianchi (2013).

Table 6
Counterfactual trade-off ratio (CTR) statistics for output over various sample periods. This table reports the median and associated [2.5%, 97.5%] percentile intervals (in brackets). The counterfactual series are obtained under the assumption that the *IT* monetary policy regime is maintained over the entire sample period. The CTR statistics measure the percentage of output to be sacrificed to lower (annualized) inflation by 1%.

Variable	1976:Q3 to 2013:Q3 (Entire sample)	1976:Q3 to 1990:Q4	1997:Q3 to 1999:Q4 (Asian fin. crisis)
Output	−0.02 [−7.39, 6.98]	−0.61 [−7.12, 5.10]	−0.06 [−1.32, 1.03]

the onset of the Asian currency crisis. If monetary policy were *IT* during this period, the hypothetical nominal interest rate should have been higher than the actual rate practiced. Output would have been lower, whereas REER would have been higher. The hypothetical inflation could have been higher and lower than the actual one. Regarding the macroeconomic consequences of the Asian currency crisis, the *IT* monetary policy stance, if conducted, could have subdued the subsequent high inflation in the late 1990s but it comes at the cost of lower output.

We then measure the trade-off between output and inflation emerged from the counterfactual analyses. To this end, we calculate the counterfactual trade-off ratio (CTR) statistic defined as

$$CTR = \sum_{t=T_0}^{T_1}\left[\hat{y}_t^{\text{actual}} - \hat{y}_t^{\text{counterfactual}}\right] / \left[\hat{\pi}_t^{\text{actual}} - \hat{\pi}_t^{\text{counterfactual}}\right].$$

We repeat this exercise for various sample periods and report the results in Table 6.[13] At first glance, the wide ranges of the CTR estimates make it difficult in evaluating the consequences of the counterfactual exercise performed here. Nevertheless, there are several findings emerged from the analysis.

Over the entire sample period, the slightly negative median estimate indicates that it was possible to achieve both higher output and lower inflation if the *IT* monetary policy were conducted. Based on the median estimate, output could have been 2 basis-points higher than the actual level while inflation would have been maintained 1% lower than the historical level. The 95th percentile interval, however, ranges from −7.39 to 6.98, which might make any policy conclusions regarding the entire sample period impetuous.

The counterfactual analysis renders a clearer answer when we focus on the pre-1990 sample. The mean CTR estimate is −0.61 which indicates that output could have been 61 basis-points higher when inflation was brought down by 1%. Moreover, the 95th percentile interval estimates take negative values more than positive ones. Given that the sample span largely overlaps with the period of the *Non-IT* monetary policy regime, these findings illustrate that it would have been possible to achieve higher output and lower inflation simultaneously if the *IT* monetary policy stance were pursued.

Finally, similar results with a tighter 95th interval emerge if the Asian currency crisis period is considered. Output could have been 6 basis-points higher than the actual level while inflation would have been subdued by 1%. The 95th percentile interval ranges from −1.32 to 1.03, which indicates that there would have been a slightly higher probability of gains in macroeconomic performance than losses.

6. Conclusions

This paper offers empirical evidences of possible regime changes in Korean economy using a SOE MS-DSGE model. We identify significant changes in monetary policy as well as shock volatilities over the sample period considered. In addition, we formally demonstrate that endowing conventional models with plausible regime switching aspects of monetary policy and shock volatilities can improve the model's performance in fitting the data better. We view that the paper's results can be useful for both researchers and policy analysts using constant coefficient DSGE models for Korean economy. The choice of how to model the nature of policy behavior and fundamental shock processes is potentially a crucial issue. Further progress on estimating DSGE models using data for Korea may require perspectives on the issues that this paper has highlighted.

Nevertheless, the model employed in this paper abstracts from several modeling features widely considered to be important for macroeconomic dynamics in emerging economies. For instance, the model omits a non-stationary component of technology shock (Aguiar and Gopinath, 2007), a consideration of sovereign debt (Uribe and Yue, 2006), and relevant financial frictions (Yun, 2013), all of which can produce a richer model dynamics consistent with the Korean economic environment. We leave these aspects for future work.

Appendix A. Log-linearized model

Let a hat (^) denotes the log deviation from the steady state. Then the log-linearized system of the DSGE model is given as follows.

• Euler equation:

$$(1+h)\hat{c}_t = h\hat{c}_{t-1} + \mathbb{E}\hat{c}_{t+1} - \frac{1-h}{\sigma}\left(\hat{i}_t - \mathbb{E}_t\hat{\pi}_{t+1}\right) + \frac{1-h}{\sigma}\left(\hat{\varepsilon}_{g,t} - \mathbb{E}_t\hat{\varepsilon}_{g,t+1}\right)$$

where $\hat{\varepsilon}_{g,t}$ is the preference shock process defined as above, e.g., $\hat{\varepsilon}_{g,t} = \rho_g\hat{\varepsilon}_{g,t-1} + \sigma_g(\xi_t^Q)\epsilon_{g,t}$, and $\epsilon_{g,t} \sim \mathbb{N}(0,1)$.

• Goods market clearing:

$$(1-\alpha)\hat{c}_t = \hat{y}_t - \alpha\eta(2-\alpha)\hat{s}_t - \alpha\eta\hat{\psi}_{F,t} - \alpha\hat{y}_t^*$$

where $\hat{\psi}_{F,t}$ is the law of one price gap defined as $\hat{\psi}_{F,t} \equiv (\hat{e}_t + \hat{p}_t^*) - \hat{p}_{F,t}$, and $\hat{s}_t$ is the terms of trade defined as $\hat{s}_t \equiv \hat{p}_{F,t} - \hat{p}_{H,t}$, so that $\hat{s}_t - \hat{s}_{t-1} = \hat{\pi}_{F,t} - \hat{\pi}_{H,t}$.

• Changes in the nominal exchange rate:

$$\Delta\hat{e}_t = \hat{q}_t - \hat{q}_{t-1} + \hat{\pi}_t - \hat{\pi}_t^*$$

• Domestic price inflation:

$$(1+\beta\delta_H)\hat{\pi}_{H,t} = \delta_H\hat{\pi}_{H,t-1} + \beta\mathbb{E}_t\hat{\pi}_{H,t+1} + \frac{(1-\theta_H)(1-\theta_H\beta)}{\theta_H}\left[\varphi\hat{y}_t - (1+\varphi)\hat{\varepsilon}_{a,t} + \alpha\hat{s}_t + \frac{\sigma}{1-h}(\hat{c}_t - \hat{c}_{t-1})\right]$$

where $\hat{\varepsilon}_{a,t}$ is the technology shock process defined as above.

• Import price inflation:

$$(1+\beta\delta_F)\hat{\pi}_{F,t} = \delta_F\hat{\pi}_{F,t-1} + \beta\mathbb{E}_t\hat{\pi}_{F,t+1} + \frac{(1-\theta_F)(1-\theta_F\beta)}{\theta_F}\hat{\psi}_{F,t} + \hat{\varepsilon}_{cp,t}$$

where $\hat{\varepsilon}_{cp,t}$ is an import cost-push shock process that follows

$$\hat{\varepsilon}_{cp,t} = \rho_{cp}\hat{\varepsilon}_{cp,t-1} + \sigma_{cp}\left(\xi_t^Q\right)\epsilon_{cp,t}, \quad \epsilon_{cp,t} \sim \mathbb{N}(0,1)$$

• Domestic CPI inflation:

$$\hat{\pi}_t = (1-\alpha)\hat{\pi}_{H,t} + \alpha\hat{\pi}_{F,t}$$

• Uncovered interest rate parity:

$$\mathbb{E}_t\hat{q}_{t+1} - \hat{q}_t = \left(\hat{i}_t - \mathbb{E}_t\hat{\pi}_{t+1}\right) - \left(\hat{i}_t^* - \mathbb{E}_t\hat{\pi}_{t+1}^*\right) + \chi\hat{a}_t + \hat{\varepsilon}_{rp,t}$$

where $\hat{\varepsilon}_{rp,t}$ denotes the risk premium shock process defined as above.

• Foreign asset budget constraint:

$$\hat{c}_t + \hat{a}_t = \beta^{-1}\hat{a}_{t-1} - \alpha\left(\hat{s}_t + \hat{\psi}_{F,t}\right) + \hat{y}_t$$

• Monetary policy:

$$\hat{i}_t = \rho_i\left(\xi_t^P\right)\hat{i}_{t-1} + \left(1-\rho_i\left(\xi_t^P\right)\right)\left[\lambda_\pi\left(\xi_t^P\right)\hat{\pi}_t + \lambda_y\left(\xi_t^P\right)\hat{y}_t + \lambda_{de}\left(\xi_t^P\right)\Delta\hat{e}_t\right] + \sigma_i\left(\xi_t^Q\right)\epsilon_{i,t}, \quad \epsilon_{i,t}\sim\mathbb{N}(0,1)$$

• Foreign interest rate process:

$$\hat{i}_t^* = \rho_{i^*}\hat{i}_{t-1}^* + \sigma_{i^*}\left(\xi_t^Q\right)\epsilon_{i^*,t}, \quad \epsilon_{i^*,t}\sim\mathbb{N}(0,1)$$

• Foreign output process:

$$\hat{y}_t^* = \rho_{y^*}\hat{y}_{t-1}^* + \sigma_{y^*}\left(\xi_t^Q\right)\epsilon_{y^*,t}, \quad \epsilon_{y^*,t}\sim\mathbb{N}(0,1)$$

• Foreign inflation process:

$$\hat{\pi}_t^* = \rho_{\pi^*}\hat{\pi}_{t-1}^* + \sigma_{\pi^*}\left(\xi_t^Q\right)\epsilon_{\pi^*,t}, \quad \epsilon_{\pi^*,t}\sim\mathbb{N}(0,1)$$

Appendix B. Data description

The model is estimated using Korean and U.S. quarterly data from 1976:Q3 to 2013:Q3. Detailed data descriptions are as follows.

$$Domestic\,Output = \log(Actual\,Domestic\,Real\,GDP/HP\,Trend) \times 100,$$

$$Real\,Effective\,Exchange\,Rate = \log(Nominal\,Exchange\,Rate \times Foreign\,CPI/Domestic\,CPI) \times 100,$$

$$(Annualized)\,Import\,Goods\,Inflation = \log(Import\,Price\,Deflator/Import\,Price\,Deflator(-1)) \times 400,$$

$$Domestic\,Nominal\,Interest\,Rate = Overnight\,Call\,Rate\,(per\,annum),$$

$$(Annualized)\,Domestic\,CPI\,Inflation = \log(Domestic\,CPI/Domestic\,CPI(-1)) \times 400,$$

$$Foreign\,Output = \log(Actual\,U.S.Real\,GDP/HP\,Trend) \times 100,$$

$$(Annualized)\,Foreign\,CPI\,Inflation = \log(U.S.CPI/U.S.CPI(-1)) \times 400,$$

$$Foreign\,Nominal\,Interest\,Rate = Federal\,Funds\,Rate\,(per\,annum),$$

where sources of the original data are:

- Domestic real GDP: real gross domestic product, seasonally adjusted, 2005 reference year, Bank of Korea's Economic Statistics System database (BOK-ECOS).
- Nominal exchange rate: the basic exchange rate of the Korean won against the U.S. dollar (the transactions volume-weighted market average of the rates applied in the previous business day's transactions between foreign exchange banks through brokers), averages of daily figures, BOK-ECOS.
- Import price deflator: import price indexes, 2010 = 100, seasonally adjusted, BOK-ECOS.
- Domestic nominal interest rate: overnight call rate, uncollateralized, percent per annum, averages of daily figures, BOK-ECOS.
- Domestic CPI: consumer price indexes, 2010 = 100, seasonally adjusted, BOK-ECOS.
- U.S. real GDP: real gross domestic product, chained dollars, billions of chained (2009) dollars, seasonally adjusted at annual rates, NIPA Table 1.1.6, line 1.
- U.S. CPI: consumer price index for all urban consumers: all items, 1982–1984 = 100, seasonally adjusted, Federal Reserve Economic Data.
- Federal funds rate: averages of daily figures, percent, Board of Governors of the Federal Reserve System.

Appendix C. Supplementary data

The estimation appendix to this article can be found online at http://dx.doi.org/10.1016/j.econmod.2015.07.020.

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