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Price and Wage Dynamics and Labor Market Conditions in Japan

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Abstract

This study estimates structural parameters characterizing price and wage dynamics in Japan by accounting for labor market tightness. To reveal the state dependency of price and wage dynamics, we first estimate a smooth transition vector autoregressive model, in which the unemployment rate is regarded as a transition variable. Thereafter, the structural parameters in each state of the labor market are estimated using the impulse response matching technique. The results are summarized as follows. First, price and wage indexation in Japan is rather low. Second, as the labor market tightens, price becomes stickier, while wage becomes less sticky. Finally, the standard theoretical model can be used to sufficiently describe Japan's economy at the period of labor market tightness. As the labor market loosens, however, the distance between the empirical and theoretical responses gradually grows.

Keywords: price and wage dynamics, ST-VAR model, DSGE model

JEL classification: **C32, E24, E31, E52**

1. Introduction

The recent literature has been actively discussing whether inflation pressure from the economic stimulus policy implemented by the Abe Cabinet has sufficiently carried on nominal wages. Obviously, real wage declines in response to demand-stimulating policies, such as fiscal and monetary policies, particularly if the response of nominal wage is less than that of price. In this case, evaluations of such policies may deem resultant benefits inadequate. From an academic viewpoint, it is important to investigate the relationship between price and wage. As is widely known, both variables have a mutual linkage through the New Keynesian Phillips Curve (NKPC) and New Keynesian Wage Phillips Curve (NKWPC) in the dynamic stochastic general equilibrium (DSGE) model. Analyzing the dynamic relationship between price and wage, therefore, can reveal not only the effectiveness of economic policies but also the structure of the real economy.

This study explores the state dependency of price and wage dynamics in a labor market situation and attempts to understand the source and pattern of variation in price and wage dynamics in Japan. To accomplish these objectives, we adopt the smooth transition vector autoregressive (STVAR) model and impulse response matching technique. The STVAR model is a vector version of the smooth transition autoregressive (STAR) model developed by Granger and Terasvirta (1993), and has recently been used in Auerbach and Gorodnichenko (2012). The model allows the coefficients and variance-covariance structure to smoothly vary by the value of a transition variable, and thus it is suitable for our analysis, which sheds light on the state dependency of price and wage dynamics. However, the results from the VAR model cannot be used to further discuss the structure of an economy (i.e., what causes a variation of dynamics in price and wage). Thus, we perform a structural estimation using impulse response matching, as proposed by Christiano et al. (2011). By estimating the DSGE model, the structural parameters in each state of the labor market are revealed, allowing us to identify the causes underpinning the changes in price and wage dynamics.

More precisely, our analysis is implemented as follows. First, we estimate the STVAR model, in which the unemployment rate is regarded as a transition variable. This first-step exercise gives us reduced-form results for the relationship between price and wage dynamics and labor market tightness. We examine the effects of supply and demand shocks on price and real wage based on long-run restrictions. Thereafter, the structural parameters are estimated to minimize the distance between the empirical and theoretical responses using the impulse response matching technique. These estimations are implemented using the Bayesian method.

The literature contains numerous studies that analyze the NKPC and NKWPC in Japan. These previous studies are categorized into two strands of approaches: one that estimates the NKPC and NKWPC as part of the DSGE model (Iiboshi et al., 2006, Sugo and Ueda, 2008) and the other does so separately (Fuchi and Watanabe, 2002, Koga and Nishizaki, 2006, Muto and Shintani, 2014). The approach in this study follows the former in that the DSGE model is completely estimated. However, this study may have certain advantages. First, a time variation of structural parameters cannot be estimated using an ordinary method to estimate the DSGE model. Although some recent studies (Farmer et al., 2011, Liu et al., 2011, Iiboshi, 2015) have estimated the Markov-switching (MS) DSGE model, this model cannot be successfully adopted if the sample size in a specific state is extremely small. On the contrary, the impulse response-matching technique can estimate the structural parameters in the DSGE model, even with a small sample size, as our method regards the impulse response functions (IRFs) as data. Compared with the latter, our estimation is desirable at the point of directly estimating the structural parameters. The coefficients in the separate estimation are formulated as a mixture of structural parameters, including price and wage stickiness and the inverse of Frisch labor elasticity. In addition, the components lying outside of the NKPC and NKWPC (e.g., a stance of monetary authority) are ignored.

This study's results are briefly summarized as follows. First, the state dependency of price and wage dynamics is observed. In particular, the response of real wage to demand shock shows a sign switch depending on labor market tightness. Second, as the labor market tightens, price becomes stickier, while wage becomes less sticky. Third, the standard DSGE model can sufficiently replicate the dynamics of the real economy when the labor market is tight. However, as the labor market loosens, the dynamics of the real economy deviates from the prediction of the theoretical model.

The remainder of this paper is organized as follows. Section 2 explains the STVAR model and reports the reduced-form empirical results. In Section 3, we construct the NK-type DSGE model, which is estimated using the Bayesian impulse response matching technique. Section 4 documents the estimation of the DSGE model using the impulse response-matching and then presents the estimates of the structural parameters. Section 5 provides concluding remarks.

2. VAR analysis

2.1. Smoothed transition VAR model

To reveal the variations in price and wage dynamics, depending on labor market tightness, we estimate the STVAR model in which unemployment rate is regarded as a transition variable. The

STVAR model is specified as follows:

$$Y_t = c(t) + \{1 - F(z_{t-1})\}B_0(L)Y_{t-1} + F(z_{t-1})B_1(L)Y_{t-1} + u_t \tag{1}$$

$$u_t \sim N(0, \Sigma_t) \tag{2}$$

$$\Sigma_t = \Sigma_0\{1 - F(z_{t-1})\} + \Sigma_1 F(z_{t-1}) \tag{3}$$

$$F(z_t) = \frac{\exp(-\gamma z_t)}{1 + \exp(-\gamma z_t)}, \gamma > 0 \tag{4}$$

where $c(t)$ is defined to include all the deterministic components of data, $B_i(L)$, $i = 0,1$ is a polynomial in the lag operator, and Σ_i, $i = 0,1$ denotes the variance–covariance matrix of the reduced-form residuals, u_t. $F(z)$, formulated as in equation (4), denotes the transition function with z as a transition variable, which allows the model to smoothly change according to the transition variable value.

As stated in Auerbach and Gorodnichenko (2012), the principal advantage of using the STVAR model, as compared to separately estimating the structural VAR (SVAR) model for each regime, is the utilization of all sample information to estimate a particular regime in which few observations are included. In addition, we also emphasize several advantages relative to the MSVAR and time-varying parameters (TVP) VAR models. Since the coefficients and variance-covariance matrix in the STVAR model smoothly vary by state, we can specify the model at the arbitrary points among the discrete states. For concreteness, we analyze the dynamics of Y_t at the mid-point of the two states by setting $F(z)$=0.5, and this cannot be done using the MSVAR model, wherein the states are perfectly separated. It is difficult to consider that the economy is divided into perfect discrete states, and thus, the STVAR model is more suitable to represent the real economy compared with the MSVAR model. On the other hand, the TVPVAR model describes continuous variations of the estimation model. However, it cannot explicitly answer what causes the time variation of the coefficients and covariance structure. In contrast, the STVAR model relates a transition variable to the variation of parameters in the model, and thus, we obtain the reason underpinning changes in the model throughout the sample period. For example, this study can reveal the relationship between price and wage stickiness and the unemployment rate, allowing for a more sophisticated micro-foundation for the theoretical model; however, this point remains beyond the focus of this study. In short, the STVAR model elucidates the state dependency of the model associated with the variable in which we are interested.

2.2. Identification of structural shocks

Based on the STVAR model mentioned above, we identify labor productivity and monetary policy shocks as supply and demand shocks using the long-run restriction. To achieve the identification of structural shocks and the estimation of structural parameters in the theoretical model, constructed in Section 3, our VAR model comprises the log first difference of real wage (w_t), price (p_t), and monetary base (m_t).[1] Real wage is a proxy of labor productivity. For simplicity, represented by the SVAR model and not the STVAR model, our 3-variables VAR system can be written in the form of a structural vector moving average as follows:

1 The source and construction of data are explained in detail in Section 2.3.

$$\begin{bmatrix} \Delta w_t \\ \Delta p_t \\ \Delta m_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) & C_{13}(L) \\ C_{21}(L) & C_{22}(L) & C_{23}(L) \\ C_{31}(L) & C_{32}(L) & C_{33}(L) \end{bmatrix} \begin{bmatrix} e_t^a \\ e_t^m \\ e_t^{non} \end{bmatrix} \tag{5}$$

where $C(L)$ is a polynomial in the lag operator. The innovations e_t^a, e_t^m, and e_t^{non}, respectively, denote labor productivity shock, monetary policy shock, and non-specified shock; these shocks are assumed to be mutually orthogonal. Following the seminal work of Blanchard and Quah (1989), we impose the restriction that only labor productivity shock has a permanent effect on real wage, that is, $C_{12}(1)=0$ and $C_{13}(1)=0$.[2] Moreover, we allow the labor productivity and monetary policy shocks to have a long-run effect on prices. In other words, a non-specified shock, which is considered as a mixture of all temporary shocks, is assumed to have no long-run effects on prices, that is, $C_{23}(1)=0$. This is based on the theoretical prediction that a monetary policy measure, such as a temporary increase in the growth rate of money or decrease in the short-term interest rate, permanently shifts price level, although it has no permanent effect on inflation rate.

2.3. Data and specification

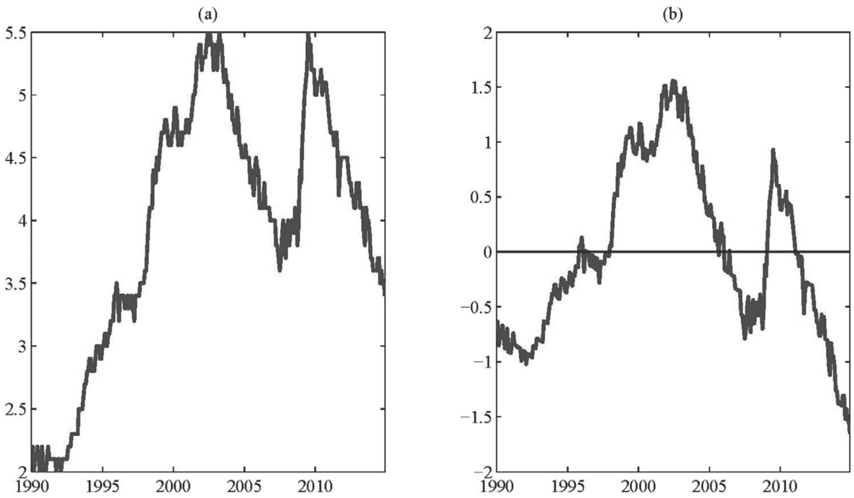
We employ monthly real wage, price, and monetary base data for 1990M1–2014M12. The series of hourly real wage is constructed as follows. First, we obtain the seasonally adjusted series of real wage indices (establishments with five employees or more and total cash earnings) and total hours worked indices (establishments with five employees or more) from the Monthly Labor Survey (Ministry of Health, Labour and Welfare). Second, each index is translated into a real value using the value of the base year. Finally, we calculate hourly real wage by dividing real wage by total hours worked. For price, the series of non-seasonally adjusted Consumer Price Index (CPI) (all items, less fresh food) is downloaded from the CPI (Ministry of Internal Affairs and Communications). We first subtract 1.4% in 1997 and 2.0% in 2014 from the year-on-year inflation rate to eliminate the effects of an increase in consumption tax.[3] Then, we recalculate the level series and perform seasonal adjustments using X-12 ARIMA. The seasonally adjusted series of the monetary base (with adjusted reserve requirement rate change) is collected from the Bank of Japan.

For transition variable z, we use the unemployment rate published in the Labour Force Survey (Ministry of Internal Affairs and Communications). As shown in Figure 1(a), the original series for the unemployment rate has an upward trend in our sample period, and thus we de-trend it by regressing on a constant and linear trend (Figure 1(b)). Furthermore, the data for period t-1 is used to take account of the possibility that labor market tightness cannot be contemporaneously reflected in wage negotiation.

2 Francis and Ramey (2005) also adopt the long-run restriction associated with real wage.

3 These values are based on a trial calculation by the Bank of Japan.

Figure 1: Unemployment rate



Note: Figure 1(a) and 1(b) denote the original series of and the detrended series of unemployment rate, respectively. The original data is obtained from Labor Force Survey. For the detrended series, the linear trend and constant term are excluded from the original series.

In the VAR model, all variables are included in the form of the first difference of the natural logarithm, and a constant term and linear trend are contained as a deterministic component. In addition, the lag length is set as two. This is because the principal aim of the VAR analysis in this study is to obtain the impulse responses necessary to estimate the DSGE model using the impulse response matching technique. Accordingly, we choose two lags as a lag length that can at least capture the dynamics of a system.

2.4. Estimation procedure

The estimation of the STVAR model is implemented using the Bayesian Markov Chain Monte Carlo (MCMC) method, particularly the random-walk Metropolis-Hastings (MH) algorithm. Basically, we follow Auerbach and Gorodnichenko (2012). Let $\tilde{Y}=\{Y_t\}_{t=1}^{T}$ be the sequence of Y_t and $\Theta=[\gamma, \Sigma_0,\Sigma_1, c(t), B_0(L), B_1(L)]$ be the parameters in which we are interested. The posterior log-likelihood for the model is then specified as follows:

$$\ln L(\Theta \mid \tilde{Y}) = const - \frac{1}{2}\sum_{t=1}^{T} \ln|\Sigma_t| - \frac{1}{2}\sum_{t=1}^{T} u_t' \Sigma_t^{-1} u_t \tag{6}$$

where $u_t=Y_t-c(t)-\{1-F(z_{t-1})\}\ B_0(L)\ Y_{t-1}-F(z_{t-1})\ B_1(L)Y_{t-1}$. Given the above, the estimation procedure of a random-walk MH algorithm is summarized as follows:

1. Derive the posterior mode $\widehat{\Theta}$ to maximize $\ln L(\Theta \mid \tilde{Y})$.

2. Set the initial value $\Theta_0 = \widehat{\Theta}$, and $n = 1$.

3. Draw $\Theta_n^{(\text{proposal})}$ from the following random-walk model:

$$\Theta_n^{(\text{proposal})} = \Theta_{n-1} + \nu_t, \nu_t \sim N(0, cH),$$

4. Calculate the acceptance rate q using $\Theta_n^{\text{proposal}}$ and Θ_{n-1} as follows:

$$q = \min\left[\frac{f\left(\Theta_n^{(\text{proposal})} \mid \tilde{Y}\right)}{f(\Theta_{n-1} \mid \tilde{Y})}, 1\right].$$

5. Accept $\Theta_n^{(\text{proposal})}$ with probability q and reject it with probability $1 - q$. Set $\Theta_n = \Theta_n^{(\text{proposal})}$ when accepted and $\Theta_n = \Theta_{n-1}$ when rejected.

6. Set $n = n + 1$ and return to Step 3 until the number of iterations reaches N times.

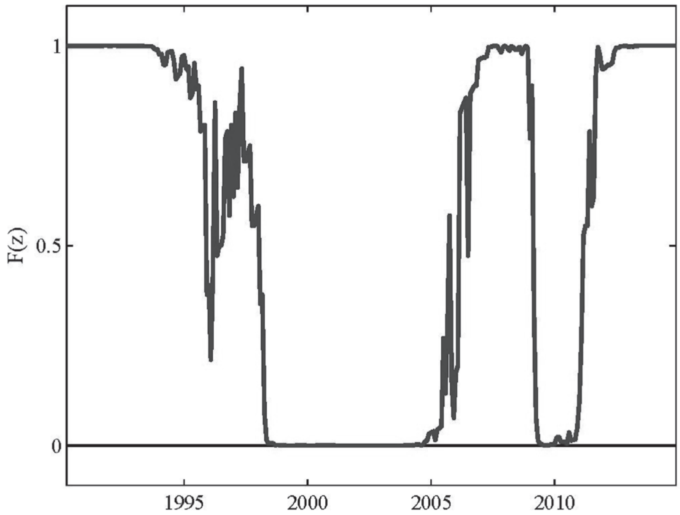
In the actual estimation, we randomly generate only Θ'=[γ, $chol(\Sigma_0)$, $chol(\Sigma_1)$] because, given Θ', the lag polynomials [B_0 (L), B_1 (L)] can be estimated with weighted least squares to maximize $\ln L(\Theta \mid \tilde{Y})$. Moreover, this simplification saves the time necessary for convergence.

In this study, N is set to be 120,000 and the first N_0 = 20,000 samples are discarded as a burn-in. To ensure the stationarity of the VAR system, the unit circle is selected only for draws within which the roots of the lag polynomials [B_0 (L), B_1 (0)] are present. With respect to the priors, the gamma distribution is associated with γ. In the benchmark, we set the mean and standard deviation of the prior to 10 and 0.1.

2.5. Impulse response function

Figure 2 shows the estimated transition probability, $F(z)$, and indicates that the higher the transition probability, the more the labor market tightness. The figure also allows us to understand labor market transition in Japan. As shown, the degree of tightness in the labor market is still high just after the economic bubble burst in the early 1990s, but it gradually declines. The state then falls $F(z) = 0$ (i.e., low tightness state) perfectly from the second half of the 1990s to the first half of the 2000s. Thereafter, the labor supply and demand situation improve by benefiting from the economic recovery in the 2000s. However, after the temporary economic deterioration by the Lehman shock, it tightens again because of the Great East Japan earthquake and the economic stimulus policy implemented by Prime Minister Shinzo Abe, or the so-called "Abenomics."

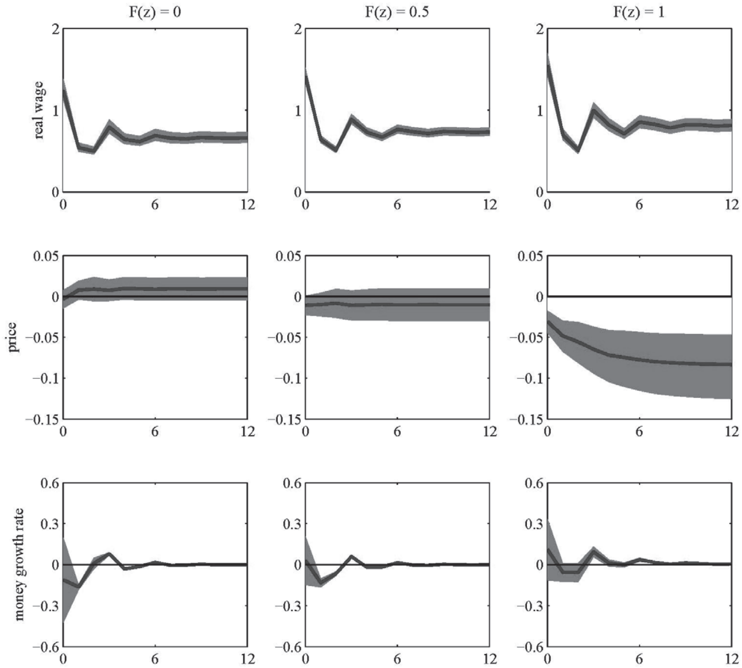
Figure 2: Estimated transition probability



Figures 3 and 4 depict the IRFs to supply and demand shocks at the state of $F(z)$=0,0.5, and 1. The solid lines and shaded areas denote the median of sampled IRFs and 90% Bayesian credible intervals. At first glance, we observe that the signs of several responses switch depending on labor market tightness. The most notable is a response of real wage to demand shock, where the response changes from negative to positive as the labor market tightens. This result indicates the possibility that the structure of price and wage dynamics, that is, the structural parameters of a real economy, alters according to labor market tightness.

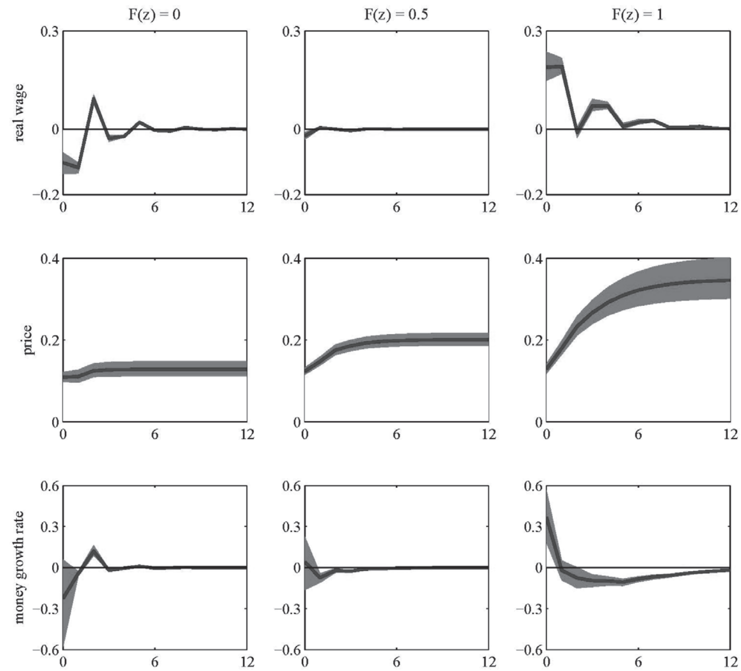
In addition, we also recognize that the sign of some responses at $F(z)$=0 is opposite to a normal prediction of the standard theoretical model. For instance, the most prominent one is a positive (but insignificant) response of price to supply shock, as shown in the middle-left chart of Figure 3. A negative response of money growth rate to both shocks also seems to be counterintuitive because the standard model, generally, presumes that both shocks stimulate consumption and, as a result, increase money demand. Consequently, as mentioned in Section 4, the calculated IRFs at $F(z)$=0 using the DSGE model do not effectively trace the VAR model. On the other hand, the responses at $F(z)$=1 are consistent with the usual expectation: a permanent labor productivity shock reduces price, while a monetary policy shock increases price and money growth rate. Furthermore, we find that a positive demand (monetary policy) shock increases real wage and price levels. The responses at $F(z)$=0.5 are in the center of the two extreme cases. The impact of responses seems weak compared with $F(z)$=0 and $F(z)$=1, especially in the responses of prices to supply shock and real wage to demand shock. Below, we estimate the value of structural parameters for each state via the impulse response matching technique using the IRFs obtained in this section.

Figure 3: Responses of variables to supply shock



Note: Solid lines and shaded areas indicate the median responses and 90% credible intervals, respectively.

Figure 4: Responses of variables to demand shock



Note: Solid lines and shaded areas indicate the median responses and 90% credible intervals, respectively.

3. Theoretical model

To interpret the (reduced-form) empirical results obtained from the VAR analysis in a structural manner, we estimate the DSGE model using the impulse response matching technique. This section presents the simple NK model employed in the Bayesian estimation.

The NK model in this study comprises Calvo-type (1983) sticky prices and wages, price and wage indexation, consumption habit, money growth rate, and no capital.[4] To capture the situation in which the Bank of Japan has been setting the policy rate to be almost zero and adopting quantitative easing (QE) after the second half of the 1990s, we employ a money growth rule as a monetary policy rule instead of the Taylor rule. Since the focus of this study is to ascertain the state dependency of price and wage dynamics, price and wage indexation is incorporated into the model. Furthermore, the stochastic process of permanent technology and transitory monetary policy shocks is introduced as a counterpart of supply and demand shocks in the VAR analysis.

3.1. Households

The economy consists of a unit mass of identical households indexed by $i \in [0,1]$. The utility maximization problem of each household can be written as follows:

$$\max U = E_0 \sum_{t=0}^{\infty} \left[\frac{\{c_{t(i)} - hc_{t-1}(i)\}^{1-\gamma}}{1-\gamma} + \frac{m_t(i)^{1-\mu}}{1-\mu} - \frac{n_t(i)^{1+\phi}}{1+\phi} \right] \tag{7}$$

subject to

$$c_t(i) + m_t(i) + b_t(i) = \frac{W_t(i)}{P_t} n_t(i) + \frac{m_{t-1}(i)}{\Pi_t} + \frac{R_{t-1}}{\Pi_t} b_{t-1}(i) + d_t(i) \tag{8}$$

where the uppercase and lowercase letters denote nominal and real variables. The household obtains positive utility from consumption c, and real money holdings m and a negative utility from hours worked n. In return, the budget constraint represents that a household allocates the sum of real labor income $\frac{W_t(i)}{P_t} n_t(i)$; inflation-adjusted money holding $\frac{m_t}{\Pi_t}$; inflation-adjusted total sum of bond and its interest payment $\frac{R_{t-1}}{\Pi_t} b_{t-1}(i)$; and real dividends d_t to consumption, money holdings, and bond. Here, P, Π, and R denote aggregate price, inflation rate, and gross nominal interest rate, respectively. Based on this optimal problem, each household chooses the optimal level of consumption and money holdings.

3.2. Labor market and wage setting

To introduce wage stickiness, we assume a continuum of differentiated labor input indexed by $l \in [0,1]$. Furthermore, it is assumed that there exist labor unions corresponding to each differentiated

4 Watanabe (2009) emphasizes the importance of the consumption habit in Japan.

labor input, and each union sets their wage rate. As adopted in Schmitt-Grohe and Uribe (2006), each household member is assumed to provide each possible type of labor input. In other words, each household belongs to every labor union. This assumption rules out the possibility that each household receives different labor income in equilibrium without adopting a contingent claim. In addition, as adopted in Gaíl et al. (2007), labor supply is assumed to be determined by labor demand (not by the optimal choice of households), given a wage fixed by the labor union.

A perfectly competitive labor-bundler firm bundles differentiated labor input $n_t(l)$ into effective labor denoted by n_t as follows:

$$n_t = \left[\int_0^1 n_t(l)^{\frac{\varepsilon_w - 1}{\varepsilon_w}} dl\right]^{\frac{\varepsilon_w}{\varepsilon_w - 1}} \tag{9}$$

where ε_w denotes the elasticity of substitution among labor inputs. As a result of the labor bundler's optimal problem, the demand function for each differentiated labor input is expressed as

$$n_t(l) = \left(\frac{W_t(l)}{W_t}\right)^{-\varepsilon_w} n_t \tag{10}$$

Then, aggregate nominal wage is equal to

$$W_t = \left[\int_0^1 W_t(l)^{1-\varepsilon_w} dl\right]^{\frac{1}{1-\varepsilon_w}} \tag{11}$$

Given the demand function for labor input l, the labor union sets its nominal wage $W_t(l)$ to maximize the lifetime utility of each household. Under the Calvo-type (1983) wage stickiness, it is assumed that the fraction $1-\rho_w$ of labor unions can reset the optimal nominal wage $W_t^*(l)$ in each period. The rest of the labor unions, which do not obtain an opportunity to re-optimize their nominal wage, sets $W_t(l)$ according to

$$W_t(l) = \left(\frac{P_{t-1}}{P_{t-2}}\right)^{\gamma_w} W_{t-1}(l) \tag{12}$$

where γ_w is a degree of wage indexation. The optimal problem for wage union l can then be written as

$$\max_{W_t^*(l)} E_t \sum_{s=0}^{\infty} \rho_w^s \Lambda_{t,t+s} \left[\frac{\{c_{t+s}(l) - hc_{t+s-1}(l)\}^{1-\gamma}}{1-\gamma} + \frac{m_{t+s}(l)^{1-\mu}}{1-\mu} - \frac{n_{t+s}(l)^{1+\phi}}{1+\phi}\right] \tag{13}$$

subject to

$$c_{t+s}(l) + m_{t+s}(l) + b_{t+s}(l)$$

$$= \int_0^1 \frac{\left(\frac{P_{t+s-1}}{P_{t-1}}\right)^{\gamma_w} W_t^*(l)}{P_{t+s}} n_{t+s}(l)\, dl + \frac{m_{t+s-1}(l)}{\Pi_{t+s}} \tag{14}$$

$$+ \frac{R_{t+s-1}}{\Pi_{t+s}} b_{t-1}(l) + d_{t+s}(l)$$

and

$$n_{t+s}(l) = \left\{ \frac{\left(\frac{P_{t+s-1}}{P_{t-1}}\right)^{\gamma_w} W_t^*(l)}{W_{t+s}} \right\}^{-\varepsilon_w} n_{t+s} \tag{15}$$

where $\Lambda_{t,t+s} = \beta^s \left(\frac{c_{t+s}}{c_t}\right)^{-1}$ denotes the stochastic discount factor. From equations (11) and (12), the law of motion of aggregate wages can be given as

$$W_t = \left[(1-\rho_w) W_t^{*1-\varepsilon_w} + \rho_w \left\{ \left(\frac{P_{t-1}}{P_{t-2}}\right)^{\gamma_w} W_{t-1} \right\}^{1-\varepsilon_w} \right]^{\frac{1}{1-\varepsilon_w}} \tag{16}$$

3.3. Firm

The production sector comprises two types of firms: monopolistically competitive firms that produce differentiated intermediate goods and perfectively competitive firms that produce single final goods using intermediate goods as input. Each intermediate goods firm indexed by $j \in [0,1]$ produces an intermediate good $y_t(j)$, and its production function is simply assumed to be a linear function:

$$y_t(j) = a_t n_t(j) \tag{17}$$

where a_t denotes the total factor productivity (TFP), which is exogenously given. As a result of the profit maximization problem for intermediate goods firms, we simply obtain real marginal cost mc_t as follows:

$$mc_t = \frac{w_t}{a_t}. \tag{18}$$

On the other hand, the final goods firms transform intermediate goods to final goods using the following Dixit–Stiglitz-type production function:

$$y_t = \left[\int_0^1 y_t(j)^{\frac{\varepsilon_p - 1}{\varepsilon_p}} dl\right]^{\frac{\varepsilon_p}{\varepsilon_p - 1}} \tag{19}$$

where ε_p is the elasticity of substitution across each type of intermediate goods. Then, the demand function for intermediate goods is obtained as

$$y_t(j) = \left(\frac{P_t(j)}{P_t}\right)^{-\varepsilon_p} y_t \tag{20}$$

and the final goods pricing rule is written as

$$P_t = \left[\int_0^1 P_t(j)^{1-\varepsilon_p} dj\right]^{\frac{1}{1-\varepsilon_p}} \tag{21}$$

3.4. Price setting

As with wage setting, intermediate goods firms set their prices according to the Calvo (1983) mechanism with price indexation. For any given period t, the fraction $1-\rho_p$ of intermediate goods firms can revise their prices to the optimal price level $P_t^*(j)$, whereas the fraction ρ_p does not have the opportunity to reset their price and set their prices as

$$P_t(j) = \left(\frac{P_{t-1}}{P_{t-2}}\right)^{\gamma_p} P_{t-1}(j) \tag{22}$$

where γ_p indicates a degree of price indexation. The optimal price $P_t^*(j)$ is determined by solving the problem:

$$\max_{P_t^*(j)} E_t \sum_{s=0}^{\infty} \rho_p^s \Lambda_{t,t+s} \left[\left(\frac{P_{t+s-1}}{P_{t-1}}\right)^{\gamma_p} P_t^*(j) y_{t+s}(j) - P_{t+s} y_{t+s}(j) mc_{t+s}(j)\right] \tag{23}$$

subject to

$$y_{t+s}(j) = \left(\frac{\left(\frac{P_{t+s-1}}{P_{t-1}}\right)^{\gamma_p} P_t^*(j)}{P_{t+s}}\right)^{-\varepsilon_w} y_{t+s}. \tag{24}$$

As in wage setting, the law of motion of aggregate prices can be written as

$$P_t = \left[\left(1-\rho_p\right)P_t^{*1-\varepsilon_p} + \rho_p\left\{\left(\frac{P_{t-1}}{P_{t-2}}\right)^{\gamma_p} P_{t-1}\right\}^{1-\varepsilon_p}\right]^{\frac{1}{1-\varepsilon_p}}. \tag{25}$$

3.5. Rest of the model

With respect to the monetary policy, we assume that the monetary authority adopts the money growth rule as follows:

$$M_t = exp\{\nu_t\}M_{t-1} \Leftrightarrow m_t = exp\{\nu_t\}\frac{m_{t-1}}{\Pi_t}. \tag{26}$$

Here, ν_t denotes the money growth rate, which depends on its own lags and inflation rate, such as

$$\hat{\nu}_t = \rho_m \hat{\nu}_{t-1} - \psi_p \hat{\pi}_t + \varepsilon_t^m \tag{27}$$

where a hat means a log deviation from a steady state value, and ε_t^m is a monetary policy shock that follows an i.i.d. normal distribution. On the other hand, the stochastic process of TFP is assumed to follow the random-walk process:

$$\ln a_t = \ln a_{t-1} + \varepsilon_t^a \tag{28}$$

Finally, the market clearing condition in this study is given by

$$y_t = c_t \tag{29}$$

because our model ignores the government sector.

3.6. Log-linearized de-trended equilibrium conditions

To eliminate a stochastic trend and ensure model stationarity, we divide real variables c_t, y_t, and W_t/P_t by the productivity a_t. Let $\tilde{x}_t$ indicate the log deviations of a stationary variable from its steady states. The log-linearized de-trended equilibrium conditions in our model are summarized as the following nine equations:

$$\frac{1}{1+h}E_t\tilde{c}_{t+1} + \frac{1-h}{\gamma(1+h)}E_t\hat{\pi}_{t+1} - \frac{1-h}{\gamma(1+h)}\hat{r}_t - \frac{h}{1-h}\hat{a}_t = \tilde{c}_t - \frac{h}{1+h}\tilde{c}_{t-1} - \frac{h}{1-h}\hat{a}_{t-1} \tag{30}$$

$$\frac{\beta}{1+\beta\gamma_p}E_t\hat{\pi}_{t+1} = \hat{\pi}_t - \frac{\gamma_p}{1+\beta\gamma_p}\hat{\pi}_{t-1} - \frac{\left(1-\rho_p\right)\left(1-\rho_p\beta\right)}{\rho_p\left(1+\beta\gamma_p\right)}\tilde{w}_t \tag{31}$$

$$\Gamma\beta E_t\tilde{w}_{t+1} + \Gamma\beta E_t\,\hat{\pi}_{t+1} + \Gamma\kappa_w\phi\hat{n}_t + \left(\Gamma\beta + \frac{\Gamma\kappa_w\gamma}{1-h} - 1\right)\hat{a}_t$$

$$= \tilde{w}_t - \left(\frac{\Gamma\kappa_w\gamma}{1-h}\right)\tilde{c}_t + \Gamma(1+\beta\gamma_w)\hat{\pi}_t - \Gamma\,\tilde{w}_{t-1} + \left(\frac{\Gamma\kappa_w\gamma h}{1-h}\right)\tilde{c}_{t-1} \tag{32}$$

$$- \Gamma\gamma_w\hat{\pi}_{t-1} - \Gamma\left(1 - \frac{\kappa_w\gamma h}{1-h}\right)\hat{a}_{t-1}$$

$$\tilde{y}_t = \tilde{c}_t \tag{33}$$

$$\tilde{y}_t - \hat{n}_t = 0 \tag{34}$$

$$\mu\hat{m}_t + \frac{1}{\mathrm{R}-1}\hat{r}_t = \frac{\gamma}{1-h}\tilde{c}_t - \frac{h\gamma}{1-h}\tilde{c}_{t-1} + \frac{\gamma}{1-h}\hat{a}_t - \frac{h\gamma}{1-h}\hat{a}_{t-1} \tag{35}$$

$$\hat{m}_t - \hat{v}_t = \hat{m}_{t-1} - E_t\hat{\pi}_{t+1} \tag{36}$$

$$\hat{v}_t = \rho_m\hat{v}_{t-1} - \psi_p\hat{\pi}_t + \varepsilon_t^m \tag{37}$$

$$\hat{a}_t = \hat{a}_{t-1} + \varepsilon_t^a \tag{38}$$

where $\Gamma=\rho_w/(1+\beta\rho_w^2)$ and $\kappa_w = (1-\beta\rho_w)(1-\rho_w)/\rho_w$.

4. Structural estimation

4.1. Methodology

The presented DSGE model is estimated using the Bayesian minimum distance technique (i.e., impulse response matching), as adopted in Altig et al. (2011), Christiano et al. (2011), and Hofmann et al. (2012). More specifically, the structural parameters in the DSGE model are derived to minimize the distance between the IRFs obtained from the VAR and DSGE models. Since we estimate the VAR model using the Bayesian method, there is no estimated point at which we can match the IRFs, calculated using the DSGE model. Hence, we regard the median of the sampled IRFs, shown in Figure 3 and 4, as reference points to minimize the distance from the theoretical responses.

Following the description of Christiano et al. (2011), let $\widehat{\Psi}$ be the stacked vector of the sampled IRFs, which has a dimension of 12 (number of horizons) times 2 (number of shocks) times 3 (number of variables). For the number of observations T, the standard asymptotic theory says that

$$\sqrt{T}\left(\hat{\psi} - \psi(\theta_0)\right) \sim N\big(0, W(\theta_0, \zeta_0)\big) \tag{39}$$

where θ_0 indicates the true values of the parameters and ζ_0 denotes those of the parameters of shocks that are in the model. As noted in Christiano et al. (2011), ζ_0 does not formally appear in this analysis. Given the above, the asymptotic distribution of $\widehat{\Psi}$ can be written as

$$\widehat{\psi} \sim N\big(\psi(\theta_0), V(\theta_0, \zeta_0, T)\big), \tag{40}$$

where

$$V(\theta_0, \zeta_0, T) \equiv \frac{W(\theta_0, \zeta_0)}{T}. \tag{41}$$

Subsequently, we regard $\widehat{\Psi}$ as data, and then the approximate likelihood of the data $\widehat{\Psi}$ as a function of θ can be written as

$$f(\widehat{\psi} \mid \theta) = \left(\frac{1}{2\pi}\right)^{\frac{N}{2}}$$

$$\mid V(\theta_0, \zeta_0, T) \mid^{-\frac{1}{2}} \tag{42}$$

$$\times \exp\left[-\frac{1}{2}\left(\widehat{\psi} - \psi(\theta_0)\right)' V(\theta_0, \zeta_0, T)^{-1} \left(\widehat{\psi} - \psi(\theta_0)\right)\right],$$

where N denotes the number of elements in $\widehat{\Psi}$. With respect to the matrix V, its estimator $\hat{V}$ is set to depend on the second moments of the sample impulse response function. More precisely, we use the matrix that has on its diagonal the diagonal elements of

$$\bar{V} = \frac{1}{M}\sum_{i=1}^{M}(\psi_i - \bar{\psi})(\psi_i - \bar{\psi})' \tag{43}$$

as $\hat{V}$ based on Christiano et al. (2011). Here, $\overline{\Psi}$ is a median of the sampled IRFs and M is the number of samples. Given the likelihood of $\widehat{\Psi}$, as represented in (43), the Bayesian posterior distribution of θ conditioned on $\widehat{\Psi}$ and $\hat{V}$ is specified by

$$f(\theta \mid \widehat{\psi}) = \frac{f(\widehat{\psi} \mid \theta)p(\theta)}{f(\widehat{\psi})} \tag{44}$$

where p(θ) indicates the priors for θ, and $f(\widehat{\Psi})$ denotes a marginal density of $\widehat{\Psi}$. Based on equation (44), we estimate the structural parameters in the DSGE model using a random-walk MH algorithm, as explained in Section 2.4.

4.2. Results

Table 1 reports the priors and posteriors of the estimated parameters for the states of $F(z)$=0, $F(z)$=0.5, and $F(z)$=1. The other parameters are fixed as follows. The discount factor is set as β= 0.999 because the interest rate in the sample period is markedly low due to the zero-interest rate policy and QE.[5] The preference parameters are chosen as γ=1.5, μ=1, and ϕ=1. The values of the fixed parameters and priors depend on previous studies.

From the result of Table 1, we first observe a drastic variation of structural parameters according to the labor market state. For example, the price stickiness parameter changes from 0.01 at $F(z)$=0 to 0.67 at $F(z)$=1, while the wage stickiness parameter changes from 0.32 to 0.01.[6] These variations suggest the necessity of considering the relationship between price and wage dynamics and labor market tightness. Accordingly, this result indicates the possibility that the labor market condition affects the price and wage decision. In return, the indexations of price and wage are estimated as considerably small and stable through all the states, and this is consistent with Muto and Shintani's (2014) results.[7] Compared with the estimated results for the United States, as reported by Hofmann et al. (2012), this degree of price and wage indexation in Japan is substantially modest. This finding (i.e., the indexation in Japan is smaller than that in the United States) also conforms to the claims in Kuroda and Yamamoto (2007) and Muto and Shintani (2014).

Subsequently, we focus on the state dependency of price and wage stickiness. As mentioned above, price stickiness tends to be high as the labor market becomes tight. For the price stickiness, if we regard the period of $F(z)$=1 as an economic boom, this result implies that it is unlikely that the price will rise in response to an increase in demand resulting from economic recovery. On the contrary, it seems to be easy to downwardly adjust a price to a decrease in demand when $F(z)$=0. This result accords with the Japanese economy's experience in the "lost two decades," that is, long-lasting deflation. On the other hand, wage is stickier when the labor supply and demand situation is loose. This is also consistent with a phenomenon called the downward rigidity of wage, as indicated in Kuroda and Yamamoto (2007).

Finally, we mention the parameters in the monetary policy rule. One can confirm that the monetary authority takes an active stance at $F(z)$=0, in which the sample from the second half of the 1990s to the 2000s is mainly included. Accordingly, reflecting the adoption of the QE policy, the monetary policy at the time was more persistent and aggressive to deflation.

5 Under β=0.999, the net nominal interest rate is assumed to equal 0.01.

6 The estimates of price and wage stickiness in this study are somewhat small compared with the results reported in previous studies (Iiboshi et al., 2006, Sugo and Ueda, 2008). It can be considered that this difference emanates from the variation in the sample period. The sample period in the extant literature is restricted to the late 1990s to avoid the period of zero lower bound on nominal interest rate. In fact, the estimated result in Muto and Shintani (2014), who adopted 1980–2013 as the sample period, is similar to ours.

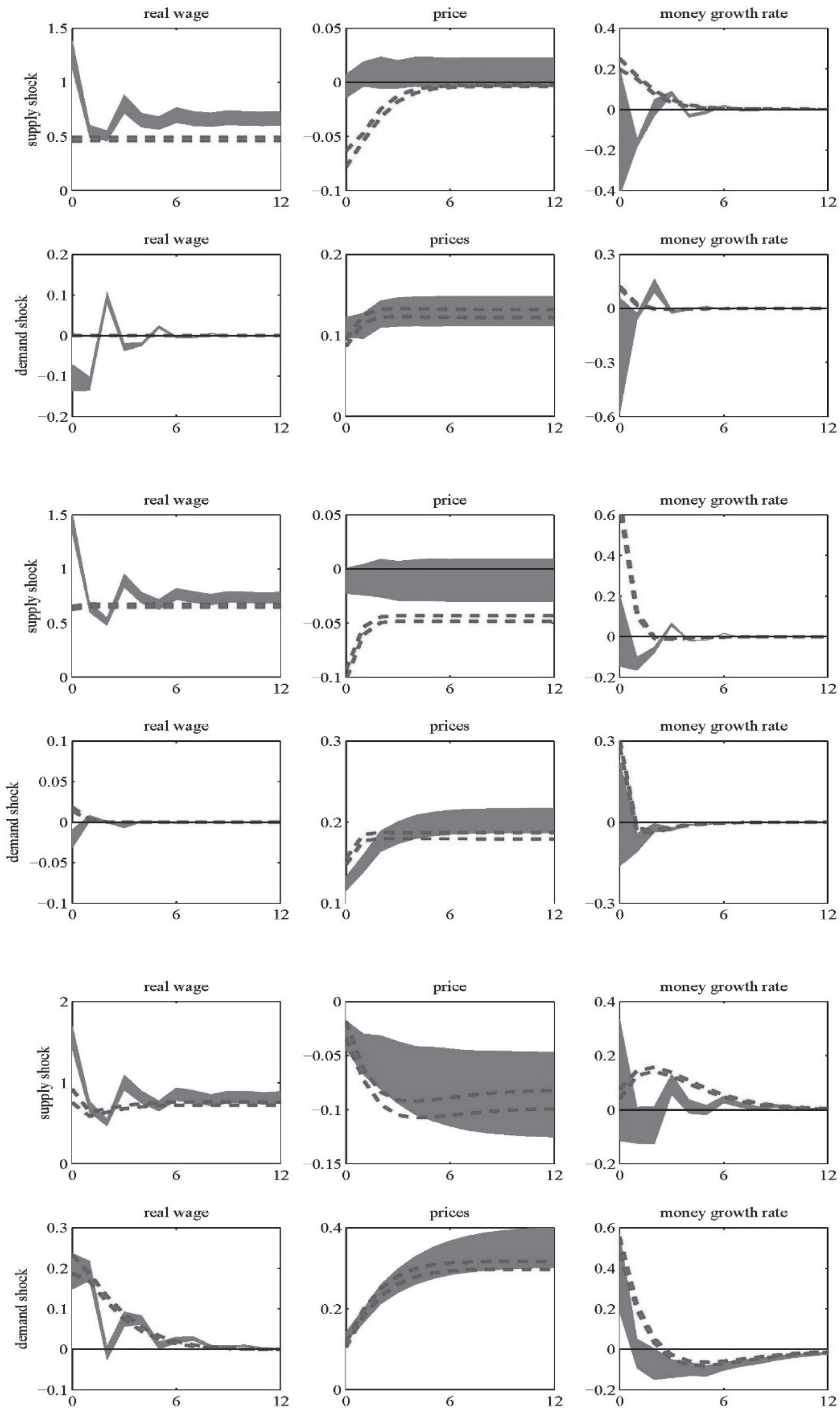
7 Although the median values of price indexation at $F(z)$=0 and wage indexation at $F(z)$=1 seem to be large, the credible intervals associated with these parameters are significantly wide. Hence, we consider these values as largely insignificant.

Table 1: Prior and posterior estimates of the DSGE model

		Prior		Posterior		
				$F(z)=0$	$F(z)=0.5$	$F(z)=1$
		Density	Mean	Median	Median	Median
		[bounds]	(Std. dev.)	[5%, 95%]	[5%, 95%]	[5%, 95%]
γ_p	Price indexation	Beta	0.5	0.30	0.02	0.01
		[0,1]	(0.2)	[0.03, 0.85]	[0.00, 0.06]	[0.00, 0.04]
γ_w	Wage indexation	Beta	0.5	0.02	0.02	0.22
		[0,1]	(0.2)	[0.00, 0.09]	[0.00, 0.06]	[0.02, 0.77]
ρ_p	Price Stickiness	Beta	0.75	0.01	0.10	0.67
		[0, 0.99]	(0.15)	[0.00, 0.34]	[0.09, 0.12]	[0.66, 0.69]
ρ_w	Wage Stickiness	Beta	0.75	0.32	0.16	0.10
		[0, 0.99]	(0.15)	[0.31, 0.34]	[0.15, 0.17]	[0.04, 0.19]
h	Consumption habit	Beta	0.7	0.57	0.12	0.59
		[0, 1]	(0.1)	[0.53, 0.59]	[0.08, 0.16]	[0.54, 0.63]
ρ_m	Persist. of MP shock	Beta	0.7	0.99	0.58	0.75
		[0, 0.99]	(0.2)	[0.98, 0.99]	[0.57, 0.60]	[0.72, 0.78]
ψ_p	MP rule	Gamma	0.5	3.18	6.80	2.23
		[0, ∞]	(0.1)	[2.91, 3.46]	[6.35, 7.30]	[2.03, 2.47]
σ_s	Std. dev. Tech. shock	Inv. Gamma	1	0.48	0.66	0.74
		[0, ∞]	(0.5)	[0.46, 0.50]	[0.64, 0.68]	[0.72, 0.76]
σ_d	Std. dev. Dem. shock	Inv. Gamma	1	0.41	1.32	0.76
		[0, ∞]	(0.5)	[0.37, 0.44]	[1.24, 1.42]	[0.69, 0.84]

Figure 5 shows the 90 percentiles of the IRFs of the DSGE model and those from the VAR analysis. The red dotted lines and shaded areas indicate the credible intervals corresponding to the DSGE and VAR models, respectively. As can be seen from Figure 5(c), the IRFs obtained from the DSGE model correspond fairly well with those from the VAR model at $F(z)$=1. However, as the labor market loosens, the deviation between theoretical and empirical responses gradually spreads. At $F(z)$=0, the theoretical responses almost completely fail to trace the empirical response, except for the response of price to demand shock. This finding indicates the possibility that it may be difficult to describe the Japanese economy from the second half of the 1990s to the first half of the 2000s using the standard DSGE model. In other words, the Japanese economy in this period is a unique situation not to be explained by the standard DSGE model. Therefore, we will have to investigate Japan-specific factors that cause the real economy to become estranged from the theoretical prediction. Unfortunately, this is beyond the focus of this study and thus, we hope to address this in the future.

Figure 5: Impulse responses of VAR and DSGE model for *F*(*z*)=0,0.5,and 1.



5. Conclusion

In this study, we analyzed price and wage dynamics in Japan by accounting for the labor market. Combining the reduced-form results derived from the STVAR model and the structural analysis using impulse response matching technique, we estimated the structural parameters in the Japanese economy based on labor market tightness. The main findings are summarized as follows.

First and most importantly, price and wage dynamics change according to labor market tightness. The sign of several IRFs in reduced-form analysis switches as the labor market tightens, and thus, the structural parameters characterizing price and wage dynamics also show a drastic variation. Second, it is found that price becomes stickier, while wage becomes less sticky as the labor market becomes tight. The former is consistent with the historical fact that it is difficult for firms to raise price in response to an increase in demand, even in an economic boom. On the other hand, the latter is in accord with the downward rigidity of wage, which is often mentioned in Japan. Related to price and wage rigidity, it also turns out that the indexation in price and wage in Japan is quite modest. Third, the standard DSGE model fairly replicates the dynamics of the Japanese economy when the labor market is tight. However, as the labor market loosens, its explanatory power to describe the real economy seems to gradually dampen.

Nevertheless, this study has certain limitations. As stated above, it reveals the state dependency of price and wage dynamics; however, the theoretical model that we estimate captures the differences in the dynamics that exogenously stem from the labor market situation. Therefore, in an extension of this study, we will attempt to construct a theoretical model in which the price and wage decision is endogenously determined based on labor market tightness. Enhancing the ability of the DSGE model to explain the real economy will allow us to better understand the structure of the Japanese economy.

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