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Abstract

For the Japanese economy, we examine whether the fiscal multiplier is higher under the effective lower bound of the nominal interest rate. Using a time-varying parameter vector autoregression model with Tobit-type nonlinearity, we calculate the fiscal multipliers under two monetary policy positions. We find that when government spending shocks are inflationary, the fiscal multiplier under the zero interest rate policy increases steadily as a result of the decrease in the real interest rate. This evidence is robust to different definitions of effective lower bound, output, and government spending.

JEL classification: E62, C32, E52

Keywords: TVP-VAR model; Fiscal multiplier; Effective lower bound; Implied rate

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1. Introduction

In response to the economic downturn after the global financial crisis in 2008, the central banks of the major industrialized economies adopted a zero interest rate policy (ZIRP). As a result, their nominal short-term interest rates have fallen to the effective lower bound (ELB). Although there remains room to implement unconventional monetary actions, such as quantitative easing and forward guidance, the central banks have lost a conventional monetary tool they would otherwise use to stimulate their economies under a liquidity trap. On the other hand, the fiscal authorities in these economies have implemented aggressive large-scale fiscal stimulus packages within this low interest rate environment (e.g., the *American Recovery and Reinvestment Action* in the United States, and the *Policy Package to Address the Economic Crisis* in Japan). The consecutive implementation of fiscal stimulus packages has ignited debate on the effectiveness of fiscal policy and, in particular, whether the fiscal multiplier is higher under the ELB of the nominal interest rate. In the aftermath of the crisis, the literature focused on answering this question within a theoretical framework, resulting in conflicting theoretical predictions. For example, some studies reported that the fiscal multiplier becomes much larger when the economy is in the ELB (Braun and Waki, 2006; Eggertsson, 2010; Christiano et al., 2011), whereas others demonstrated that it does not necessarily become larger in the ELB (Erceg and Lindé, 2014; Boneva et al., 2016). This theoretical conflict motivated us to examine this question empirically.

Using a structural time-varying parameter vector autoregression (TVP-VAR) model with a censored variable, we analyze whether the Japanese government spending multiplier becomes larger under the ELB. Long samples from the period near the zero lower bound in Japan are suitable for our study. It is important that we separate the contribution of the ZIRP from those of all other possible factors that might affect the fiscal multiplier. For instance, Auerbach and Gorodnichenko (2012) and Fazzari et al. (2015) demonstrate that fiscal policy is more effective when the economy is in a recession. However, because the low interest rate environment virtually coincides with a recession, it is difficult to evaluate the contribution of the ZIRP separately by comparing the multipliers calculated in the ELB and non-ELB periods, as in Ramey and Zubairy (2018) and Miyamoto et al. (2018). In other words, it is probable that the effects of the low interest rate and the recession are mixed in the size of the multipliers estimated after dividing the sample period. Our model includes three notable features to clarify the role of the ZIRP in determining the size of the fiscal multiplier. First, the time-varying specification allows us to estimate the policy effect in each period without dividing the sample period. Second, we employ Iwata and Wu's (2006) Tobit specification, which treats the short-term interest rate as a censored variable. Third, the interest rate equation in our VAR model is formulated as a structural Taylor rule, not a reduced-form equation. This formulation has been used in some empirical studies (Stock and Watson, 2001; Hayashi and Koeda, 2019), and is adopted here to accord with the theoretical model mentioned above. Exploiting the time-varying structure and the Tobit specification, we calculate two impulse responses (IRs) and fiscal multipliers at each period, one of which is derived from the lower bound value of the interest rate (referred to as the benchmark), while the other uses the implied rate obtained from the Tobit model (referred to as the counterfactual). Because the difference between the two stems only from the responses of the interest rate, we can regard it as the contribution of the monetary policy stance to the fiscal multiplier. Hence, our analysis differs from those of Ramey and Zubairy (2018) and Miyamoto et al. (2018) in that we eliminate the possibility of factors other than the ZIRP affecting the size of the multipliers.

As mentioned above, we conduct a counterfactual simulation in which hypothetical responses are drawn by changing only the coefficients in the interest rate equation. It is well known that this sort of simulation is likely to conflict with Lucas's (1976) argument that the coefficients of other equations in the VAR system also change if the monetary policy rule is changed. This occurs because rational agents immediately adjust their behavior in response to the policy changes. To resolve this problem, Hayashi and Koeda (2019) estimate the reduced-form coefficients for each monetary policy regime, and then use these to compute the counterfactual responses. This is the conventional solution to the above problem in a VAR analysis. However, in practice, it is not clear whether the behavior of the private sector is reversible when a policy regime switch occurs. In the Japanese economy in particular, private agents are unlikely to act as before if the central bank lifts the zero interest rate, because the economy has been in the ELB since the mid-1990s. Therefore, we conduct our counterfactual simulation following Sims and Zha (2006), who argue that the nonlinearities stemming from a regime change, as emphasized by Lucas (1976), matter only as far as private agents change their beliefs on the policy regime. Leeper and Zha (2003) also show that the hypothetical path conditioning on invariant coefficients in the nonpolicy sector could sufficiently approximate the true effects, unless the policy intervention shifts agents' beliefs. Based on this argument, we run our counterfactual simulation assuming that the private sector and the fiscal authority cannot instantly shift their beliefs on the prevailing monetary policy rule when the monetary authority changes this rule. Indeed, the Japanese policy rate has fluctuated only slightly with daily transactions in the short-term money market during the ELB period. Thus, it seems reasonable to assume that agents cannot detect a change in the monetary policy rule in the short run. On the other hand, our counterfactual is naturally subject to the Lucas critique in the long run, because agents can adjust their beliefs accordingly. Therefore, in the actual estimation below, we calculate the hypothetical responses for two years after the shock, during which the agents' beliefs on monetary policy are unlikely to be adjusted adequately.

The main finding of this study is that when a government spending shock is inflationary, the fiscal multiplier under the zero interest rate policy increases steadily as a result of the decrease in the real interest rate. This finding is consistent with the mechanisms presented in Eggertsson (2010) and Christiano et al. (2011). In addition, we reveal substantial variations over time in the Taylor rule coefficients and in the IRs of each variable to a government spending shock, which supports our time-varying specification of the model. Moreover, the estimated implied interest rate takes a negative value with active fluctuations during the ELB period. This implies that the interest rate may have changed in response to structural shocks had the Bank of Japan (BOJ) not successfully controlled the short-term interest rate.

The remainder of this paper is organized as follows. In Section 2, we explain our empirical framework. Specifically, we describe the structural TVP-VAR model with Tobit nonlinearity, the Bayesian technique, and our data. Section 3 presents the estimated results for the implied rate, IRs, and fiscal multipliers. In the robustness check in Section 4, we change the definitions of the ELB period, output measure, and government spending measure, and then re-estimate the model. Finally, Section 5 concludes the paper.

2. Empirical model

2.1. TVP-VAR model with censored variables

To examine the effect of a government spending shock under the ELB of the nominal short-term interest rate, we employ a four-variable TVP-VAR model with a censored variable. The VAR model incorporates the quarter-on-quarter growth rate of government spending (g_t), inflation rate (p_t), GDP growth rate (x_t), and nominal interest rate, in this order. The notable feature of our VAR model is that we impose the Tobit specification on the nominal short-term interest rate. As in Iwata and Wu (2006), we distinguish the observed nominal interest rate (r_t) and latent rate (r_t^*). Here, we suppose that the observed rate is equal to the latent rate if it exceeds the lower bound of the nominal interest rate, and is equal to the lower bound otherwise. Denoting the lower bound of the nominal interest rate as c, the relationship between r_t and r_t^* is specified as

$$r_t = \begin{cases} r_t^* & \text{if } r_t^* > c \\ c & \text{if } r_t^* \leq c. \end{cases} \tag{1}$$

Other than adopting time-varying parameters and a censored variable, our VAR model is essentially the same as those employed by Stock and Watson (2001) and Hayashi and Koeda (2019), in the sense that the first three equations are reduced-form equations and the fourth equation is a structural Taylor rule.

Defining $Y_t = [g_t, p_t, x_t]'$ and $\tilde{Y}_t = [g_t, p_t, x_t, r_t]'$, the three reduced-form equations are formulated as

$$\begin{gathered} Y_t = B_{0,t} + B_{1,t}\tilde{Y}_{t-1} + \cdots + B_{s,t}\tilde{Y}_{t-s} + A_t^{-1}\Sigma_t\varepsilon_t, \\ \varepsilon_t \sim N(0, I). \end{gathered} \tag{2}$$

In addition, we define $\theta_t = [vec(B_{0,t})', \cdots, vec(B_{s,t})']'$ and $X_t = I \otimes \left(1, \tilde{Y}_{t-1}, \cdots, \tilde{Y}_{t-s}\right)$, where the vec operator creates a column vector from $B_{i,t} (i = 1, \cdots, s)$ by stacking the column vectors of $B_{i,t}$, and $\otimes$ denotes the Kronecker product. Exploiting these newly defined variables, eq. (2) can be rewritten as

$$Y_t = X_t\theta_t + A_t^{-1}\Sigma_t\varepsilon_t. \tag{3}$$

Here, $\varepsilon_t = [\varepsilon_{g,t}, \varepsilon_{p,t}, \varepsilon_{x,t}]'$ is a vector of structural shocks, which are assumed to follow mutually independent standard normal distributions; A_t is a matrix of time-varying coefficients specifying the simultaneous relationship between the endogenous variables; and

Σ_t denotes the time-varying volatilities of the structural shocks in the diagonal elements. We assume A_t and Σ_t take the following form:

$$A_t = \begin{pmatrix} 1 & 0 & 0 \\ a_{gp,t} & 1 & 0 \\ a_{gx,t} & a_{px,t} & 1 \end{pmatrix}, \quad \text{and} \quad \Sigma_t = \begin{pmatrix} \sigma_{g,t} & 0 & 0 \\ 0 & \sigma_{p,t} & 0 \\ 0 & 0 & \sigma_{x,t} \end{pmatrix}, \tag{4}$$

respectively, where $a_{kl,t}$ denotes the simultaneous response of variable l to the shock $\varepsilon_{k,t}$ at time t, and $\sigma_{l,t}$ denotes the time-varying volatility of the structural shock $\varepsilon_{l,t}$. In other words, the structural shocks are identified by a recursive restriction, assuming that government spending is the most exogenous variable not simultaneously affected by other structural shocks.[1]

The fourth equation is the time-varying censored Taylor rule with interest rate smoothing, specified as

$$\begin{gathered} r_t = \max\left[c, r_t^*\right], \\ r_t^* = (1-\rho_t)\left[\phi_{0,t} + \phi_{\pi,t}\pi_t + \phi_{\chi,t}\chi_t\right] + \rho_t r_{t-1} + \varepsilon_{r,t}, \\ \varepsilon_{r,t} \sim N(0, \sigma_{r,t}^2). \end{gathered} \tag{5}$$

The implied interest rate is assumed to be a function of the contemporaneous values of year-on-year inflation ($\pi_t \equiv p_t + p_{t-1} + p_{t-2} + p_{t-3}$), year-on-year GDP growth rate ($\chi_t \equiv x_t + x_{t-1} + x_{t-2} + x_{t-3}$), and the lagged observed interest rate. Following the literature on fiscal policy analysis (e.g., Blanchard and Perotti, 2002; Galí et al., 2007), we employ the GDP growth rate rather than the more usual output gap (e.g., Taylor, 1993; Clarida et al., 1998) as an output measure in the Taylor rule.[2] We also incorporate inflation and the GDP growth rate into the equation as year-on-year rates, assuming that the central bank adjusts the policy rate in response to the slow movement of the

[1]Perotti (2002) discusses the possibility of a simultaneous response of real government spending to price. That is, real government spending may fluctuate in response to inflation when government spending is contracted in nominal terms. In the time-varying specification, we encounter difficulty when we consider such simultaneity in our estimation of the time-varying contemporaneous matrix. Fortunately, as noted by Kato (2003), in a constant-parameter VAR model, the qualitative effect on output remains unchanged independently of the contemporaneous elasticity of government spending to price. Although we recognize the importance of this simultaneous relationship, we adopt a specification in which price has no contemporaneous effect on government spending.

[2]In the robustness check, we estimate the model after replacing the GDP growth rate with the GDP gap in order to reconcile with the standard Taylor rule specification.

economy. In addition, the error term in eq. (5) is assumed to be orthogonal to the errors in eq. (2). As mentioned above, the VAR system employed in this study comprises three reduced-form equations (2) and a structural Taylor rule (5). Note that the implied interest rate appears only on the left-hand side of the system in order to be consistent with the specification of Iwata and Wu (2006).

Following the seminal work of Primiceri (2005) on the TVP-VAR model, the time-varying parameters in the proposed model are assumed to follow a random-walk process. Let α_t be a stacked vector of the lower triangular elements in A_t, $h_t = (h_{g,t}, h_{p,t}, h_{x,t}, h_{r,t})'$ with $h_{l,t} = \ln(\sigma^2_{l,t})$, and $\varphi = (\varphi_{0,t}, \varphi_{\pi,t}, \varphi_{\chi,t}, \rho_t)'$ with $\varphi_m = (1-\rho_t)\phi_{m,t}$. Then, the dynamic process of the model's time-varying parameters is given by

$$\begin{aligned} \theta_{t+1} &= \theta_t + u_{\theta,t}, \\ \alpha_{t+1} &= \alpha_t + u_{\alpha,t}, \\ \varphi_{t+1} &= \varphi_t + u_{\varphi,t}, \\ h_{t+1} &= h_t + u_{h,t}, \end{aligned} \qquad \begin{pmatrix} u_{\theta,t} \\ u_{\alpha,t} \\ u_{\varphi,t} \\ u_{h,t} \end{pmatrix} \sim \left(0, \begin{pmatrix} \Sigma_\theta & O & O & O \\ O & \Sigma_\alpha & O & O \\ O & O & \Sigma_\varphi & O \\ O & O & O & \Sigma_h \end{pmatrix} \right), \tag{6}$$

for $t = s+1, \cdots, T$, where $\theta_{s+1} \sim N(\mu_{\theta,0}, \Sigma_{\theta,0})$, $\alpha_{s+1} \sim N(\mu_{\alpha,0}, \Sigma_{\alpha,0})$, $\varphi_{s+1} \sim N(\mu_{\varphi,0}, \Sigma_{\varphi,0})$, and $h_{s+1} \sim N(\mu_{h,0}, \Sigma_{h,0})$. The model innovations in eq.(6) are assumed to be mutually uncorrelated, and the variance-covariance structure for $(\Sigma_\theta, \Sigma_\alpha, \Sigma_\varphi, \Sigma_h)$ is assumed to be a diagonal matrix.

2.2. MCMC estimation

As widely employed in the literature, we estimate the TVP-VAR model with a censored variable using the Gibbs sampler of the Bayesian Markov Chain Monte Carlo (MCMC) method. Because our model includes latent variables in addition to many parameters, the posterior distribution is too complicated to calculate analytically. Thus, the MCMC method is appropriate for our estimation.

Define $\theta = \{\theta_t\}_{t=s+1}^T$, $\alpha = \{\alpha_t\}_{t=s+1}^T$, $\varphi = \{\varphi_t\}_{t=s+1}^T$, $h = \{h_t\}_{t=s+1}^T$, $r^* = \{r_t^*\}_{t=s+1}^T$, $\Omega = (\Sigma_\theta, \Sigma_\alpha, \Sigma_\varphi, \Sigma_h)$, and $y = \{\tilde{Y}_t\}_{t=1}^T$. Given the data y and prior density function $\pi(\Theta)$, where $\Theta = (\theta, \alpha, \varphi, h, r^*, \Omega)$, the samples from the posterior distribution $\pi(\Theta|y)$ are obtained as follows:

1. Set initial values of $\theta^{(0)}$, $\alpha^{(0)}$, $\varphi^{(0)}$, $h^{(0)}$, $r^{*(0)}$, $\Omega^{(0)}$, and $j = 0$.
2. Draw $\varphi^{(j+1)}$ from $\pi\left(\varphi \mid \Sigma_\varphi^{(j)}, r^{*(j)}, h^{(j)}, y\right)$.
3. Draw $\Sigma_\varphi^{(j+1)}$ from $\pi\left(\Sigma_\varphi \mid \varphi^{(j+1)}\right)$.

4. Draw $\theta^{(j+1)}$ from $\pi\left(\theta \mid \alpha^{(j)}, h^{(j)}, \Sigma_\beta^{(j)}, y\right)$.
5. Draw $\Sigma_\theta^{(j+1)}$ from $\pi\left(\Sigma_\theta \mid \theta^{(j+1)}\right)$.
6. Draw $\alpha^{(j+1)}$ from $\pi\left(\alpha \mid \theta^{(j+1)}, h^{(j)}, \Sigma_\alpha^{(j)}, y\right)$.
7. Draw $\Sigma_\alpha^{(j+1)}$ from $\pi\left(\Sigma_\alpha \mid \alpha^{(j+1)}\right)$.
8. Draw $h^{(j+1)}$ from $\pi\left(h \mid \theta^{(j+1)}, \alpha^{(j+1)}, \varphi^{(j+1)}, \Sigma_h^{(j)}, y\right)$.
9. Draw $\Sigma_h^{(j+1)}$ from $\pi\left(\Sigma_h \mid h^{(j+1)}\right)$.
10. Draw $r^{*(j+1)}$ from $\pi\left(r^* \mid \varphi^{(j+1)}, h^{j+1}, y\right)$.
11. Return to step 2 until N iterations have been completed.

For the above process, N is set to 50,000. However, the first 20,000 samples are discarded as a burn-in. Then, every 10th draw is saved in order to alleviate the autocorrelation between the draws. Thus, the inference exhibited below is implemented based on 3,000 samples. Moreover, we only select draws in which the roots of the VAR coefficients are inside the unit circle throughout the sample period. This ensures the stationarity of the VAR system.

The sampling of φ, θ and α in steps 2, 4, and 6 are conducted using the Kalman filter and smoother, and the stochastic volatility h in step 8 is sampled using the multi-move sampler of Watanabe and Omori (2004), whose algorithm generates samples from the exact conditional posterior of h. The inverse of variances in Ω are generated from Gamma distribution under conjugate priors. The implied interest rate in the ELB period is generated from a truncated normal distribution.

2.3. Data and specification

Our empirical analysis uses quarterly data on Japan's government spending, consumer price index (CPI; all items, less fresh food), GDP, and nominal interest rate for the period 1985Q3 to 2019Q1. Figure 1 shows the time series for these data. The data on government spending and GDP are obtained from *the System of National Accounts*, published by the Cabinet Office, Government of Japan. We define the sum of government consumption and public investment as government spending. The CPI data are released by the Statistics Bureau of Japan. This price index is effectively a policy target of the BOJ. The nominal short-term interest rate is the uncollateralized overnight call rate (end of month), which was the policy rate in Japan before the adoption of the ZIRP. Because this series is available only after July 1985, our sample period starts from 1985Q3. All data other than the short-term interest rate are seasonally adjusted, and government spending and GDP are transformed into per capita values by dividing them by the total population.[3] The lag length in the TVP-VAR model, denoted by s in eq. (2), is set to three, following the lag structure in the Taylor rule eq. (5), which includes the year-on-year inflation and GDP growth rates.

[Figure 1 about here.]

The ELB periods in this study essentially adhere to the definition of Hayashi and Koeda (2019), in which a period is part of the ELB if the net policy rate (the call rate minus the rate paid on reserves) is below 50 basis points.[4] Based on this definition, we regard the periods from 1999Q1 to 2000Q1, 2001Q1 to 2006Q2, and 2008Q4 to date as the ELB periods. The BOJ has set the rate paid on the reserve to 0.1% since November 2008; thus, the periods after 2008Q4 are also categorized as ELB periods, even though the call rate is above 0.05%. In addition, the BOJ has adopted a negative interest rate policy since January 2016, imposing a negative interest rate of 0.1% on part of the excess reserve. Taking into account this transition of the BOJ's monetary policies, the lower bound of the nominal interest rate, denoted as c in eq. (1), is set to 0.05 for the periods 1999Q1 to 2000Q2 and 2001Q1 to 2006Q2, 0.1 for the period from 2008Q4 to 2015Q4, and -0.1 for the period after 2016/Q1. During the ELB period, the implied interest rate is generated from a normal distribution truncated between $-\infty$ and c, based on eq. (5).

The priors for the inverse of the ith diagonal element of $\Sigma_j (j = \theta, \alpha, \varphi, h)$ are assumed to follow an independent gamma distribution, as follows:

$$\begin{aligned} &(\Sigma_\theta)_i^{-2} \sim \text{Gamma}(10, 0.01), \quad (\Sigma_\alpha)_i^{-2} \sim \text{Gamma}(2, 0.01), \\ &(\Sigma_\varphi)_i^{-2} \sim \text{Gamma}(2, 0.01), \quad (\Sigma_h)_i^{-2} \sim \text{Gamma}(2, 0.01). \end{aligned} \tag{7}$$

[3]In addition to seasonally adjusting the data, we remove the effect of the increase in the consumption tax from the CPI. Specifically, we first perform an X-12-ARIMA, with level shift dummies that take the value one at April 1989, April 1997, and April 2014 in the original monthly CPI series. Then, we eliminate the coefficient associated with each dummy from the seasonally adjusted series, and construct the quarterly series by taking the average over three months.

[4]Iwata and Wu (2006) and Nakajima (2011) set the threshold between the ELB and non-ELB periods to 50 basis points of the nominal short-term interest rate, without paying attention to the interest rate payment on the bank reserve. Here, we follow the stricter definition adopted in Hayashi and Koeda (2019).

These priors are set based on Nakajima et al. (2011), who also estimate a four-variable TVP-VAR model with quarterly frequency for the Japanese economy. For the initial state of the time-varying parameters, we set the following priors:

$$\begin{aligned} &\theta_{s+1} \sim N(0, I), \quad \alpha_{s+1} \sim N(0, I), \\ &\varphi_{s+1} \sim N(\bar{\varphi}, I), \quad h_{s+1} \sim N(0, I). \end{aligned} \tag{8}$$

With the exception of the initial state of the coefficients in the Taylor rule, we adopt flat priors rather than the ordinary least square estimators obtained in the pre-sample, which is an alternative way of selecting the initial state of the time-varying parameters (e.g., Primiceri, 2005; Hofmann et al., 2012). We do so because we cannot exploit a sufficient size of pre-sample to estimate the model, owing to the aforementioned lack of data. On the other hand, the initial state of the parameters in the Taylor rule, denoted by $\bar{\varphi}$, is set based on the results of Clarida et al. (1998), who estimate the Japanese Taylor rule for the period April 1974 to November 1994.[5]

3. Estimation results

3.1. Parameter estimates

First, we check the convergence of our MCMC algorithm. Following Primiceri (2005), we show (a) the 20-th-order sample autocorrelation, (b) the inefficiency factor (IF), and (c) Raftery and Lewis's (1992) total number of runs for the hyperparameters (50 free elements of Σ_θ, Σ_α, Σ_φ, and Σ_h) and volatilities (four sets of volatility states for 132 periods); see Figure 2.

[Figure 2 about here.]

In Figure 2(a), most of the 20-th-order autocorrelation values are around zero, except for the volatilities of the shocks in the Taylor rule between points 447 and 578. However, even in the volatilities of the shocks in the interest rate equation, the values are all less than 0.4. Related to the autocorrelation, Figure 2(b) plots the IF for the posterior estimates of the parameters, which are defined by $1 + 2\sum_{\tau=1}^{\infty} \rho_\tau$, where ρ_τ is the sample autocorrelation at lag τ. In Figure 2(b), the IF estimates are approximated by summing

[5] Clarida et al. (1998) report the following estimates for the Taylor rule: $[\phi_0, \phi_\pi, \phi_\chi, \rho_t] = [1.21, 2.04, 0.08, 0.93]$.

the autocorrelations until lag 100. The IF is a measure of efficiency, indicating the number of samples per MCMC sample needed to obtain the same variance of the sample mean from uncorrelated draws. Most of the IFs are below 40, and even the largest is far less than 100. Thus, we consider our MCMC algorithm to be satisfactory for inference, as noted in Primiceri (2005). Furthermore, the Raftery and Lewis (1992) diagnostic of the total number of runs required to achieve a certain precision, shown in Figure 2(c), is always below the total number of iterations in our estimation.[6] In summary, the MCMC algorithm converges sufficiently and is reasonable for posterior inference.

Figure 3 reports the time profiles for the estimated structural parameters in the Taylor rule throughout the sample period. Thick and thin solid lines indicate median and 68% credible intervals, respectively, for the estimates, and the shaded areas denote ELB periods. In contrast to prior studies that estimate the Japanese Taylor rule, our estimation includes the period after the BOJ adopted the ZIRP.[7] During the ELB spells, the Tobit specification allows us to exploit the implied rate and the implied monetary policy rule the monetary authority would adopt if there were no ELB on the interest rate. In the top-left chart in Figure 3, we show the variation of the time-varying intercept, denoted by $\psi_{0,t}$. As originally stated in Taylor (1993), the intercept in the Taylor rule stands for the equilibrium real interest rate, and is set to be consistent with the growth rate in the steady state, that is, the potential growth rate of the economy. Considering the dynamics of $\psi_{0,t}$ from the viewpoint of Taylor's (1993) original formulation, our estimate can be interpreted as reflecting the decrease in the total factor productivity growth rate documented by Hayashi and Prescott (2002). In Figures 3(b)–(d), we plot the dynamics of the coefficients for inflation and output and the speed adjustment parameters, respectively. The ranges of the estimated parameters are essentially compatible with the estimates reported in the literature. Specifically, as summarized in Miyazawa (2011), previous studies estimate the inflation coefficient, output coefficient, and inertial parameter as lying in the ranges of [1.33, 3.84], [0.17, 1.001], and [0.55, 0.93], respectively; these ranges are indicated

[6] Following Primiceri (2005), the parameters used to specify Raftery and Lewis's (1992) diagnostics set are as follows: quantile = 0.025; desired accuracy = 0.025; and required probability of attaining the required accuracy = 0.95.

[7] As noted in Miyazawa (2011), the sample periods of most studies stop before the adoption of the ZIRP. One exception is the work of Hayashi and Koeda (2019), whose sample period covers 1992 to 2012.

by dotted lines in Figure 3. In addition, Hayashi and Koeda (2019) find that the inflation and output coefficients are 0.69 and 0.05, respectively, and the speed of adjustment is estimated to be 9.8% (i.e., $\rho = 0.902$) for the sample period 1992–2012.

Note that our results exhibit substantial variations in the values of the structural parameters over time, which are not considered in previous studies. The estimation results for $\psi_{\pi,t}$, $\psi_{x,t}$, and ρ_t show trend changes around the mid-1990s, when the interest rate falls to near zero. The inflation coefficient then becomes small and stable, while the output coefficient becomes more volatile. A possible reason for this result is that the variation in inflation is relatively smaller than that of output during the (two) lost decades in Japan. In contrast, it is easy to interpret the dynamics of the inertial parameter. The parameter takes the bottom value around 1995 when the BOJ frequently cut the policy rate to cope with the economic slowdown stemming from the burst of the asset price bubble in the early 1990s. Overall, the estimation results emphasize the variations of the parameters over time in the Taylor rule, and thereby support the time-varying formulation of our estimation.

[Figure 3 about here.]

Figure 4 depicts the estimated series of the implied interest rate (solid line), as well as the actual rate (dotted line). The figure shows that the policy rate takes negative values if there is room to cut the rate in the ELB period. In other words, this result implies that the interest rate may fluctuate in response to structural shocks in the ELB period if there were no lower bound. For comparison purposes, we plot the shadow rate estimated by Krippner (2015) (dot–dash line). The negative interest rate in this study is constructed using the Taylor rule, whereas Krippner (2015) estimates the rate using the term structure model and the information included in the yield curve.[8] To highlight the differences between the estimation methods and to emphasize that our estimated rate is the hypothetical rate set by the central bank if no lower bound exists, as in Iwata and Wu (2006), we refer to our estimated negative interest rate as the implied rate rather than the shadow rate. In spite of the methodological differences, the variations of our implied rate are fairly close to the shadow rate estimated in Krippner (2015), until the second ELB period. However, the gap between the two gradually becomes wider in the

[8]Ueno (2017) also estimates the shadow rate in Japan using the term structure model.

third ELB period. Our implied rate decreases rapidly in response to the significant drop in the GDP growth rate just after the global financial crisis; however, no such decrease is observed in the estimates of Krippner (2015). Moreover, in contrast to our estimates, the shadow rate of Krippner (2015) describes a persistent downward trend in the third ELB period, which coincides with the BOJ's adoption of quantitative and qualitative easing with yield curve control, aimed at shifting the entire yield curve downward. Although structural models are often used to estimate the shadow rate under the lower bound of the nominal interest rate (Krippner, 2015; Ueno, 2017), Iwata and Wu (2006) and Nakajima (2011) estimate the implied rate using the Tobit model and the (reduced-form) monetary policy rule. As stated above, the primary purpose of this study is to empirically check whether the theoretical predictions presented in Eggertsson (2010) and Christiano et al. (2011) hold. Thus, we regard our implied rate estimated from the structural Taylor rule as a counterpart of the policy rates discussed in Eggertsson (2010) and Christiano et al. (2011), and use this estimate in our analysis.

[Figure 4 about here.]

3.2. Impulse responses

To examine the effect of a government spending shock on output, we derive the IRs from the estimated TVP-VAR model. Owing to the nonlinearity of the TVP-VAR model with a censored variable, the IRs in this study are calculated using the generalized IR developed by Koop et al. (1996). In general, the generalized IRs in period $t+j$ in response to the shock that occurred in period t are defined as the difference $E[y_{t+j} \mid \Omega_{t-1}, \tilde{\varepsilon}_t^1] - E[y_{t+j} \mid \Omega_{t-1}, \tilde{\varepsilon}_t^0]$, where Ω_{t-1} is the information set in period $t-1$, $\tilde{\varepsilon}_t^1$ is the nonzero shock of interest, and $\tilde{\varepsilon}_t^0$ denotes a zero shock in period t. In our case, the counterpart of y_{t+j} is $\tilde{Y}_{t+j} = [g_{t+j}, p_{t+j}, x_{t+j}, r^*_{t+j}]'$, the explanatory variables in the TVP-VAR model and in the Taylor rule in period t (i.e., X_t in eq. (3), and π_t, χ_t, and r_{t-1} in eq. (5), respectively) correspond to the initial condition Ω_{t-1}, and we set $\tilde{\varepsilon}_t^1 = [\varepsilon_t^{1\prime}, \varepsilon_{r,t}^1]' = (1,0,0,0)'$ and $\tilde{\varepsilon}_t^0 = [\varepsilon_t^{0\prime}, \varepsilon_{r,t}^0]' = (0,0,0,0)'$, because we are interested in the effect of a government spending shock. For each outcome of the Gibbs sampler, we randomly generate future shocks, $\tilde{\varepsilon}_{t+j}(j=1,\cdots,h)$, from $N(0,I)$, and then calculate y_{t+j}^1 conditional on $(\Omega_{t-1}, \tilde{\varepsilon}_{t+j}, \tilde{\varepsilon}_t^1)$, and y_{t+h}^0 conditional on $(\Omega_{t-1}, \tilde{\varepsilon}_{t+j}, \tilde{\varepsilon}_t^0)$. Repeating this process 100 times and averaging the outcomes, we obtain estimates of the conditional expectation, denoted by $\hat{y}^1_{t+j}$ and $\hat{y}^0_{t+j}$, where we regard the differences between $\hat{y}^1_{t+j}$ and $\hat{y}^0_{t+j}$ as estimates of the generalized IRs.

Specifically, the calculation for the generalized IRs in period $t + j$ to the shock in period t is implemented as follows. Given the outcome of the Gibbs sampler and the initial condition in period t, we first obtain a pair of values for $[\Delta g_t, p_t, x_t]'$ under both nonzero and zero shocks, based on the VAR system. Second, the implied interest rates (r^*_t) under nonzero and zero shocks are derived from the Taylor rule, using χ_t and π_t, respectively, obtained in the first step, where $\chi_t = x_t+x_{t-1}+x_{t-2}+x_{t-3}$ and $\pi_t \equiv p_t+p_{t-1}+p_{t-2}+p_{t-3}$ are constructed using x_t, p_t, and the initial condition. Then, we check whether the computed implied rate is larger than the lower bound (c). The observed interest rate (r_t) is set to the implied rate (r^*_t) if the implied rate exceeds the lower bound, and is set to the lower bound (c) otherwise. Subsequently, a pair of values of endogenous variables in period $t+1$ are computed using the values in period t obtained above. We repeat this process until the j-period ahead to obtain the responses in period $t + j$.

The generalized IRs are defined as the differences between the paths of the endogenous variables under the two situations (i.e., nonzero and zero shocks), such that the estimated response of the interest rate (r_t) is zero if the observed rates in both situations are set to c, given that $r_t = \max[c, r^*_t]$. When the implied rate is replaced by c, the responses of other endogenous variables are considered to be those observed in an economy in which the ELB constraint is valid (i.e., $r_t = c$). This is because the VAR system described in eq. (3) includes the implied rate and the observed rate on the right-hand side to calculate the dynamics of the variables. Indeed, the Japanese economy has fallen within the ELB for some time; thus, we regard the IRs calculated under the condition $r_t = \max[c, r^*_t]$ as a benchmark, in the sense of capturing the actual dynamics of the Japanese economy. On the other hand, as a counterfactual simulation, we calculate the hypothetical IRs using the implied interest rate rather than the lower bound of the nominal interest rate, even if the ELB restriction is binding. In contrast to the benchmark, the responses of the interest rate in this case are not zero, even when the economy hits a lower bound, because the IRs are defined as the differences between the calculated implied rates of the nonzero and zero shocks. Comparing the economy with a government spending shock (i.e., a nonzero shock) to that without a shock (i.e., a zero shock), there may be a difference between the values of the implied rates, both of which are replaced in c if we follow the condition $r_t = \max[c, r^*_t]$.

Moreover, we calculate the responses of the other variables by incorporating the implied rate, rather than the lower bound value, on the right-hand side of the equation. This counterfactual analysis reveals the effect of a government spending shock on the interest rate, and the effect of the variation in the interest rate on macroeconomic variables, which are hidden by the ZIRP. In summary, the IRs in the benchmark model represent the actual situation of the ELB period, in which the interest rate is set to the lower bound. In contrast, the IRs in the counterfactual simulation describe the hypothetical dynamics of the variables in the sense that we allow the interest rate to fluctuate beyond the lower bound. Comparing the IRs derived from the benchmark with those derived from the counterfactual, we extract the role of the ELB in the effectiveness of a government spending shock on the economy.

Note that our counterfactual simulation is conducted under the condition that the coefficients in the three reduced-form equations are invariant, even in the face of a changing monetary policy stance. In the strict sense, this kind of simulation is incompatible with the Lucas critique; however, it is also true that such simulations are widely employed in the literature to evaluate the effects of policy changes (e.g., Sims and Zha, 2006; Baumeister and Benati, 2013; Ramey, 2013). Nevertheless, we calculate the hypothetical paths for each variable up to at most two-years after the shock, based on the notion that this sort of counterfactual simulation is justified only when agents' beliefs are unchanged, as discussed above.

In what follows, we show the fiscal multiplier rather than the IRs in order to align the size of the effect of the government spending shock on output in each period. The k-period-ahead fiscal multiplier to the shock at period t is computed as

$$\text{fiscal multiplier}_{k,t} = \frac{\sum_{j=0}^{k} \text{IR of } x_{t+j} \times \text{GDP}_t}{\sum_{j=0}^{k} \text{IR of } g_{t+j} \times \text{gov. spending}_t}, \tag{9}$$

because the data on the GDP and government spending are included in the first difference of the natural logarithm within the VAR model. If the multipliers derived from the benchmark are greater than those from the counterfactual, this implies that a government spending shock has a larger effect on the economy when the ELB constraint is binding, and vice versa. In addition, the responses of the interest rate and price are normalized to 1% of the government spending shock at the impact period.

We begin by examining the contemporaneous responses of each variable to a government spending shock. Figure 5 illustrates the time profiles of the contemporaneous responses of the (a) fiscal multiplier, (b) price, (c) nominal interest rate, and (d) ex post real interest rate,[9] where solid thick and thin lines denote the median and 68% credible intervals, respectively, corresponding to the benchmark. Because we assume that the interest rate has no contemporaneous effects on government spending, inflation, or GDP, there is no difference between the benchmark and the counterfactual at the impact period. The counterfactual responses of the interest rates, for which the differences emerge instantly, are drawn by dashed lines. First, we can see from Figure 5 that the estimated fiscal multiplier and responses exhibit remarkable variation over time throughout the sample period, supporting our application of the TVP-VAR model. The variation of the multiplier shows a rapid drop after the economic bubble burst around 1990, a continuing sluggishness around the first-half of the 2000s, and a recovery after the second-half of the 2000s. Although several researchers have pointed out that the effectiveness of fiscal policy in Japan has diminished since the bubble burst, its recent revival is a noteworthy finding of this study. With regard to price, we observe that the government spending shock induces inflation (the standard theoretical view), but that exceptional periods exist, during which the price decreases to the shock (e.g., the end of the second and third ELB periods). The price's negative response to a government spending shock seems contrary to the conventional view that a government spending shock is inflationary because of the demand shock. However, empirical evidence shows a decrease of the price in response to the shock (e.g., Mountford and Uhlig, 2009; Fatas and Mihov, 2001).[10] As shown in Figure 5(c) and (d), such exceptional price behavior causes the negative response of the nominal implied rate. Furthermore, it creates a situation in which the response of the real interest rate in the counterfactual becomes lower than that derived in the benchmark. However, given this exception, the result suggests that monetary tightening would occur in response to an inflationary government spending shock if no ELB constraint exists. Therefore, the ZIRP plays a significant role in suppressing the real interest rate in such circumstances.

[Figure 5 about here.]

[9] The response of ex post real interest rate is defined as the difference between the response of the nominal interest rate and current inflation.

[10] Jørgensen and Ravn (2018) recently proposed a theoretical model that replicates a drop in price as a result of a government spending shock by incorporating variable technology utilization in the standard New Keynesian model.

Figure 6(a) displays the time profiles of the fiscal multipliers for (i) one-quarter, (ii) one-year, and (iii) two-year horizons. As in Figure 5, the solid lines correspond to the benchmark, represented by $r_t = \max[c, r_t^*]$. The dashed lines correspond to the results from the counterfactual, represented by $r_t = r_t^*$, and indicate ELB periods as shaded areas. To highlight the role of the ELB in the magnitude of the fiscal multiplier, we plot the point estimates for the fiscal multipliers under the two cases in Figure 6(a). Instead of plotting the credible intervals, Figure 6(b) shows the posterior ratio such that the size of the multiplier in the benchmark is larger than that in the counterfactual simulation for each horizon in each period. We interpret this as follows. The difference between the fiscal multipliers in the two simulations becomes statistically more significant as the probability increases. Accordingly, a high probability implies that it is more plausible that the effectiveness of a government spending shock on output is enhanced by keeping the interest rate within the lower bound. Because the same values of future shocks $\tilde{\varepsilon}_{t+j}$ are shared when calculating the generalized responses in both cases, the multiplier in the benchmark coincides perfectly with that in the counterfactual, except for the ELB period. Thus, the posterior probabilities are calculated for the ELB period only. In addition, Figure 7 shows the responses of the (a) nominal interest rate, (b) price, and (c) real interest rate at each horizon, where the response of the price is defined as the cumulative response of inflation.

We first observe from Figures 6 and 7 that the estimated multipliers in the benchmark are larger (smaller) than those in the counterfactual when the response of the real interest rate in the former is smaller (larger) than that in the latter.[11] For example, the positive gap between the two multipliers, which is more evident after one and two years, is observed in the period just after the global financial crisis. During this period, the responses of the real interest rate in the benchmark are considerably suppressed compared with those of the counterfactual. This proves empirically that the theoretical mechanism of Eggertsson

[11] At first glance, our finding seems to support those of Miyamoto et al. (2018), who find that the multiplier in the ELB period is larger than that in the non-ELB period. However, there is a methodological difference between the work of Miyamoto el al. (2018) and this study. The former compares the multipliers derived from the non-ELB period (1980Q1–1994Q4) and the ELB period (1995Q1–2014Q1). Here, we estimate two values of the fiscal multipliers in the same period. The advantage of our method is that we control for factors other than the monetary policy stance, because the difference between the benchmark and the counterfactual lies only in how we determine the interest rate when we calculate the IRs.

(2010) and Christiano et al. (2011) holds in practice. Because the dynamics of the real interest rate stem from those of the nominal interest rate and inflation, this finding can be interpreted as follows. The enhancement in the fiscal multiplier is driven by the absence of monetary tightening in an ELB period, where the upward pressure on the nominal interest rate accompanied with the positive responses of price and output is neutralized by the ZIRP. Therefore, we conclude that, as predicted by theory, the decrease in the real interest rate is crucial to increasing the effect of a government spending shock on output during an ELB period. In addition to the fiscal multiplier (see Figure 7(b)), the price responses in the benchmark model are located above those in the counterfactual, as long as the positive response of the nominal interest rate observed in the counterfactual is offset by fixing the interest rate to zero. In this respect, our result agrees with the standard theoretical prediction that an increase in the nominal interest rate mitigates an increase in price.

Finally, note that there is a statistical vulnerability in the difference between the benchmark and the counterfactual. The posterior probabilities shown in Figure 6(b) take values around 0.6 and, at most, less than 0.7, even when the median of the benchmark exceeds that of the counterfactual. Compared with the significance level adopted in a standard inference, the ratio obtained in this study is not sufficient to reject the null hypothesis that a statistical difference between them does not exist. In other words, it is possible that the credible intervals for each IR, which are not drawn in Figure 6(a), overlap. This statistical weakness arises from our TVP-VAR specification, because the TVP-VAR model contains many parameters that need to be estimated from a limited sample of time-series data. Thus, the credible intervals associated with the IRs tend to be wider. Because the variation of the economy over time, captured by the TVP-VAR model, seems to be supported by the results thus far, we prioritize the time-varying specification over statistical accuracy.

[Figure 6 about here.]

[Figure 7 about here.]

4. Robustness check

4.1. Different definition of the ELB period

To check that our result is robust, without depending on the definition of the ELB, we divide the sample period into non-ELB and ELB periods based on the definition adopted by Miyamoto et al. (2018). Thus, we employ 1% as the threshold value to separate the sample period. Following this definition, we estimate the model by resetting the lower bound to 1% (i.e., $c = 1$). In what follows, we refer to the results in Section 3 as the baseline.

Figure 8 shows the implied interest rate estimated based on the new definition, and that of the baseline. The median and 68% credible intervals of the new model are described by dot-dashed and thin lines, respectively; the estimate for the baseline is denoted by a solid line. In this robustness check, the period after 1995Q4 is categorized as the ELB period, as indicated by shaded area, and the implied interest rate during the ELB period is drawn from a truncated normal distribution between $-\infty$ and 1. Compared with the baseline estimate, the newly estimated implied rate tends to show values that are more negative. However, its behavior is similar to that of the baseline model.

[Figure 8 about here.]

In addition to the implied rate, Figure 9 shows the time profiles of the fiscal multiplier and the posterior probability, as calculated in the baseline. For comparison, Figure 9(b) plots the posterior probabilities obtained in the baseline and in the robustness check. Obviously, there is no quantitative or qualitative difference between these results. Therefore, our main result, that the multiplier derived from the benchmark economy is larger than that derived from the counterfactual if a government spending shock is inflationary, is robust, even if we change the definition of the ELB period following that of Miyamoto et al. (2018).

[Figure 9 about here.]

4.2. Different measure of output

Following the literature on fiscal policy (e.g., Blanchard and Perotti, 2002), in the baseline formulation, we use the GDP growth rate in the VAR model and, thus, the year-on-year GDP growth rate in the Taylor rule. On the other hand, the standard specification of the Taylor rule contains the output gap, not the GDP growth rate (Taylor, 1993; Clarida et al., 1998; Hayashi and Koeda, 2019). Following the literature on monetary policy, we conduct an alternative estimation using the GDP gap as the output measure to confirm the robustness of our results. To do so, we slightly modify the specification of the Taylor rule, as follows:

$$r_t^* = (1 - \rho_t)[\psi_{0,t} + \psi_{\pi,t}\pi_t + \psi_{z,t}z_t] + \rho_t r_{t-1} + \varepsilon_{r,t}. \tag{10}$$

Here, we directly incorporate the GDP gap, denoted by z_t, into the right-hand side of the equation. The GDP gap is a series of log-differentials between the actual and potential GDP; we use the GDP gap released by the Cabinet Office, Government of Japan, as shown in Figure 10. Similarly to the year-on-year GDP growth rate, the GDP gap exhibits slow-moving variation. Therefore, we use the GDP gap directly in the Taylor rule.

[Figure 10 about here.]

Figure 11 shows the variation in the fiscal multipliers over time and the posterior probability that the value of the multiplier in the counterfactual exceeds that of the benchmark in each period; this is the counterpart of Figure 6. The difference between the multipliers of the benchmark and the counterfactual is clearly evident compared with the results in Figure 6. As a result, the posterior probabilities displayed in Figure 11(b) take values greater than those in the baseline for most of the ELB period. This suggests that the difference between the benchmark and counterfactual is likely to be more reliable when using the GDP gap as an output measure. However, although we acknowledge the quantitative differences in the posterior probabilities, depending on the output measure, the variations over time are qualitatively the same between the baseline and the robustness check. Moreover, the positional relationships between the multipliers described in Figure 11(a) are similar to those in the baseline estimation. That is, the multiplier in the benchmark is larger than that in the counterfactual when a government spending shock causes inflation. Hence, we conclude that the main results do not depend on the output measure.

[Figure 11 about here.]

4.3. Effect of a public investment shock

In terms of an economic stimulus package, the effectiveness of a public investment shock is more relevant in practice than that of overall spending, because fiscal policy aimed at an economic stimulus is implemented mainly through an increase in public investment.[12] In addition, the theoretical size of the multiplier associated with public investment under the ELB is controversial, as discussed in Albertini et al. (2014) and Bouakez et al. (2017). Because public investment is productive spending in the sense that it contributes to the accumulation of public capital, it has supply-side effects of reducing the real marginal cost and inflation. When the nominal interest rate is fixed to a lower bound under the ELB, deflation causes an increase in the real interest rate, and then attenuates the effect of the fiscal policy shock. Albertini et al. (2014) reports that the multiplier in the ELB decreases steadily as the proportion of productive spending increases. Bouakez et al. (2017) shows that the multiplier associated with public investment may become large when the time-to-build delay is comparatively long. This is because the demand-side inflationary effect stemming from an increase in public investment exceeds the supply-side deflationary pressure, which is brought about by an increase in public capital. In order to address the above-mentioned academic and practical concerns, we replace government spending with public investment in the following estimation.

Similarly to the baseline estimation, Figures 12, 13, and 14 illustrate the time profiles of the contemporaneous responses of each variable, the fiscal multiplier and posterior probability, and the responses of the nominal interest rate, price, and real interest rate, respectively. For the comparison with the baseline model, we plot the posterior probability obtained in the baseline estimation using dashed lines in Figure 13 (b). In addition, the IRs of the price and nominal and real interest rates are normalized such that the response of public investment becomes 1% at the impact period.

[Figure 12 about here.]

[Figure 13 about here.]

[Figure 14 about here.]

Although the overall behavior of the multiplier, as shown in Figures 12 and 13, exhibits a U-shape, as in the baseline, two features differ from the baseline: the negative values of the multiplier in the period from the end of 1990s to the mid-2000s, and a sharp

[12] Morita (2017) documents the amounts of public investment in the recent Japanese fiscal stimulus packages in detail.

increase in the multiplier during the third ELB period. These findings suggest that the sluggishness in the multiplier during the lost decades and its recent revival observed in the baseline are mainly attributed to the effect of public investment shocks. This emphasizes the importance of a public investment shock as an economic stimulus policy. Focusing on the difference between the benchmark ($r_t = \max[c, r_t^*]$) and the counterfactual ($r_t = r_t^*$) during the ELB period, Figure 14(b) shows that a public investment shock is more inflationary than an overall government spending shock (Figure 7(b)). This empirical finding supports the theoretical prediction of Bouakez et al. (2017) rather than that of Albertini et al. (2014). More precisely, the demand-side inflationary effect of a public investment shock seems to dominate the supply-side deflationary pressure, at least in Japan. This inflationary effect generates a positive gap in the multiplier for most of the ELB period. This includes the end of the sample period, during which a negative gap emerges in the baseline owing to the negative response of inflation to the shock. In other words, the effectiveness of a public investment shock is amplified under the ELB. At the same time, the crowding-out effect observed in the lost decades is mitigated, owing to its inflationary effect. Indeed, as shown in Figure 14(c), a public investment shock reduces the real interest rate in the benchmark compared with that in the counterfactual. This indicates that the necessary condition for enhancing the fiscal multiplier under a low interest rate is also satisfied if we use public investment rather than total government spending. However, note that the credible intervals associated with the contemporaneous fiscal multiplier shown in Figure 12 are wider than those in the baseline estimation. This implies that the fiscal multiplier when using public investment entails greater uncertainty than when using overall spending. Consequently, as shown in Figure 13(b), the posterior probabilities, which highlight the role of the ZIRP in the effectiveness of fiscal policy, takes comparatively lower values than those of the baseline results.

5. Conclusion

This study uses a TVP-VAR model with an implied interest rate to empirically analyze whether the effect of a government spending shock on output depends on the monetary policy in Japan. To achieve this purpose, we compute two types of IRs based on the estimation results: a benchmark, which calculates the IRs after replacing the interest rate with the lower bound value if the implied rate hits the ELB; and a counterfactual, which computes the IRs using the implied interest rate as an endogenous variable, even if the rate takes a negative value. The gap between the two IRs isolates the role of the ZIRP in the effect of a fiscal policy shock, because the difference depends only on how we determine the interest rate under the ELB.

The results reveal that the fiscal multiplier becomes larger as a result of the decrease in the real interest rate when we eliminate the implied fluctuation of the nominal interest rate by remaining within the lower bound. In other words, the multiplier derived from the benchmark is greater than that derived from the counterfactual when a government spending shock is inflationary. Moreover, this main result remains qualitatively unchanged for different definitions of the ELB period, output, and fiscal variable. Therefore, we conclude that the effectiveness of a government spending shock is enhanced when the central bank adopts the ZIRP. Furthermore, this empirical evidence supports the theoretical findings of Eggertsson (2010) and Christiano et al. (2011).

In concluding this paper, we mention the foreseeability of fiscal policy, as emphasized by Ramey (2011).[13] This study does not consider fiscal foresight, because its primary aim is to examine whether the theoretical prediction of the effect of fiscal policy under the ELB is plausible. In practice, it is difficult to incorporate additional variables, which can capture a fiscal news shock, into our TVP-VAR model, given the estimation accuracy. However, Fisher and Peters (2010) propose that industries' excess stock returns related to government spending can be regarded as a proxy for a fiscal news shock. Nonetheless, it is important to check whether the results presented here are robust after considering fiscal foresight. This topic is left to future work.

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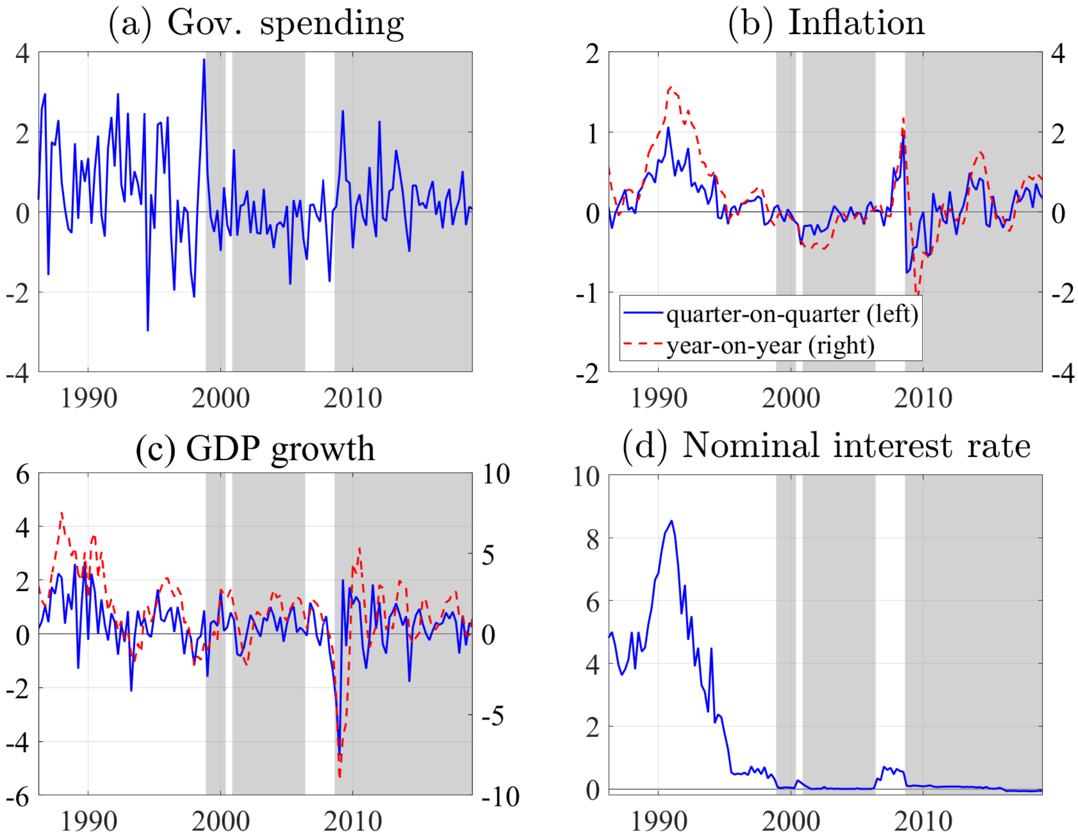

Figure 1: Data description

Note: Except for the short-term interest rate, solid lines indicate the quarter-on-quarter growth rate of each variable. The dashed lines in Figures 1(b) and (c) present the year-on-year growth rates of the GDP and CPI, respectively. In addition, the shaded areas correspond to the ELB period.

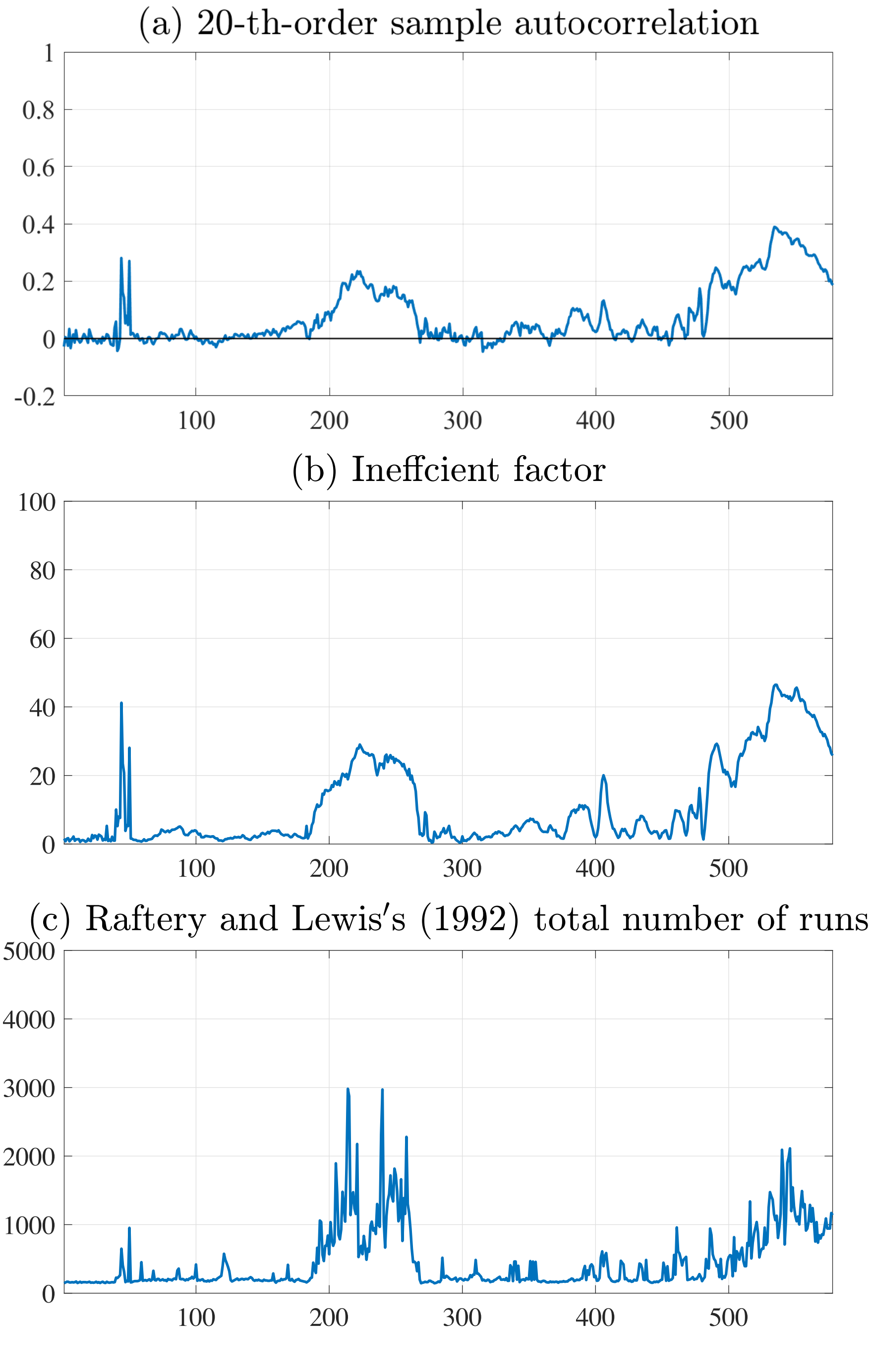

Figure 2: Convergence diagnostics for hyperparameters and volatilities

Notes: This figure shows (a) the 20-th-order sample autocorrelation, (b) the inefficiency factor, and (c) Raftery and Lewis's (1992) total number of runs.

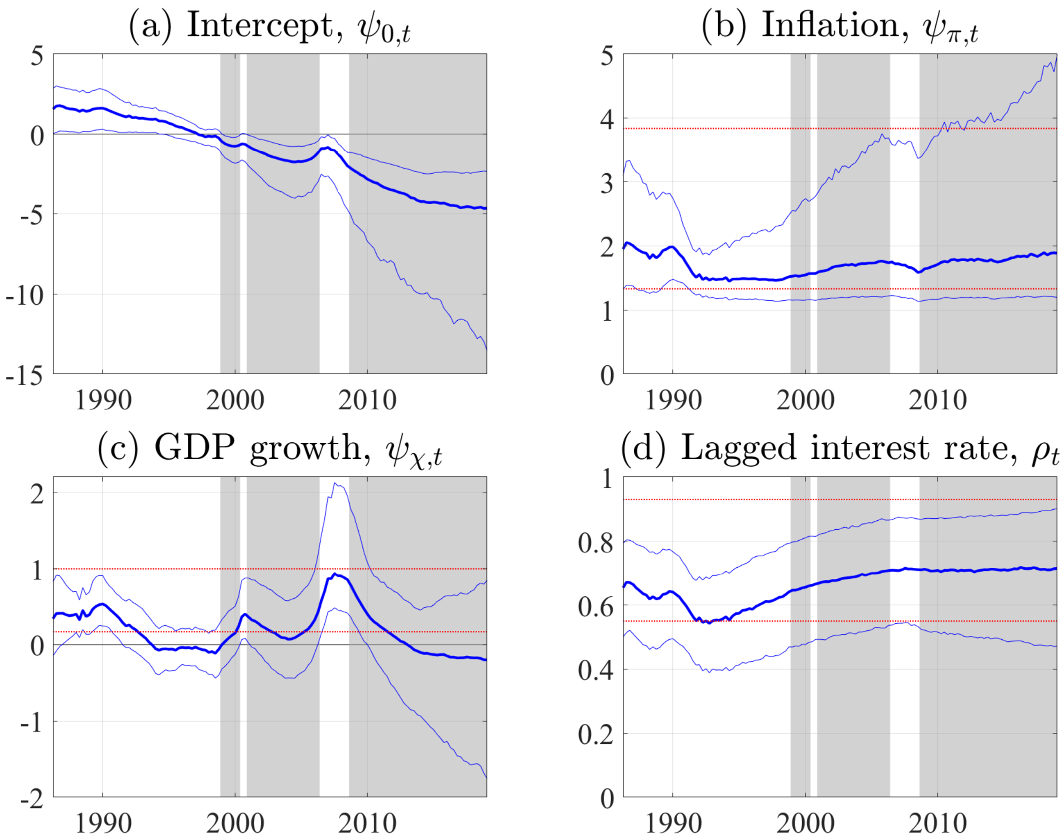

Figure 3: Estimated Taylor rule parameters

Notes: This figure shows the time profiles of the structural parameters in the Taylor rule in Eq. (5). Solid thick and thin lines correspond to the median estimates and 68% credible intervals, respectively. In addition, dotted lines represent the ranges of the estimated parameters reported in previous studies.

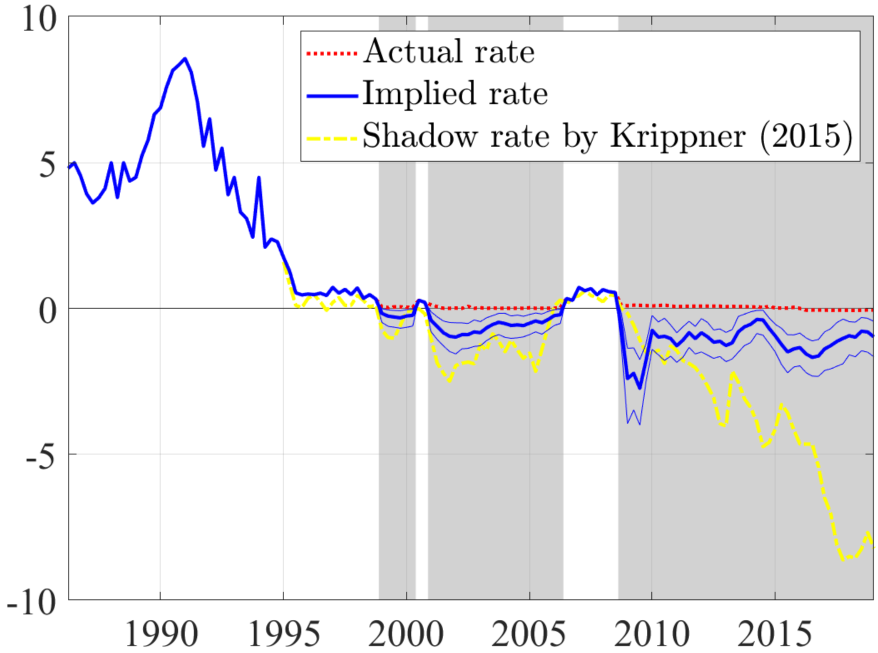

Figure 4: Estimated implied rate

Notes: The figure shows the actual short-term interest rate (dotted line), estimated implied rate (solid thick line), and 68% credible intervals associated with the implied rate (solid thin line). For comparison purposes, we also plot the shadow rate (dot-dashed line) estimated by Krippner (2015).

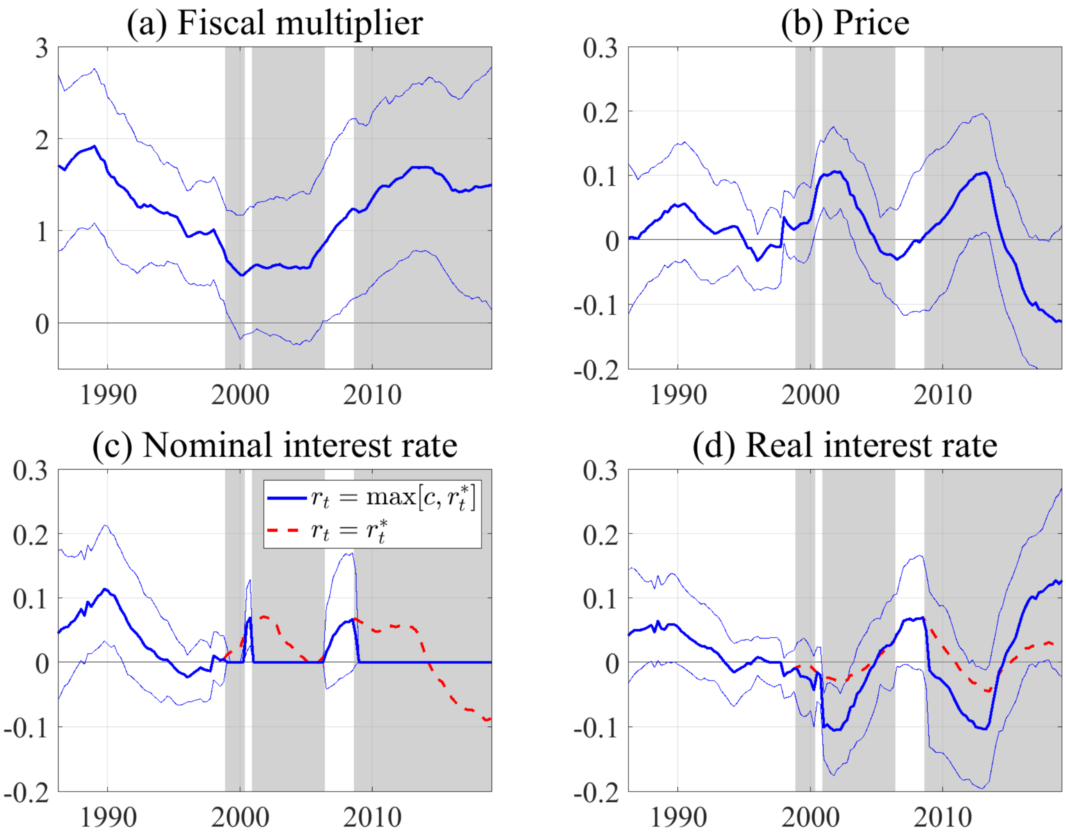

Figure 5: Time-varying responses of fiscal multiplier, price, and nominal and real interest rates at impact period

Notes: The figures show the time profiles of the (a) fiscal multiplier, as well as the responses of the (b) price, (c) nominal interest rate, and (d) ex post real interest rate at the impact period. The ex post real interest rate is calculated by subtracting the response of inflation from the nominal interest rate. The solid thick and thin lines correspond to the median and 68% credible intervals, respectively, in the benchmark model. The dashed lines in Figure 5(c) and (d) indicate the responses in the counterfactual simulation.

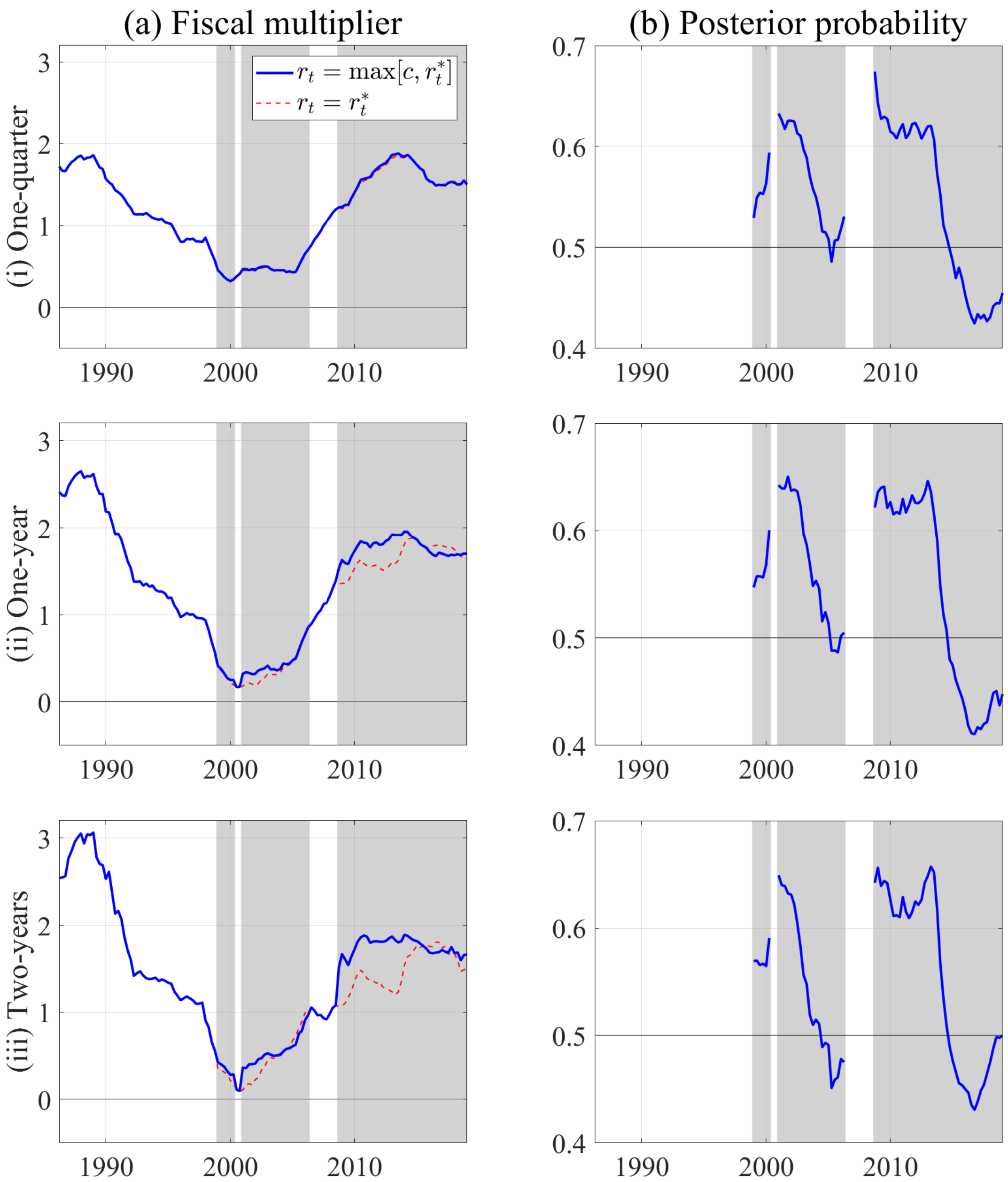

Figure 6: Fiscal multiplier and posterior probability

Notes: Figure 6(a) shows the time profiles of the fiscal multipliers derived from ELB periods (solid lines) and non-ELB periods (dashed lines) at (i) one quarter, (ii) one year, and (iii) two years after the shock. Figure 6(b) shows the corresponding posterior probabilities defined as the ratio of such samples as the multiplier in an ELB period is greater than that in a non-ELB period. Because the differences between the ELB and non-ELB multipliers only happen in the ELB periods (by construction), the time profiles of the probabilities are depicted for the ELB periods only.

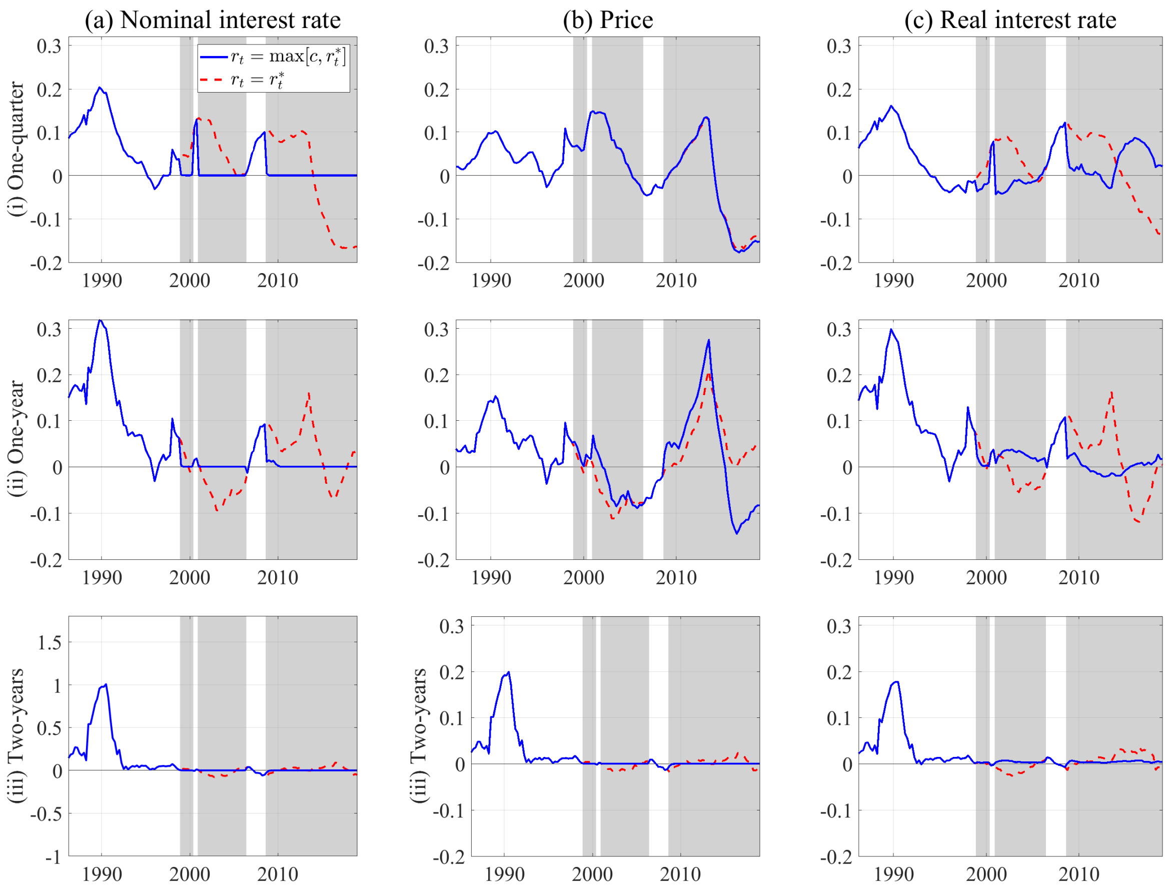

Figure 7: Time-varying responses of nominal and real interest rates and price

Notes: This figure shows the time profiles of the responses of the (a) nominal interest rate, (b) price, and (c) ex post real interest rate to a government spending shock. Here, the response of the price is defined as the accumulated response of inflation. The ex post real interest rate is calculated by subtracting the response of inflation from that of the nominal interest rate. In each figure, the solid and dashed lines denote responses derived from the benchmark and counterfactual, respectively.

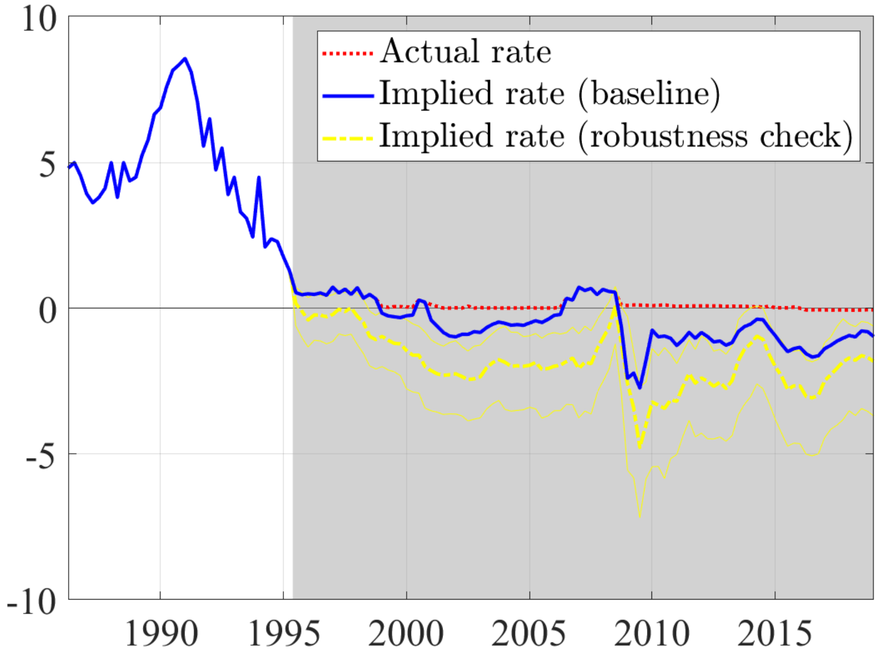

Figure 8: Estimated implied rate: $c = 1\%$

Notes: This figure plots the estimated implied rate under various definitions of the ELB. The dot-dashed line is the implied rate when we regard 1% as the threshold value of the ELB.

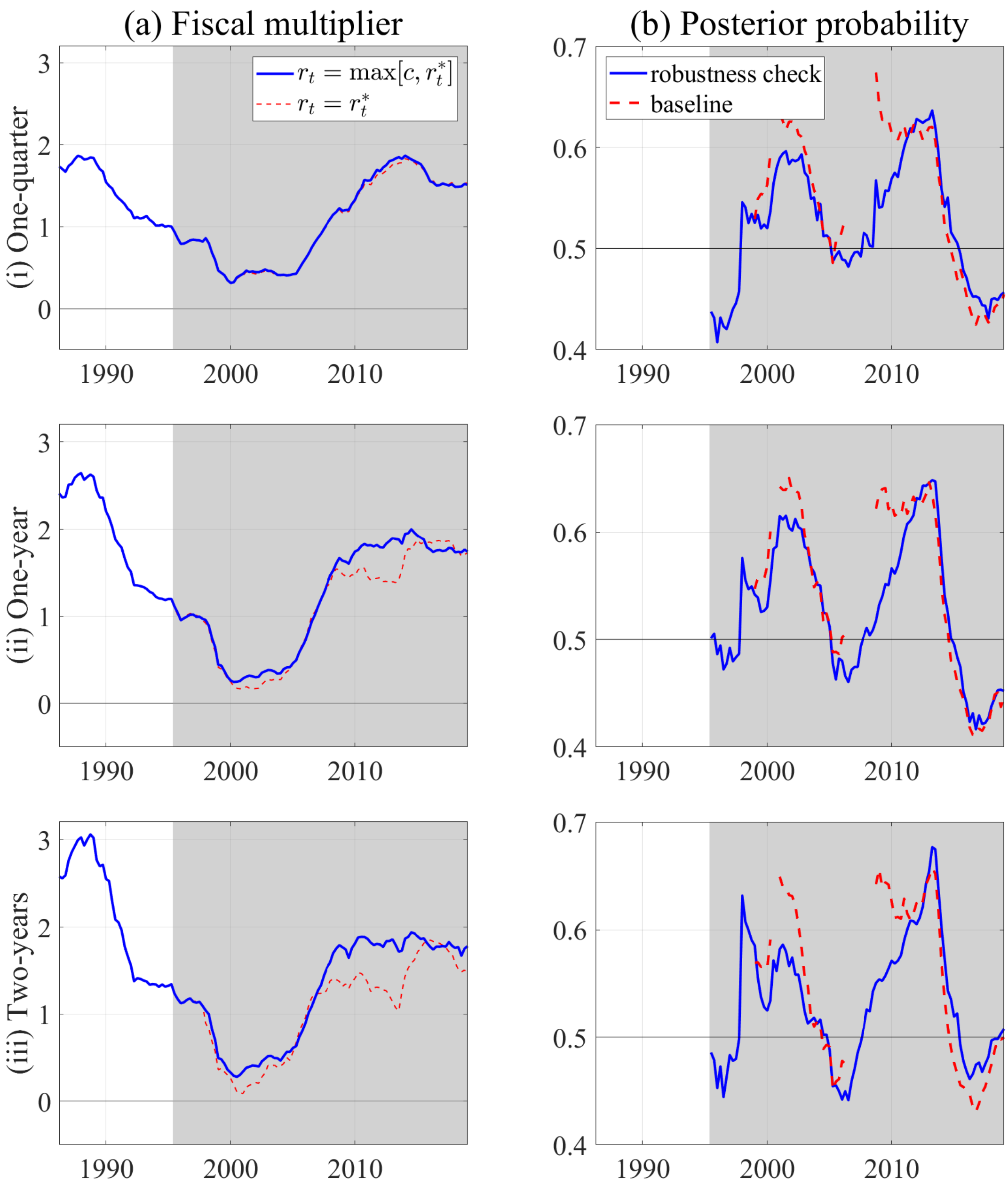

Figure 9: Fiscal multiplier and posterior probability: $c = 1\%$

Notes: This figure shows the time profiles of the fiscal multipliers and the posterior probabilities for c = 1%.

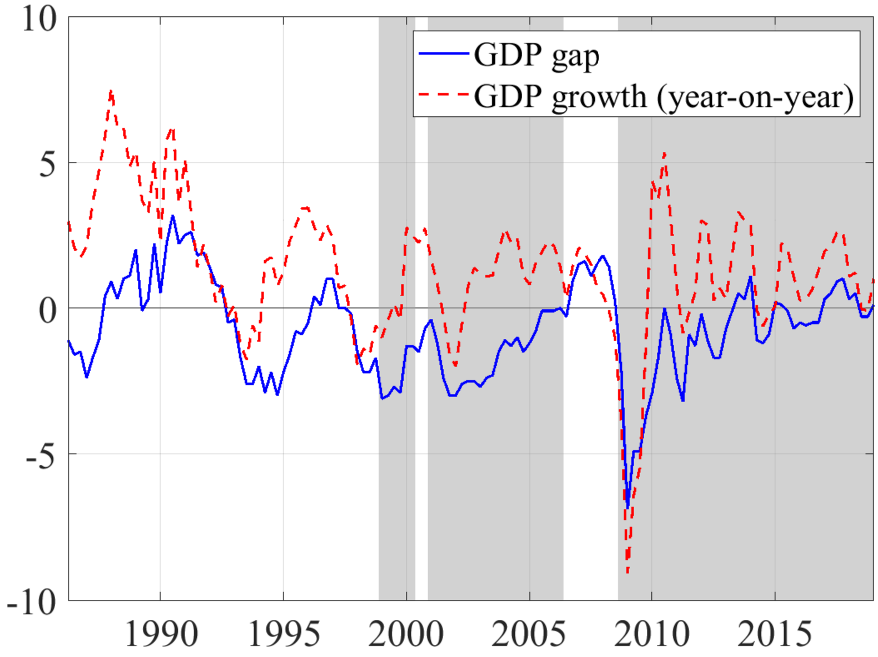

Figure 10: Data for the GDP gap

Notes: The solid line shows the evolution of the GDP gap reported by the Cabinet Office, Government of Japan. For comparison purposes, we plot the year-on-year GDP growth rate used in the baseline (dashed line). In the robustness check, we estimate the model by replacing the GDP growth rate with the GDP gap.

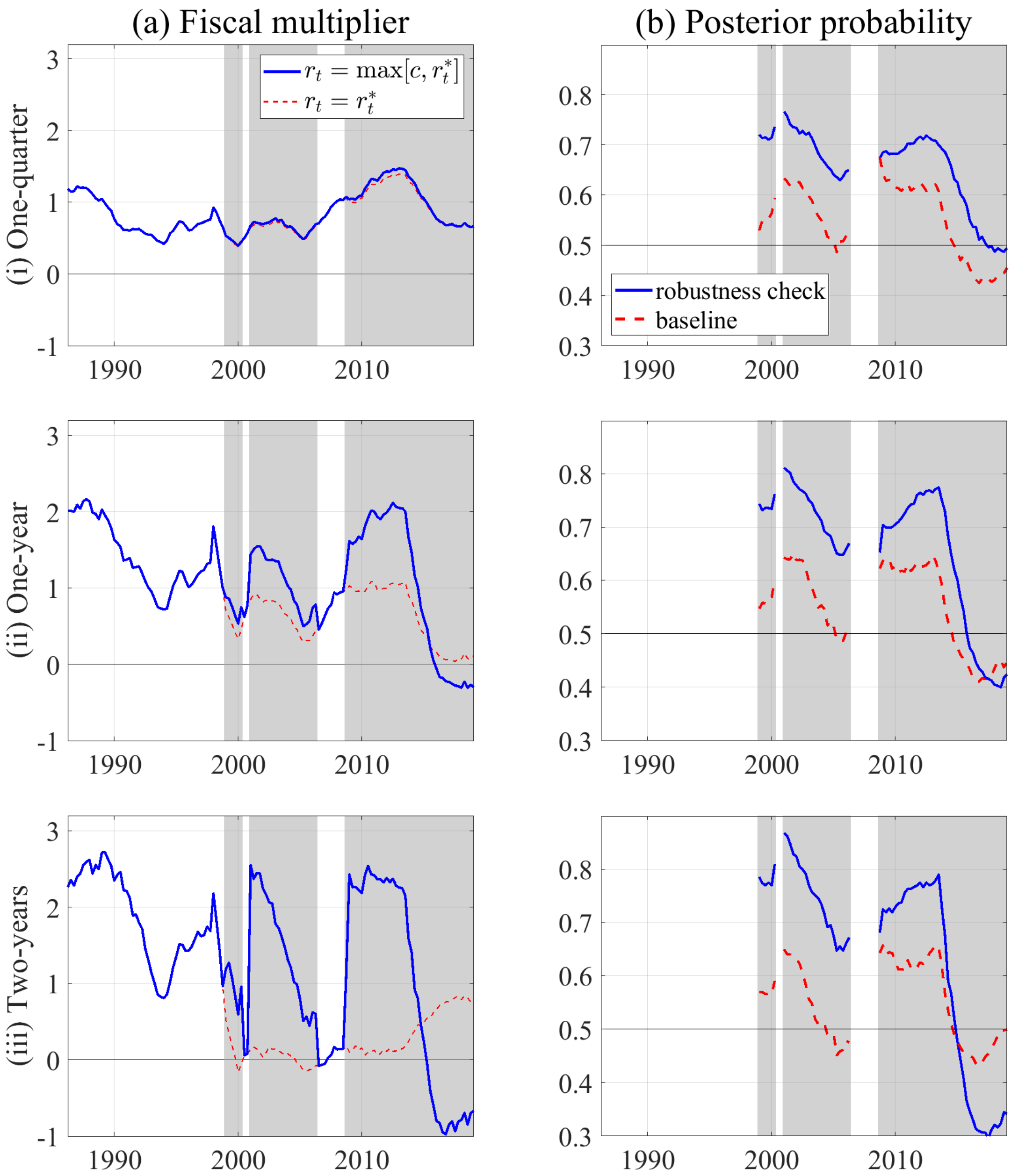

Figure 11: Fiscal multiplier and posterior probability: GDP gap

Notes: This figure shows the time profiles of the fiscal multipliers and the posterior probabilities when we use the GDP gap rather than the GDP growth rate.

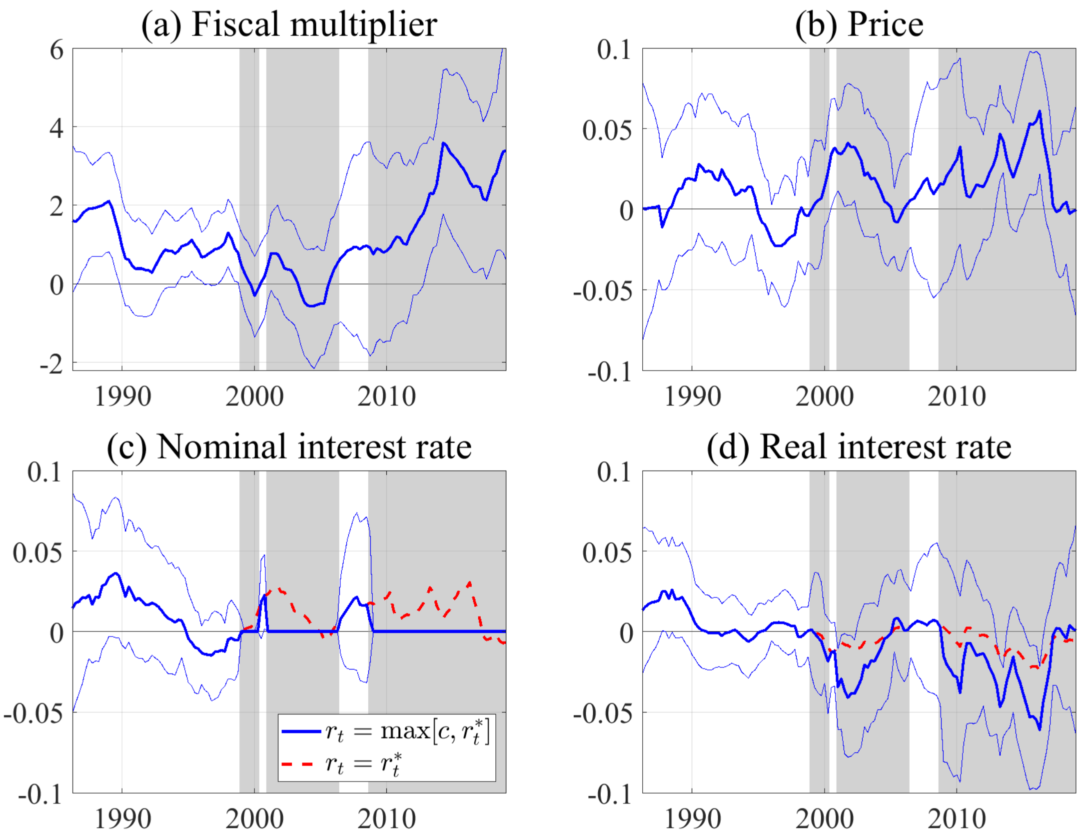

Figure 12: Time-varying responses for the fiscal multiplier, price, and nominal and real interest rates at the impact period: public investment

Notes: This figure shows the time-varying responses of the fiscal multiplier, price, and nominal and real interest rates at the impact period when we use public investment rather than total government spending.

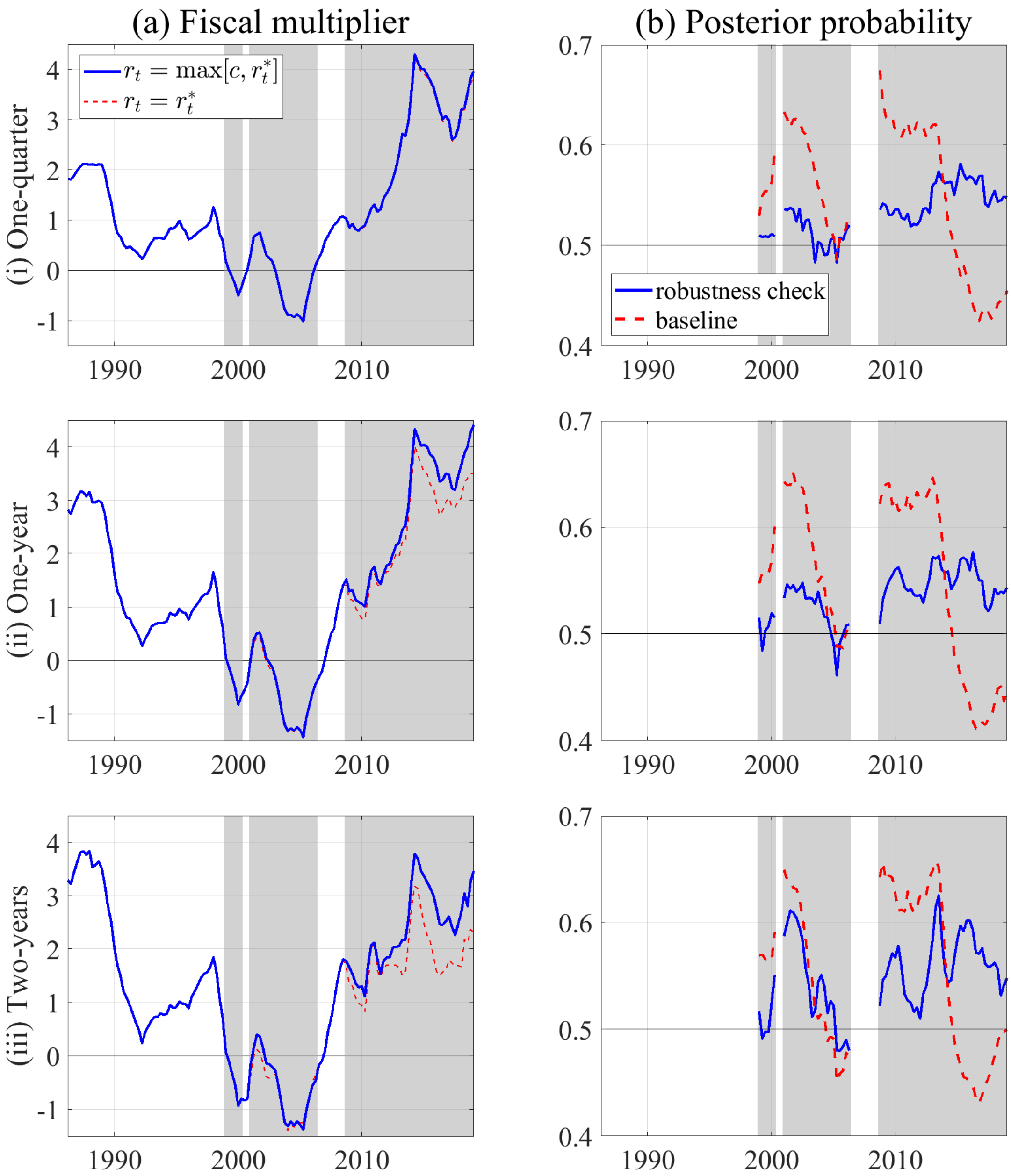

Figure 13: Fiscal multiplier and posterior probability: public investment

Notes: This figure shows the time profiles of the fiscal multipliers and the posterior probabilities when we use public investment rather than total government spending.

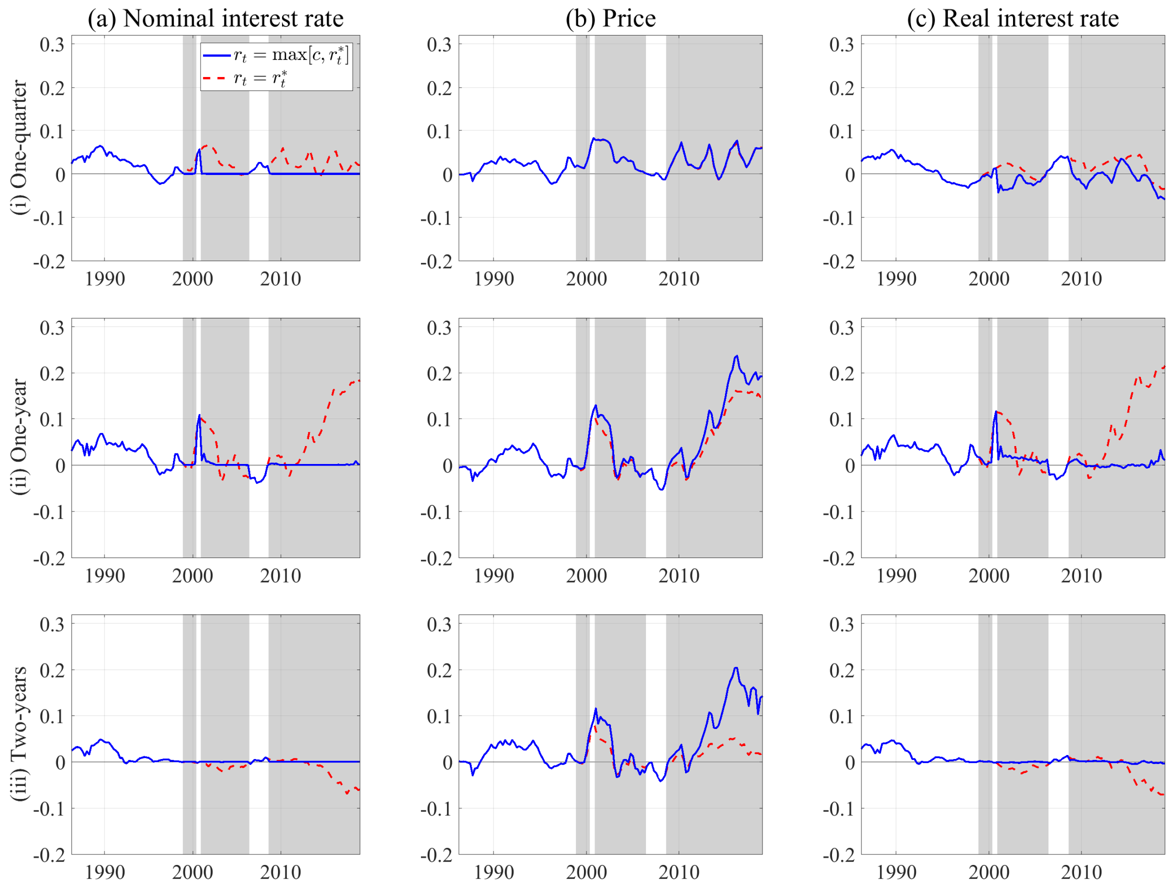

Figure 14: Time-varying responses of nominal and real interest rates and price: public investment
Note: This figure shows the time profiles of the responses of the (a) nominal interest rate, (b) price, and (c) ex post real interest rate to a public investment shock.