Analysis on the Macroeconomic Effects of Fiscal Policy and Business Cycles in Japan

by

Hiroshi Morita

Submitted in partial fulfillment of the requirements
for the degree of
Doctor of Philosophy
in
Economics

Graduate School of Economics
Hitotsubashi University
2013

Acknowledgment

This dissertation is compiled from research papers that I wrote at the Graduate School of Economics of Hitotsubashi University. The thesis is comprised of five chapters on the analysis of the macroeconomic effects of fiscal policy and business cycles in Japan, and is organized as follows. Chapter 1 summarizes this dissertation. Chapter 2, 3, and 4 explore the effects of fiscal policy in Japan in the framework of vector autoregression (VAR) model. In Chapter 5, I try to clarify the source of business cycles in Japan by taking external shocks into account. Among the five chapters, Chapter 2 is based on Morita (2013a) "The Effects of Anticipated Fiscal Policy on Macroeconomic Effects in Japan," Global COE hi-stat discussion paper series No. 219, and Chapter 5 is based on Morita (2013b)"External Shocks and Japanese Business Cycles: Evidence from a Sign-Restricted VAR," Global COE hi-stat discussion paper series No. 277.

I am deeply grateful to my supervisor, Professor Etsuro Shioji, for valuable and helpful comments. Without his continuing support and encouragement, I would never have completed this dissertation. I appreciate Professors Naohito Abe and Toshiaki Watanabe for their valuable comments and instructions. I also acknowledge Professors Takashi Kano, Tatsuyoshi Okimoto, Makoto Saito and all the participants of my presentations for their constructive comments. Their suggestions reflect the improvement of this dissertation. I also thank my family for their enduring financial support.

Finally, my recherches have been supported by Hitotsubashi University Global COE Program "Research Unit for Statistical and Empirical Analysis in Social Sciences".

December 2013

Hiroshi Morita

Contents

Chapter 1

Overview

1.1 Introduction

Since the collapse of the bubble economy in the early 1990s, Japan has fallen into a period of low economic growth. Furthermore, recent financial crises and some natural disasters have also exerted a serious negative influence on the Japanese economy. In the face of this economic downturn, the fiscal and monetary authorities in Japan have implemented a large number of economic stimulus packages. At the same time, however, concerns about the effectiveness of these economic policies are increasing. Although many researchers have tried to analyze this from an academic perspective, there has not been a great deal of consensus on this topic.[1] Facing such a situation, the principal goal of this dissertation is to reevaluate and clarify the macroeconomic effects of fiscal and monetary policy in Japan (in the framework of a time series analysis), by using the vector autoregression (VAR) model. Moreover, beyond the analysis of the effects of the macroeconomic policy, we also explore the sources of business fluctuations in the recent Japanese economy. Related to the Lehman shock, which created a huge negative impact in Japan, we focus on the role of external shocks, such as risk premiums and foreign

[1] Shioji (2000) and Miyao (2000) are seen as representative of research on monetary policy in Japan, while Watanabe et al. (2011) are referred to as representative of its fiscal policy.

demands. In this dissertation, we empirically investigate the effects of fiscal and monetary policy in the former part (Chapters 2, 3, and 4) and the source of business cycles in the latter part (Chapter 5).

The contributions of our analysis are summarized as follows. First, we identify anticipated fiscal policy shocks, that is, fiscal news shocks, as well as unanticipated fiscal policy shocks. The foreseeability of fiscal policy has so far been disregarded in Japan's macroeconomic analysis. However, as Ramey (2011) points out, there is a possibility that the implementation of fiscal policy is subject to two lags (i.e., decision and implementation lags); in this light, we must estimate the actual effects of fiscal policy.

Second, we adopt new and advanced estimation techniques in our analysis. Throughout this dissertation, we identify structural shocks by sign restrictions originally proposed by Uhlig (2005). In addition, we achieve model-based identification that cannot be done by a recursive restriction using Cholesky decomposition, in terms of deriving sign restrictions from the Dynamic Stochastic General Equilibrium (DSGE) model. In addition, the Markov Switching model and the time-varying parameters VAR model are employed in Chapters 3 and 4. The Markov Chain Monte Carlo (MCMC) method is also used in these estimations.

Third, in Chapter 4, we present a new identification method for unconventional monetary policies conducted under the zero lower bound (ZLB) of the nominal interest rate.

Finally, this dissertation conducts analysis of current topics by using new data. For instance, Chapter 2 investigates the quantitative effects of Abenomics, which is an economic stimulus package implemented by Prime Minister Shinzo Abe, on macroeconomic variables by using historical decomposition. Chapter 5 clarifies exactly which structural shock can explain the rapid fall of output in light of the Lehman shock and the Great East Japan Earthquake.

1.2 Overview: Chapter 2

Chapter 2 investigates the effects of anticipated fiscal policy shock in Japan. Since Ramey (2011) indicated the importance of fiscal policy predictability (or so-called fiscal foresight), a large number of studies have analyzed the effects of anticipated fiscal policy for the U.S. economy (e.g., Mountford and Uhlig 2009; Mertens and Ravn 2010; Tenhofen and Wolff 2010). As stated by Ramey (2011), we cannot estimate the true effects of fiscal policy if we fail to capture the timing of fiscal policy shocks. On the other hand, in Japan, there have been few analyses that take fiscal foresight into account. The principal purpose throughout this dissertation, especially in Chapters 2, 3 and 4, is to reevaluate the effects of fiscal policy of the Japanese fiscal policy based on the ideas of Ramey (2011).

In Chapter 2, we identify anticipated fiscal policy shocks by combining the approach presented in Fisher and Peters (2010) and model-based robust sign restrictions. Fisher and Peters (2010) identify anticipated government military spending shock by regarding innovations occurring in excess stock returns in the military contractors as a proxy of fiscal news shock. We apply their identification scheme to the relationship between government spending and the construction industry in Japan. However, a problem exists regarding the direct application of Fisher and Peters (2010) method to the Japanese economy. To be specific, not all variations in stock returns of the construction industry are driven by the fiscal news shock, because construction firms deal not only with the public sector but also the private sector. To cope with this problem, we introduce the sign restrictions developed by Uhlig (2005) to identify anticipated fiscal policy shock. In particular, the robust sign restrictions, in which restrictions are derived from the dynamic stochastic general equilibrium (DSGE) model, are employed as in Dedola and Neri (2007) and Pappa (2009). In Chapter 2, we build a New Keynesian (NK) model that is a variant of Gali et al. (2007), and then derive the sign restrictions characterizing anticipated fiscal policy shock from the theoretical model. In addition, Chapter 2 performs the historical decomposition (HD) in the period of Abenomics to evaluate the impact of the recent large fiscal news shock on macroeconomic dynamics.

The estimated results in Chapter 2 reveal that anticipated and unanticipated fiscal policy shocks have a positive effect on consumption, but a negative effect on investment. It also turns out that a large part of forecast error variances in private consumption are explained by anticipated fiscal policy shocks. The results of the HD focus on Abenomics indicate that the fiscal news shock occurred in 2013Q3 and 2013Q4, and that they positively contributed to private consumption.

1.3 Overview: Chapter 3

Chapter 3 aims to ascertain whether the macroeconomic effects of fiscal policy are affected by the existence of rule-of-thumb (ROT) households. This is motivated by the theoretical findings shown by Galí et al. (2007), in which they incorporate the households faced with liquidity constraints into the NK model and theoretically demonstrate the positive response of consumption to government spending shock observed in empirical works.

To achieve our purpose, Chapter 3 adopts the following process. First, we extend the consumption function presented by Campbell and Mankiw (1989) to the Markov Switching (MS) model. Based on the MS-consumption function, the sample period is divided into high and low ROT household states depending on the share of ROT households. Then, we estimate the VAR model for each sample period, and calculate the impulse response of consumption to fiscal policy shocks. Chapter 3 also uses sign restrictions and identifies unanticipated and anticipated fiscal policy shock.

The results of Chapter 3 are summarized as follows. The high ROT household states are observed in the periods following large negative economic shocks, such as the oil shock and the economic bubble burst in the early 1990s. The estimated parameter values in MS-consumption function indicate that the average growth rate of consumption is low, but that variance is elevated in the high ROT household period. The main finding in Chapter 3 is that consumption in a high ROT state responds more significantly to unanticipated fiscal policy shocks compared to consumption in a low ROT state. This finding empirically supports the theoretical results of Galí et al. (2007).

1.4 Overview: Chapter 4

In Chapter 4, we estimate the TVP-VAR model with stochastic volatility, developed by Primiceri (2005), and analyze time-varying effects for fiscal and monetary policy in Japan. One of the contributions of Chapter 4 is to present a new identification strategy for unconventional monetary policy, such as the zero interest rate policy (ZIRP) or quantitative easing (QE). We characterize monetary policy shock under the ZLB of the nominal interest rate by combining zero restrictions for this equation adopted by Nakajima (2011), with sign restrictions for unconventional monetary policy as discussed in Franta (2011). More precisely, we regard a shock that lowers the short-term interest rate and increases the monetary base as a policy shock, and we further assume that the short-term interest rate cannot respond to structural shocks under the ZLB period.

Moreover, the principal aim of Chapter 4 is to assess whether the effects of fiscal policy increase during the ZLB period, as theoretically shown in Christiano et al. (2011) and Eggertsson (2011). To accomplish this, we calculate the time-varying fiscal multiplier based on the estimated impulse response functions (IRFs).

The findings of Chapter 4 are as follows. In the ZLB period, the estimated volatility of the short-term interest rate is small, while that of the monetary base is large. In addition, both monetary and fiscal policies have positive effects on output throughout the sample period. Finally, we cannot obtain the evidence that fiscal policy effects increase during the period of the ZLB.

1.5 Overview: Chapter 5

Unlike the above chapters, Chapter 5 focuses on the effects of external shocks on the Japanese economy, such as risk premium shocks and foreign demand shocks. More precisely, we investigate the sources of Japanese business fluctuations since the 1990s, taking into account both external shocks and domestic supply and demand shock. The recent global financial crisis, which caused a large reduction in output, reconfirms the fact that the Japanese economy is heavily influenced by economic situations in foreign countries.

The methodology adopted in Chapter 5 is similar to Chapter 2, that is, we use the model-based robust sign restrictions. In this Chapter, we construct a small open economy NK model that is a variant of the one presented by Leeper et al. (2011) to determine the features of the structural shocks. Moreover, we perform the FEVD and HD as well as IRF analysis in order to evaluate the effects of external shocks not only qualitatively but also quantitatively. In particular, we clarify which shocks contribute to the rapid fall of output in the Lehman shock and the Great East Japan Earthquake.

The results obtained in Chapter 5 show that approximately 30% to 50% of the forecast error variances in output can be explained by external shocks. Furthermore, we demonstrate that supply shock is the main influencing factor in Japanese business fluctuations throughout the sample period and that the roles of external shocks have been growing in the post-Lehman shock period, including the effect of the Great East Japan Earthquake.

Chapter 2

The Effects of Anticipated Fiscal Policy Shock on Macroeconomic Dynamics in Japan

2.1 Introduction

Amidst the recent global financial crisis, large-scale fiscal stimulus packages (e.g., the American Recovery and Reinvestment Act) have been implemented worldwide. This situation continues to provoke debate on the macroeconomic effects of fiscal policy throughout the world, and has prompted the key question: Does fiscal policy stimulate economic activity? In order to answer this question, this study uses a Vector Autoregression (VAR) model to evaluate the effects of fiscal policy by using a macroeconomic time series of the Japanese economy for the sample period 1980Q1 – 2013Q2. Since the collapse of the bubble economy in the early 1990s, the Japanese economy has suffered a long depression; also, a large number of fiscal stimulus packages have been implemented throughout the sample period. In this sense, the Japanese economy is an attractive subject for this research.

There exists extensive literature (e.g., Bayoumi 2001; Kuttner and Posen 2002; Watanabe, Yabu and Ito 2011) that has already investigated the effects of fiscal policy in Japan using VAR analysis. However, these previous works overlook the important fact that changes in fiscal policy might be anticipated. As noted by Blanchard and Perotti (2002) and Ramey (2011), fiscal policy is subject to two lags: the decision lag and the implementation lag. The former denotes a period between the time when a regulation is submitted and the time when it is enacted; the latter refers to the period from the enactment of the regulation to its actual enforcement. Owing to the existence of the implementation lag, there is the possibility that agents are aware of the change in the fiscal policy and can react to it immediately. Therefore, if such an increase in government spending is identified as an unanticipated shock, we may fail to capture the true effects of fiscal policy. This study examines the effects of fiscal policy while taking into account the possibility that the fiscal policy is anticipated. The foreseeability of fiscal policy is called fiscal foresight. In this study, we investigate the effects of both anticipated and unanticipated fiscal policy shocks.

The Japanese economy has recently experienced a news shock regarding a large fiscal stimulus package– the economic recovery policy put forth by Prime Minister Shinzo Abe called "Abenomics". In fact, stock prices in Japan (the Nikkei average) soared to more than ten thousand yen after Shinzo Abe won the general election, in spite of his not yet having implemented any policies. Moreover, most fiscal stimulus packages in the 1990s and 2000s also affected stock prices, as discussed in Fukuda and Yamada (2011). This evidence presented in Fukuda and Yamada (2011) demonstrates that fiscal foresight is the key to understanding the true effects of fiscal policy, at least in Japan. That such findings have been repeatedly observed is the another motivation for analyzing the Japanese economy, in addition to the Abenomics policy itself being an important research target. Of course, the empirical methodology conducted in this study is applicable to the economies of other countries.

Based on the approach of Fisher and Peters (2010), we identify the anticipated fiscal policy shocks. Their approach is as follows. If the financial market is efficient and agents are forward looking, the asset prices reflect the information that is currently available. Hence, news regarding changes in fiscal policies causes fluctuations in the prices of stocks of companies that are affected by the fiscal policy. On the basis of this assumption, they identified government (military) spending shocks as innovations on the excess stock returns of large U.S. military contractors. In this study, we apply this identification strategy to the relationship between government spending and the construction industry in Japan. We do this because fiscal policy aimed at an economic stimulus in Japan is usually performed through public works, and those public works are undertaken by the construction industry.

However, there are some caveats regarding the direct application of the Fisher and Peters (2010) method to our analysis. First, as Fisher and Peters (2010) also states, not all variations in stock returns are due to fiscal policy news, because firms sell not only to the public sector but also to the private sector. In addition, stock prices are also influenced by the workings of the entire economy. To resolve this problem, they identify the top three military contractors and determine the excess stock returns as defined by the deviation of the stock returns of these military contractors from the market returns. While we also adopt the excess stock returns concept, the unavailability of data makes it impossible to identify the construction firms that undertake the largest amount of public works projects in Japan. Therefore, to overcome this problem, we adopt the sign-restriction VAR methodology developed by Uhlig (2005). Since the sign-restricted VAR model identifies structural shocks by imposing restrictions on the shape of the impulse response function (IRF), this approach enables us to isolate the changes in stock returns that result from anticipated fiscal policy shocks involving future increases in government spending. Therefore, we identify anticipated fiscal policy shocks by imposing the condition that excess stock returns rise in the first period, and are followed by increased government spending.

A second problem that is uniquely associated with this study stems from using the stock returns of the construction industry as a variable in the analysis. In general, the construction industry tends to be affected by a health of the Japanese economy more than other industries, and thus, the influence of business fluctuations might persist, despite using the excess stock returns. In other words, there is a possibility that stock returns of the construction industry also respond to other shocks. In order to confront this problem, we adopted the robust sign restriction that is employed in Dedola and Neri (2007) and Pappa (2009). In this methodology, sign restrictions are derived from the theoretical model. More precisely, the theoretical IRF for each structural shock is calculated under a sufficiently wide range of parameters, and the features that are common to any combination of parameter values are adopted as robust sign restrictions. Based on the theoretical prediction, we are able to determine the characteristics that distinguish anticipated fiscal policy shocks from other shocks.

After the calculation of the IRFs, we examine the relative importance of fiscal policy shocks in the fluctuation of Japanese businesses, and investigate the role of fiscal policy shocks–especially anticipated fiscal shocks–in the recent Abenomics period. For these purposes, we perform forecast error variance decomposition (FEVD) and a historical decomposition (HD), as well as an IRF analysis. These methods enable us to evaluate the effect of fiscal policy shocks both qualitatively and quantitatively.

Ramey (2011) pointed out that the standard VAR analysis without fiscal foresight fails to capture the true effects of fiscal policy. Subsequently, several studies attempted to estimate the effects of anticipated fiscal policy in the U.S. (e.g., Mountford and Uhlig 2009; Tenhofen and Wolff 2010; Fisher and Peters 2010; Mertens and Ravn 2010). To the best of my knowledge, this is the first study that estimates the effects of anticipated fiscal policy shocks in Japan. Also, this study makes a theoretical contribution by extending the model of Galí et al. (2007), which is often employed in the literature, to incorporate fiscal news shocks. Moreover, the method presented in this study is widely applicable to the analysis of the economies of other countries.

Our analysis reveals the following. First, by comparing the estimated series of anticipated fiscal policy shocks and announcement dates of fiscal stimulus packages, as reported in Fukuda and Yamada (2011), our identification strategy seems likely to correctly capture the fiscal news shock. Secondly, the IRF analysis reveals that fiscal policy shocks have positive effects on consumption but negative effects on investment. In particular, it turns out that consumption persistently responds positively to anticipated fiscal policy shocks, while investment is crowded out by unanticipated fiscal policy shocks. Thirdly, the results of FEVD also show that anticipated fiscal shocks play a large role in the variation of consumption. Finally, from the results of HD–particularly focusing on Abenomics–we confirm that the fiscal news shocks that occurred in 2013Q3 and 2013Q4 positively contributed to consumption.

The rest of this study is organized as follows. In Section 2.2, we present the theoretical model and determine the robust sign restrictions by calculating the theoretical IRFs. Section 2.3 explains the estimation method of the sign-restricted VAR model and then describes the data series and the specifications of the estimation model employed in this study. Section 2.4 presents the estimation results of the IRF, FEVD, and HD analyses. Section 2.5 presents the conclusions.

2.2 Theoretical Model

In order to determine the robust sign restrictions that characterize a fiscal policy shock, we adopt the methodology used in Dedola and Neri (2007) and Pappa (2009). In line with their method, we first construct a New Keynesian (NK) model similar to Galí et al. (2007), and we then determine the common features of the dynamics induced by fiscal policy shocks by calculating the IRFs under various parameterizations.

To replicate the positive response of consumption to an unanticipated government spending shock, as seen in the results of the VAR analysis, the model of Galí et al. (2007) has the following prominent characteristics: price stickiness, rule-of-thumb households, debt financing, and wage unions. In addition to the above features, we incorporate wage stickiness, public capital, and the news process of fiscal policy shocks, and we extend the fiscal and monetary policy rules to include reactions to the output gap. These extensions allow us to obtain the IRFs corresponding to the model of a wider class and thus mitigate the problem of misspecification in the theoretical model. Moreover, the structural shocks other than fiscal policy shocks (i.e., technology, monetary policy, and labor preference shocks) are also included in our model to ensure that the dynamics that characterize fiscal policy shocks are unique and cannot be generated by other shocks.

2.2.1 Households

The economy is populated by a continuum of households indexed by $i \in [0, 1]$, and each household is comprised of a continuum of members. The households are divided into two types: optimizing or Ricardian (R) households that have access to the capital market, and rule-of-thumb or non-Ricardian (N) households that face liquidity constraints and consume all of their disposable income in each period. As in Galí et al. (2007) and Colciago (2011), we assume that a fraction $\mu \in [0, 1]$ of the population are non-Ricardian households, and the remaining population $1 - \mu$ are Ricardian households.

Labor market structure

We first explain the labor market structure. Following Colciago (2011), we assume a continuum of differentiated labor input indexed by $l \in [0, 1]$. There exist labor unions that correspond to each differentiated labor input, and each union sets their wage rate. As adopted in Schmitt-Grohe and Uribe (2006) and Colciago (2011), it is assumed that each member of a household provides each possible type of labor input, i.e., each household belongs to every labor union. Furthermore, as stated in Galí et al. (2007), labor supply is assumed to be determined by labor demand (not by the optimal choice of households), given the wage fixed by labor union. The labor supply to a differentiated labor input l is given by

$$n_t(l) = \left(\frac{W_t(l)}{W_t}\right)^{-\varepsilon_w} n_t^d \tag{2.1}$$

where ε_w is the elasticity of substitution between labor inputs. Here, n_t^d is the labor demand for an effective labor, and W_t is the aggregate nominal wage. The formal derivation of this equation is denoted in a section devoted to wage setting. Moreover, we assume that the members in each household are distributed uniformly across unions without depending on the type of households. This assumption ensures that the labor demand for differentiated labor input l is spread uniformly across the households, and thus there is no difference in labor supply across the household type, i.e., $n_t(i,l) = n_t^R(i,l) = n_t^N(i,l)$. Since each household supplies each type of differentiated labor input $n_t(i,l)$ and obtains the nominal wage $W_t(l)$, the total labor income for household i is given by $\int_0^1 W_t(l)n_t(i,l)dl$. This is common across households.

Ricardian households

Let $c_t^R(i)$ be the real consumption of Ricardian households. Then, the lifetime utility of Ricardian households is written by

$$U = E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{c_t^R(i)^{1-\gamma} - 1}{1-\gamma} - \chi_t \frac{n_t(i)^{1+\lambda}}{1+\lambda}\right], \tag{2.2}$$

where $n_t(i) = \int_0^1 n_t(i,l)dl$, and as stated above, we omit a superscript R from hours worked. They maximize this lifetime utility function subject to the budget constraint

$$P_t c_t^R(i) + P_t i_t^R(i) + B_t^R = \int_0^1 W_t(l) n_t(i,l)dl + P_t r_t^k k_{t-1}^R(i) + R_{t-1}B_{t-1}^R(i) + D_t(i) - P_t \tau_t^R(i), \tag{2.3}$$

and the capital accumulate equation,

$$k_t^R(i) = (1-\delta)k_{t-1}^R(i) + \left\{1 - S\left(\frac{i_t^R(i)}{i_{t-1}^R(i)}\right)\right\} i_t^R(i), \tag{2.4}$$

where χ_t captures the labor preference that follows an exogenous process, and β, γ, and λ denote the discount rate, risk aversion, and inverse of the Frisch labor elasticity, respectively. Uppercase letters denote nominal variables. P_t, B_t, and R_t are the aggregate price level, a one-period riskless nominal bond, and the gross nominal return on a bond, respectively. Since firms that produce intermediate goods face a monopolistic competition and make excess profits, Ricardian households receive dividends $D_t^R(i)$. $\tau_t^R(i)$ denotes the lump-sum taxes paid by Ricardian households. $i_t^R(i)$ and $k_t^R(i)$, respectively, denote real investment and real capital stock, and r_t^k is the real rental rate on capital. Unlike Galí et al. (2007), we assume that the adjustment costs of investment are proportional to the rate of change in the investment, as in Christiano, Eichenbaum, and Evans (2005), where $S(1) = S'(1) = 0$, and $S''(1) > 0$. As mentioned below, most studies using the dynamic stochastic general equilibrium (DSGE) model for Japan adopt this type of investment adjustment cost. Hence, we employ it to more easily calibrate our model.

Non-Ricardian households

Non-Ricardian households simply consume their current disposable income in each period. By denoting the consumption of non-Ricardian households as $c_t^N(i)$, their budget constraints are written as

$$P_t c_t^N(i) = \int_0^1 W_t(l) n_t(i,l) dl - P_t \tau_t^N(i), \tag{2.5}$$

where $\tau_t^N(i)$ denotes the lump-sum taxes paid by non-Ricardian households. As in Galí et al. (2007), we assume lump-sum tax paid by non-Ricardian households differ from those for Ricardian households in order to equate steady state consumption across household types.

2.2.2 Wage setting

As discussed above, each household provides a differentiated labor input indexed by $l \in [0,1]$, and each household belongs to every labor union l. A perfectly competitive labor-bundling firm bundles a differentiated labor input $n_t(l)$ into an effective labor n_t, according to

$$n_t = \left[\int_0^1 n_t(l)^{\frac{\varepsilon_w - 1}{\varepsilon_w}} dl \right]^{\frac{\varepsilon_w}{\varepsilon_w - 1}}. \tag{2.6}$$

As a result of the labor-bundling problem, the demand function for each differentiated labor input is expressed as Eq.(2.1) for all l. The aggregate nominal wage is equal to

$$W_t = \left[\int_0^1 W_t(l)^{1-\varepsilon_w} dl \right]^{\frac{1}{1-\varepsilon_w}}. \tag{2.7}$$

With respect to wage setting, we follow the model used in Galí et al. (2007) and Colciago (2011), in which each labor union l sets its nominal wage $W_t(l)$ to maximize the weighted average of the lifetime utility of Ricardian and non-Ricardian households. In each period, a labor union resets the optimal nominal wage $W_t^*(l)$ with a probability $1-\rho_w$. Thus, the problem for a labor union l is written as

$$\max_{W_t^*(l)} E_t \sum_{s=0}^{\infty} \rho_w \Lambda_{t,t+s} \left[(1-\mu) \frac{c_{t+s}^R(l)^{1-\gamma} - 1}{1-\gamma} + \mu \frac{c_{t+s}^N(l)^{1-\gamma} - 1}{1-\gamma} - \chi_t \frac{n_{t+s}(l)^{1+\lambda}}{1+\lambda} \right], \tag{2.8}$$

subject to (2.3), (5.4), and (2.1), where $\Lambda_{t,t+s} = \beta^s (c_{t+s}^R / c_t^R)^{-1}$ denotes the stochastic discount factor. In the symmetric equilibrium, the first-order condition can be expressed as

$$W_t^*(l) = \frac{\varepsilon_w}{\varepsilon_w - 1} \frac{E_t \sum_{s=0}^{\infty} \rho_w^s \Lambda_{t,t+s} \chi_t n_{t+s}^{1+\lambda}}{E_t \sum_{s=0}^{\infty} \rho_w^s \Lambda_{t,t+s} \left[(1-\mu) \frac{n_{t+s}}{P_{t+s} c_{t+s}^{R\ \gamma}} + \mu \frac{n_{t+s}}{P_{t+s} c_{t+s}^{N\ \gamma}} \right]}. \tag{2.9}$$

Combining (2.9) with (2.7), the evolution of the aggregate nominal wage is given by

$$W_t = [(1-\rho_w) W_t^{*\,1-\varepsilon_w} + \rho_w W_{t-1}^{1-\varepsilon_w}]^{\frac{1}{1-\varepsilon_w}}. \tag{2.10}$$

Then, the log-linearization of (2.9) and (2.10) around the steady state yields the dynamic equation of the real wage as

$$\hat{w}_t = \Gamma \hat{w}_{t-1} + \Gamma \beta E_t \hat{w}_{t+1} + \Gamma \beta E_t \hat{\pi}_{t+1} - \Gamma \hat{\pi}_t + \kappa_w \Gamma \gamma \hat{c}_t + \kappa_w \Gamma \lambda \hat{n}_t + \kappa_w \Gamma \hat{\chi}_t, \tag{2.11}$$

where a hat denotes the deviation from the steady-state value, $\Gamma = \rho_w / (1 + \beta \rho_w^2)$, and $\kappa_w = (1-\beta\rho_w)(1-\rho_w)/\rho_w$.

2.2.3 Firms

The production sector consists of two types of firms: the monopolistically competitive firms that produce differentiated intermediate goods, and the perfectly competitive firms that produce single final goods by using intermediate goods as the input. Each intermediate-goods firm, indexed by $j \in [0,1]$, produces an intermediate good $y_t(j)$, and its production function is assumed to be the Cobb-Douglas form:

$$y_t(j) = z_t n_t(j)^{1-\alpha} k_{t-1}(j)^{\alpha} {k_{t-1}^{g}}^{\alpha_g} \tag{2.12}$$

where $k_{t-1}(j)$ and $n_t(j)$, respectively, denote the capital stock and the labor input used by firm j, and k_{t-1}^{g} denotes the public capital stock. Also, z_t indicates the total factor productivity (TFP), which is given exogenously.

Although the market for intermediate goods is monopolistically competitive, the factor market faced by intermediate-goods firms is assumed to be competitive. As a result of the cost minimization problem for intermediate-goods firms, the real marginal cost mc_t is given by

$$mc_t = \frac{w_t}{(1-\alpha)A_t {k_t^{g}}^{\alpha_g}} \left(\frac{(1-\alpha)r_t^k}{\alpha w_t} \right)^{\alpha}. \tag{2.13}$$

The final-goods firm has the following (CES) technology to produce the final goods y_t:

$$y_t = \left[\int_0^1 y_t(j)^{\frac{\varepsilon_p - 1}{\varepsilon_p}} dj \right]^{\frac{\varepsilon_p}{\varepsilon_p - 1}}, \tag{2.14}$$

where ε_p is the elasticity of substitution across each type of intermediate goods. By solving a profit maximization problem for the final-goods firm, the demand function for the intermediate goods is obtained as

$$y_t(j) = \left(\frac{P_t(j)}{P_t} \right)^{-\varepsilon_p} y_t, \tag{2.15}$$

and the final-goods pricing rule is written as

$$P_t = \left[\int_0^1 P_t(j)^{1-\varepsilon_p} dj\right]^{\frac{1}{1-\varepsilon_p}}. \tag{2.16}$$

2.2.4 Price setting

As in a wage union, intermediate-goods firms set their prices according to the Calvo (1983) mechanism. An intermediate-goods firm j can change its price with probability $1-\rho_p$, and thus the optimal price $P_t^*(j)$ is determined by solving the problem:

$$\max_{P_t^*(j)} E_t \sum_{s=0}^{\infty} \rho_p^s \Lambda_{t,t+s} \left[P_t^*(j) y_{t+s}(j) - P_{t+s} y_{t+s}(j) mc_{t+s}\right] \tag{2.17}$$

subject to the demand function (2.15). The laws of motion for the optimal and aggregate prices are written in way similar to that of wages:

$$P_t^* = \frac{\varepsilon_p}{\varepsilon_p - 1} \frac{E_t \sum_{s=0}^{\infty} \rho_p^s \Lambda_{t,t+s} P_{t+s} y_{t+s}(j) mc_{t+s}}{E_t \sum_{s=0}^{\infty} \rho_p^s \Lambda_{t,t+s} y_{t+s}(j)} \tag{2.18}$$

and

$$P_t = \left[(1-\rho_p) P_t^{*1-\varepsilon_p} + \rho_p P_{t-1}^{1-\varepsilon_p}\right]^{\frac{1}{1-\varepsilon_p}}. \tag{2.19}$$

The New-Keynesian Phillips curve is obtained by log-linearization of (2.18) and (2.19), as follows:

$$\hat{\pi}_t = \beta E_t \hat{\pi}_{t+1} + \kappa_p \hat{mc}_t \tag{2.20}$$

where $\kappa_p = (1-\beta\rho_p)(1-\rho_p)/\rho_p$.

2.2.5 Fiscal policy and monetary policy

The government budget constraint is

$$P_t \tau_t + B_t = P_t g_t + R_{t-1} B_{t-1} \tag{2.21}$$

where g_t denotes the real government spending. Also, we assume a tax rule of the form

$$\hat{\tau}_t = \phi_b \hat{b}_{t-1} + \phi_g \hat{g}_t + \phi_y \hat{y}_t \tag{2.22}$$

where $\hat{\tau}_t \equiv (\tau_t - \tau)/y$, $\hat{b}_t \equiv (\frac{B_t}{P_t} - \frac{B}{P})/y$, and $\hat{g}_t \equiv (g_t - g)/y$. Furthermore, the public capital simply accumulates as follows;

$$k_t^g = (1-\delta) k_{t-1}^g + g_t. \tag{2.23}$$

On the other hand, the monetary authority is assumed to set the nominal interest rate according to a simple Taylor rule:

$$\hat{r}_t = \psi_\pi \hat{\pi}_t + \psi_y \hat{y}_t + u_t^m, \tag{2.24}$$

where a hat denotes the log deviation from the steady-state value, and u_t^m denotes a monetary policy disturbance, which is assumed to be exogenous.

2.2.6 Aggregate and Market clearing

Aggregate consumption, lump-sum taxes, capital, investment, bond, and dividends are given, respectively, by

$$c_t = (1-\mu)c_t^R + \mu c_t^N;\ k_t = (1-\mu)k_t^R;\ b_t = (1-\mu)b_t^R$$
$$\tau_t = (1-\mu)\tau_t^R + \mu\tau_t^N;\ i_t = (1-\mu)i_t^R;\ d_t = (1-\mu)d_t^R.$$

The clearing conditions of the factor and goods market are expressed as

$$n_t = \int_0^1 n_t(j)dj;\ k_t = \int_0^1 k_t(j)dj;$$
$$y_t = c_t + i_t + g_t.$$

2.2.7 Dynamics of exogenous variables

The dynamics of the exogenous variables g_t, z_t, u_t^m, and χ_t are assumed to be, respectively,

$$\hat{g}_t = \rho_g \hat{g}_{t-1} + \varepsilon_t^g + \xi_{t-q}^g \tag{2.25}$$

$$\hat{z}_t = \rho_z \hat{z}_{t-1} + \varepsilon_t^z + \xi_{t-q}^z \tag{2.26}$$

$$\hat{u}_t^m = \rho_m \hat{u}_{t-1}^m + \varepsilon_t^m + \xi_{t-q}^m \tag{2.27}$$

$$\hat{\chi}_t = \rho_\chi \hat{\chi}_{t-1} + \varepsilon_t^\chi + \xi_{t-q}^\chi \tag{2.28}$$

where ε_t^k, $k \in [g, z, m, \chi]$ denotes an unanticipated shock in period t. ξ_{t-q}^k,$k \in [g, z, m, \chi]$ denotes an anticipated shock that is realized in period t but that was announced in period $t-q$.

2.2.8 Parameter range

The sign restrictions imposed on our VAR model are set on the basis of the IRFs derived from the above theoretical model. In order to determine the robust sign restrictions, the IRFs are computed by giving the ranges for certain parameters. More precisely, the theoretical IRFs in this study are calculated as follows. Our process for finding the robust sign restrictions follows Pappa (2009). Let Θ denote the parameters in the interval $[\theta_l, \theta_u]$. Θ is assumed to be uniformly distributed in the interval $[\theta_l, \theta_u]$, namely $\Theta \sim U(\theta_l, \theta_u)$. We randomly draw Θ and calculate the IRFs. Repeating this process sufficiently many times provides the range of IRFs that corresponds to the combination of the various parameter values. Only robust signs are then adopted as the restrictions imposed on the empirical model. The values of the parameters are selected based on the results estimated in previous studies or on the values used in the exercises.

In this study, a quarter is chosen as the unit period in order for it to match the frequency of the data used in the empirical analysis. Previous studies have estimated the degree of risk aversion γ to be in the range from 1.25 (Sugo and Ueda, 2008) to 1.91 (Iiboshi, Nisiyama and Watanabe, 2008). Thus, we restrict γ to the interval $[1, 2]$. The inverse of labor supply substitution λ was estimated to be 2.08 in Iiboshi, Nisiyama, and Watanabe (2008) and 2.15 in Sugo and Ueda (2008), while Galí et al. (2007) adopted 0.5. Thus, in this study, we restrict λ to the interval $[0.5, 2]$.

The parameter μ indicates the proportion of non-Ricardian households and determines the dynamics of aggregate consumption, as presented in Galí et al. (2007). Previous analysis using macro time-series data (Hatano, 2004; Iwata, 2009) have estimated the value of μ for Japan to be 0.3; other analyses using micro data (Kohara and Horioka, 2006) have these values to be $0.08 - 0.15$. In consideration of these results, we limit μ to the interval $[0.1, 0.4]$. The lower and upper bounds of the stickiness of prices and wages are chosen to be 0.2 and 0.9, respectively. The upper bound is set to be somewhat larger than the value estimated in Iiboshi, Nisiyama, and Watanabe (2008) and Sugo and Ueda (2008), while the lower is set to be consistent with the value in Pappa (2009). Also, the interval for the investment adjustment cost κ is set to be $[0, 0.3]$, which is centered on the value of 0.15 used in Sugo and Ueda (2008).

The elasticity of taxes to government spending and bonds (i.e., ϕ_g,ϕ_b) are chosen on the basis given in Galí et al. (2007). However, the calibration parameter values in Galí et al. (2007) were based on the U.S. economy. Thus, we adopt wider intervals for these parameters to mitigate this problem. To be specific, the interval for the elasticity of tax to government spending ϕ_g was $[-0.25, 0.25]$ and that to debt ϕ_b is $[0, 0.5]$.[1] The elasticity of tax to output is set to be $[0.125, 0.175]$ based on the results presented in Watanabe, Yabu, and Ito (2011). The parameters in the monetary policy rule are set as follows. The response of the interest rate to inflation was limited to the interval $[1.01, 1.5]$. This range fulfills the Taylor principle, and therefore it is often used in calibration exercises. The intervals of the coefficient on the output gap is set to be $[0, 0.2]$ based on the results in Iiboshi, Nisiyama, and Watanabe (2006) and Sugo and Ueda (2008).

The parameters for the persistency of the exogenous variables and the elasticity of substitution in production and labor are assumed to be in the ranges $[0.8, 0.95]$ and $[6, 11]$, respectively. The remaining parameters are then fixed to particular values. All parameter values and intervals are displayed in Table 2.1.

2.2.9 Sign restrictions

Figure 2.1 and Figure 2.2 display the 68% probability bands for the responses of government spending, deficit, and output to unanticipated (Figure 2.1) and anticipated (Figure 2.2) fiscal

[1] In Galí et al. (2007), ϕ_g and ϕ_b were set to be 0.13 and 0.33, respectively.

Table 2.1: Calibration parameters

Parameter	Value	Description
β	0.99	Subjective discount factor
δ	0.025	Depreciation rate
α	0.3	Share of capital
γ	$[1, 2]$	Risk aversion
λ	$[0.5, 2]$	Inverse labor supply elasticity
α_g	$[0, 0.2]$	Productivity of public capital
κ	$[0, 0.3]$	Investment adjustment cost
μ	$[0.1, 0.4]$	Share of non-Ricardian households
ϵ_p	$[6, 11]$	Elasticity of substitution in production
ϵ_w	$[6, 11]$	Elasticity of substitution in labor
ρ_p	$[0.2, 0.9]$	Calvo parameter on prices
ρ_w	$[0.2, 0.9]$	Calvo parameter on wages
ϕ_g	$[-0.25, 0.25]$	Elasticity of tax to government spending
ϕ_b	$[0, 0.5]$	Elasticity of tax to bond
ϕ_y	$[0.125, 0.175]$	Elasticity of tax to output
ψ_π	$[1.01, 1.5]$	Monetary policy response of inflation
ψ_y	$[0, 0.2]$	Monetary policy response of output
ρ_k	$[0.8, 0.95]$	persistency of exogenous shocks

Note: The range of parameter value are basically set up according to the previous studies.

policy, technology, monetary policy, and labor preference shocks. Fiscal policy and technology shocks indicate a 1% increase in government spending and TFP, while monetary policy and labor preference shocks indicate a 1% decrease in interest rate and labor preference. In the benchmark calibration, the foresight period p is assumed to be 3, namely, the news announced in the first period is realized in the fourth period. In this exercise, the number of random draws is $10,000$.

In Figure 2.1, we observe that a positive government spending shock that is financed by a deficit will immediately increase output. This is caused by a negative wealth effect, namely, households increase their labor supply because an increase in government spending lowers their lifetime income and raises output. Therefore, fiscal policy shocks increase deficit and output. By contrast, other structural shocks we considered raised output but decreased the deficit because an increase in the output increased the tax revenue through the tax rule. Thus, as stated in

Table 2.2: Sign restriction

	gov. spending	stock returns	deficit	GDP
unanticipated	> 0 for 1Q		> 0 for 1Q	> 0 for 1Q
anticipated	$= 0$ for 1Q > 0 for 4Q-6Q	> 0 for 1Q	> 0 for 4Q-6Q	> 0 for 4Q-6Q

Note: Sign restrictions adopted in this study are based on the theoretical IRF shown in Figure 2.1 and Figure 2.2.

Pappa (2009), the responses of a deficit enable us to distinguish fiscal policy from other shocks. Similar to unanticipated shocks, anticipated fiscal policy shocks can be distinguished from other shocks by using the response of a deficit. As seen in Figure 2.2, the output indicates there were positive responses to four anticipated structural shocks during the period in which the news was realized. However, the deficit responded positively only to anticipated fiscal policy shocks.

Based on these results and the discussion in Section 2.1, we impose sign restrictions to the path of government spending, excess stock returns, the deficit, and GDP. Table 2.2 summarizes the sign restrictions that are adopted in the VAR analysis. For the unanticipated fiscal policy shocks, the sign restrictions are imposed during the impact period. On the other hand, the restrictions characterizing anticipated fiscal policy shocks are imposed on government spending, the deficit, and the GDP, from the 4th to the 6th period. This is because there is a possibility that the duration of time from the announcement of the news until the change in policy is realized will differ for each fiscal package. As in Peersman (2005), a sign restriction is only imposed on stock returns during the impact period since stock returns are financial variables. We also assume that government spending is exogenous. In other words, we assume that government spending is affected by unanticipated fiscal policy shocks only at the impact period. This assumption is widely employed in most of the previous studies (e.g., Blanchard and Perotti 2002, Galí et al. 2007). In our case, this restriction eliminates the situation in which government spending immediately responds to an anticipated fiscal policy shock. If government spending rises during the period in which the news is announced, this increase in government spending is unanticipated. Therefore, this restriction on government spending to anticipated fiscal policy shocks plays an important role in distinguishing between unanticipated and anticipated shocks.

2.3 Estimation methodology

2.3.1 Sign-restricted VAR model

The sign-restricted VAR model is estimated by the following process. First, we estimate the reduced-form VAR model:

$$Y_t = B_1 Y_{t-1} + B_2 Y_{t-2} + \cdots + B_p Y_{t-p} + u_t \tag{2.29}$$

$$u_t = A_0 \varepsilon_t \tag{2.30}$$

$$u_t \sim N(0, \Sigma),\ \ \varepsilon_t \sim N(0, I) \tag{2.31}$$

where Y_t is a vector of endogenous variables that include (in this order) government spending, excess stock returns of the construction industry, the deficit, GDP, private consumption, and nonresidential investment; $B = [B_1, \cdots, B_p]$ are matrices with coefficients; and u_t is a vector of the reduced-form residuals of the variance-covariance matrix, Σ. A_0 is a lower triangular matrix obtained from the Cholesky decomposition of Σ, and ε_t is a vector of structural shocks that are mutually independent and normalized to be of variance 1.

Next, we draw random samples of B and Σ from their posterior distributions. Using the noninformative Normal-Wishart family as the prior distribution, the posterior distributions of

$vec(B)$ and Σ^{-1}, respectively, become $N(vec(\hat{B}), \hat{\Sigma} \otimes (X'X)^{-1})$, and $W(\hat{\Sigma}^{-1}/T, T)$, where $\hat{B}$ and $\hat{\Sigma}$ are OLS estimators, X is the matrix of the explanatory variables, and T is the sample size. For each draw, we then calculate the structural shocks and the matrix of contemporaneous relations among the endogenous variables. In this step, we randomly generate the orthogonal matrix Q, such that $Q'Q = I$. Using this matrix, (2.30) can be rewritten as

$$u_t = A_0 Q' Q \varepsilon_t = A \tilde{\varepsilon}_t. \tag{2.32}$$

Then,

$$E\left[A\tilde{\varepsilon}_t \tilde{\varepsilon}_t' A'\right] = E\left[A_0 Q' Q \varepsilon_t \varepsilon_t' Q' Q A_0'\right] = A_0' A_0 = \Sigma. \tag{2.33}$$

Therefore, we construct a new set of structural shocks $\tilde{\varepsilon}_t$ and contemporaneous relations A that maintain the variance-covariance structure. Some values of the IRFs can be calculated from the set of (B, Σ) by randomly generating a Q matrix.

In addition, we impose an assumption on the elements of Q, as follows:

$$Q = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & q_1 & q_2 & q_3 & q_4 & q_5 \\ 0 & q_6 & q_7 & q_8 & q_9 & q_{10} \\ 0 & q_{11} & q_{12} & q_{13} & q_{14} & q_{15} \\ 0 & q_{16} & q_{17} & q_{18} & q_{19} & q_{20} \\ 0 & q_{21} & q_{22} & q_{23} & q_{24} & q_{25} \end{bmatrix}. \tag{2.34}$$

By using this form of the Q matrix, government spending is affected by only unanticipated fiscal policy shock at the impact period. The new contemporaneous relation A obtained by multiplying Q by the lower triangular matrix A_0 becomes

$$A = A_0 Q' = \begin{bmatrix} a_1 & 0 & 0 & 0 & 0 & 0 \\ a_2 & a_3 & a_4 & a_5 & a_6 & a_7 \\ a_8 & a_9 & a_{10} & a_{11} & a_{12} & a_{13} \\ a_{14} & a_{15} & a_{16} & a_{17} & a_{18} & a_{19} \\ a_{20} & a_{21} & a_{22} & a_{23} & a_{24} & a_{25} \\ a_{26} & a_{27} & a_{28} & a_{29} & a_{30} & a_{31} \end{bmatrix}. \tag{2.35}$$

Under this form of A, the exogeneity of government spending is always assured.

Furthermore, the present study generates a Q matrix using the following procedure. Let q be defined as

$$q = \begin{bmatrix} q_1 & q_2 & q_3 & q_4 & q_5 \\ q_6 & q_7 & q_8 & q_9 & q_{10} \\ q_{11} & q_{12} & q_{13} & q_{14} & q_{15} \\ q_{16} & q_{17} & q_{18} & q_{19} & q_{20} \\ q_{21} & q_{22} & q_{23} & q_{24} & q_{25} \end{bmatrix}. \tag{2.36}$$

This matrix is also an orthogonal matrix similar to Q. In order to generate q, we use the method adopted in Peersman (2005). For a 5×5 Givens matrix Q_{12}, Q_{13}, Q_{14}, Q_{15}, Q_{23}, Q_{24}, Q_{25}, Q_{34}, Q_{35}, Q_{45}, where $Q_{ij}(\theta)$ is a matrix with $cos(\theta)$ as the (i,i)-th and (j,j)-th elements, $-sin(\theta)$ as the (j,i)-th element, and $sin(\theta)$ as the (i,j)-th element. The diagonal element of this matrix is one, while the off-diagonal elements are equal to zero. Then, the Q matrix is defined as

$$Q = \prod_{i,j} Q_{ij}(\theta_k)$$

where $\theta_k, k = 1, \cdots, 10$, are drawn randomly from a uniform distribution $U(0, 360)$.

Finally, the IRFs are calculated based on each draw (B, Σ, A). If they satisfy the sign restriction in Table 2.2, they are valid IRF candidates and are preserved; otherwise, they are discarded. By repeating the above processes, the IRFs that are consistent with the sign restriction imposed in Table 2.2 are obtained. In this study, there were 300 random draws for generating B and Σ, and 500 for generating A.

2.3.2 Data and Specification

We used quarterly data of government spending, GDP, private consumption, nonresidential investment, tax revenue, and excess stock returns of the construction industry for the period 1980Q1-2013Q2. The government spending is defined as the sum of government consumption and public investment. The series, except for the excess stock returns, are real, seasonally adjusted per capita, and logarithmized. The first five variables are downloaded from the System of National Accounts (SNA) database in Japan. We use "taxes and stamp revenues" in the *Monthly Finance Review* issued by the Ministry of Finance, Japan, as the series of tax revenue. It does not cover a subsidy or the transfer. The deficit in the VAR model is defined as the log differential of government spending and tax revenue. The data on excess stock returns is calculated as follows. First, the quarterly average of the stock prices in both the construction industry and the whole market are computed by using the closing prices of the daily data. These data are taken from the Nikkei NEEDS-Financial QUEST database. We then obtain the series of stock returns as the growth rate of stock prices, based on this quarterly data. Finally, following Fisher and Peters (2010), we construct the excess stock returns by subtracting the returns of the construction industry from the returns of the whole market. Similar to the method in Fisher and Peters (2010), we use the accumulated excess stock returns in our analysis. The data used in this study are described in Figure 2.3.

The estimated system is a six-variable VAR that includes linear and quadratic time trends as well as a constant term. In order to detrend the data series so that it is consistent with the DSGE model presented above, we incorporate the time trends into the estimated system. Based on the claim of Sims, Stock, and Watson (1990) that taking the differences of a time series may result in the loss of important information, the estimation is carried out in levels. Although the Akaike information criterion suggests two lags, we set the number to four in order to sufficiently capture the dynamic effects of fiscal policy. In this framework, we identify two types of fiscal policy shocks: unanticipated and anticipated.

2.4 Empirical results

2.4.1 Estimated structural shock

Figure 2.4 shows a plot of the estimated series of anticipated fiscal policy shocks; the solid blue line and dotted red line denote the estimated series and one standard deviation, respectively. We first confirm that our identification method correctly captures the news shock with respect to fiscal policies.

Fukuda and Yamada (2011) reported the announcement dates of fiscal stimulus packages between 1992 and 2010. We compare the series of anticipated fiscal policy shocks and the dates of announcement. In the estimated series of anticipated fiscal policy shocks, between 1992 and 2010, there are eleven positive shocks that exceed one standard deviation. Among them, five shocks (i.e., 1992Q1, 1998Q1, 2008Q4, 2010Q3, and 2010Q4) correspond to the date reported in Fukuda and Yamada (2011), and two shocks (i.e., 1993Q4 and 2002Q3) only roughly correspond to the reported dates. Moreover, for these latter two shocks, the fiscal stimulus packages were announced in the period after these shocks were observed. Since the date reported in Fukuda and Yamada (2011) is the day that the outline of a fiscal package was published in a newspaper, there is a possibility that people foresaw the implementation of the fiscal policy prior to that day. Taking this possibility into account, we can consider that the latter two shocks also correctly capture the news shock with respect to a fiscal policy. Therefore, it seems that our identification method adequately captures anticipated fiscal policy shocks, as far as the estimated series of the shock is observed.

2.4.2 Impulse response function

The impulse response functions for unanticipated and anticipated fiscal policy shocks are displayed in Figure 2.5 and Figure 2.6, respectively. The solid blue lines and shaded areas indicate the median of sampled IRFs and the 68 percent credible intervals, respectively. In addition, we plot the IRFs that were the closest (in terms of minimizing the sum of the squares of the differences) to the median responses among those obtained in each admissible rotation, as shown by the red dotted line, in order to overcome the following problem. As pointed out by Fry and Pagan (2011) and Inoue and Kilian (2011), the median responses in the sign-restricted VAR model summarize the information from different structural models because each IRF is computed based on a different rotation matrix. In other words, the response that fully corresponds to the median might not exist in the set of admissible structural models, and thus the inferences using the median response might fail to lead to correct results. However, in our case, this problem is deemed to be less serious because the median responses and the closest ones share similar dynamics. In addition, the ratio of valid draws to all draws is 1.62% in this estimation.

In Figure 2.5, we see that government spending and the deficit both rose persistently beyond the first period in which the restrictions are imposed. This means that government spending is financed by issuing the debt. According to the sign restrictions, output has a positive sign during the impact period, but there is no persistency in its responses. The response of consumption also indicates a positive sign in response to unanticipated fiscal policy shocks. Contrary to output, however, consumption rises persistently in response to unanticipated fiscal policy shocks. It increases significantly for about one year, except for the third period. These positive responses of consumption agree with the findings for the U.S. that are presented in Blanchard and Perotti (2002) and Galí et al. (2007). In return, investment is crowded out by an increase in government spending, as predicted by the theory. It is considered that the reductions in investment cause the transitory effects in output. In addition, the excess stock returns rise significantly for the first two quarters.

Figure 2.6 shows the impulse response functions for anticipated fiscal policy shocks. Unlike the case of unanticipated fiscal policy shocks, government spending gradually increases and peaks during the sixth period, and thus describes a hump-shaped response. The deficit and stock returns are confirmed follow the sign restrictions. During the impact period, consumption rises but investment falls, and thus output hardly changes in response to the shock. Output has a positive sign from the fourth to the sixth periods following the restriction, and then it rises significantly after the sixth period. The response of consumption to an anticipated fiscal policy shock is most interesting. In response to good news regarding a future fiscal policy, consumption rises immediately and remains significantly positive. Compared to an unanticipated fiscal policy shock, the persistency of response in consumption is high. These findings are consistent with those in Fisher and Peters (2010), but they are inconsistent with those in Ramey (2011). Finally, investment falls, as is also seen in unanticipated fiscal policy shocks. However, its responses are insignificant except for during the third and fourth periods. Unlike the case with consumption, therefore, it seems that unanticipated fiscal policy shock plays an important role in the dynamics of investment.

2.4.3 Forecast error variance decomposition

The results of FEVD are summarized in Table 2.3 and Table 2.4. The estimated time horizon is 10 quarters, as in the earlier IRF analysis. These tables show the relative importance of each shock in terms of the variations in each variable. To overcome the problem mentioned above, we performed the FEVD for every admissible rotation and then took their medians. Therefore, the presented results indicate the median value across all decompositions.

Table 2.3: Forecast error variance decomposition -unanticipated fiscal policy shock-

	gov. spending	stock returns	deficit	GDP	consumption	investment
1	100.0	1.2	5.3	2.3	2.4	3.6
2	94.0	1.9	4.1	2.1	4.4	5.4
3	83.7	1.7	5.7	3.4	4.1	8.0
4	72.3	2.1	6.0	3.0	4.1	8.2
5	65.8	2.3	7.1	2.9	4.1	8.6
6	61.3	2.8	7.6	2.7	3.7	8.4
7	58.1	3.3	8.1	2.7	3.7	8.9
8	55.3	4.1	8.4	2.6	3.7	9.3
9	53.4	5.1	8.7	2.5	3.5	9.2
10	52.9	6.3	9.3	2.4	3.6	9.0

Note: This table denotes the results of forecast variance decomposition. The values in this table indicate the ratio explained by unanticipated fiscal policy shock.

Table 2.4: Forecast error variance decomposition -anticipated fiscal policy shock-

	gov. spending	stock returns	deficit	GDP	consumption	investment
1	0.0	9.3	1.1	1.6	21.6	4.0
2	2.1	8.0	3.5	2.1	17.7	4.8
3	6.3	7.1	3.5	2.3	17.1	4.6
4	10.7	7.1	3.9	2.9	19.0	4.6
5	14.0	7.1	3.9	3.6	18.5	4.1
6	16.3	7.9	3.8	4.9	17.8	3.8
7	17.4	8.4	3.6	6.3	17.7	3.6
8	18.0	8.9	3.6	7.0	16.6	3.7
9	18.3	9.5	3.8	7.2	15.8	3.8
10	18.1	9.7	4.1	7.3	15.1	3.8

Note: This table denotes the results of forecast variance decomposition. The values in this table indicate the ratio explained by unanticipated fiscal policy shock.

The principal finding is that anticipated fiscal policy shocks explain approximately 15% to

20% of the forecast error variances in consumption. Anticipated shocks play a greater role, compared to unanticipated shocks, in the variation of consumption. As with consumption, the variation of output is also explained by anticipated shock rather than by unanticipated shock. In particular, this tendency becomes clearer when the news shock is realized (i.e., after the fourth period). Investment is affected by unanticipated shocks. This is consistent with the result of the IRFs, in which unanticipated fiscal policy shocks significantly crowd-out investments.

Most of the variation in government spending can be explained by unanticipated fiscal policy shocks, but the ratio explained by anticipated shocks increases after the news shock has been realized. Finally, about 10% of the variation in stock returns is explained by anticipated fiscal policy shocks. In other words, not all of the variations in stock returns are caused by news shocks regarding fiscal policy. Therefore, this result justifies our identification method, that combines sign restrictions with the ideas presented in Fisher and Peters (2010).

2.4.4 Historical decomposition

Finally, we present the results of a HD in order to clarify the effects of Abenomics, especially with anticipated fiscal policy shocks. Figure 2.7 shows the HD of the data before 2012Q2 for government spending, stock returns, the deficit, the GDP, consumption, and investment conditioning. Similar to the FEVD, the presented results are the median values across all decompositions. In Figure 2.7, the blue and red bars indicate the contribution of the unanticipated and anticipated shocks, respectively.

First of all, we observe the positive contributions of anticipated fiscal policy shocks to the stock returns during 2012Q3 and 2012Q4, when Prime Minister Shinzo Abe won for the Liberal Democratic Party (LDP) presidential election and the general election in the Lower House. Not surprisingly, government spending was not largely affected by the anticipated shock during this time. Moreover, the anticipated fiscal policy shocks that occurred in 2012Q3 and 2012Q4 positively and immediately contributed to the dynamics of consumption. On the other hand, the effect on output is more gradual, and the positive contribution was observed in 2013Q2. The results that are observed in both consumption and output conform to the findings in the IRF and FEVD. Investment shows negative effects of influence during the estimated periods. Finally, we find that the variation in the deficit is affect by unanticipated fiscal policy shocks in the same way as in government spending and investment.

2.5 Conclusion

This study analyzes the effects of fiscal policy in Japan and, in particular, sheds light on the effects of anticipated fiscal policy shocks. We identify anticipated increases in government spending by using excess stock returns in the construction industry and by employing the robust sign restrictions based on a variant of the NK model of Galí et al. (2007).

We summarize the main results as follows. First, the identification method presented in this study is likely to capture anticipated fiscal policy shocks because the estimated series of anticipated shocks roughly corresponds to the date of announcement for fiscal packages. Second, we discover that consumption persistently responds positively to anticipated fiscal policy shocks. This finding is evidence that people react to fiscal news shocks immediately, and it emphasizes the importance of taking fiscal foresight into consideration. Third, the results of FEVD also show the importance of anticipated fiscal policy shocks in the movement of consumption. In addition, the result that not all variations in stock returns can be explained by that shock justifies our hybrid method that incorporates the ideas in Fisher and Peters (2010) into a sign-restricted VAR. Finally, the results of the HD featuring the period referred to as Abenomics reveal that a positive anticipated fiscal policy shock occurred in 2012Q3 and 2012Q4, and that it immediately contributed to consumption. Moreover, that shock had a positive effect on the

GDP in 2013Q2.

In conclusion, our results imply that anticipated fiscal policy shocks have expansionary effects on consumption. Also, previous works that ignored fiscal foresight may have underestimated the effects of fiscal policy. Therefore, future evaluations of macroeconomic policy should take fiscal foresight into consideration in order to obtain the true effects. Finally, a detailed investigation of the effects of Abenomics using more sample periods remains as an area of future research.

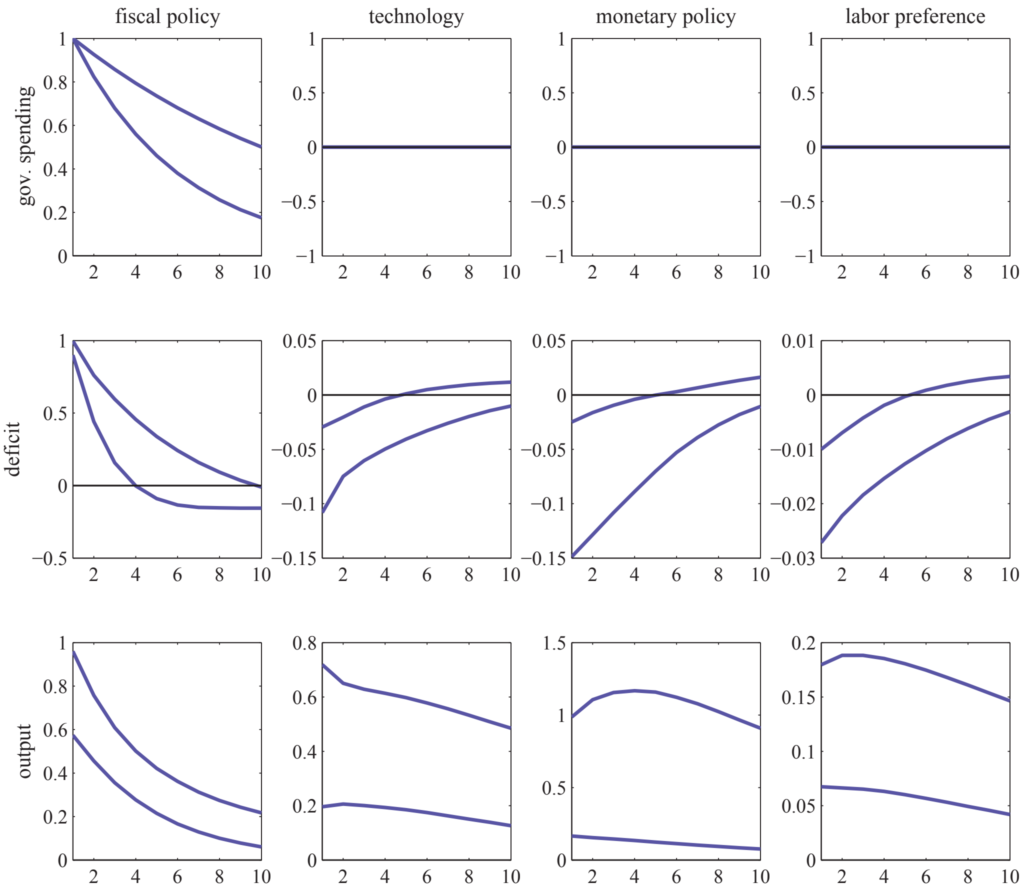

Figure 2.1: The 68% bands for the theoretical responses to unanticipated fiscal policy, technology, monetary policy, and labor preference shocks

Note: Each row and each column correspond to each variable and each shock, respectively. Variables are placed in order of government spending, deficits and GDP from the top, and the shocks are placed in order of fiscal policy, technology, and monetary policy and labor preference shocks from the left.

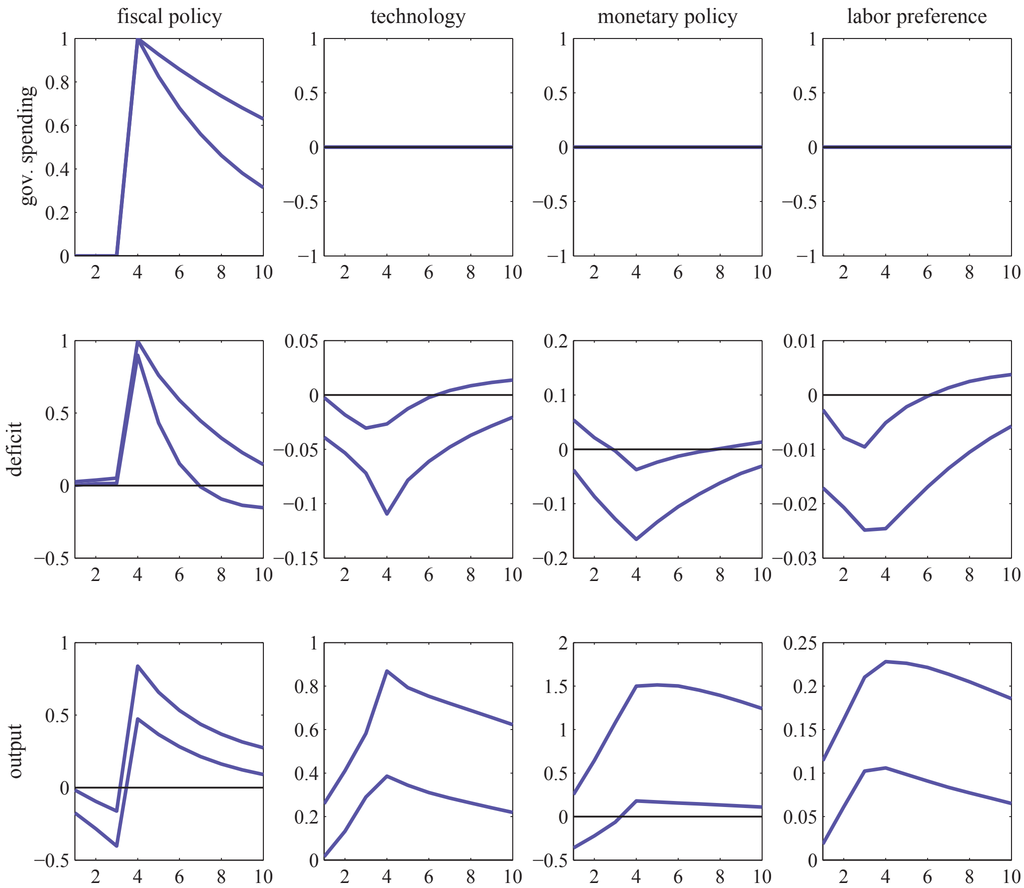

Figure 2.2: The 68% bands for the theoretical responses to anticipated fiscal policy, technology, monetary policy, and labor preference shocks

Note: Each row and each column correspond to each variable and each shock, respectively. Variables are placed in order of government spending, deficits and GDP from the top, and the shocks are placed in order of fiscal policy, technology, and monetary policy and labor preference shocks from the left.

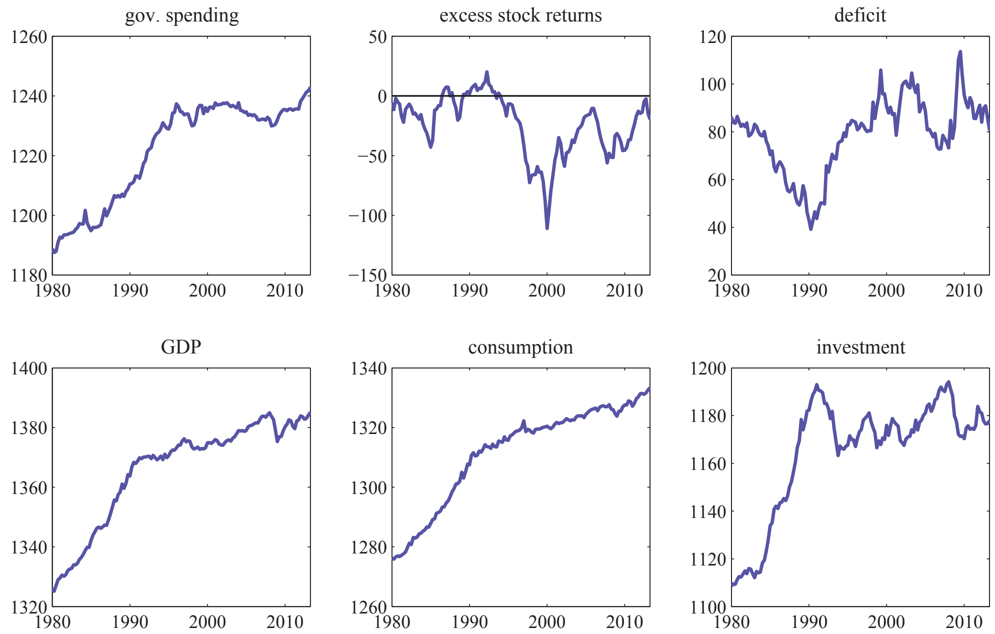

Figure 2.3: The data used in VAR model

Note: Except for excess stock returns, the series are taken to their logarithmic values and multiplied by 100 in order to interpret the results in percentages.

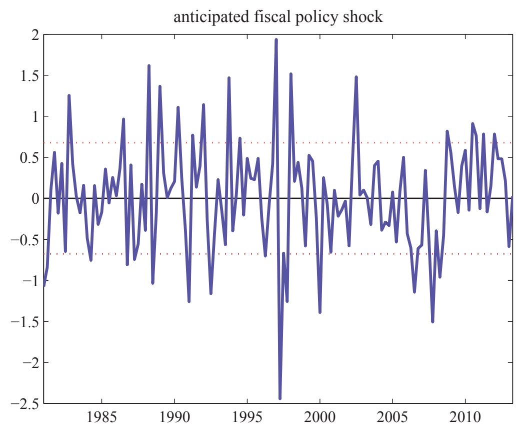

Figure 2.4: The series of structural shocks

Note: Blue line indicates the median of sampled series for anticipated fiscal policy shock, and the dotted red line indicates one standard deviation.

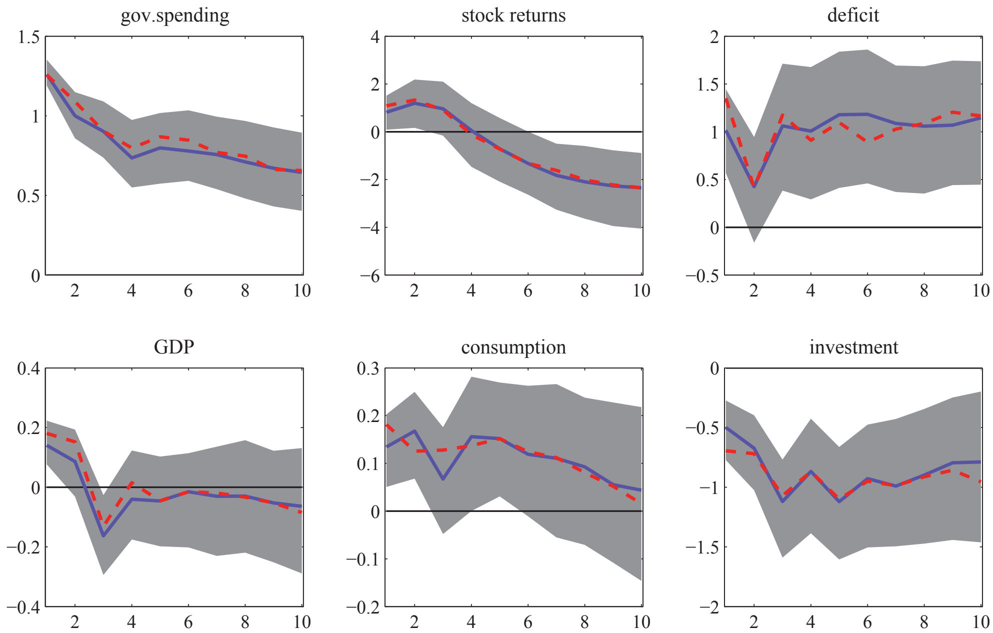

Figure 2.5: Responses to unanticipated fiscal policy shock

Note: This figure shows the estimated IRFs to unanticipated fiscal policy shock. The solid blue lines and shaded area indicate the median of sampled IRFs and the 68% credible intervals, respectively. The red dotted lines indicate the IRFs that are the closest to the median responses among those obtained in each admissible rotation.

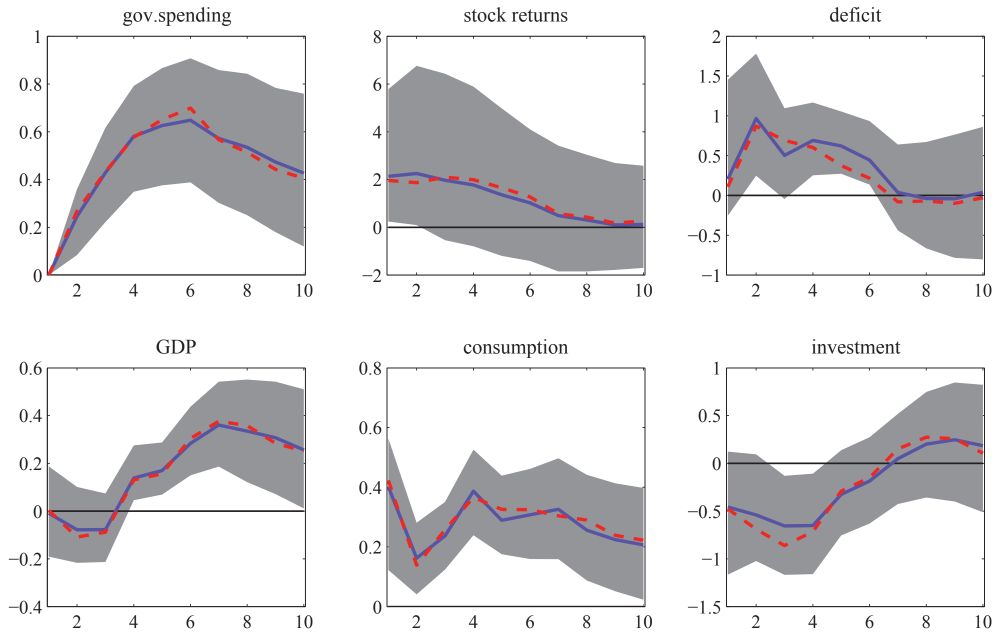

Figure 2.6: Responses to anticipated fiscal policy shock

Note: This figure shows the estimated IRFs to anticipated fiscal policy shock. The solid blue lines and shaded area indicate the median of sampled IRFs and the 68% credible intervals, respectively. The red dotted lines indicate the IRFs that are the closest to the median responses among those obtained in each admissible rotation.

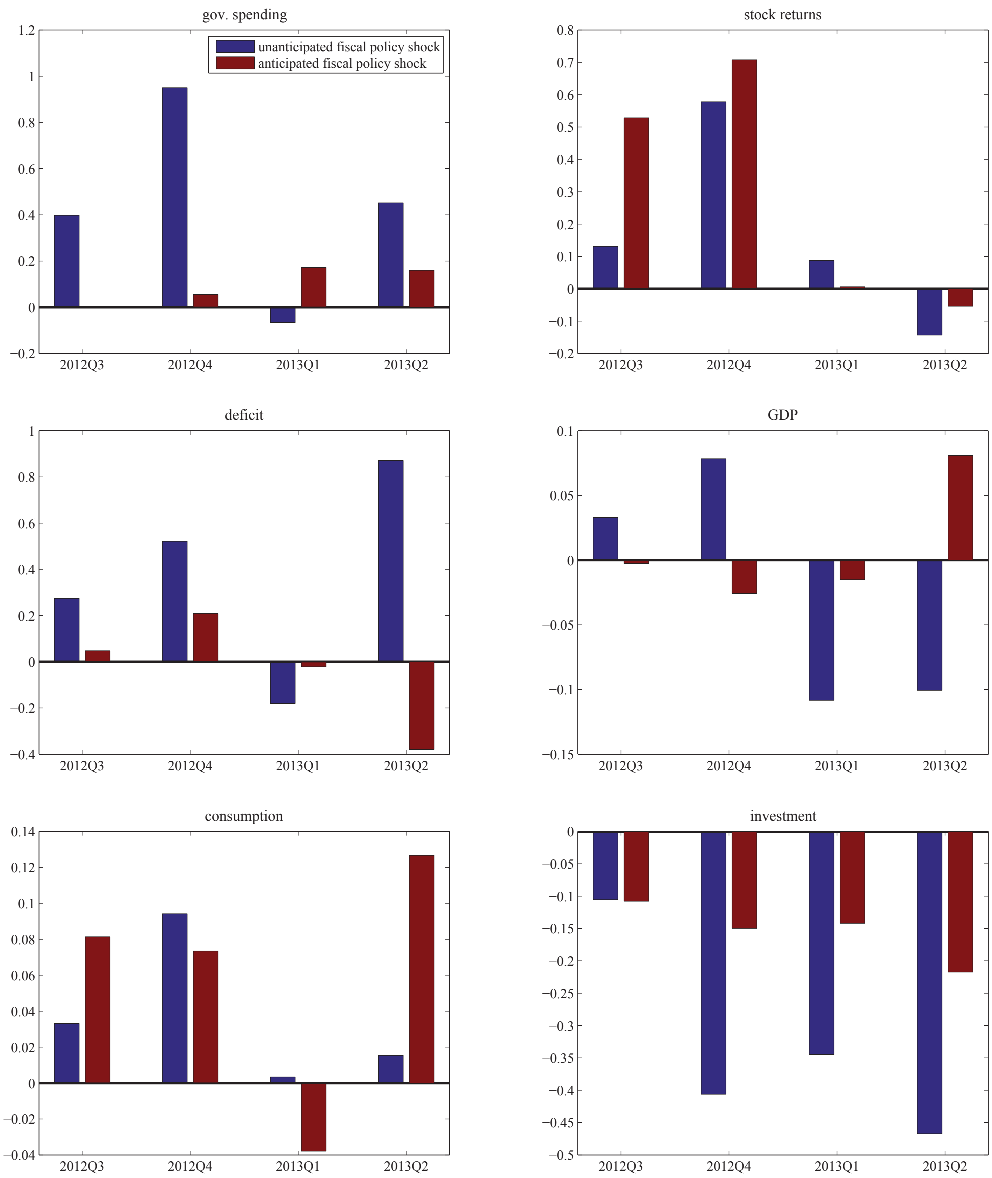

Figure 2.7: Historical decomposition conditioning on the data before 2012Q2

Note: This figure shows the results of historical decomposition conditioning on the data before 2012Q2. The blue bar and the red bar respectively present the contribution of unanticipated and anticipated fiscal policy shock to each variable.

Chapter 3

State-Dependent Effects of Fiscal Policy in Japan: Do Rule-of-thumb Households Increase the Effects of Fiscal Policy?

3.1 Introduction

Since the recent economic downturn stemming from the financial crisis in the United States and the European debt crisis, fiscal policy has been adopted as the key instrument to stimulate economic activity. Large fiscal packages have been implemented worldwide, and have led to debates on the effectiveness of fiscal policy. In this study, we empirically investigate the effectiveness of fiscal policy in reinvigorating economic activity, and specifically focus on the relationship between the effects of fiscal policy and the share of rule-of-thumb (ROT) households.[1]

[1] The ROT households are defined as households whose current consumption equals current disposable income because of borrowing constraints, discussed in Galí et al. (2007).

We examine the Japanese economy from 1972 to 2013 for a number of following reasons. One is that the fiscal authority in Japan has implemented a large number of fiscal packages–as compared with other countries–after the economic bubble burst in early the 1990s. Another is that concerns regarding the effects of fiscal policy have increased recently in response to large scale fiscal package implemented by Prime Minister Shinzo Abe. For the above reasons, we choose the Japanese economy for this study. However, it is possible to apply the analysis conducted in this study to other countries.

The analysis in this study is motivated by two strands of literature. The theoretical findings of Galí et al. (2007) is the first one. While most empirical studies represented by Blanchard and Perotti (2002) report positive effects of fiscal stimulus on consumption, the standard Real Business Cycle model (e.g., Baxter and King 1993) predicts that an increase in government spending lowers private consumption. This is because agents with rational expectation regard an increase in government spending as an increase in tax. This conflict between theoretical and empirical results is known as the government spending puzzle. Galí et al. (2007) attempt to solve this puzzle by incorporating ROT households into a New Keynesian (NK) type Dynamic Stochastic General Equilibrium (DSGE) model. Their model predicts that the effects of fiscal policy on consumption are enhanced due to the increase in shares of ROT households in the economy. In this study, we ascertain whether an empirical analysis pertaining to this real-world phenomenon in Japan supports the theoretical prediction of Galí et al. (2007); therefore, this study contributes to the policy-making significance of their model.

This study also depends on literature that investigates the state dependent effects of fiscal policy; for example, studies conducted by Auerbach and Gorodnichenko (2012), Fazzari et al. (2013) and Owyang et al. (2013). These studies use time-series data to estimate a smooth transition or threshold Vector Autoregression (VAR) model that is able to distill state-dependent effects of fiscal policy. We adopt a similar approach and methodology in this study to analyze the impacts of government spending on the Japanese economy. However, this study is different from previous studies in the following points. First, we do not focus on a change in the state of economy, but rather the change in the share of ROT households. Thus, we provide a suggestion with respect to the economic structure for the change in the effectiveness of fiscal policy.

Among existing studies, Tagkalakis (2008) is mentioned as the most similar study to our own. Tagkalakis (2008) also focuses on the relationship between the effects of fiscal policy and liquidity constraints, and reports that fiscal policy becomes more effective in a recession when liquidity constraints bind for a large fraction of households. In his work, the maximum ratio of loan to the value of house in housing mortgages (LTV ratio) is regarded as a proxy of the degree of credit constraints, and he estimates the static single equation including LTV ratio. In contrast to Tagkalakis (2008), our study directly estimates the share of ROT households on the basis of the consumption function presented by Campbell and Mankiw (1989). This share corresponds to the degree of borrowing constraints employed in Galí et al. (2007).[2] Furthermore, we will clarify the dynamic effects of fiscal policy on consumption using a VAR model.

The empirical results obtained in this study are briefly summarized as follows. The share of ROT households increases after relatively large negative economic shocks such as oil shock and the bubble burst. Moreover, as predicted in the theoretical model of Galí et al. (2007), unanticipated fiscal policy becomes more effective in the state of high ROT households.

The rest of this study is organized as follows. Section 3.2 presents the theoretical predictions concerning the effects of fiscal policy and the share of ROT households. In Section 3.3, we explain the empirical model in detail. Section 3.4 describes the data used and the specification of the empirical model for the empirical analysis. Estimation results are given in Section 3.5, where we show the estimated smoothed probability, the estimated parameter values, and the impulse

[2] Galí et al. (2007) set the share of ROT households in their calibration based on the estimate in Campbell and Mankiw (1989).

response functions (IRFs). Finally, we conclude in Section 3.6.

3.2 Theoretical prediction

We first demonstrate how the effects of fiscal policy depend on the share of ROT households, based on the theoretical model presented in Chapter 2. Here, we fix the parameter values except for the share of ROT households in order to focus on the relationship between the share of ROT households and the stimulus effects of fiscal policy in consumption. The parameter values follow those of Galí et al. (2007) and Colciago (2011) and are presented in Table 3.1. As in Chapter 2, anticipated fiscal policy shock is assumed to be realized three quarters after news is announced.

Table 3.1: Calibration parameters

Parameter	Value	Description
β	0.99	Subjective discount factor
δ	0.025	Depreciation rate
α	0.33	Share of capital
γ	1	Risk aversion
λ	0.2	Inverse labor supply elasticity
α_g	0	Productivity of public capital
κ	0.15	Investment adjustment cost
μ	[0.1 or 0.5]	Share of non-Ricardian households
ϵ_p	6	Elasticity of substitution in production
ϵ_w	6	Elasticity of substitution in labor
ρ_p	0.75	Calvo parameter on prices
ρ_w	0.55	Calvo parameter on wages
ϕ_g	0.1	Elasticity of tax to government spending
ϕ_b	0.33	Elasticity of tax to bond
ϕ_y	0	Elasticity of tax to output
ψ_π	1.5	Monetary policy response of inflation
ψ_y	0	Monetary policy response of output
ρ_k	0.9	persistency of exogenous shocks

Note: The notation correspond with the model presented in Chapter 2. The parameter values are set based on Galí et al. (2007) and Colciago (2011), and thus some parameters (e.g., α^g, ϕ_y, and ψ_y) are set to be zero unlike Chapter 2.

Figure 3.1 shows the responses of consumption to both unanticipated (left row) and anticipated (right row) fiscal policy shock. Blue and red lines indicate the responses corresponding to the share of ROT households being equal to 0.1 and 0.5, respectively. One can see that consumption indicates quite different reactions depending on the share of ROT households. In the case of low ROT households, consumption declines in response to both types of government spending shocks as a result of dominating the negative wealth effect. In contrast, aggregate consumption increases in the case of high share of ROT households. This is because, under the assumption of monopolistic labor supply, the real wage rises in response to a government spending shock; thus, ROT households increase their consumption as much as disposable income increases. From the presented exercise, we can confirm the theoretical relationship that the effects of fiscal policy on consumption enhance as the share of ROT households increases.

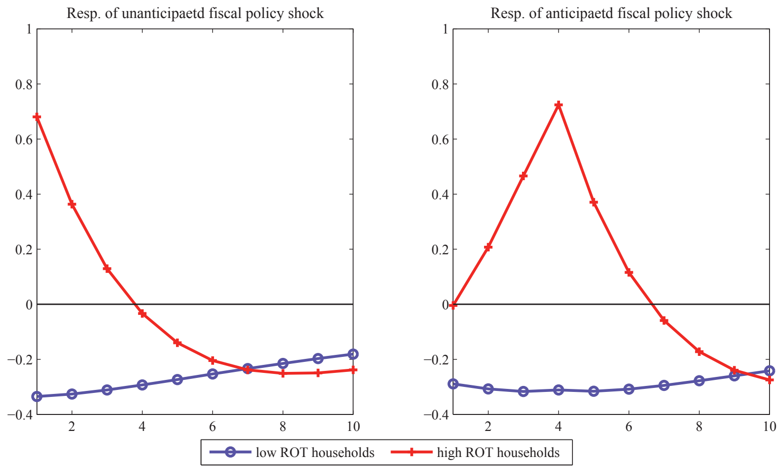

Figure 3.1: Theoretical impulse response function

Note: This figure shows the theoretical impulse responses of consumption to unanticipated (left) and anticipated (right) fiscal policy shocks. Blue and red lines correspond to the responses in case of low and high ROT household, respectively.

3.3 Estimation Methodology

3.3.1 Outline

The aim of this study is to ascertain empirically whether the theoretical prediction presented above is observed in the real economy. For this purpose, we undertake the following methodological steps. First, we estimate Campbell and Mankiw (1989)'s consumption function extended to the Markov switching model, and use the results to divide the sample period on the basis of the share of ROT households. Second, for each sample period, we estimate a VAR model and investigate the effects of fiscal policy using impulse response function (IRF) analysis. The second step leads to the identification of two types of fiscal policy shocks: unanticipated and anticipated. Unanticipated fiscal policy shocks, defined by surprise changes in government spending, have been analyzed in many previous studies. Existing studies (e.g., Ramey 2011) point to the importance of fiscal foresight by analyzing the macroeconomic effects of anticipated fiscal shocks. These two fiscal shocks are identified by the method presented in Chapter 2. We conduct a series of estimations using the Bayesian Markov Chain Monte Carlo (MCMC) method.

3.3.2 Estimation model

In order to divide our sample period into two parts depending on the ratio of ROT households, we extend the consumption function developed by Campbell and Mankiw (1989) to the Markov-switching model as follows:

$$\Delta c_t = [\mu_0 + \lambda_0 \Delta y_t^d](1 - S_t) + [\mu_1 + \lambda_1 \Delta y_t^d] S_t + u_t \tag{3.1}$$

$$u_t \sim N(0, \sigma_0^2(1 - S_t) + \sigma_1^2 S_t) \tag{3.2}$$

where Δc_t and Δy_t^d denote the growth rate of consumption for nondurables and services and disposable income, and $S_t = 0, 1$ is the latent variable indicating a regime at period t. We assume a Markov process for the dynamics of S_t, that is, the dynamics of regime change follow the transition probability below:

$$P = \begin{bmatrix} p_{00} & 1 - p_{11} \\ 1 - p_{00} & p_{11} \end{bmatrix} \tag{3.3}$$

where the (i, j) element of P is $Pr[S_t = i \mid S_{t-1} = j]$. According to Campbell and Mankiw (1989), the coefficients of Δy_t^d (i.e., λ_0 and λ_1) indicate the share of ROT households. By imposing the restriction $\lambda_1 > \lambda_0$, the regime $S_t = 1$ represents a period that has a large share of ROT households. Similar to the coefficients on disposable income, the constant term and the variance are also allowed to vary depending on a regime shift.

As noted in Campbell and Mankiw (1989), the error term u_t is not necessarily orthogonal to Δy_t^d, and thus the instrument variables (IVs) are needed to estimate (3.1). Following Campbell and Mankiw (1989), this study uses lags $2 - 4$ of disposable income and consumption growth and lag 2 of the log consumption-disposable income ratio as IVs. However, Kim (2004) points out that the estimation of the Markov-switching model using the Hamilton filter is not valid in the case of the existing problem of endogeneity, and instead presents a transformed model that can directly employ the Hamilton filter. In line with Kim (2004), we transform (3.1) as follows. First, the endogenous regressors Δy_t^d are formulated as

$$\Delta y_t^d = z_t' \delta + v_t \tag{3.4}$$

$$v_t \sim N(0, \sigma_v^2) \tag{3.5}$$

where z_t is a vector of IVs, and v_t is correlated with u_t. We also assume that the covariance matrix of $[v_t, u_t]$ is changed by a current regime, that is,

$$Cov(v_t, u_t) = C_{vu,S_t} = (1 - S_t)C_{vu,0} + S_t C_{vu,1} \tag{3.6}$$

Then, by applying the Cholesky decomposition to this covariance matrix, $[v_t, u_t]'$ can be written as a function of two independent shocks:

$$\begin{bmatrix} v_t \\ u_t \end{bmatrix} = \begin{bmatrix} b_{11,S_t} & 0 \\ b_{21,S_t} & b_{22,S_t} \end{bmatrix} \begin{bmatrix} \omega_{1t} \\ \omega_{2t} \end{bmatrix}, \quad \begin{bmatrix} \omega_{1t} \\ \omega_{2t} \end{bmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} 1 & 0 \\ 0 & 1 \end{pmatrix} \right) \tag{3.7}$$

Eq.(3.7) allows us to rewrite Eqs.(3.1) and (3.4) as

$$\Delta c_t = [\mu_0 + \lambda_0 \Delta y_t^d](1 - S_t) + [\mu_1 + \lambda_1 \Delta y_t^d]S_t + b_{21,S_t}\omega_{1t} + b_{22,S_t}\omega_{2t} \tag{3.8}$$

$$\Delta y_t^d = z_t'\delta + b_{11,S_t}\omega_{1t}. \tag{3.9}$$

Solving (3.9) to ω_{1t} and substituting into (3.8), we obtain

$$\Delta c_t = [\mu_0 + \gamma_0 \hat{v}_t + \lambda_0 \Delta y_t^d](1 - S_t) + [\mu_1 + \gamma_1 \hat{v}_t + \lambda_1 \Delta y_t^d]S_t + e_t \tag{3.10}$$

$$e_t \sim N(0, \sigma_{e0}^2(1 - S_t) + \sigma_{e1}^2 S_t) \tag{3.11}$$

where $\gamma_i, i = 0, 1$, v_t and e_t correspond to $b_{21,S_t}/b_{11,S_t}, S_t = i$, $\Delta y_t^d - z_t'\delta$ and $b_{21,S_t}\omega_{2t}$, respectively, and $\hat{\delta}$ is OLS estimator of δ. We apply the Hamilton filter to this transformed model and estimate the series of the latent regime S_t.

Given the sequence of S_t, we estimate the VAR model as follows:

$$Y_t = c(t) + B_0(L)(1 - S_t)Y_{t-1} + B_1(L)S_tY_{t-1} + \nu_t, \tag{3.12}$$

$$\nu_t \sim N(0, \Sigma_0(1 - S_t) + \Sigma_1 S_t), \tag{3.13}$$

where $B_i(L), i = 0, 1$ is a polynomial in the lag operator, $c(t)$ is defined to include all deterministic components of the data, and $\Sigma_i, i = 0, 1$ denotes the variance-covariance matrix of the reduced form residuals ν_t. This model allows the propagation of structural shock to change in both dynamic and contemporaneous ways. Y_t is a vector of endogenous variables that consists of government spending, excess stock returns of the construction industry, deficit, output, and private consumption. Under this VAR system, we identify both unanticipated and anticipated fiscal policy shock by using sign restrictions. As discussed below, the first four variables contained in the VAR model are necessary to achieve our identification.

3.3.3 Identification

As in Chapter 2, we identify two types of fiscal policy shocks: unanticipated and anticipated. First, unanticipated fiscal policy shock is identified under the standard assumption that government spending is the most exogenous variable and thus is affected only by this shock at the impact period. This assumption is based on the fact that fiscal policy cannot react to changes in the state of the economy immediately due to the decision lag. This fact has also been accounted for by most studies in the reviewed literature (e.g. Blanchard and Perotti 2002, Galí et al. 2007). On the other hand, anticipated fiscal policy shock is identified based on the robust sign restrictions presented in Chapter 2. In other words, we combine the approach presented by Fisher and Peters (2010) and the model-based sign restrictions.

In summation, we identify both unanticipated and anticipated fiscal policy shock by imposing sign restrictions summarized in Table 4.1.

Table 3.2: Sign restrictions

	gov. spending	stock returns	deficit	GDP
unanticipated fiscal shock	> 0 for 1Q		> 0 for 1Q	> 0 for 1Q
anticipated fiscal shock	$= 0$ for 1Q > 0 for 4-6Q	> 0 for 1Q	> 0 for 4-6Q	> 0 for 4-6Q

Note: This table shows sign restrictions imposing on VAR model. Blank spaces denote that no restrictions are imposed on this variable at this period.

3.3.4 Bayesian estimation

We adopt the Bayesian MCMC method to estimate the Markov-switching consumption function and the VAR model discussed above. In the Bayesian estimation, the joint posterior distribution of parameters is calculated from the prior distribution and the observed data, and statistical inferences are them performed based on its distribution. However, our estimation model includes regime switching and two equations (i.e., the consumption function and VAR model) are estimated simultaneously. This makes it difficult to calculate the joint posterior distribution analytically. In such cases, the samplings of unknown parameters are performed by the MCMC method. Even if the joint posterior distribution is too complicated to calculate analytically, the MCMC method enables us to sample the parameters because the parameters are drawn from the full conditional posterior distribution that is based on known information and all the other available parameters.

We describe our estimation procedure via the MCMC method as follows. First, we define $\boldsymbol{\beta} = [\mu_0, \mu_1, \gamma_0, \gamma_1]$, $\boldsymbol{p} = [p_{00}, p_{11}]$, $\boldsymbol{x} = \left[\{\Delta c_t\}_{t=1}^{T}, \{\Delta y_t^d\}_{t=1}^{T}, \{\boldsymbol{Y_t}\}_{t=1}^{T}\right]$, where T is the number of observations. Second we set the prior distributions for each parameter, and choose the normal distribution for $\boldsymbol{\beta}$ and the inverse gamma distribution for $\sigma_{ei}^2, i = 0, 1$. Furthermore, in order to satisfy the restrictions $0 < \lambda_0 < \lambda_1 < 1$, the prior distribution for λ_1 is set to be the truncated normal distribution whose density is zero unless $0 < \lambda_1 < 1$, the prior for λ_0 is set to be the truncated normal distribution whose density is zero unless $0 < \lambda_0 < \lambda_1$. For the VAR model, we set the independent Normal-Wishart distributions for the VAR coefficient and variance-covariance matrix. The priors of transition probabilities are the beta distributions.[3] Given this information, we can generate random samples from $\pi(\boldsymbol{\beta}, \lambda_0, \lambda_1, \sigma_{ei}^2, \{S_t\}_{t=1}^T, \boldsymbol{p}, B_i(L), \Sigma_i \mid \boldsymbol{x})$ using the MCMC method as follows;

1. Set initial values in $\boldsymbol{\beta}^{(0)}$, $\lambda_0^{(0)}$, $\lambda_1^{(0)}$, $\sigma_{ei}^{2\,(0)}$, $\{S_t\}_{t=1}^{T\ (0)}$, $\boldsymbol{p}$, $B_i(L)^{(0)}$, $\Sigma_i{}^{(0)}$ and $j = 0$.

2. Generate $\{S_t\}_{t=1}^{T\ (j+1)}$ from $\pi(\{S_t\}_{t=1}^T \mid \boldsymbol{\beta}^{(j)}, \lambda_0^{(j)}, \lambda_1^{(j)}, \sigma_{ei}^{2\,(j)}, \boldsymbol{p}^{(j)}, \boldsymbol{x})$.

3. Generate $\boldsymbol{p}^{(j+1)}$ from $\pi(\boldsymbol{p} \mid \boldsymbol{\beta}^{(j)}, \lambda_0^{(j)}, \lambda_1^{(j)}, \sigma_{ei}^{2\,(j)}, \{S_t\}_{t=1}^{T\ (j+1)}, \boldsymbol{x})$.

4. Generate $\boldsymbol{\beta}^{(j+1)}$ from $\pi(\boldsymbol{\beta} \mid \lambda_0^{(j)}, \lambda_1^{(j)}, \sigma_{ei}^{2\,(j)}, \{S_t\}_{t=1}^{T\ (j+1)}, \boldsymbol{p}^{(j+1)}, \boldsymbol{x})$.

5. Generate $\lambda_1^{(j+1)}$ from $\pi(\lambda_1 \mid \boldsymbol{\beta}^{(j+1)}, \lambda_0^{(j)}, \sigma_{ei}^{2\,(j)}, \{S_t\}_{t=1}^{T\ (j+1)}, \boldsymbol{p}^{(j+1)}, \boldsymbol{x})$.

6. Generate $\lambda_0^{(j+1)}$ from $\pi(\lambda_0 \mid \boldsymbol{\beta}^{(j+1)}, \lambda_1^{(j+1)}, \sigma_{ei}^{2\,(j)}, \{S_t\}_{t=1}^{T\ (j+1)}, \boldsymbol{p}^{(j+1)}, \boldsymbol{x})$.

7. Generate $\sigma_{ei}^{2\,(j+1)}$ from $\pi(\sigma_{ei}^2 \mid \boldsymbol{\beta}^{(j+1)}, \lambda_0^{(j+1)}, \lambda_1^{(j+1)}, \{S_t\}_{t=1}^{T\ (j+1)}, \boldsymbol{p}^{(j+1)}, \boldsymbol{x})$ for $i = 0, 1$.

8. Divide $\{\boldsymbol{Y_t}\}_{t=1}^T$ into two parts depending on $\{S_t\}_{t=1}^{T\ (j+1)}$.

9. Generate $\boldsymbol{B_i(L)}^{(j+1)}$ form $\pi(\boldsymbol{B_i(L)} \mid \boldsymbol{\Sigma_i}^{(j)}, \{\boldsymbol{Y_{it}}\}_{t=1}^T)$ for $i = 0, 1$.

10. Generate $\boldsymbol{\Sigma_i}^{(j+1)}$ form $\pi(\boldsymbol{\Sigma_i} \mid \boldsymbol{B_i(L)}^{(j+1)}, \{\boldsymbol{Y_{it}}\}_{t=1}^T)$ for $i = 0, 1$.

11. Identify the structural shocks based on $\boldsymbol{B_i(L)}^{(j+1)}$ and $\boldsymbol{\Sigma_i}^{(j+1)}$ by using sign restrictions.

[3] More precisely, the priors are the following:

$$\boldsymbol{\beta} \sim N(0, \boldsymbol{I}), \sigma_{ei}^2 \sim IG(5/2, 1/2), p_{i,j} \sim beta(19, 1)$$
$$\lambda_1 \sim N(0.5, 1)I(0 \leq \lambda_1 \leq 1), \lambda_0 \sim N(0.5, 1)I(0 \leq \lambda_0 \leq \lambda_1)$$
$$\boldsymbol{B_i(L)} \sim N(0, \boldsymbol{I} \times 0.01), \boldsymbol{\Sigma_i} \sim IW(100/2, 0.1/2).$$

where $I(\cdot)$ is the indicator function that is equal to one if the condition in the parameters is satisfied, and zero otherwise.

12. Return to step 2 until the number of iteration reaches N times.

In this study, N is set to 25000 and the first $N_0 = 5000$ samples are discarded as burn-in.

To obtain reliable results, we impose some restrictions on the above sampling process. With respect to drawing the latent variable $\{S_t\}_{t=1}^{T}$ in step 2, we impose the restriction that the regime continues for at least four periods after a regime change occurs. Furthermore, in the identification step, we repeatedly generate the orthogonal matrix Q until the calculated IRFs satisfy the sign restrictions or the number of iterations reaches 100. If the iterations reach 100 and the valid draw cannot be obtained, we return to step 2.[4]

3.4 Data and specification

We estimate the presented model by using quarterly data from Japan for the period 1972Q2 - 2013Q1. Because the stock price of the construction industry has only been released since 1972, the sample period starts in 1972Q2. For the consumption function, we obtain data for the consumption of nondurables and services, and disposable income from SNA database. Disposable income in this study is *National disposable income (Households (including private unincorporated enterprises))*. While both real and nominal consumption data regarding nondurables and services are available in the SNA database, only a nominal data series has been published on disposable income. Therefore, we calculate the consumption deflator by using nominal and real consumption data, and then construct the real series of disposable income by dividing nominal values by this deflator. Furthermore, we adjust both data series for seasonality by using X-12-ARIMA. Each data series in the VAR model is obtained from the same sources used in Chapter 2.

For the data from the SNA database, we combine the series from the 68SNA and the 93SNA

[4] The form of the Q matrix is assumed to be the same as in Chapter 2, and the Q matrix is generated by the same method in Chapter 2.

in the first quarter of the year 1980 using the growth rate of 68SNA because the data of 93SNA exists only after 1980. Moreover, all data in both the consumption function and the VAR model are converted to their per capita values using total population data from the Japanese Population Census (Ministry of Internal Affairs and Communication). With the exception of stock returns, all data are converted to their logarithmic values and multiplied by 100 in order to interpret the results in terms of percentages.

For the specification of the VAR model, we perform the estimation in levels, and include the constant term as well as linear and quadratic time trends into the estimation model. As is common when using quarterly data, we set four lags in the VAR model.

3.5 Empirical Results

3.5.1 Estimated parameter values

In Figure 4.2, we show the sample autocorrelation function, the sample paths, and the posterior densities for the parameters in the consumption function. The sample paths for each of the parameters look stable, and the sample autocorrelation function damps stably. Furthermore, Table 3.3 shows the estimates for posterior means, standard deviations, the 95 percent credible intervals, and the p-value of the convergence diagnostics (CD) of Geweke (1992) for the parameters in the consumption function. Based on the CD statistics, the null hypothesis of the convergence to the posterior distribution is not rejected at the 5 percent significance level. These results suggest that the samples in our estimation are efficiently generated and adequately converge.

In the following, we define the high ROT households period and the low ROT households period as high state and low state, respectively. With respect to the estimated parameter values, the result first indicates that the share of ROT households is 0.2 in the low state and 0.35 in

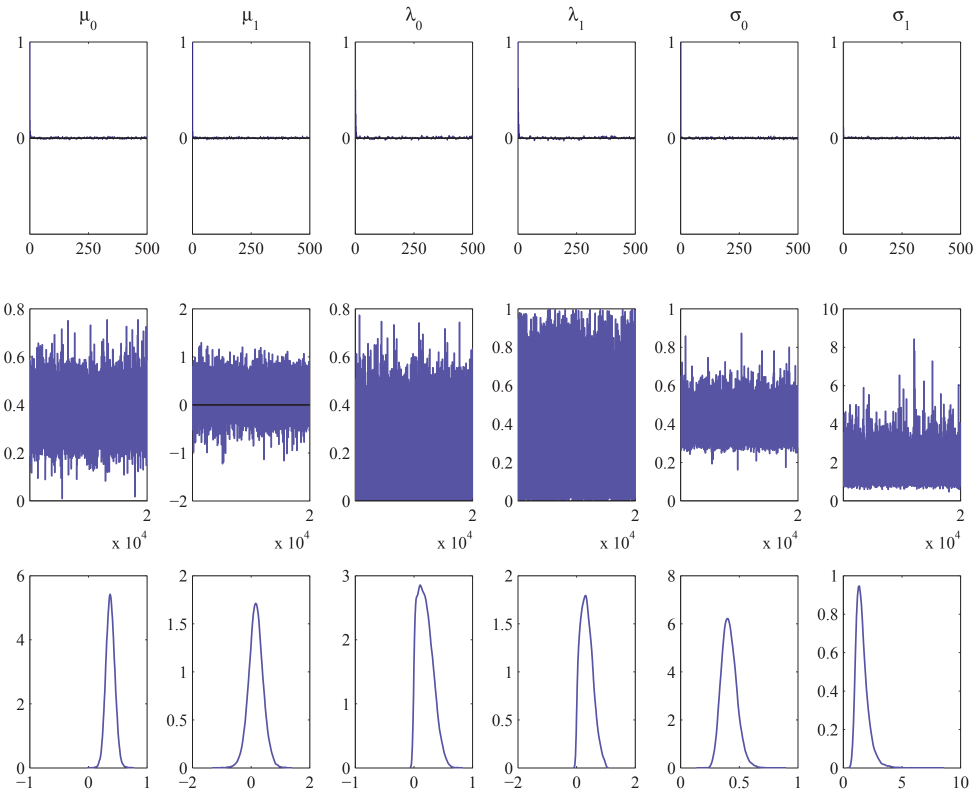

Figure 3.2: Estimation results of consumption function

Note: This figure shows sample autocorrelations (top), sample path (middle), and posterior densities (bottom) of estimation parameters in consumption function.

the high state. Moreover, the estimation results reveal the following. The constant term in the high state is smaller than in the low state. Since the constant term denotes the average growth rate in consumption, this result means that the growth rate of consumption has slowed down in the high ROT state. In terms of variance, our result shows that the high ROT state is very volatile as compared with the low state.

3.5.2 Smoothed probability

Figure 3.3 displays the smoothed probability of the high state (i.e., $S_t = 1$). The shaded areas in this figure indicate the periods where its probability exceeds 0.5. On examining this

Table 3.3: Estimation results of consumption function

parameter	mean	Stdev.	95 percent intervals	CD
μ_0	0.37	0.09	[0.52, 0.23]	0.23
μ_1	0.16	0.28	[0.67, -0.34]	0.17
λ_0	0.19	0.23	[0.48, 0.01]	0.52
λ_1	0.35	0.35	[0.81, 0.02]	0.83
σ^2_{e0}	0.41	0.06	[0.56, 0.30]	0.69
σ^2_{e1}	1.61	0.64	[2.97, 0.89]	0.45

Note: This table shows the estimated parameter value in consumption function. CD denotes the p-value of the convergence diagnostics of Geweke (1992). Bandwidth of a Parzen window is set to be 500.

figure, we find that the Japanese economy has fallen into the high ROT household period in the second half of the 1970s and also in the 1990s.[5] These periods correspond to the recessions that stemmed from oil shock and the collapse of bubble economy, respectively. Moreover, the probability also rises in the recession period after the Lehman shock although it does not exceed 50%. These findings suggest the possibility that a negative economic shock caused the increase in the share of ROT households. This is consistent with what Tagkalakis (2008) states. In fact, the estimated results presented in Table 3.3 also show that the high ROT state is a period of high consumption volatilities with the average growth rate being low (see μ_1 and σ^2_{e1}). However, liquidity constraints seem to bind only after large shocks because not all recession periods dated by Cabinet Office of Japan is classified as the high state.

3.5.3 Impulse response functions

Figure 3.4 shows the IRFs of government spending and private consumption to both unanticipated (the first column) and anticipated (the second column) fiscal policy shocks. The blue and red lines respectively indicate the median of sampled IRFs in the low and high states, and

[5]The result that the share of ROT households increase in the second half of the 1970s but decrease in the 1980s is consistent with the fact presented in Ogawa et al. (1986) and Hatano (2004). However, the increase of the share of ROT households in the 1990s is not reported in Hatano (2004). This difference between Hatano (2004) and our result might be generated from the methodological difference that Hatano (2004) does not take the time variations of volatility into account. With respect to this point, we consider that further investigation is necessary.

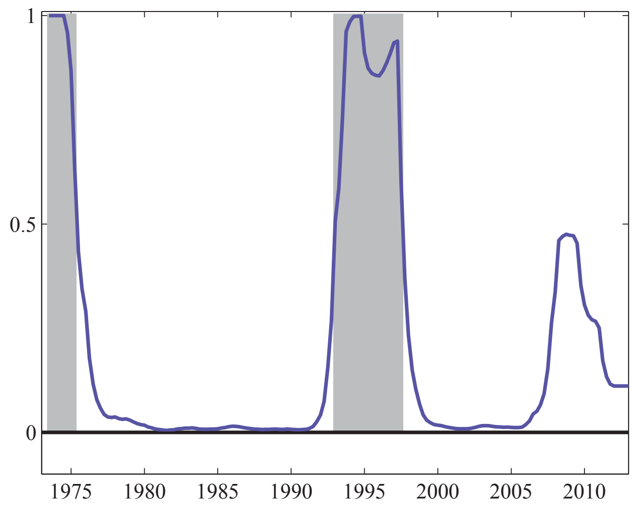

Figure 3.3: Smoothed probability of high ROT households period

Note: This figure shows the smoothed probability of high ROT households period. The periods when its probability exceeds 0.5 are colored by the shaded area.

the shaded areas show 68 percent credible intervals that correspond to the high state. While the IRFs to unanticipated fiscal policy shock are normalized so that a response of government spending at period 1 becomes JPY10000, the IRFs to anticipated fiscal policy shock are normalized so that an accumulated response of government spending from period 1 to period 4 becomes JPY10000.

For both types of fiscal policy shocks, the responses of government spending show almost the same path in each state. Therefore, the differences in response of consumption, shown below, stem from differences in the share of ROT households and not in the dynamics of government spending.

In response to an unanticipated fiscal policy shock, the IRF of private consumption shows a positive sign in both states, but its response is much larger in the case of the high state. A median response in the low state in the impact period is outside the 68 percent credible intervals of the high state, and its response is 0.42 against 1.10 in the high state. Unlike

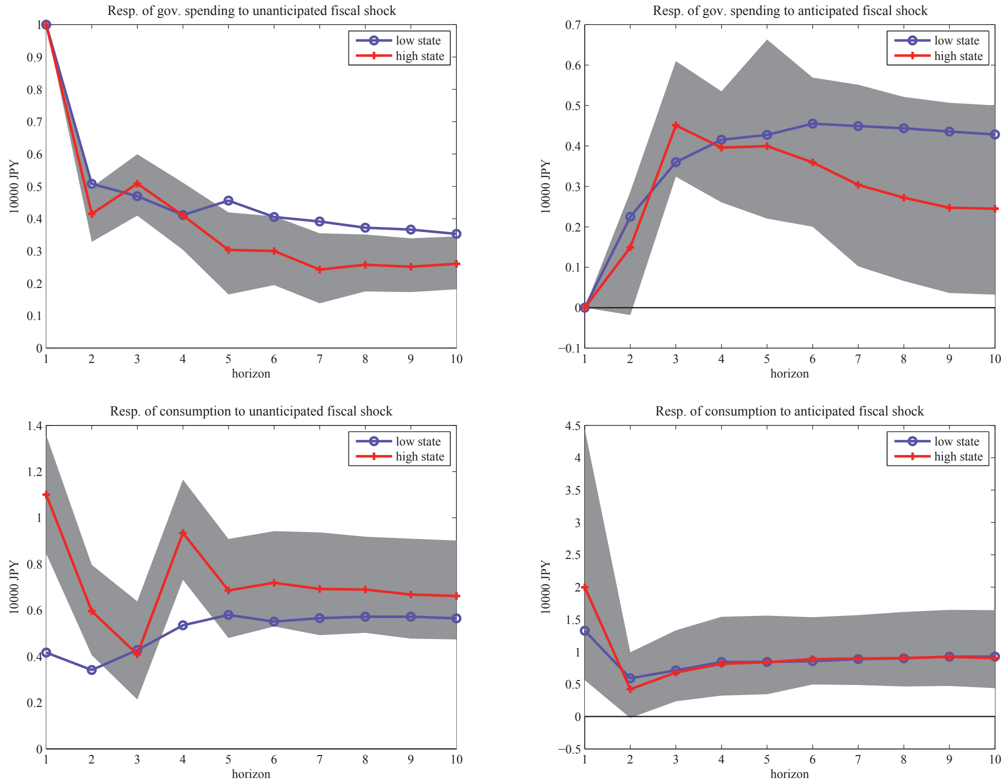

Figure 3.4: Impulse responses

Note: This figure show the IRFs of government spending and private consumption to unanticipated (left) and anticipated (right) fiscal policy shocks. In this figure, the blue and red lines indicate the responses in low and high ROT household period, respectively. And, the shaded area denotes the 68% credible intervals corresponding to the high ROT households period.

the case of an unanticipated fiscal policy shock, there is almost no difference in the response of private consumption although consumption positively responds to anticipated fiscal policy shock. A median response in the low state is contained within the credible intervals of the high state, and the responses seem to be almost the same in both states.

Following Nakajima et al. (2011), we also compute the ratio of the MCMC draws where the IRF of the high state exceeds that of the low state to evaluate the statistical difference in the IRF. Table 3.4 reports the difference in the IRFs of private consumption to both types of fiscal policy shocks. The posterior probability that the response to unanticipated fiscal policy shock in the high state is larger than that in low state is 70% for the impact period, 66.2% for

Table 3.4: Ratio of sample that the IRF of high state exceeds that of low state

horizon	unanticipated fiscal policy shock	anticipated fiscal policy shock
1	70.0	61.1
2	66.2	43.9
3	49.1	49.6
4	73.2	50.0
5	57.0	50.6
6	62.4	53.3
7	59.6	51.9
8	58.9	51.6
9	57.5	51.4
10	57.6	50.2

Note: This table shows posterior probability for the difference in IRFs between each state.

the two-quarters ahead period, 49.1% for the three-quarters ahead period, and 73.2% for the four-quarters ahead period. Except for the three-quarters ahead period, it could be said that the IRF in the high state exceeds that in the low state for one year after the shock. On the other hand, the probability for anticipated fiscal policy shock is 61.1% for the impact period, 43.9% for the two-quarters ahead period, 49.6% for the three-quarters ahead period, and 50.0% for four-quarters ahead period. Therefore, we consider the evidence that the response of consumption to anticipated fiscal policy shock is difference between each state is weak.

From the above exercises, we find that both fiscal policy shocks have positive effects on private consumption. Furthermore, we conclude that the effects of unanticipated fiscal policy are enhanced in the high share of ROT household period as predicted by Galí et al. (2007).

3.6 Conclusion

In this study, we have analyzed the relationship between the effects of fiscal policy and the share of ROT households. In particular, we focus on the empirical effects of fiscal policy on private consumption depending on the share of ROT households. In order to accomplish our aim, we estimate Campbell and Mankiw (1989)'s consumption function extended to Markov switching and divide the sample period into two parts depending on the share of ROT households. Then, we estimate a VAR model for divided sample periods and investigate the effects of fiscal policy on consumption using IRF analysis. In our VAR model, we identify two types of fiscal policy shocks, unanticipated and anticipated.

The main results are summarized as follows. First, the smoothed probability reveals that the Japanese economy experienced twice the period when the liquidity constraints bind: the recession periods after the oil shock and the collapse of the bubble economy. Our results also suggest the possibility that the share of ROT households has increased after the Lehman shock. Second, the estimated parameter values indicate that the average growth rate of consumption is low and the variance is high in the state of high ROT households. Accounting for those findings, we can consider that the high ROT households' periods correspond to the periods of recession in Japan. However, it seems that only large shock, such as oil shock and the bubble burst, becomes a trigger that causes the regime shift from a low state to a high state. Finally, the IRFs obtained in the VAR model show that private consumption in a high ROT state responds much more to unanticipated fiscal policy shock as compared to that in a low ROT state. However, such differences in IRFs between the two states is not observed in the case of anticipated fiscal policy shock.

In conclusion, our results imply that the share of ROT households rises during recession, and subsequently (unanticipated) fiscal policy stimulates private consumption more effectively. This finding is similar to the results of Fazzari et al. (2013), Auerbach and Gorodnichenko (2012), and Owyang et al. (2013) that report that the fiscal multiplier in the US increases in recession. Moreover, the theoretical prediction of Galí et al. (2007) with regard to the consumption impacts of fiscal policy is empirically supported by our results, at least, regarding unanticipated fiscal policy shock. Therefore, we conclude that incorporating ROT households into DSGE models to replicate the positive sign of consumption to fiscal policy shock is not incorrect.

Chapter 4

Time-Varying Effects of Fiscal and Monetary Policy in Japan: New Identification for Monetary Policy at the Zero Lower Bound

4.1 Introduction

After recent financial crises (e.g., the Lehman shock and the Europe debt crisis) the authorities of major industrialized economies have been simultaneously implementing aggressive fiscal policies and monetary easing. In particular, the central banks of those economies have conducted quantitative easing (QE) at the zero lower bound (ZLB) of the nominal interest rate as a part of the unconventional monetary policy. There is continued active discussion and a growing literature on the effectiveness of those economic policies. The aim of this study is to evaluate the effects of these fiscal and monetary policies in Japan, by using the time-varying parameters vector autoregression (TVP-VAR) model developed by Primiceri (2005). As pointed out in several previous studies, Japan is "a front-runner of unconventional monetary policy (Kimura and Nakajima 2013)" and has experienced a zero interest rate for a sufficiently long time. In addition, a large number of fiscal stimulus packages have been implemented in Japan since the collapse of the bubble economy in the early 1990s. Therefore, the Japanese economy is an appropriate and interesting subject for our research. Additionally, the TVP-VAR model enables us to estimate the time-varying effects of these economic policies.

This study contributes to the existed literature as follows. First, we present a new method for identifying monetary policy at the ZLB of the nominal interest rate. There are several VAR analyses about Japanese monetary policy after QE.[1] For example, Honda et al. (2007) identifies monetary policy shock by regarding the current account balances as the monetary policy instrument, and reports that QE increases output through the stock price channel. However, their sample period only covered the years 2001 to 2006 when the Bank of Japan conducted QE policy. On the other hand, Fujiwara (2006) and Inoue and Okimoto (2008) estimate a Markov-Switching VAR model using both conventional and unconventional monetary policy periods. They conclude that the regime change occurred in the late 1990s, and that the effectiveness of monetary policy seems to decrease after the structural change. More recently, Hayashi and Koeda (2013) estimates the two-regime structural VAR model and incorporate the exit condition from zero interest rate policy, in which monetary policy regime changes when a certain condition about inflation rate is satisfied. They also found the expansionary effects of monetary policy on inflation and output. Similar to the studies using the TVP-VAR model, Nakajima (2011), Franta (2011), Nakajima et al. (2011) and Kimura and Nakajima (2013) also estimate the time-varying effect of monetary policy in Japan.

Although each previous study gives a sufficient attention to the identification of monetary

[1] Shioji (2000) and Miyao (2002) analyze Japanese monetary policy before QE.

policy, the characterization of the monetary policy under the ZLB of nominal interest rate can be improved. With exception to Franta (2011), monetary policy shock is commonly identified as a shock that lowers the interest rate (e.g., Fujiwara 2006, Inoue and Okimoto 2008, Nakajima 2013). Also, oftentimes interest rate shock and monetary base shock are identified separately (e.g., Hayashi and Koeda 2013, Kimura and Nakajima 2013). It seems to be inappropriate to identify monetary policy shock in this manner because there is no room to lower the interest rate in the ZLB period. In addition, a cut in the policy rate is usually performed through the supply of monetary base to the market, and thus monetary policy shock should be characterized by both monetary variables. The next point is related to the dynamics of the short-term interest rate in the ZLB. As noted in Nakajima (2011), the effects of structural shock are unlikely to work through the interest rate channel at the ZLB period because the policy rate falls to zero. Nevertheless, previous studies, with exception to Nakajima (2011) and Kimura and Nakajima (2013), allow the short-term interest rate to vary in response to structural shocks even in the ZLB. To resolve these difficulties, we incorporate zero restrictions into the short-term interest rate equation based on that of Nakajima (2011), and identify monetary policy shock as a combination of the interest rate and the monetary base, following Franta (2011).

More precisely, by using sign restrictions, monetary policy shock is characterized as a shock that lowers the short-term interest rate and raises the monetary base, and a nonnegativity constraint is imposed on the short term interest rate. Intuitively, the nonnegativity constraint eliminates the possibility that the interest rate falls more than the observed rates in response to monetary policy shock. On the other hand, zero restrictions on the coefficient and contemporaneous relations in the interest rate equation allow the interest rate to be fixed during the ZLB period. By combining the above two methodologies, we are able to identify the effects of monetary policy under the ZLB of nominal interest rate.

In addition to presenting the new identification for monetary policy, this study also contributes to the literature by analyzing the effects of fiscal policy. Specifically, we focus on the effects of fiscal policy at the ZLB. This is motivated by the theoretical prediction presented by Braun and Waki (2006), Christiano et al. (2011), and Eggertsson (2010), in which it is theoretically shown that the effects of fiscal policy are enhanced at the ZLB. We investigate whether this theoretical prediction can be observed empirically. Although Gerba and Hauzenberger (2013) tried to estimate the fiscal and monetary interactions in the US by using a TVP-VAR model, they did not specifically identify monetary policy under the ZLB. To my knowledge, there are no studies regarding the Japanese economy that analyze the time-varying effects of fiscal policy using the TVP-VAR model. As in Chapter 2 and Chapter 3, this chapter also identifies both unanticipated and anticipated fiscal policy shocks based on the robust sign restrictions.

The results obtained in this study are summarized as follows. First, in the ZLB period, the volatility of the short-term interest rate is estimated to be fairly small and the volatility of the monetary base is quite large. Second, the results of an impulse response function (IRF) analysis show that both monetary and fiscal policy persistently has positive effects on output, and these effects change during the sample period. Third, we were unable to confirm that the effects of fiscal policies were enhanced during the ZLB period, as was theoretically predicted by Christiano et al. (2011) and Eggertsson (2010).

The rest of this study is organized as follows. In Section 4.2, we explain the empirical framework adopted in this study. Specifically, we describe the TVP-VAR model with sign restrictions, the identification strategy, and the Bayesian technique used in this study. Section 4.3 presents the data and the specifications employed in this study. Section 4.4 shows the estimated results of the volatility, the IRFs, and the fiscal multiplier. Section 4.5 presents our conclusions.

4.2 Empirical methodology

4.2.1 TVP-VAR model

This study employs the TVP-VAR model developed by Primiceri (2005). For a k-dimensional vector of endogenous variables, $y_t = (y_{1t}, \cdots, y_{kt})'$, the structural-form VAR model with time-varying parameters is formulated as

$$A_t y_t = B_{1t} y_{t-1} + \cdots + B_{st} y_{t-s} + u_t, t = s+1, \cdots, T, \tag{4.1}$$

$$u_t \sim N(0, \Sigma_t \Sigma_t'), \tag{4.2}$$

where the $B_{it} (i = 1, \cdots, s)$ are $k \times k$ matrices of time-varying coefficients and A_t is a $k \times k$ matrix of time-varying coefficients specifying the contemporaneous relations among the endogenous variables. Moreover, A_t is assumed to be a lower-triangular matrix, denoted by

$$A_t = \begin{pmatrix} 1 & 0 & \cdots & 0 \\ a_{21t} & 1 & \cdots & \vdots \\ \cdots & \ddots & \ddots & 0 \\ a_{k1t} & \cdots & a_{k,k-1,t} & 1 \end{pmatrix}. \tag{4.3}$$

The disturbance u_t denotes the vector of structural shocks, and is distributed according to a k-dimensional normal distribution with mean 0 and time-varying covariance matrix $\Sigma_t \Sigma_t'$, where

$$\Sigma_t = \begin{pmatrix} \sigma_{1t} & 0 & \cdots & 0 \\ 0 & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma_{kt} \end{pmatrix}. \tag{4.4}$$

Multiplying both sides of (4.1) by A_t^{-1} yields the reduced-form VAR model:

$$y_t = C_{1t} y_{t-1} + \cdots + C_{st} y_{t-s} + A_t^{-1} \Sigma_t \varepsilon_t, \tag{4.5}$$

$$\varepsilon_t \sim N(0, I_k), \tag{4.6}$$

where $C_{it} = A_t^{-1} B_{it} (i = 1, \cdots, s)$ and ε_t is a vector of structural shocks that are normalized to have variance 1. Defining $\beta_t = [vec(C_{1t})', \cdots, vec(C_{st})']' (k^2 s \times 1)$ and $X_t = I_k \otimes (y_{t-1}', \cdots, y_{t-s}')$, where $\otimes$ denotes the Kronecker product, (4.5) can be rewritten as

$$y_t = X_t \beta_t + A_t^{-1} \Sigma_t \varepsilon_t. \tag{4.7}$$

Subsequently, we set the process for the time-varying parameters. Let a_t be a stacked vector of the lower triangular elements in A_t and $h_t = (h_{1t}, \cdots, h_{kt})'$ with $h_{it} = \ln(\sigma_{it}^2)$. As suggested by Primiceri (2005), we assume that the time-varying parameters evolve according to a random walk process:

$$\beta_{t+1} = \beta_t + u_{\beta t}, \tag{4.8}$$

$$\alpha_{t+1} = \alpha_t + u_{\alpha t}, \tag{4.9}$$

$$h_{t+1} = h_t + u_{ht}, \tag{4.10}$$

for $t = s+1, \cdots, T$. Moreover, the model innovations are assumed to follow a joint normal distribution:

$$\begin{pmatrix} \varepsilon_t \\ u_{\beta t} \\ u_{\alpha t} \\ u_{ht} \end{pmatrix} \sim N \left(0, \begin{pmatrix} I_k & 0 & 0 & 0 \\ 0 & \Sigma_\beta & 0 & 0 \\ 0 & 0 & \Sigma_\alpha & 0 \\ 0 & 0 & 0 & \Sigma_h \end{pmatrix} \right) \tag{4.11}$$

where the variance-covariance structure for the innovations of the time-varying parameters $(\Sigma_\beta, \Sigma_\alpha, \Sigma_h)$ is assumed to be a diagonal matrix.

4.2.2 Identification

This study combines two identification methods to identify an unconventional monetary policy shock. One is the TVP-VAR-ZLB model presented by Nakajima (2011); the other is the sign-restricted identification method that was originally developed by Uhlig (2005) and Mountford and Uhlig (2009) and was later adapted to the TVP-VAR model by Canova and Gambetti (2009) and Franta (2011).

TVP-VAR-ZLB model

Nakajima (2011) estimated a four-variable TVP-VAR model that included the inflation rate, the output gap, the short-term interest rate, and the medium-term interest rate. Nakajima (2011) considered the situation in which the short-term interest rate did not react to structural shocks at the zero lower bound (ZLB) of the nominal interest rate, and estimated the TVP-VAR model by imposing the restriction that the coefficients and simultaneous parameters related to the short-term interest rate were equal to zero during the ZLB period.

More precisely, we assume that the variations in the time-varying parameters (i.e., $u_{\beta t}$ and $u_{\alpha t}$) in the interest rate equation vanish at the ZLB. In other words, the relevant elements β_t and α_t remain at their most recent values when the nominal interest rate hits the lower bound because variations in $u_{\beta t}$ and $u_{\alpha t}$ never occur. Based on this assumption, the values of these parameters are replaced by zero. By doing so, Nakajima (2011) incorporates the dynamic property of the short-term interest rate at the ZLB into the TVP-VAR model and calls the result the TVP-VAR-ZLB model.

However, there is one shortcoming in this study. Although the medium-term interest rate is contained in the estimation model, monetary policy shocks are only characterized by variations in the short-term interest rate. Therefore, a monetary policy that increases the monetary base, such as QE, cannot be captured explicitly.

Sign restrictions for unconventional monetary policy

Contrary to Nakajima (2011), the TVP-VAR model employed by Franta (2011) comprises of both the short-term interest rate and the monetary base, in addition to industrial production and the inflation rate. Furthermore, Franta (2011) uses sign restrictions to distinguish in an intuitive way a monetary policy shock at the ZLB. Throughout the sample period, a monetary policy shock is characterized as a shock that lowers the short-term interest rate while raising the monetary base and the output. In addition, a nonnegativity constraint is imposed on the short-term interest rate when the interest rate falls close to the ZLB. To be specific, in response to a monetary policy shock, the interest rate is assumed not to fall more than 1 basis point for the QE period, 5 basis points for a zero interest rate policy (ZIRP) period, and 50 basis points for a very low call rate period.

Although monetary policy is identified by a combination of interest rate and monetary base–thus taking QE into consideration–there yet remains the possibility that the short-term interest rate fluctuates in response to structural shocks, in spite of having fallen to the ZLB. Furthermore, the nonnegativity constraints imposed on the interest rate are not strictly adhered to because the minimum value of a call rate is 0.1 basis points during the sample period. In fact, even during the ZLB period, the time-varying volatility of the reduced-form residuals in the interest rate equation is estimated to be comparatively large, and this can make it difficult to sample a valid draw that strictly satisfies nonnegativity constraint. Therefore, besides the nonnegativity constraints, it is necessary to impose zero restrictions on the time-varying coefficients in order to correctly identify an unconventional monetary policy shock.

Identification strategy

Based on sign restrictions employed in Mountford and Uhlig (2009), this study identifies four types of structural shocks: business cycle, monetary policy, unanticipated fiscal policy and anticipated fiscal policy. To do so, the TVP-VAR model in this study consists of government spending, the short-term interest rate (call rate), excess stock returns of construction industry, the monetary base, and GDP. Regarding monetary policy shock, we extend the TVP-VAR-ZLB model presented by Nakajima (2011) with sign restrictions to discern monetary policy shock under the ZLB of nominal interest rate. Moreover, anticipated fiscal policy shock is characterized by incorporating the approach using stock returns as presented in Fisher and

Peters (2010) into "news shock" restrictions for government spending. The identification of business cycle and unanticipated fiscal policy shocks is performed according to the restrictions adopted in Mountford and Uhlig (2009). As shown in Section 4.2.3, each shock is distilled to satisfy orthogonality.

A business cycle shock is discerned as a shock which improves GDP. As described below, it is important for identifying anticipated fiscal policy shock to distinguish a business cycle shock which is firstly orthogonal to other structural shocks.

A monetary policy shock is a shock that lowers the short-term interest rate and increases monetary base. This assumption is widely accepted in previous studies that examine the effects of monetary policy by using sign restrictions. In this study, additional restrictions are imposed in order to characterize the dynamic property of the short-term interest rate under the ZLB. As noted in Nakajima (2011), the monetary policy does not work through the interest rate channel when the interest rate falls to near zero. In other words, it seems that the short-term interest rate does not change in response to monetary policy shock during the ZLB. Following Nakajima (2011), this study introduces the restriction that the coefficients related to the short-term interest rate equals zero when the interest rate hits the ZLB, thus removing the transmission effects for the interest rate of the structural shocks. Furthermore, regarding the contemporaneous relation during the ZLB period, we calculate the impulse vector of monetary policy shock–where the interest rate does not respond to monetary policy shock–by applying 'zero' restriction of Mountford and Uhlig (2009). This is further explained in Section 4.2.3. Combining the above two restrictions, we try to capture the dynamic property of monetary policy under the ZLB of nominal interest rate in the framework of a TVP-VAR model. In addition, the nonnegativity constraints are also incorporated to rule out the possibility that the interest rate falls more than the observed rate in response to monetary policy shock, as in Franta (2011).

With respect to fiscal policy shocks, unanticipated fiscal policy shock is simply defined as a shock which persistently increases government spending after the shock. Anticipated fiscal policy shock is identified by incorporating the approach presented in Fisher and Peters (2010) into sign restriction from Mountford and Uhlig (2009). Fisher and Peters (2010) consider that fiscal news fluctuate the current stock prices of firms that are related to fiscal policy. They capture the anticipated increase in U.S. government military spending by regarding the stock returns of military contractors as a proxy for news shock. By applying this identification strategy to the relationship between government spending and the Japanese construction industry, we attempt to identify anticipated fiscal policy shock in Japan.[2] Here, it is emphasized that business cycle shocks have been already identified to be orthogonal to fiscal policy shocks. As Fisher and Peters (2010) point out, not all variations in stock returns are due to fiscal policy news because firms sell not only to the public sector but also to the private sector. Moreover, stock prices are also influenced by the workings of the entire economy. This study resolves this problem to distinguish business cycle shock in advance. In addition to using stock returns, as in Mountford and Uhlig (2009), we impose the sign restriction that government spending only rises after fiscal news is realized but that it does not react beforehand. By doing so, we isolate the changes in stock returns that result from anticipated fiscal policy shocks involving future increases in government spending.

The sign restrictions employed in this study are provided in Table 4.1. Except for the responses of government spending and GDP to anticipated fiscal policy shocks, sign restrictions are imposed for one quarter. As in Chapter 2 and Chapter 3, we impose positive signs on the responses of GDP to monetary policy and both fiscal policy shocks.

[2] We do this because fiscal policy aimed at an economic stimulus in Japan is usually performed through public works, and those works are undertaken by construction industry.

Table 4.1: Sign restrictions

	Gov. spending	Interest rate	Stock returns	Monetary base	GDP
Non-fiscal shocks					
Business cycle					> 0 for 1Q
Monetary policy		≤ 0 for 1Q		> 0 for 1Q	> 0 for 1Q
		$\geq -\text{rate}(t)$ for 1Q			
Fiscal policy shock					
Unanticipated	> 0 for 1Q				> 0 for 1Q
Anticipated	$= 0$ for 1Q		> 0 for 1Q		> 0 for 4-6Q
	> 0 for 4-6Q				

4.2.3 Bayesian estimation

The estimation of the TVP-VAR model is carried out via a Bayesian approach using Markov-chain Monte Carlo (MCMC) method. This section briefly explains the process of estimating the TVP-VAR model.

Let us define $\beta = \{\beta_t\}_{t=s+1}^{T}$, $\alpha = \{\alpha_t\}_{t=s+1}^{T}$, $h = \{h_t\}_{t=1}^{T}$, $\omega = (\Sigma_\beta, \Sigma_\alpha, \Sigma_h)$, and $y = \{y_t\}_{t=s+1}^{T}$. Given the data y and the prior density function $\pi(\Theta)$, where $\Theta = \beta, \alpha, h, \omega$, the random samples from the posterior distribution $\pi(\Theta \mid y)$ are obtained as follows:

1. Set initial values of $\beta^{(0)}$, $\alpha^{(0)}$, $h^{(0)}$, $\omega^{(0)}$, and $j = 0$.

2. Sample $\beta^{(j+1)}$ from $\pi(\beta \mid \alpha^{(j)}, h^{(j)}, \Sigma_\beta^{(j)}, y)$.

3. Sample $\Sigma_\beta^{(j+1)}$ from $\pi(\Sigma_\beta \mid \beta^{(j+1)})$.

4. Sample $\alpha^{(j+1)}$ from $\pi(\alpha \mid \beta^{(j+1)}, h^{(j)}, \Sigma_\alpha^{(j)}, y)$.

5. Sample $\Sigma_\alpha^{(j+1)}$ from $\pi(\Sigma_\alpha \mid \alpha^{(j+1)})$.

6. Sample $h^{(j+1)}$ from $\pi(h \mid \beta^{(j+1)}, \alpha^{(j+1)}, \Sigma_h^{(j)}, y)$.

7. Sample $\Sigma_h^{(j+1)}$ from $\pi(\Sigma_h \mid h^{(j+1)})$.

8. Identify the structural shock based on $\beta^{(j+1)}$, $\alpha^{(j+1)}$ and $h^{(j+1)}$ by using sign restrictions.

9. Return to step.2 until N iterations have been completed.

For the above, N is set at 30000 and the first $N_0 = 20000$ samples are discarded as burn-in.

The detailed implementation for identifying restrictions in step 8 is performed following the work of Mountford and Uhlig (2009). Given the random samples of β, α and h in each iteration, we first generate a $k \times 1$ vector $q^{(1)}$ with the length being unity for each period, and calculate the impulse response vector associated to the business cycle shock at period as follows:

$$a_t^{(1)} = A_t^{-1}\Sigma_t q^{(1)}. \tag{4.12}$$

Subsequently, the impulse response vector to monetary policy shock $a_t^{(2)}$ is calculated by using $q^{(2)}$, where the restrictions $q^{(2)\prime}q^{(2)} = 1$, $q^{(2)\prime}q^{(1)} = 0$ are imposed on $q^{(2)}$ to satisfy the orthonormality. Furthermore, during the ZLB period the restriction that the short-term-interest rate does not react to monetary policy shock is additionally imposed on $q^{(2)}$. This restriction can be written as

$$0 = Cq^{(2)}, \tag{4.13}$$

where C is a $1 \times k$ matrix of the form

$$C = [c_{j1}(0), \cdots, c_{jk}(0)]. \tag{4.14}$$

Here, c_{ji} is the response of the j-th variables (which is the short-term interest rate) to the i-th column of $A_t^{-1}\Sigma_t$.

Likewise, anticipated fiscal policy shock is identified using $q^{(3)}$ which satisfies the orthonormality $[q^{(1)}, q^{(2)}, q^{(3)}]' \times [q^{(1)}, q^{(2)}, q^{(3)}] = [0, 0, 1]'$ and the 'news shock' restriction,

$$0 = Rq^{(3)}, \tag{4.15}$$

where R is a $3 \times k$ matrix, denoted as

$$R = \begin{bmatrix} r_{l1}(0) & \cdots & r_{lk}(0) \end{bmatrix}. \tag{4.16}$$

Similar to the element of C, r_{li} is the response of the l-th variables (which is government spending) to the i-th column of $A_t^{-1}\Sigma_t$. By this restriction, the response of government spending for the first period is set to zero. Finally, $q^{(4)}$ is generated to be orthogonal to $q^{(i)}, i = 1, 2, 3$, and $a^{(4)}$ is calculated to identify unanticipated fiscal policy shock.

In each period of each iteration, the above procedure is repeated until the impulse responses satisfy the sign restrictions, or the draws of q reaches 15. If the sign restrictions are satisfied at a period t, the new contemporaneous relation calculated by q is saved as a valid draw at a period t. If a valid draw is not provided in 15 draws at period t, we save the result obtained in the previous iteration, instead.

4.3 Data and specification

We employ the quarterly data of government spending, the short-term interest rate, excess stock returns of the construction industry, the monetary base, and GDP for the period 1985Q3-2013Q3. Government spending is defined as the sum of government consumption and public investment. The series, except for the call rate and excess stock returns, are seasonally adjusted per capita, and are logarithmically transformed. Following Fisher and Peters (2010), the excess stock returns are defined as the difference between the whole market returns and the returns of the construction industry. The short-term interest rate in this study is the uncollateralized overnight call rate which corresponds to the policy rate in Japan.

The data for government spending and GDP are obtained from the System of National Accounts (SNA) database in Japan, and released by the Cabinet Office of the Government of

Japan. We combine the series of *Quarterly Estimates of GDP (Reference year = 2005): Jul.-Sep. 2013 (The 2nd preliminary)* and *Provisional Estimates of GDP for Benchmark Year 2005.* The monetary data (i.e., the short-term interest rate and monetary base) are obtained from the Bank of Japan. We use *Call Rates, Uncollateralized Overnight (a)/Average(b)* as the short-term interest rate, and *Monetary Base (Reserve Requirement Rate Change Adjusted)/Seasonally Adjusted (X-12-ARIMA)/Average Amounts Outstanding* as a monetary base. The data for excess stock returns were calculated as follows. First, the quarterly average of the stock prices in both the construction industry and the whole market were computed by using the data for the daily closing price. The data of stock prices was downloaded from the Nikkei NEEDS-Financial QUEST database. Then, we obtained the series of stock returns as the growth rate of stock prices based on this quarterly series. Finally, following Fisher and Peters (2010), we constructed the excess stock returns by subtracting the whole market returns from the returns of the construction industry. Moreover, except for the interest rate and stock returns, the data were converted to their per capita value using total population data from the Japanese Population Census (Ministry of Internal Affairs and Communication).

The estimated VAR model includes the first differences of government spending, monetary base, GDP, and the level of excess stock returns and the short-term interest rate as well as a constant term. The number of lags is set to be three.

With respect to the ZLB period, this study follows the definition adopted by Hayashi and Koeda (2013). Accordingly, we regard the period in which the net policy rate, calculated by subtracting the rate paid on reserve (0.1% since 2008Q4) from the call rate, falls below 50 basis points as the ZLB period. Figure 4.1 shows a plot of the series of the short-term rate; the shaded areas indicate the ZLB period. These correspond to the periods from 1999Q2 to 2000Q2, 2001Q2 to 2006Q2 and 2008Q4 to date.

Following Nakajima (2011), the priors are assumed to be

$$\omega_{\beta i}^2 \sim IG(40, 0.02), \quad \omega_{\alpha i}^2 \sim IG(4, 0.02), \quad \omega_{hi}^2 \sim IG(4, 0.02), \tag{4.17}$$

where $\omega_{ji}^2, j = \beta, \alpha, h$ indicates the i-th element of $\Sigma_j, j = \beta, \alpha, h$. For the initial state of the time-varying parameters, we set the rather flat priors as follows:

$$\beta_0 \sim N(0, 10I), \quad \alpha_0 \sim N(0, 10I), \quad h_0 \sim N(0, 10I). \tag{4.18}$$

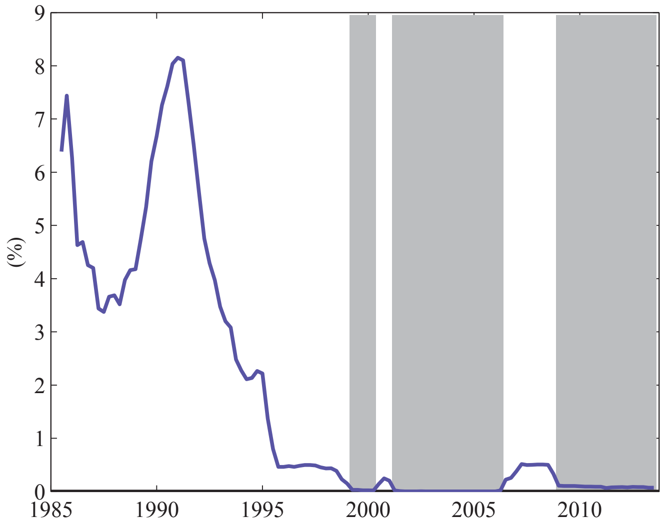

Figure 4.1: The short-term interest rate

Note: This figure shows the short-term interest rate (the uncollateralized overnight call rate) for the period 1985Q3-2013Q3. The shaded areas indicate the ZLB period in which the net policy rate, calculated by subtracting the rate paid on reserve (0.1% since 2008Q4) from the call rate, falls below 50 basis points.

4.4 Estimated results

4.4.1 Estimated parameter values

In Figure 4.2, we show the sample autocorrelation function, the sample paths, and the posterior densities for selected parameters. The sample paths for each parameter look stable, and the sample autocorrelation function damps stably. Furthermore, Table 4.2 shows the estimates for posterior means, standard deviations, the 95 percent credible intervals, and the p-value of the convergence diagnostics (CD) of Geweke (1992) for selected parameters. Based on the CD statistics, the null hypothesis of the convergence to the posterior distribution is not rejected at the 5 percent significance level. These results suggest that the samples in our estimation are efficiently generated and adequately converge.

Table 4.2: parameter values

Parameters	Mean	St. dev	95% intervals	CD
$\beta_{1,50}$	0.1004	0.2274	[−0.3463 0.5488]	0.409
$\alpha_{1,50}$	−0.0074	0.0321	[−0.0712 0.0536]	0.131
$h_{1,50}$	0.3676	0.2565	[−0.1198 0.8819]	0.081
$\omega_{\beta,1}$	0.0225	0.0018	[0.0194 0.0263]	0.885
$\omega_{\alpha,1}$	0.0276	0.0027	[0.0229 0.0334]	0.291
$\omega_{h,1}$	0.0813	0.0232	[0.0482 0.1379]	0.273

Note: This table shows the estimation result for selected parameters. CD denotes the p-value of the convergence diagnostics of Geweke (1992). Bandwidth of a Parzen window is set to be 500. The estimates of $\omega_{\beta,1}$ and $\omega_{\alpha,1}$ are multiplied by 100.

4.4.2 Volatility

Figure 4.3 shows the data and the posterior estimates of the volatilities of the reduced-form residuals (i.e., $\sigma_{it}^2 = \exp(h_{it}/2)$). The solid line and the shaded areas indicate the median of the sampled volatility and the 68% credible intervals, respectively.

The remarkable finding in Figure 4.3 is that the volatility of the short-term interest rate is estimated to be almost zero during the ZLB period. This stems from the zero restrictions

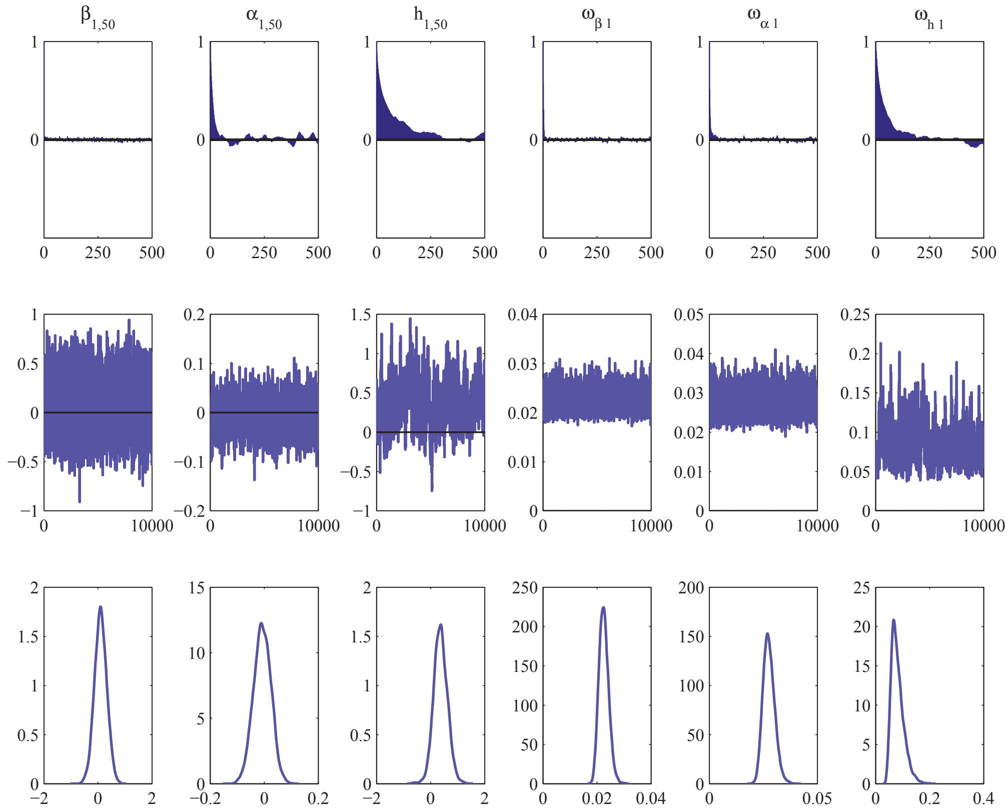

Figure 4.2: Estimation results for selected parameters

Note: This figure shows sample autocorrelations (top), sample paths (middle), and posterior densities (bottom).

on the coefficients and the contemporaneous relation of the interest rate equation. In fact, the estimated volatility presented by Franta (2011) does not exhibit such a small value. This small estimated volatility enables us to sample valid draws even when imposing a strict nonnegativity constraint on the short-term interest rate.

On the other hand, the volatility of the monetary base rose during the period in which an unconventional monetary policy was adopted, as seen in Figure 4.3. Furthermore, it soared in response to the bold monetary easing by Abenomics. This implies that the monetary policy shock played an important role in the variations of the monetary base.

In addition, we observe that the volatility of government spending in the 1990s was relatively

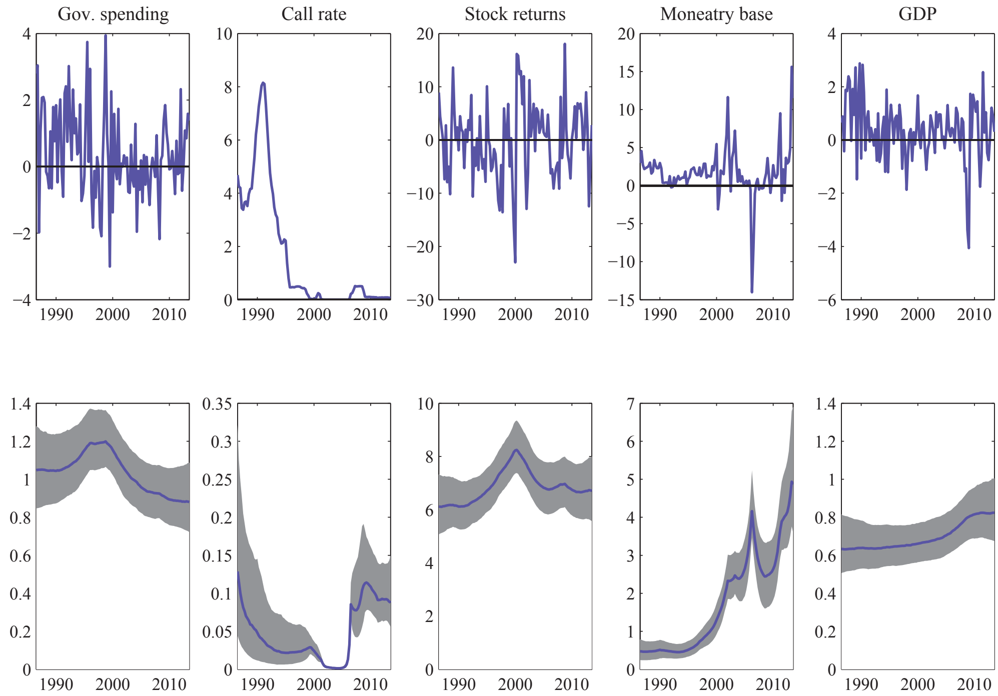

Figure 4.3: Data and standard deviations of reduced-form residuals

Note: This figure shows the data (top) and the estimated stochastic volatilities of reduced-form residuals (bottom). In the bottom charts, blue lines indicates the median of estimated values and the shaded areas indicates 68% credible intervals.

high compared with that after the 2000s. This is thought to reflect the aggressive fiscal policy implemented in the 1990s and the passive fiscal policy in the 2000s. Subsequently, stock returns became more volatile after the 2000s. The volatility of GDP was stable throughout the sample period.

4.4.3 Impulse response function

The IRFs for unanticipated and anticipated fiscal policy shocks and monetary policy shocks are plotted in Figures 4.4, 4.5, and 4.6, respectively. The figures show time series of the IRFs during the impact period and at one- and two-year horizons. In the figures, the shaded areas indicate the ZLB period.

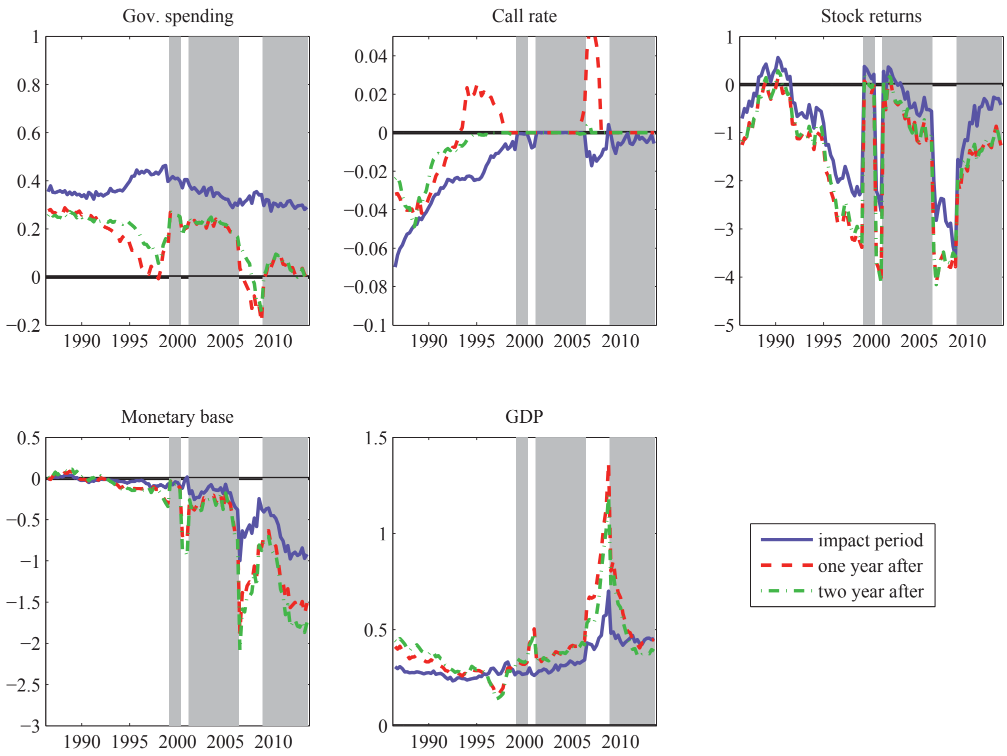

Figure 4.4: Responses to unanticipated fiscal policy shocks

Note: This figure shows the time-varying impulse responses to unanticipated fiscal policy shocks. The impact (blue), one-year (red), and two-year (green) horizons are plotted. The shaded periods correspond to the period of the ZLB of nominal interest rate.

Figure 4.4 shows the responses to an unanticipated fiscal policy shock. During the impact period, unanticipated fiscal policy shock raises government spending and output. Throughout the sample period, this shock has persistently positive effects on output (see the bottom-right chart). From the viewpoint of the fiscal and monetary interaction, we find that the response of the short-term interest rate shows a negative sign to an unanticipated fiscal policy shock except during the ZLB period. This negative reaction is very notable until the ZIRP period, which began in 1999Q2. Accordingly, this result indicates that the monetary authority takes an accommodative stance when positive fiscal policy is enacted.

Figure 4.5 shows the responses to the anticipated fiscal policy shock. Unlike the case of

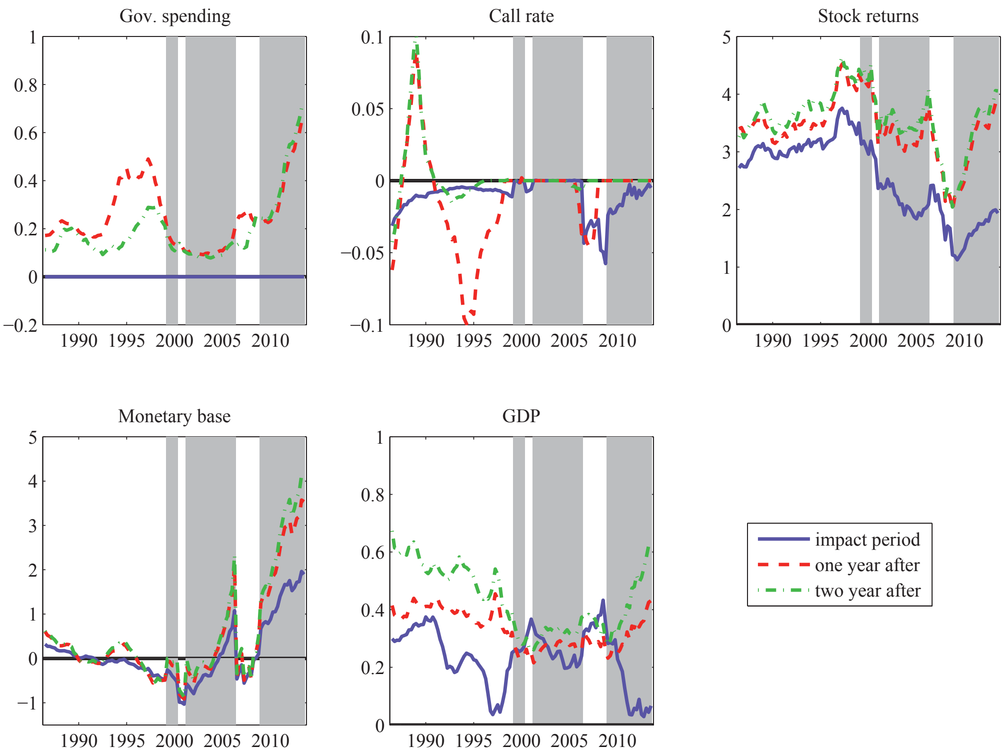

Figure 4.5: Responses to anticipated fiscal policy shocks

Note: This figure shows the time-varying impulse responses to anticipated fiscal policy shocks. The impact (blue), one-year (red), and two-year (green) horizons are plotted. The shaded periods correspond to the period of the ZLB of nominal interest rate.

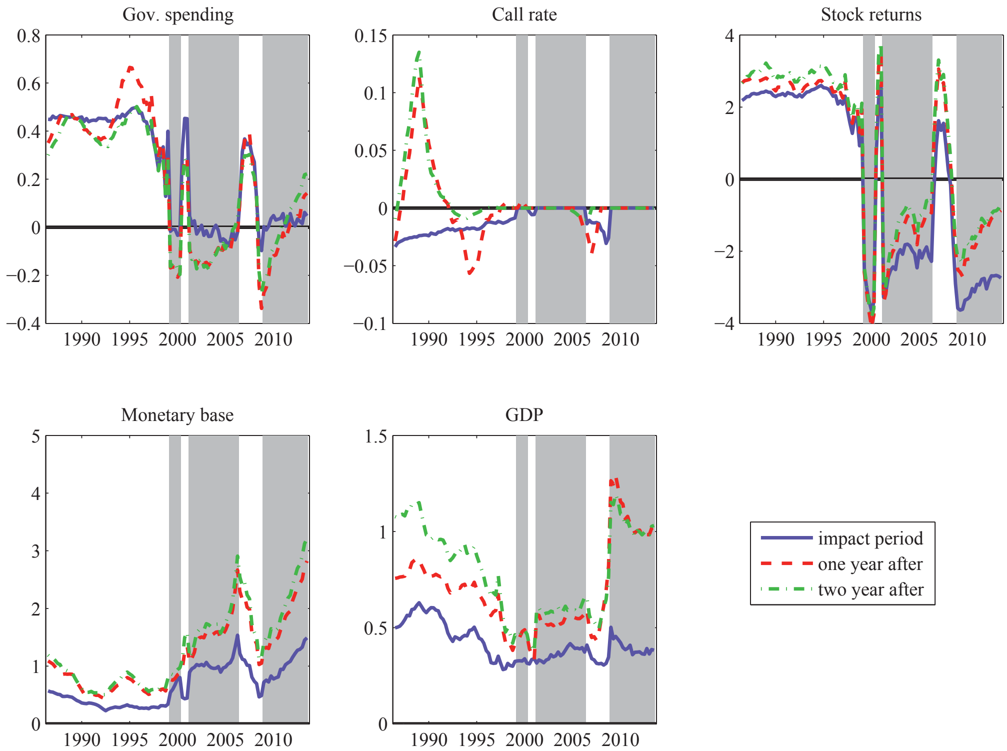

Figure 4.6: Responses to monetary policy shocks

Note: This figure shows the time-varying impulse responses to monetary policy shocks. The impact (blue), one-year (red), and two-year (green) horizons are plotted. The shaded periods correspond to the period of the ZLB of nominal interest rate.

the unanticipated fiscal policy shock, government spending does not respond during the impact period; however, it does gradually increase. Government spending shows quite large responses at the one- and two-year horizons in recent years, reflecting the demand for reconstruction after the Great East Japan Earthquake and the recent Abenomics. The responses of output vary throughout the sample period. Regarding monetary policy instruments, we note that the short-term interest rate barely reacts to the shock at first impact. In return, the monetary base increases simultaneously with the fiscal news shock, and this response is clearly observed in the ZLB period and the recent QE period.

Finally, we discuss the effects of monetary policy, which are shown in Figure 4.6. For the short-term interest rate, we observe that the strict nonnegativity constraint worked well during the ZLB period, especially in the QE period. Furthermore, it is found that the accommodative monetary policy shock had an expansionary effect on the monetary base during the QE and the recent Abenomics period. Also, we can see from the figure that output rose in response to an expansionary monetary policy shock. Although the positive response during the impact period was imposed by a sign restriction, a persistent positive effect is observed throughout the sample period. In addition, stock returns also indicate positive responses at the one- and two-year horizons.

4.4.4 Fiscal multiplier

In the figures shown above, we cannot compare the sizes of effects of the fiscal policy on output in each period because the size of the fiscal policy shock varies over time. Figure 4.7 shows the estimates of fiscal multipliers. Since the endogenous variables in the VAR model are within the logarithms, the fiscal multiplier is computed as follows:

$$\text{fiscal multiplier}_t = \frac{\text{IRF of GDP}}{\text{IRF of Gov. spending}} \times \frac{\text{GDP}_t}{\text{Gov. spending}_t}$$

This study calculates two types of fiscal multipliers: "impact" and "cumulative". The former uses impulse responses during the impact period, while the latter uses the sum of the impulse responses for four quarters. For the anticipated fiscal policy shock, only the cumulative multiplier is computed because government spending does not change during the first period.

Under the influence of the prior distribution and the restrictions imposed on the output, it seems that the fiscal multipliers are somewhat large. Therefore, this study focuses only on the time variation of the multipliers and not their values. For the unanticipated fiscal policy shock, both the impact and accumulated multipliers show the same historical pattern. The multipliers

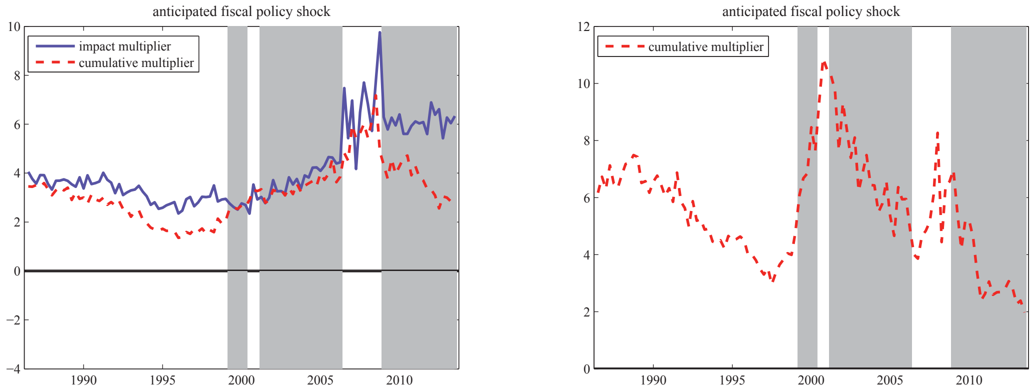

Figure 4.7: Fiscal multiplier

Note: Note: This figure shows the time-variations in th fiscal multipliers that are calculated based on the results of time-varying impulse responses. The shaded areas correspond to the period of the ZLB of nominal interest rate.

seem to be stable in the 1990s, and they gradually increase after the 2000s. On the other hand, the multiplier corresponding to the anticipated fiscal policy shock soars after the second half of the 1990s, but thereafter falls consistently.

For the fiscal policy under the ZLB, we observe that the value of the multipliers for unanticipated fiscal policy shock trends upwards during the first and second ZLB period (i.e., the periods of the ZIRP and the QE). With respect to anticipated fiscal policy shock, it seems that the multiplier increases only in the ZIRP period. Moreover, there is no evidence that the multipliers increase in the recent ZLB period. On the contrary, they decrease in comparison with a value of the non-ZLB period from the early 2000s. In conclusion, we do not see any evidence that the effects of fiscal policy were enhanced at the ZLB of the nominal interest rate.

4.5 Conclusion

This study has investigated the effects of fiscal and monetary policy shocks in Japan by using the TVP-VAR model and accounting for the ZLB of the nominal interest rate. In particular, we focused on the effects of the macroeconomic policies under the ZLB of the nominal interest rate. We proposed a new identification method for unconventional monetary policies, in which we combined the methods presented by Nakajima (2011) and Franta (2011). Furthermore, we calculated the fiscal multiplier throughout the sample period to confirm whether the effects of the fiscal policy increased during the period of the ZLB, as predicted by Christiano et al. (2011) and Eggertsson (2010).

We summarize the main results as follows. First, the TVP-VAR-ZLB model can replicate the fact that the volatility of the short-term interest rate becomes quite small during the ZLB period. This plays an important role in being able to sample valid draws that satisfy nonnegativity constraints on the short-term interest rate. Additionally, during the period of the QE and recent Abenomics, our estimated results capture the rise of volatility. Furthermore, the calculated IRFs imply that the effects of the fiscal and monetary policy shocks change over the sample period. The time-varying effects are observed in both the magnitude of the shock and the transmission mechanism. Also, the economic policy shocks identified in this study had positive effects on output in the medium and long term. Finally, as one of the principal findings of this study, we found there is no evidence that the effects of fiscal policy increase during a ZLB period in Japan.

Since several developed countries have experienced quite low policy rates, the estimation methodology presented in this study will be useful and applicable for analyzing those countries. In particular, it is important to clarify and reconcile the observed and theoretically predicted effects of fiscal policies during a ZLB. Thus, it remains to be determined if our results are only specific to the particular economy of Japan.

Chapter 5

External Shocks and Japanese Business Cycles: Evidence from a Sign-restricted VAR Model

5.1 Introduction

The Japanese economy is greatly influenced by external factors such as exchange rates, foreign demand, and oil prices. Recently, the global financial crisis that stemmed from the Lehman shock caused a large reduction in output in Japan. This event not only reconfirmed the perception that the Japanese economy is heavily influenced by the economic health of foreign countries, but also became a turning point for researchers to pay closer attention to external shocks. For example, Kawai and Takagi (2009) and Shioji and Uchino (2012) examined the effect of the Lehman shock and emphasized the finding that such external shocks play an important role in the Japanese economy.[1] However, Kawai and Takagi (2009) estimated the VAR model and included only GDP in Japan, emerging Asia, and the US and Europe–the post-Lehman shock period was not accounted for in their VAR analysis. Moreover, Shioji and Uchino (2012) only focused on the influence of the global crisis on the automobile industry.

Given the dearth of similar analyses in the literature, this study investigates macroeconomic roles of external shocks in Japanese business cycles based on the VAR model. Specifically, this study contributes to the body of knowledge on this topic by identifying risk premium shock and foreign demand shock, in addition to domestic supply and demand shocks, as external influences on the Japanese economy since the 1990s. As explained in more detail in Section 5.2, risk premium shock changes the exchange rate by varying the interest rate of a foreign bond. Similarly, foreign demand shock, as defined in this study, indicates a change in demand for the home product in foreign countries and thus captures business fluctuations overseas. In other words, foreign demand shock is the demand shock on the exports of home country goods rather than the demand shock that occurs in foreign countries; thus, it comprises both demand and supply shocks in foreign countries. Accordingly, we can capture most of the influences of external factors on the Japanese economy by identifying both risk premium and foreign demand shocks.[2]

Our methodology employs the sign-restricted VAR model–developed by Uhlig (2005) and recently used by Peersman (2005), Dedola and Neri (2007), Mountford and Uhlig (2009), and Pappa (2009)–in order to overcome the problems associated with the identification of structural shocks and misspecification of the model that have limited previous studies. A sign-restricted VAR model identifies shocks by restricting the shape of impulse response functions (IRFs), mak-

[1] Similarly, Shioji et al. (2011) indicate the importance of external shocks on the Japanese economy by estimating the open Dynamic Stochastic General Equilibrium (DSGE) model using the Bayesian MCMC method.

[2] Although oil price shock is also one of the most important external factors, this study ignores its effect for simplicity. However, a part of the effect might be identified as a foreign demand shock because oil shock impacts the world economy simultaneously.

ing the consideration of the order of exogeneity among the variables unnecessary. In addition, this particular type of model clarifies what kind of shock is identified because the restrictions are based on theoretical models. Given these advantages, the structural shocks described in this study are more correctly identified than in previous studies that have adopted the Cholesky decomposition. Although estimating the model using Bayesian methods allows structural shocks to be explicitly defined in the theoretical model thereby simplifying their interpretation, the correct results can fail to be estimated if the model is built inaccurately. However, the sign-restricted VAR can overcome this problem.

In this study, the sign restrictions imposed on the VAR model are derived from the theoretical model, namely, the Real Business Cycle (RBC) and New Keynesian (NK) models. In particular, we examine the robust sign restrictions based on the methodology presented in Dedola and Neri (2007) and Pappa (2009). More precisely, the theoretical IRF for each structural shock is calculated under sufficiently wide ranges for parameters, and the features that are common to any combination of parameter values are adopted as robust sign restrictions. Through this approach, the study mitigates the possibility of misspecification in the estimation model.

Specifically, this study examines the degree to which business fluctuations in Japan can be explained by external shocks, and investigates which shocks play important roles in each phase of the business cycle. For that purpose, we perform forecast error variance decomposition (FEVD) and historical decomposition (HD) as well as IRF analysis. These methods enable us to evaluate the effect of external shocks qualitatively and quantitatively. Although a number of previous studies have used these tools to analyze Japanese business cycles (e.g., Miyao, 2000, 2006), to our knowledge this study is the first to perform a HD accounting specifically for external shock.

In addition, this study has the following two advantages. First, the frequency of the data used in this study is monthly. This data frequency allows us to understand clearly what happened in each time period and to increase the sample size. For instance, we can observe precisely what occurred during the Lehman shock by using the monthly data from August 2008. Second, in addition to the Lehman shock, the Great East Japan Earthquake is contained in our sample period. Because this earthquake also caused a large reduction in output and exports, the present study is able to analyze the macroeconomic influences of this natural disaster.

The results of the FEVD indicate that 30% to 50% of the variation in output can be explained by external shocks. This result concurs with the findings of previous studies by confirming the importance of external factors. Furthermore, the HD reveals that supply shocks mainly explain the historical variations in output throughout the sample period. However, the findings also demonstrate that external shocks play a major role in altering the phases of business fluctuations. In particular, foreign demand shocks have contributed to the economic recovery since 2002, while risk premium shocks explain a large part of the reduction in output since the Lehman shock. Moreover, both these external shocks, in addition to domestic shocks, play a large role in the period after the Great East Japan Earthquake.

The rest of this study is organized as follows. In Section 5.2, the theoretical model is built in order to ascertain the robust sign restrictions. The theoretical IRFs are also drawn under various parameterizations, and restrictions are derived from the common IRF features. In Section 5.3, we describe the empirical methods and data used in this study. Section 5.4 presents the estimation results of the IRF, FEVD, and HD analyses. Section 5.5 concludes.

5.2 Theoretical model

This section derives the sign restrictions from a theoretical model. First, we construct a small open economy NK model that is a variant of the one presented by Leeper et al. (2011). Their model is a medium-scale small open economy model that has real and nominal rigidities and was built to analyze the effect of fiscal policy. Therefore, non-Ricardian households that are subject to liquidity constraints as well as debt financing are incorporated into the model. Furthermore, a local currency pricing approach is adopted; in other words, domestic intermediate goods firms set their export prices in a foreign currency unit.

The most important feature of their model is that the benchmark NK model nests the RBC model. By contrast, in previous studies that have adopted the sign-restricted VAR model (e.g., Braun and Shioji, 2007; Dedola and Neri, 2007; Pappa, 2009), sign restrictions have been based on the common features of the IRFs generated from the RBC and NK models. On this point, Leeper et al.'s (2011) model has advantageous characteristics for the present analysis.

However, this study does not estimate the model using Bayesian methods; it rather employs a simple model compared with that of Leeper et al. (2011). Specifically, several of the settings used to capture the properties of the actual data (e.g., habit formation and capital utilization rate) are omitted from the model proposed herein. Instead, we explicitly introduce import goods firms that set their imported goods' prices under Calvo (1983)-type price stickiness in order to examine how the degree of pass-through influences such prices.

5.2.1 Households

The economy is populated by a continuum of households indexed by $i \in [0, 1]$, and each household is comprised of a continuum of members, as in Chapter 2. The households are divided into two types: optimizing or Ricardian (R) households that have access to the capital market, and rule-of-thumb or non-Ricardian (N) households that face liquidity constraints and consume all of their disposable income in each period. As in Galí et al. (2007) and Colciago (2011), we assume that a fraction $\mu \in [0, 1]$ of population are non-Ricardian households, and the remaining population $1 - \mu$ are Ricardian households. The labor market structure in this chapter is also the same in Chapter 2. Accordingly, each member of the households provides each differentiated labor input.

Ricardian households

Let $c_t^R(i)$ be the real consumption of Ricardian households. Then, the lifetime utility of Ricardian households is written by

$$U = E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{c_t^R(i)^{1-\gamma} - 1}{1-\gamma} - \chi_t \frac{n_t(i)^{1+\lambda}}{1+\lambda} \right], \tag{5.1}$$

where $n_t(i) = \int_0^1 n_t(i,l)dl$, and, as noted in Chapter 2, we omit a superscript R from hours worked. They maximize this lifetime utility function subject to the budget constraint

$$\begin{aligned} &P_t c_t^R(i) + P_t i_t^R(i) + B_t^R(i) + S_t F_t^R(i) + P_t \tau_t^R(i) \\ &= R_{t-1} B_{t-1}^R(i) + R_{t-1}^* S_t F_{t-1}^R(i) + \int_0^1 W_t(l) n_t(i,l) dl + P_t r_t^k k_{t-1}^R(i) + D_t^R(i) \end{aligned} \tag{5.2}$$

and the capital accumulation equation

$$k_t^R(i) = (1-\delta) k_{t-1}^R(i) + \left\{ 1 - s\left(\frac{i_t^R(i)}{i_{t-1}^R(i)} \right) \right\} i_t^R(i) \tag{5.3}$$

where χ_t captures the labor preference that follows an exogenous process, and β, γ, and λ denote the discount rate, risk aversion, and inverse of Frisch labor elasticity, respectively. Uppercase letters denote nominal variables, and P_t is the price of a final good. Ricardian households receive interest payments from domestic and international one-period risk-free nominal bonds denoted by B_t and F_t, respectively as well as wage income $\int_0^1 W_t(l) n_t(i,l) dl$, rental income from capital $r_t^k k_t$, and dividends D_t. They consume c_t^R and invest i_t^R, and they pay a lump-sum tax denoted by τ_t^R. As noted in Chapter 2, the labor market structure rules out differences in labor income among each household.

The gross nominal interest rates paid for domestic and international bonds are R_t and R_t^*. Since the international bond and its interest payment are denominated in foreign currency units, they are converted into home currency units by a multiplicative nominal exchange rate S_t. Concerning capital accumulation, we assume an investment adjustment cost $s(\cdot)i^R$ as in Christiano et al. (2005), where $s(1) = s'(1) = 0$ and $s''(1) > 0$.[3]

Non-Ricardian households

On the other hand, non-Ricardian households simply consume their current disposable income in each period. By denoting the consumption of non-Ricardian households as $c_t^N(i)$, their budget constraints are written as

$$P_t c_t^N(i) = \int_0^1 W_t(l) n_t(i,l) dl - P_t \tau_t^N(i), \tag{5.4}$$

where $\tau_t^N(i)$ denotes the lump-sum taxes paid by non-Ricardian households. As in Galí et al. (2007), we assume the lump-sum tax paid by non-Ricardian households differ from those for Ricardian households in order to equate steady state consumption across household types.

5.2.2 Wage setting

As mentioned above, each household provides a differentiated labor input $n_t(l)$ for intermediate goods firms. A perfectly competitive labor force produces a composite effective labor n_t according to

$$n_t = \left[\int_0^1 n_t(l)^{\frac{\varepsilon_w - 1}{\varepsilon_w}} dl \right]^{\frac{\varepsilon_w}{\varepsilon_w - 1}} \tag{5.5}$$

[3] In this study, we define investment adjustment cost as $\kappa \equiv 1/s''(1)$.

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

$$n_t(l) = \left(\frac{W_t(l)}{W_t}\right)^{-\varepsilon_w} n_t^d, \tag{5.6}$$

for all l, and the aggregate nominal wage is equal to

$$W_t = \left[\int_0^1 W_t(l)^{1-\varepsilon_w} dl\right]^{\frac{1}{1-\varepsilon_w}}. \tag{5.7}$$

Here, we assume that labor demand is uniformly distributed regardless of households types.

With respect to wage setting, we follow the modeling used in Galí et al. (2007) and Colciago (2011), in which each labor union l sets its nominal wage $W_t(l)$ to maximize the weighted average of the lifetime utility of Ricardian households and non-Ricardian households. In each period, a labor union resets the optimal nominal wage $W_t^*(l)$ with a probability $1-\rho_w$. Thus, the problem for a labor union l is written as

$$\max_{W_t^*(l)} E_t \sum_{s=0}^{\infty} \rho_w \Lambda_{t,t+s} \left[(1-\mu)\frac{c_{t+s}^R(l)^{1-\gamma}-1}{1-\gamma} + \mu\frac{c_{t+s}^N(l)^{1-\gamma}-1}{1-\gamma} - \chi_t \frac{n_{t+s}(l)^{1+\lambda}}{1+\lambda}\right] \tag{5.8}$$

subject to (5.2), (5.4), and (5.6), where $\Lambda_{t,t+s} = \beta^s (c_{t+s}^R / c_t^R)^{-1}$ denotes the stochastic discount factor. In the symmetric equilibrium, the first-order condition can be expressed as

$$W_t^*(l) = \frac{\varepsilon_w}{\varepsilon_w - 1} \frac{E_t \sum_{s=0}^{\infty} \rho_w^s \Lambda_{t,t+s} \chi_t n_{t+s}^{1+\lambda}}{E_t \sum_{s=0}^{\infty} \rho_w^s \Lambda_{t,t+s} \left[(1-\mu)\frac{n_{t+s}}{P_{t+s} {c_{t+s}^R}^{\gamma}} + \mu \frac{n_{t+s}}{P_{t+s} {c_{t+s}^N}^{\gamma}}\right]}. \tag{5.9}$$

Combining this with (5.7), the evolution of the aggregate nominal wage is given by

$$W_t = [(1-\rho_w)W_t^{*1-\varepsilon_w} + \rho_w W_{t-1}^{1-\varepsilon_w}]^{\frac{1}{1-\varepsilon_w}}. \tag{5.10}$$

Then, the log-linearization of (5.9) and (5.10) around the steady state yields the dynamic equation of the real wage as

$$\hat{w}_t = \Gamma\hat{w}_{t-1} + \Gamma\beta E_t\hat{w}_{t+1} + \Gamma\beta E_t\hat{\pi}_{t+1} - \Gamma\hat{\pi}_t + \kappa_w\Gamma\gamma\hat{c}_t + \kappa_w\Gamma\lambda\hat{n}_t + \kappa_w\Gamma\chi_t \tag{5.11}$$

where a hat denotes the deviation from the steady-state value, $\Gamma = \rho_w/(1+\beta\rho_w^2)$, and $\kappa_w = (1-\beta\rho_w)(1-\rho_w)/\rho_w$.

5.2.3 Firms

Final goods firms

Final goods firms are assumed to face perfectly competitive goods markets. They produce final goods $y_{n,t}$ by combining a bundle of domestic intermediate goods $y_{h,t}$ and a bundle of imported intermediate goods $y_{m,t}$ according to the following CES production function:

$$y_{n,t} = \left[\omega^{\frac{1}{\eta}} y_{h,t}^{\frac{\eta-1}{\eta}} + (1-\omega)^{\frac{1}{\eta}} y_{m,t}^{\frac{\eta-1}{\eta}}\right]^{\frac{\eta}{\eta-1}} \tag{5.12}$$

where η is the elasticity of substitution between domestic and imported intermediate goods, and $\omega \in [0,1]$ denotes the degree of home bias in producing final goods. Similar to final goods, bundles of domestic and imported intermediate goods are produced according to the following technology:

$$y_{h,t} = \left[\int_0^1 y_{h,t}(j)^{\frac{\theta_p-1}{\theta_p}} dj\right]^{\frac{\theta_p}{\theta_p-1}}, y_{m,t} = \left[\int_0^1 y_{m,t}(m)^{\frac{\theta_p-1}{\theta_p}} dm\right]^{\frac{\theta_p}{\theta_p-1}} \tag{5.13}$$

where θ_p is the elasticity of substitution between differentiated intermediate goods.

Domestic intermediate goods firms

Domestic intermediate goods firms $j \in [0,1]$ that are faced with a monopolistically competitive market produce differentiated intermediate goods according to the Cobb–Douglas production function:

$$y_{h,t}(j) = A_t k_{t-1}(j)^{\alpha} n_t(j)^{1-\alpha}. \tag{5.14}$$

where A_t denotes the total factor productivity (TFP), which is given exogenously. Intermediate goods are assumed to be not only purchased by final goods firms in the home country, but also exported abroad.

Although the market for intermediate goods is monopolistically competitive, the factor market faced by intermediate goods firms is assumed to be competitive. As a result of the cost minimization problem for intermediate goods firms, the real marginal cost mc_t is given by

$$mc_t = \frac{w_t}{(1-\alpha)A_t}\left(\frac{(1-\alpha)r_t^k}{\alpha w_t}\right)^{\alpha}. \tag{5.15}$$

5.2.4 Price setting

Domestic intermediate goods firms

As noted above, domestic intermediate goods firms sell their goods in both home and foreign markets. These prices are charged separately according to the demand in each market. In the home market, demand for firm j's output $y_{h,t}(j)$ is given by

$$y_{h,t}(j) = \omega\left(\frac{P_{h,t}(j)}{P_{h,t}}\right)^{-\theta_p}\left(\frac{P_{h,t}}{P_t}\right)^{-\eta} y_{n,t}, \tag{5.16}$$

where $P_{h,t}(j)$ is the output price in the home market set by firm j and $P_{h,t}$ is an aggregate price index of domestic intermediate goods, which is derived from a profit maximization problem of final goods firms and intermediate goods bundlers. Similarly, demand for firm j's output $y_{x,t}(j)$ in the foreign market is assumed to be

$$y_{x,t}(j) = \left(\frac{P^*_{x,t}(j)}{P^*_{x,t}}\right)^{-\theta_x} y_t^{foreign} \tag{5.17}$$

where $P^*_{x,t}(j)$ and $P^*_{x,t}$ are the export price charged by firm j and an aggregate exported goods price index, respectively. Both the export price and the aggregate index are denominated in the foreign currency unit, while $y_t^{foreign}$ is aggregate foreign demand, which is given exogenously.

According to the demand functions (5.16) and (5.17), intermediate goods firms set their domestic and export prices under a Calvo (1983) mechanism. Therefore, the problem of domestic intermediate goods firms j setting $P_{h,t}(j)$ is written as

$$\max_{P_{h,t}(j)} E_t \sum_{s=0}^{\infty} \rho_h^s \Lambda_{t,t+s} \left[P_{h,t}(j) y_{h,t+s}(j) - P_{t+s} y_{h,t+s}(j) mc_{t+s}\right] \tag{5.18}$$

subject to (16), where ρ_h denotes the probability that they cannot reoptimize their domestic prices in given period. Likewise, the problem of setting export prices $P^*_{x,t}(j)$ is

$$\max_{P^*_{x,t}(j)} E_t \sum_{s=0}^{\infty} \rho_x^s \Lambda_{t,t+s} \left[P^*_{x,t}(j) S_{t+s} y_{x,t+s}(j) - P_{t+s} y_{x,t+s}(j) mc_{t+s}\right] \tag{5.19}$$

subject to (5.17), where ρ_x also denotes the probability that they cannot reoptimize their export prices.

By solving these problems, a log-linearized NK Phillips curve (NKPC) is obtained as follows:

$$\hat{\pi}_{h,t} = \beta E_t \hat{\pi}_{h,t+1} + \frac{(1-\rho_p)(1-\beta\rho_p)}{\rho_p}[\hat{mc}_t - \hat{P}_{h,t} + \hat{P}_t] \tag{5.20}$$

$$\hat{\pi}^*_{x,t} = \beta E_t \hat{\pi}^*_{x,t+1} + \frac{(1-\rho_x)(1-\beta\rho_x)}{\rho_x}[\hat{mc}_t - \hat{S}_t - \hat{P}^*_{x,t} + \hat{P}_t] \tag{5.21}$$

where a hat denotes the log deviation from the steady state.

Imported goods firms

Imported goods firms $m \in [0,1]$ import differentiated foreign goods $y_{m,t}(m)$ at prices $P^*_{m,t}$ denominated in the foreign currency unit and sell them to final goods firms at prices $P_{m,t}(m)$ denominated in the home currency unit. It is assumed that imported goods firms are also faced with a monopolistically competitive market and thus set their prices $P_{m,t}(m)$ under Calvo (1983)-type price stickiness. Thus, in each period they reset their prices with a probability $1-\rho_m$ in order to maximize the discounted sum of profit flows

$$\max_{P_{m,t}(m)} E_t \sum_{s=0}^{\infty} \rho_m^s \Lambda_{t,t+s} \left[P_{m,t}(m) y_{m,t+s}(m) - P_{t+s} y_{m,t+s}(m) S_{t+s} P^*_{m,t+s}\right] \tag{5.22}$$

subject to demand for imported goods $y_{m,t}(m)$

$$y_{m,t}(m) = (1-\omega)\left(\frac{P_{m,t}(m)}{P_{m,t}}\right)^{-\theta_p}\left(\frac{P_{m,t}}{P_t}\right)^{-\eta} y_{n,t} \tag{5.23}$$

where $P_{m,t}$ is the aggregate import price index. As for the problem of domestic intermediate goods firms, the NKPC is given by

$$\hat{\pi}_{m,t} = \beta E_t \hat{\pi}_{m,t+1} + \frac{(1-\rho_m)(1-\beta\rho_m)}{\rho_m}[\hat{P}^*_{m,t} + \hat{S}_t - \hat{P}_{m,t} + \hat{P}_t]. \tag{5.24}$$

5.2.5 Fiscal policy and monetary policy

The government budget constraint is given by

$$B_t + P_t\tau_t = P_t g_t + R_{t-1}B_{t-1} \tag{5.25}$$

where g_t denotes real government spending and is regarded as an exogenous variable. As shown in Galí et al. (2007) and Leeper et al. (2011), tax is assumed to respond to the debt-to-GDP ratio. Hence, a fiscal rule is written by

$$\hat{\tau}_t = \phi_b \hat{b}_{t-1} + \phi_g \hat{g}_t \tag{5.26}$$

where $\hat{\tau}_t \equiv (\tau_t - \tau)/y$, $\hat{b}_t \equiv (B_t/P_t - B/P)/y$, and $\hat{g}_t \equiv (g_t - g)/y$.

In return, the monetary authority sets the nominal interest rate r_t following the Taylor rule

$$\hat{r}_t = \phi_\pi \pi_t + \phi_s \hat{S}_t + u_t^{mp} \tag{5.27}$$

where $\hat{r}_t$ denotes the log deviation of the net nominal interest rate from the steady-state value. Furthermore, the nominal interest rate is also assumed to respond to the log deviation of an exchange rate $\hat{S}_t$ by following Lubik and Schorfheide (2005) and Adolfson et al. (2007). Finally, u_t^{mp} denotes monetary policy disturbance, which is assumed to be exogenous.

5.2.6 Rest of the model

Risk premium and foreign assets

As described in Shioji et al. (2011), the interest rate of a foreign bond denominated in the foreign currency unit r_t^* is linked to a constant world interest rate r^w through the uncovered interest rate parity condition with a risk premium term that depends on the net foreign asset position (S_tF_t/P_ty_t)

$$r_t^* = r^w + \psi\left\{\exp\left(-\frac{S_tF_t}{P_ty_t}\right) - 1\right\} + u_t^{risk} \tag{5.28}$$

where u_t^{risk} denotes the risk premium following an exogenous process. In this specification, if the home country is a net borrower (i.e., $F_t < 0$), it has to pay an interest payment higher than the world interest rate to the foreign country.

In addition, the dynamics of the net foreign asset are written as

$$S_tF_t = R_{t-1}^* S_tF_{t-1} + S_tP_{x,t}^* y_{x,t} - P_{m,t}y_{m,t}. \tag{5.29}$$

This implies that the trade balance equals the capital account balance in each period.

Market clearing condition and resource constraint

Aggregate consumption, lump-sum taxes, capital, investment, domestic and international bonds, and dividends are given by

$$c_t = (1-\mu)c_t^R + \mu c_t^N; \quad k_t = (1-\mu)k_t^R; \quad B_t = (1-\mu)B_t^R;$$

$$\tau_t = (1-\mu)\tau_t^R + \mu\tau_t^N; \quad i_t = (1-\mu)i_t^R; \quad F_t = (1-\mu)F_t^R;$$

$$D_t = (1-\mu)D_t^R.$$

The clearing conditions in the factor and goods markets are expressed as

$$n_t = \int_0^1 n_t(j)dj; \quad k_t = \int_0^1 k_t(j)dj;$$

$$y_{n,t} = c_t + i_t + g_t$$

and the resource constraint is

$$y_t = y_{h,t} + y_{x,t}.$$

Exogenous variables

With respect to the evolution of the exogenous variables u_t^{risk}, $y_t^{foreign}$, A_t, χ_t, g_t, and u_t^{mp}, we assume the following AR(1) process

$$\hat{u}_t^{risk} = \rho_{risk}\hat{u}_{t-1}^{risk} + e_t^{risk} \tag{5.30}$$

$$\hat{y}_t^{foreign} = \rho_{foreign}\hat{y}_{t-1}^{foreign} + e_t^{foreign} \tag{5.31}$$

$$\hat{A}_t = \rho_A \hat{A}_{t-1} + e_t^A \tag{5.32}$$

$$\hat{\chi}_t = \rho_\chi \hat{\chi}_{t-1} + e_t^\chi \tag{5.33}$$

$$\hat{g}_t = \rho_g \hat{g}_{t-1} + e_t^g \tag{5.34}$$

$$\hat{u}_t^{mp} = \rho_{mp}\hat{u}_{t-1}^{mp} + e_t^{mp} \tag{5.35}$$

where e_t^{risk}, $e_t^{foreign}$, e_t^A, e_t^χ, e_t^g, and e_t^{mp} represent risk premium, foreign demand, TFP, labor preference, fiscal policy, and monetary policy shocks, respectively.

5.2.7 Parameter range

The model period is assumed to be one month in order to match the frequency of the data used in the empirical analysis. To find robust sign restrictions, IRFs are computed by specifying the ranges for certain parameters. Specifically, we calculate the theoretical IRFs as follows. First, we find the robust sign restrictions by following the process presented by Pappa (2009). Let Θ denote the parameters in the interval $[\theta_l, \theta_u]$. Θ is assumed to be uniformly distributed in the interval $[\theta_l, \theta_u]$, namely $\Theta \sim U(\theta_l, \theta_u)$. Then, we randomly draw Θ and calculate the IRFs. Repeating this process sufficiently provides the range of IRFs that corresponds to the combination of various parameter values. Then, only robust signs are adopted as the restrictions imposed on the empirical model. Finally, the values of the calibration parameters are selected based on the results estimated in previous studies, or based on the values used in the calibration exercises.

In previous analyses that have used macro-level data, the share of non-Ricardian households in Japan has been estimated to be approximately 30% (Hatano, 2004; Iwata, 2009). Conversely, Kohara and Horioka (2006) used micro-level data, and estimated the same value to be between 8% and 15%. In this study, we set the interval $[0, 0.5]$ for the share of non-Ricardian households in order to include the values estimated by previous studies. In regard to the stickiness of price and wage, the lower and upper bounds are chosen to be 0.2 and 0.9, respectively. The upper bound is set to be a somewhat larger value than that estimated by Iiboshi et al. (2008) and Sugo and Ueda (2008), while the lower bound is set with reference to Pappa (2009). Furthermore, the interval for the investment adjustment cost is set to be $[0.15, 5]$. The lower value of this interval is based on the findings of Sugo and Ueda (2008), while the higher value is chosen as a level at which investment can be carried out smoothly.

The parameters in the fiscal policy rule (i.e., ϕ_b, ϕ_g) are chosen based on the findings of Galí et al. (2007). However, because the calibration parameter values in Galí et al. (2007) were set to fit the U.S. economy and a calibration period of one quarter, we adopt wider intervals for these parameters in order to generalize their application. Specifically, the interval for the elasticity of tax to government spending is $[-0.05, 0.15]$ and that to debt is $[0, 0.25]$. In order to take account of the possibility of reducing tax in response to increases in government spending, the interval of the elasticity of tax to government spending includes negative values.

The parameters in the monetary policy rule are set as follows. The response of the interest rate to inflation is limited to the interval $[1.01, 1.5]$. This range fulfills the Taylor principle and therefore it is often used in calibration exercises. However, the feedback parameter for the exchange rate under the Taylor rule is controversial. First, it is uncertain whether the central bank determines the interest rate in response to the exchange rate. Moreover, the interest rate in Japan since the second half of the 1990s has fallen into a static state because of the country's zero interest rate policy. However, some studies have found that the interest rate reacts to the exchange rate. For example, Nakazawa et al. (2002) estimate a VAR model and report that the call rate negatively responds to depreciation shock in the exchange rate. Nakazawa (2002) also confirms that monetary policy reacts to a variation in the exchange rate by estimating the Taylor rule directly.

However, the values of the feedback parameter for the exchange rate estimated by Nakazawa (2002) are small or not significant in the sample period used (1987M7–2001M3). Therefore, this study adopts the interval $[0, 0.5]$ for this parameter. Moreover, the parameters for risk premium and the persistency of exogenous variables are assumed to be in $[0.01, 0.03]$ and $[0.8, 0.95]$, respectively. The remaining parameters are then fixed to a particular set of values. All parameter values and intervals are displayed in Table 5.1.

5.2.8 Sign restrictions

Figure 5.1 displays 68% bands for the responses of the nominal exchange rate, real exports, output, and price for six types of exogenous shocks. In this exercise, the number of random draws is 10,000.

The first row in Figure 5.1 depicts the responses to risk premium shock. As we can see from the figure, risk premium shock raises the nominal exchange rate (i.e., depreciation) during the first two periods and increases export quantities through a fall in the export price. Although the magnitude of response changes depends on the degree of price rigidity, the IRF of exports show a positive sign during the first three periods, while the response of output shows a negative sign, at least in the examined period. The final goods price takes various signs according to the

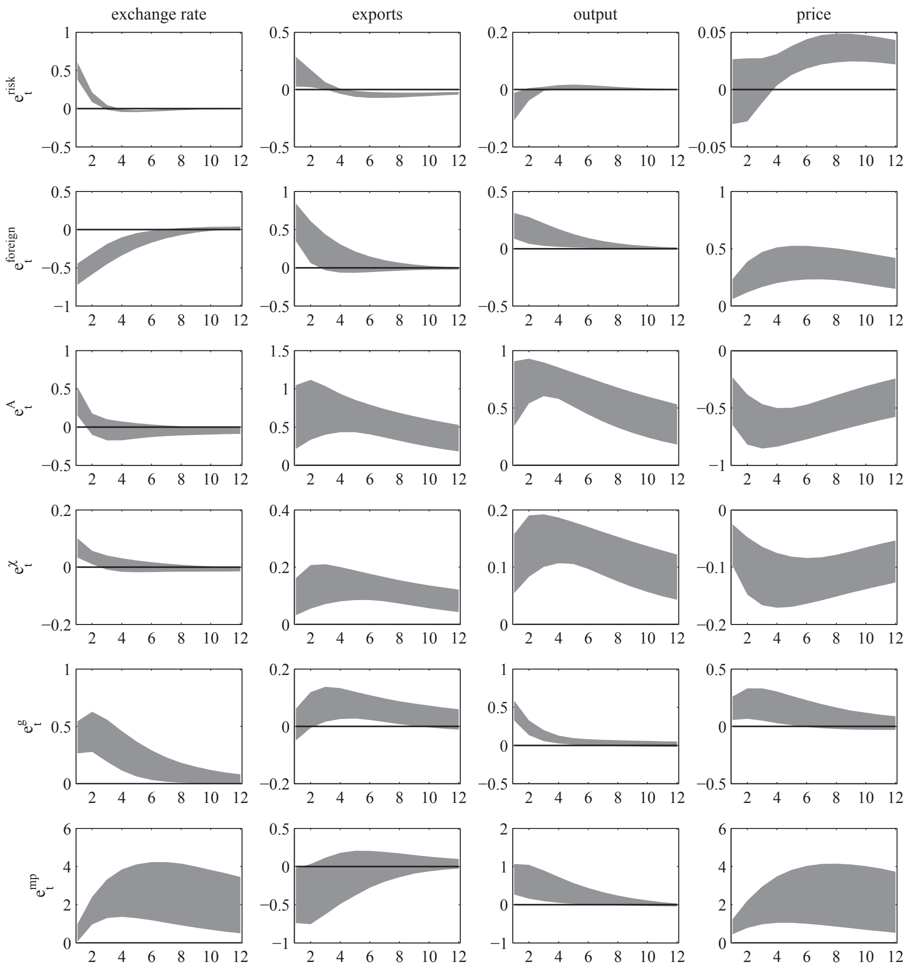

Figure 5.1: Impulse Response Function (Theoretical model)

Note: Figure 5.1 displays the 68% bands for the nominal exchange rate, real exports, output, and price for six types of exogenous shocks.

Table 5.1: Calibration Parameters

Parameter	Value	Description
β	0.9975	Subjective discount rate
δ	0.01	Depreciation rate
α	0.3	Share of capital
γ	1.5	Risk aversion
λ	2	Inverse labor supply elasticity
ω	0.88	Degree of home bias
μ	[0, 0.5]	Share of non-Ricardian households
κ	[0.15, 5]	Investment adjustment cost
ρ_p, ρ_x	[0.2, 0.9]	Calvo parameters on domestic and exported prices
ρ_f	[0.2, 0.9]	Calvo parameters on imported prices
ρ_w	[0.2, 0.9]	Calvo parameters on wages
θ_p	6	Elasticity of substitution in production
ε_w	6	Elasticity of substitution in labor input
θ_x	1	Price elasticity of exported goods
η	1	Elasticity of substitution bet. domestic and imported goods
ψ	[0.01, 0.03]	Parameter on the risk premium
s_g	0.2	Steady-state share of government spending
ϕ_g	[-0.05, 0.15]	Elasticity of tax to government spending
ϕ_b	[0, 0.25]	Elasticity of tax to debt
ϕ_π	[1.01, 1.5]	Monetary policy response of an inflation rate
ϕ_S	[0, 0.5]	Monetary policy response of an exchange rate
ρ_i	[0.8, 0.95]	persistency of exogenous shocks

Notes: The values of parameters are basically set up according to the previous studies. With respect to the value of ω, it is set on the basis of the import inducement coefficient in the Input-Output Table as in Shioji et al. (2011).

parameter values in the short run; however, specific qualitative features cannot be ascertained.

The responses to foreign demand shock are shown in the second row. Export quantities naturally increase and output also rises in response to foreign demand shock. However, similar to risk premium shock, the persistency of the response in exports and output depends on the degree of price stickiness; thus, the sign for exports becomes negative in the third period when the degree of price stickiness is low. In response to foreign demand shock, the nominal exchange rate appreciates as expected, and the IRF of the final goods price also shows a positive sign throughout the study period.

The third and fourth rows show the responses to TFP and labor preference shocks, respectively. We classify these two types of shocks as supply shocks because each variable describes a similar path in response to them. As seen in traditional theoretical analysis, a supply shock reduces aggregate price and increases output through the fall in the marginal cost. Moreover, a decrease in the marginal cost lowers the export price. Hence, export quantities positively respond to a supply shock. Furthermore, the nominal exchange rate depreciates because of a falling domestic price.

Finally, the fifth and sixth rows in Figure 5.1 show the responses to fiscal and monetary policy shocks, which are categorized as demand shocks. Except for exports, the signs of the responses to each variable are the same. Similar to supply shock, positive responses to price and output are obtained in line with traditional theory. In addition, the nominal exchange rate shows a persistent positive response to demand shock. With respect to export quantities, no particular sign can be determined because the responses are different for each shock. Based on these results, we can set the sign restrictions summarized in Table 5.2.

Table 5.2: Sign restrictions

		exchange rate			exports			output			price		
		1	2	3	1	2	3	1	2	3	1	2	3
risk premium shock		+	+		+	+	+	−					
foreign demand shock		−	−	−	+	+		+	+	+	+	+	+
supply shock	TFP shock	+			+	+	+	+	+	+	−	−	−
	labor preference shock	+	+		+	+	+	+	+	+	−	−	−
	summary	+			+	+	+	+	+	+	−	−	−
demand shock	fiscal policy shock	+	+	+			+	+	+	+	+	+	+
	monetary policy shock	+	+	+	−			+	+	+	+	+	+
	summary	+	+	+				+	+	+	+	+	+

Notes: Sign restrictions are set to be consistent with the theoretical IRFs. Blank spaces denote that no sign restrictions are imposed. The thick letters mean that signs there are used to impose on the IRFs in the empirical analysis.

These sign restrictions are generally imposed for the first three months. However, if the same signs are not observed during these three periods, as seen in the response of output to risk premium shock, restrictions are imposed only on the period when the robust restrictions were obtained. The short-run restrictions shown in Table 5.2 can be used to identify the four structural shocks. Risk premium shock is distinguished from other shocks by the negative response in output in the examined period. Furthermore, foreign demand shock is the only shock that causes the exchange rate to appreciate. Moreover, supply and demand shocks are identified by a price response.

For these sign restrictions, we refer to the J-curve effect, which implies that the response of trade balance to a variation in the exchange rate shows the opposite sign to that expected in the short run. Gupta-Kapoor and Ramakrishnan (1999) and Hsing (2005) have pointed out the existence of the J-curve effect in Japan. However, our results hardly change even if no restrictions are imposed on the response of exports to risk premium shock in the first period. Hence, we present the results based on the restrictions in Table 5.2 as the benchmark.

5.3 Estimation model and data

5.3.1 Sign-restricted VAR model

The sign-restricted VAR model is estimated by the following process. First, we estimate the reduced form VAR model:

$$Y_t = B_0 + B_1 Y_{t-1} + B_2 Y_{t-2} + \cdots + B_p Y_{t-p} + u_t \tag{5.36}$$

$$u_t = A_0 \varepsilon_t \tag{5.37}$$

$$u_t \sim N(0, \Sigma), \ \varepsilon_t \sim N(0, I) \tag{5.38}$$

where Y_t is a vector of endogenous variables, $B = [B_0, B_1, \cdots, B_p]$ is a vector of constants and matrices with coefficients, and u_t is a vector of reduced form residuals with the variance-covariance matrix, Σ. A_0 is a lower triangular matrix given by the Cholesky decomposition of Σ and ε_t is a vector of structural shocks that are mutually independent and normalized to be of variance 1.

Second, we draw random samples of B and Σ from their posterior distributions. Using the non-informative Normal-Wishart family as the prior distribution, the posterior distributions of $vec(B)$ and Σ^{-1} respectively become $N(vec(\hat{B}), \hat{\Sigma} \otimes (X'X)^{-1})$, and $W(\hat{\Sigma}^{-1}/T, T)$, where $\hat{B}$ and $\hat{\Sigma}$ are OLS estimators, X is the matrix of the explanatory variables, and T is the sample size. Third, the structural shocks and matrix of contemporaneous relations among the endogenous variables are calculated from each draw. In this step, we randomly generate the orthogonal matrix Q, namely $Q'Q = I$. Using this matrix, (5.34) can be rewritten as

$$u_t = A_0 Q' Q \varepsilon_t = A \tilde{\varepsilon}_t. \tag{5.39}$$

Then,

$$E\left[A \tilde{\varepsilon}_t \tilde{\varepsilon}_t' A'\right] = E\left[A_0 Q' Q \varepsilon_t \varepsilon_t' Q' Q A_0'\right] = A_0' A_0 = \Sigma \tag{5.40}$$

Therefore, we construct a new set of structural shocks $\tilde{\varepsilon}_t$ and contemporaneous relations A while maintaining the variance-covariance structure. Some numbers of the IRFs can be calculated from the set of (B, Σ) by generating a Q matrix randomly. As explained by Peersman (2005) and Pappa (2009), the present study generates a Q matrix using the following procedure. In a four-variable VAR model, we use a 4×4 Givens matrix Q_{12}, Q_{13}, Q_{14}, Q_{23}, Q_{24}, and Q_{34}, where $Q_{ij}(\theta)$ is a matrix with $\cos(\theta)$ in the (i, i) and the (j, j) elements, $-\sin(\theta)$ in the (j, i) element, and $\sin(\theta)$ in the (i, j) element. The diagonal element of this matrix is one, and the off-diagonal elements are equal to zero. Then, the Q matrix is defined as

$$Q = Q_{12}(\theta_1) Q_{13}(\theta_2) Q_{14}(\theta_3) Q_{23}(\theta_4) Q_{24}(\theta_5) Q_{34}(\theta_6)$$

where $\theta_k, k = 1, \cdots, 6$ are drawn randomly from a uniform distribution $U(0, 360)$.

Finally, the IRFs are calculated based on each draw (B, Σ, A). If they satisfy the sign restrictions in Table 5.2, they are candidates of valid IRFs and are reserved; otherwise, they are discarded. We repeat this process until we have 300 valid IRFs as our final samples.

5.3.2 Data and specification

The endogenous variables in our VAR model are the log of the nominal effective exchange rate (NEER), real exports, and the indices of all industry activity (IIA), which is regarded as a proxy of GDP, and inflation in CPI (all items, less fresh food).[4] The data are all monthly and the sample period is 1990M1–2013M5 (see Figure 5.2). Except for the NEER, the series are seasonally adjusted.[5] In the actual estimation, because an increase in the NEER implies depreciation, we multiply it by -1 to make it consistent with the specification presented in the theoretical model. Hereafter, the term "output" indicates the IIA.

The estimated system contains linear and quadratic time trends as well as a constant term and a consumption tax dummy (1997M4). By incorporating time trends into the system, we detrend the data series to be consistent with the DSGE model presented above. In addition, it is estimated in levels because taking a difference may lose important information contained in the original series, as pointed out by Sims et al. (1990). The number of lags is chosen to be three as suggested by the Akaike information criterion. In this framework, based on the sign restrictions derived from the theoretical model, we identify four types of structural shocks, namely risk premium, foreign demand, aggregate supply, and aggregate demand ($\varepsilon_t' = [\varepsilon_t^{risk}\ \varepsilon_t^{foreign}\ \varepsilon_t^{supply}\ \varepsilon_t^{demand}]$).

[4]Data on the NEER and real exports are downloaded from the Bank of Japan website. Data on the CPI and IIA are from the Ministry of Internal Affairs and Communication and the Ministry of Economy, Trade, and Industry, respectively. These data are taken from the Nikkei NEEDS-Financial QUEST database.

[5]With respect to real exports and the IIA, seasonally adjusted series can be obtained from the data source. By contrast, because the series of the CPI–especially before 2000–is not seasonally adjusted in this data source, we perform a seasonal adjustment using X-12-ARIMA.

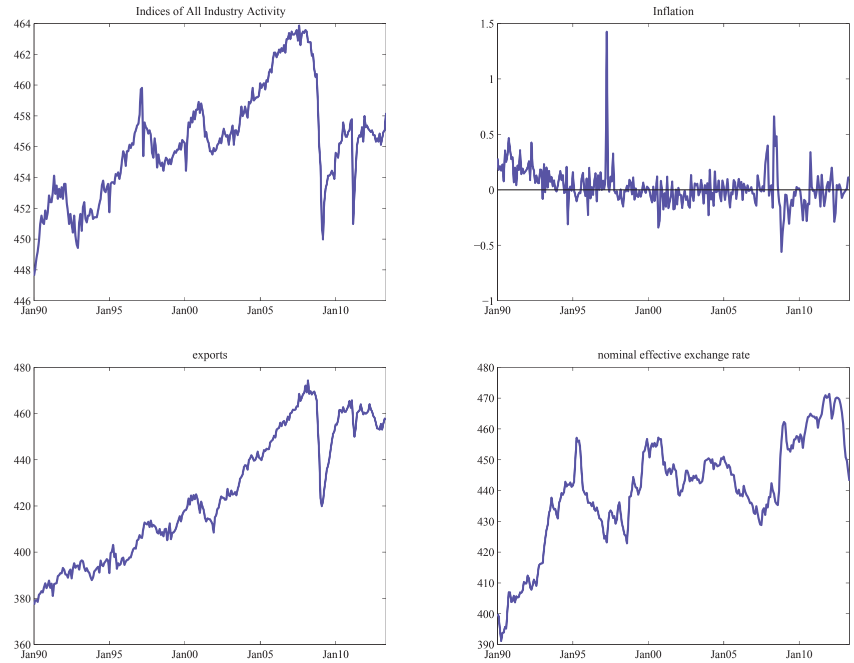

Figure 5.2: Data

Note: Except for inflation rate, the series are taken to the logarithm and multiplied by 100. The inflation rate is defined as the log first difference of consumer price index (all items, less fresh food).

5.4 Empirical results

5.4.1 Impulse response function

Figure 5.3 depicts the estimated IRFs. The solid blue lines and shaded areas indicate the median of sampled IRFs and 68 percent credible intervals, respectively. In addition, we plot the IRFs that are the closest (in terms of minimizing the sum of the squares of differences) to the median responses among those obtained in each admissible rotation, as shown by the red dotted line. As pointed out by Fry and Pagan (2011) and Inoue and Kilian (2011), the median responses in the sign-restricted VAR model summarize the information from different structural models because each IRF is computed based on a different rotation matrix. In other words, the response that fully corresponds to the median might not exist in the set of admissible structural

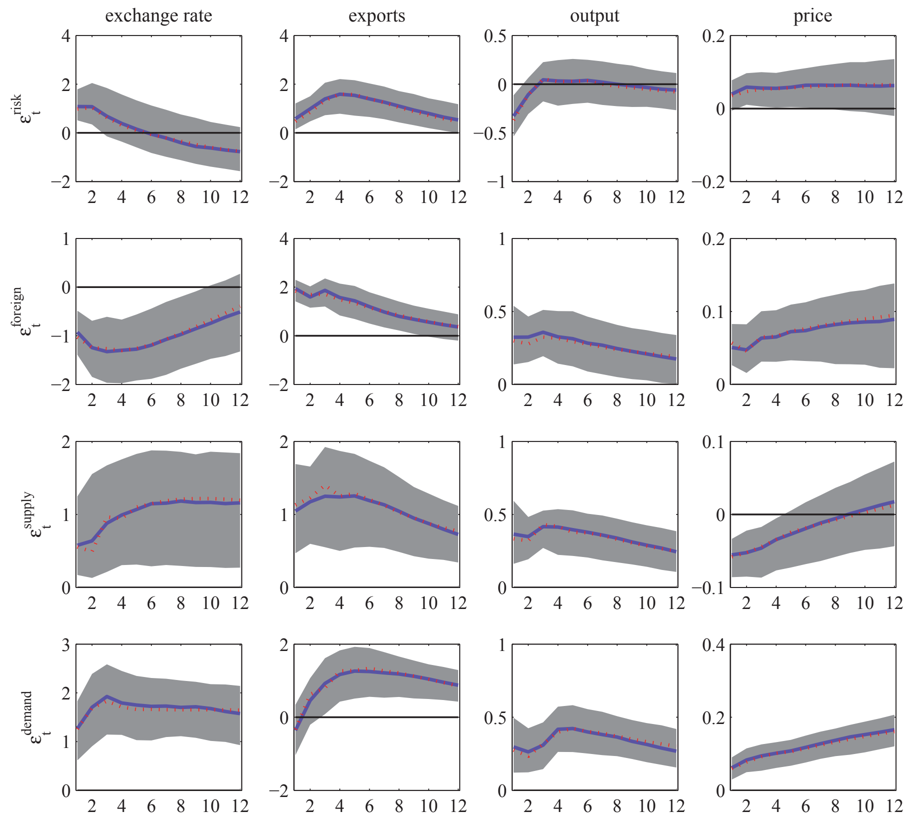

Figure 5.3: Impulse Response Function (Empirical model)

Note: This figure shows the estimated IRFs. The solid blue lines and shaded areas indicate the median of sampled IRFs and 68% credible intervals, respectively. In addition, the red dotted line indicates the IRFs that are the closest (in terms of minimizing the sum of the squares of differences) to the median responses among those obtained in each admissible rotation.

models, and thus the inferences using the median response might fail to lead to correct results. However, in our case, this problem is deemed to be less serious because the median responses and the closest ones share similar dynamics.

As seen in Figure 5.3, we confirm that the IRFs follow the sign restriction. The responses of price are described as the accumulated response of inflation. It is thus worthwhile exploring the responses for which no restrictions are imposed. First, price responds to risk premium shock positively and significantly in the medium run. This finding is consistent with the response observed in the theoretical prediction. Second, demand shock increases export quantities in the medium and long run, as expected by government spending shock in the theoretical model.

The estimated series of the structural shock, which is the median values of the sampled structural shocks, are plotted in Figure 5.4. At first glance, we can observe the large negative shocks in risk premium and foreign demand during the period when the Lehman shock hit, implying that this external shock appreciated the exchange rate and decreased export quantities in the presented analysis. Moreover, the large negative shocks in supply and demand as well as foreign demand are observed in March 2011, and this is considered to reflect the influence of the Great East Japan Earthquake. In addition, demand shock takes positive values persistently after December 2012 (except for March 2013). Although this period is at the end of our sample, it might capture the effects of "Abenomics", the economic recovery policy implemented by Prime Minister Shinzo Abe.

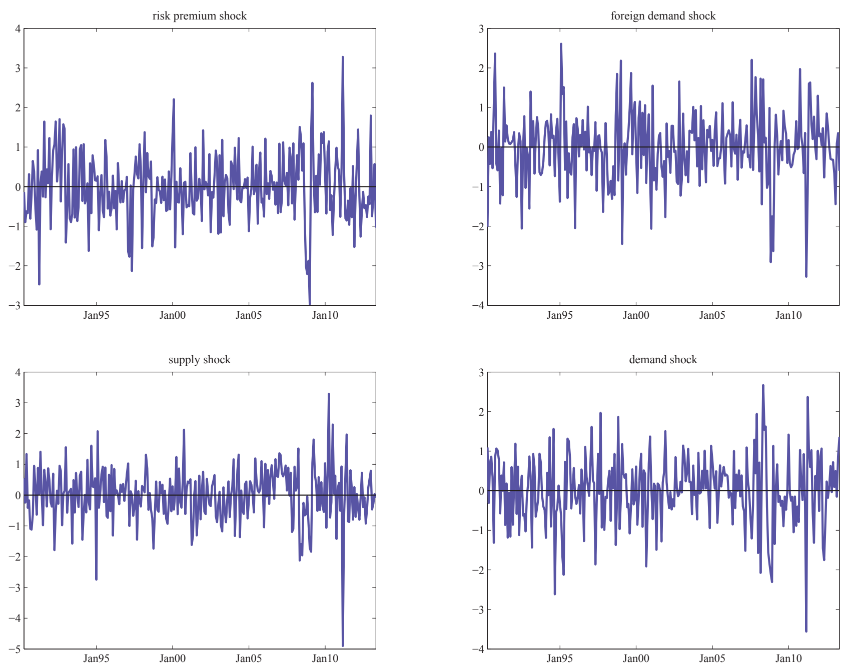

Figure 5.4: The series of structural shocks

5.4.2 Forecast error variance decomposition

The results of the FEVD are presented in Figure 5.5. The estimated time horizon is 12 months as in the earlier IRF analysis. This figure shows the relative importance of each shock in terms of the variations in each variable. To overcome the problem mentioned above, we perform the FEVD for every admissible rotation and retain their medians. Therefore, the presented results indicate the median value across all decompositions.

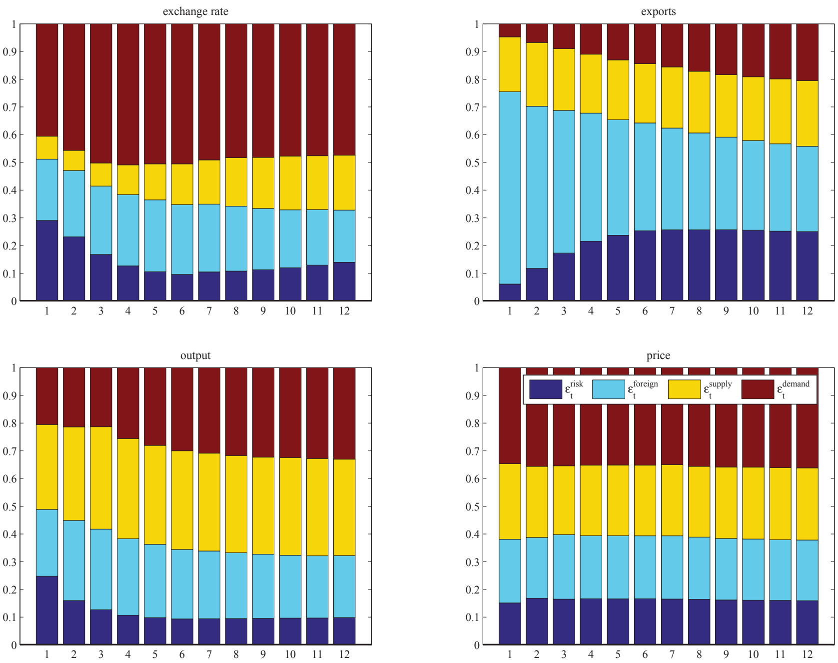

Figure 5.5: Forecast error variance decomposition

Note: This figure shows the results of the forecast variance decomposition. In this figure, colored regions indicate the ratio explained by risk premium, foreign demand, supply and demand shock in an order from bottom.

The primary finding is that external shocks (i.e., risk premium and foreign demand shocks) explain approximately 30% to 50% of the forecast error variances in output, as shown in the bottom right-hand figure. This result is somewhat larger than the values reported by Shioji et al. (2011). The difference might be considered to stem from differences in the estimation method used as well as the theoretical model used on which the empirical model is based.

Additionally, a large part of the variation in real exports can be explained by external shocks, especially foreign demand shock. Throughout the sample period, foreign demand shock continues to play a major role in the variations in real exports. The bottom right-hand figure shows that 60% or more of variations in price are explained by domestic shocks (i.e., supply and demand shocks). These results concur with the findings of Shioji et al. (2011). With regard to the nominal exchange rate, both external and domestic factors influence its variation. This finding implies that the exchange rate is determined by the relative economic situations in both the home and the foreign country.

The results of the FEVD thus show the importance of external shocks in the variations in Japanese output. Hence, it could be concluded that the Japanese economy is subject to external factors.

5.4.3 Historical decomposition

The results of the HD of output are shown in Figure 5.6. Similar to the FEVD, the presented results are the median values across all decompositions. In Figure 5.6, the solid blue line denotes the estimated output series, while the dotted red line denotes the decomposed series of each shock. Furthermore, the shaded areas show the recessions dated by the Cabinet Office of the Government of Japan.

First, we find that the supply shock series roughly traces the movement in output throughout the sample period. This finding implies that supply shock (e.g., variation in TFP) plays an important role in Japanese business cycles, as previously pointed by Hayashi and Prescott (2002) and Miyao (2006). In the next step, we investigate the sources of these business fluctuations by dividing the whole sample period into three sub-periods: the depression of the 1990s after the collapse of the bubble economy, the recovery since 2002, and the period after the Lehman shock.

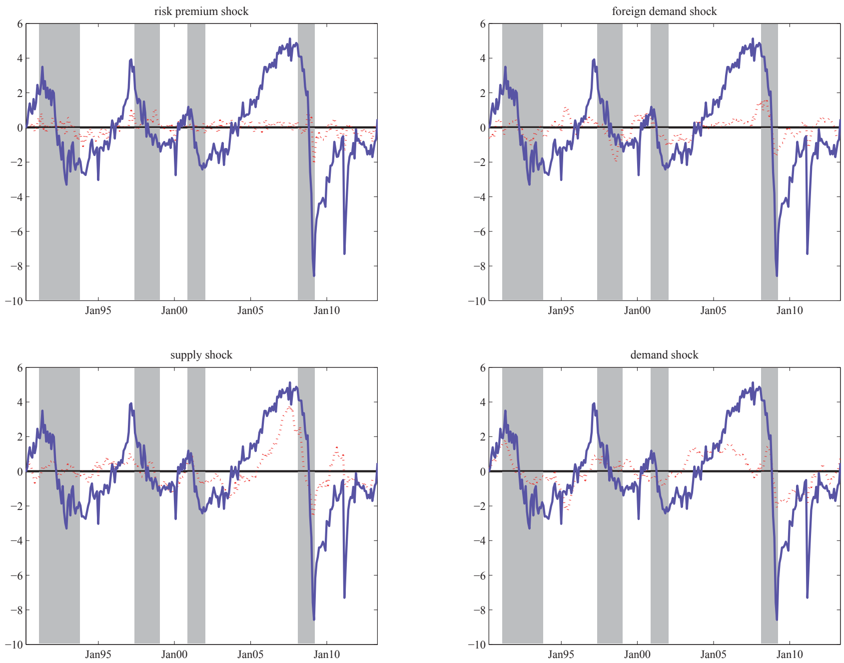

Figure 5.6: Historical decomposition in output

Note: This figure shows the result of the historical decomposition in output. A solid blue line and a red dotted line denote the estimated series of output and the portion explained by each structural shock. Furthermore, the shaded areas show the recessions dated by Cabinet Office, Government of Japan.

Two factors are mainly recognized as having caused the depression in the 1990s: a shortage of aggregate demand and a downturn in productivity. This study supports both these hypotheses. In the bottom chart on the right-hand side of Figure 5.6, the contribution of demand shock shows a downward trend in the recession that started in March 1991, which explains the movement in output. Additionally, we also observe the negative contribution of demand shock after September 1994. By contrast, supply shock does not seem to have played a large role in this period, as observed from the slight downward trend in the recession during the second half of the 1990s. However, this finding also suggests that there is no remarkable positive supply shock in the 1990s. These results imply that the long depression in the 1990s was caused by low productivity growth. Furthermore, foreign demand shock denotes the negative contribution around 1997, which reflects the onset of the Asian currency crisis. The role of risk premium shock on change in output is small, although a negative contribution is observed in the mid-1990s when the appreciation of the yen caused an economic depression.

Second, supply shock contributes greatly to the economic recovery beginning in 2002. This is consistent with the results of Inaba (2007). Moreover, as shown in the bottom chart on the right-hand side of Figure 5.5, demand shock also plays an important role in economic recovery, especially in the first half of this period. This economic recovery is often referred to as the export-led recovery. In our estimation, the upper trends in the contribution of foreign demand shock are observed in 2003 and from the second half of 2007 to the first half of 2008. Therefore, our results also support the view that risk premium shock has not yet played an important role in this phase.

Finally, we focus on the period after the Lehman shock that contains the Great East Japan Earthquake as well as the recession that stemmed from the collapse of Lehman Brothers. We present the results of the HD in output conditioning on the data before 2008 in order to clarify what happened within this period (Figure 5.7).

The results presented in this study indicate that all the structural shocks analyzed herein contribute to the great fall in output caused by the Lehman shock. Indeed, the contribution of supply shock had already shown a downward trend before the Lehman shock occurred (i.e., the recession had already started), as suggested by the shaded area in Figure 5.6. The other shocks then additionally contribute to the decline in output with lags after August 2008. In this phase, both of the external shocks investigated in this study led to large business fluctuations.

Moreover, our results show the recovery process following the Lehman shock. The trend in supply shock recovers persistently around February 2009, while the contributions of the other shocks show almost no influence (risk premium and demand shocks) or slow recovery (foreign demand shock). This result (i.e., the slow recovery in foreign demand) might be considered to reflect the fact that the shock originally occurred in foreign countries.

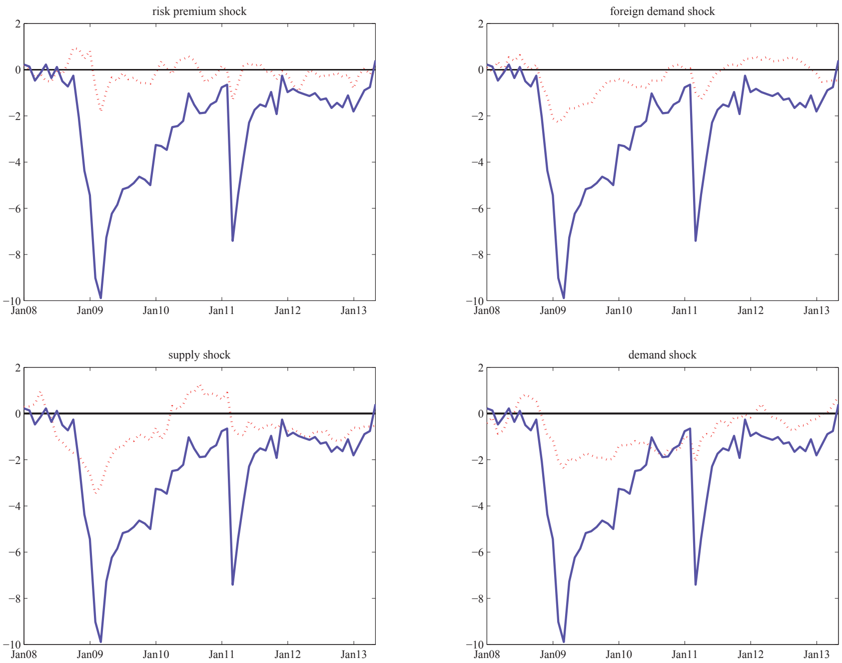

Figure 5.7: Historical decomposition in output conditioning on the data before 2008

Note: This figure shows the result of the historical decomposition in output conditioning on the data before January 2008. A solid blue line and a red dotted line denote the estimated series of output and the portion explained by each structural shock.

In March 2011 when the Great East Japan Earthquake struck, a sharp decline in output is observed. Similar to the Lehman shock, all the structural shocks assessed herein negatively affect output at this time. In particular, the supply shock indicates a negative contribution that might reflect the destruction of the supply chain after the earthquake occurred. Moreover, it is also observed that the fall in foreign demand bottomed out several months after the earthquake. This result might reflect the influence of the subsequent disaster at the Fukushima nuclear power plant. By contrast, demand shock shows a positive contribution after the earthquake, which might capture the reconstruction demand in the disaster zone. However, risk premium shock seems to have only a temporary effect, while the positive contribution of demand shock at the end of the sample period might capture the effects of Abenomics.

Summarizing the results of the HD, the long depression throughout the 1990s was caused by negative demand shock and stagnant TFP growth, while demand and supply shock in addition to foreign demand shock played important roles in the recovery during the 2000s. Furthermore, all the structural shocks analyzed in this study contributed to the great drop in output following the Lehman shock and Great East Japan Earthquake. Therefore, we can conclude that business fluctuations were caused by different shocks in each period and that the role of external shocks has grown in recent years.

5.5 Conclusion

This study has investigated the sources of Japanese business fluctuations by taking account of external shocks such as risk premium and foreign demand shock. For this purpose, an NK open macroeconomic model was first constructed in order to ascertain the features of each structural shock. Then, we estimated the sign-restricted VAR model based on theoretical IRFs, and conducted FEVD and HD as well as IRF analysis. From these analyses, the effects of external shocks were evaluated qualitatively and quantitatively.

The results of the FEVD reveals that 30% to 50% of the forecast error variances in output can be explained by external shocks; these results support the well-recognized notion that the Japanese economy is greatly influenced by external factors. This finding confirms the importance of taking account of external shocks when examining the Japanese economy.

In addition, the results of the HD allow us to draw four main conclusions. First, throughout the sample period, supply shock explains the variation in output more than any other shock. Second, the long depression in the 1990s was caused by a combination of negative demand shock and the absence of positive supply shock. The contribution of supply shock in the 1990s was almost nonexistent. Third, the economic recovery since 2002 has stemmed from an increase in foreign and domestic demand, and an improvement in productivity. Finally, the factors behind the recent two rapid falls–namely, the Lehman shock and Great East Japan Earthquake–included both domestic and external shocks. The two external shocks played an important role in this final phase.

These findings are somewhat consistent with the results presented by previous studies. The finding that both external and domestic shocks contribute to the reduction in output at the time of the Lehman shock is similar to that stated by Shioji and Uchino (2012).[6] Compared with other previous studies, our analysis has the advantage of identifying structural shocks by using theoretical restrictions based on the DSGE model. In addition, this study reveals the macroeconomic role of each shock at the point of the Great East Japan Earthquake in the VAR analysis framework by using the latest monthly data.

We also showed that the role of external shocks is larger in recent business fluctuations, especially the period after the Lehman shock. This finding implies that the Japanese economy has been increasingly affected by opening the country's borders to global trade partners. Therefore, future analyses based on an open economy framework in Japan are crucial for verifying or predicting the effects of economic policy.

[6] Shioji and Uchino (2012) conclude that the great decline in exports and production in the Japanese automobile industry cannot be explained only by a large external shock.

Bibliography

[1] Adolfson, M., Laséen, S., Lindé, J., Villani, M., 2007. Bayesian estimation of an open economy DSGE model with incomplete pass-through. Journal of International Economics 72(2), 481-511.

[2] Auerbach, A.J., Gorodnichenko, Y., 2012. Measuring the output responses to fiscal policy. American Economic Journal: Economic Policy 2012, 4(2), 1-27.

[3] Baxter, M., King, R.G., 1993. Fiscal policy in general equilibrium. American Economic Review 83(3), 315-334.

[4] Bayoumi, T., 2001. The morning after: explaining the slowdown in Japanese growth in the 1990s. Journal of International Economics 53(2), 241-259.

[5] Blanchard, O.J., Perotti, R., 2002. An empirical characterization of the dynamic effects of changes in government spending and taxes on output. Quarterly Journal of Economics 117(4), 1329-1368.

[6] Braun, R.A., Shioji, E., 2007. Investment specific technological changes in Japan. Seoul Journal of Economics 20(1), 165-200.

[7] Braun, R.A., Waki, Y., 2006. Monetary policy during Japan's lost decade. The Japanese Economic Review 57(2), 324-344.

[8] Calvo, G.A., 1983. Staggered prices in a utility-maximizing framework. Journal of Monetary Economics 12(3), 383-398.

[9] Campbell, J.Y., Mankiw, N.G., 1989. Consumption, income and interest rates: Reinterpreteing the time series evidence. NBER Macroeconomics Annual 1989(4), 185-246.

[10] Canova, F., Gambetti, L., 2009. Structural changes in the US economy: Is there a role for monetary policy? Journal of Economic Dynamics and Control 33(2), 477-490.

[11] Christiano, L., Eichenbaum, M., Evans, C.L., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. Journal of Political Economy 113(1), 1-45.

[12] Christiano, L., Eichenbaum, M., Rebelo, S., 2011. When is the government spending multiplier large? Journal of Political Economy, 119(1), 78-121.

[13] Colciago, A., 2011. Rule-of-thumb consumers meet sticky wages. Journal of Monetary, Credit and Banking 43(2)-(3), 325-353.

[14] Dedola, L., Neri, S., 2007. What does a technology shock do? A VAR analysis with model-based sign restrictions. Journal of Monetary Economics 54(2), 512-549.

[15] Eggertsson, G.B., 2011. What fiscal policy is effective at zero interest rate? NBER Macroeconomic Annual 2010, University of Chicago Press.

[16] Fazzari, S.M., Morley, J., Panovska, I., 2013. State-dependent effects of fiscal policy. UNSW Australian School of Business Research Paper No. 2012-ECON27B.

[17] Fisher, J.D.M., Peters, R., 2010. Using stock returns to identify government spending shocks. Economic Journal 120(544), 414-436.

[18] Franta, M., 2011. Identification of monetary policy shocks in Japan using sign restrictions within the TVP-VAR framework. IMES Discussion Paper No. 2011-E-13.

[19] Fry, R., Pagan, A., 2011. Sign restrictions in structural vector autoregressions: a critical review. Journal of Economic Literature 49(4), 938-960.

[20] Fujiwara, I., 2006. Evaluating monetary policy when nominal interest rates are almost zero. Journal of the Japanese and International Economies 20(3), 434-453.

[21] Fukuda, S., Yamada, J., 2011. Stock price targeting and fiscal deficit in Japan: Why did the fiscal deficit increase during Japan's lost decades? Journal of the Japanese and International Economies 25(4), 447-464.

[22] Galí, J., López-Salido, D., Vallés, J., 2007. Understanding the effects of government spending on consumption. Journal of the European Economic Association 5(1), 227-270.

[23] Gerba, E., Hauzenberger, K., 2013. Estimating US fiscal and monetary interactions in a time varying VAR. University of Kent School of Economics Discussion Papers, KDPE 1303.

[24] Geweke, J., 1992. Evaluating the accuracy of sampling-based approaches to the calculation of posterior moments. J.M. Bernardo, J.O. Berger, A.P. Dawid, A.F.M. Smith (Eds.), Bayesian Statistics, vol. 4Oxford University Press, New York (1992), 169-188.

[25] Gupta-Kapoor, A., Ramakrishnan, U., 1999. Is there a J-curve. a new estimation for Japan. International Economic Journal 13(4), 71-79.

[26] Hatano, T., 2004. The macroeconomic effects of deficit. Financial Review 74, 65-91 (In Japanese).

[27] Hayashi, F., Koeda, J., 2013. A regime-switching SVAR analysis of quantitative easing. CARF Working paper, CARF-F322.

[28] Hayashi, F., Prescott, E.C., 2002. The 1990s in Japan: A lost decade. Review of Economic Dynamics 5(1), 206-235.

[29] Honda, Y., Kuroki, Y., Tachibana, M., 2007. An injection of base money at zero interest rates: Empirical evidence from the Japanese experience 2001-2006. OSIPP Discussion Paper, 07-08.

[30] Hsing, H-M., 2005. Re-examination of J-curve effect for Japan, Korea, Taiwan. Japan and the World Economy 17(1), 43-58.

[31] Iiboshi, H., Nishiyama, S., Watanabe, T., 2008. An estimated dynamic stochastic general equilibrium model for the Japanese economy: A Bayesian analysis. Mimeo.

[32] Inaba, M., 2007. Business cycle accounting for the Japanese economy using the parameterized expectations algorithm. RIETI discussion paper series 07-E-061.

[33] Inoue, A., Kilian, L., 2011. Inference on impulse response functions in structural VAR model. CEPR Discussion Papers 8419.

[34] Inoue, T., Okimoto, T., 2008. Were there structural breaks in the effects of Japanese monetary policy? Re-evaluating policy effects of the lost decade. Journal of the Japanese and International Economies 22(3), 320-342.

[35] Iwata, Y., 2009. Fiscal policy in an estimated DSGE model of Japanese economy: Do non-Ricardian households explain all? ESRI Discussion Paper Series No.216.

[36] Kawai, M., Takagi, S., 2009. Why was Japan hit so hard by the global financial crisis? ADBI working paper series No. 153.

[37] Kim, C-J., 2004. Markov-switching models with endogenous explanatory variables. Journal of Econometrics 122(1), 127-136.

[38] Kimura, T., Nakajima, J. 2013. Identifying conventional and unconventional monetary policy shocks: A latent threshold Approach. Bank of Japan Working Paper Series, No. 13-E-7.

[39] Kohara, M., Horioka, C.Y., 2006. Do borrowing constraints matter? An analysis of why the permanent income hypothesis does not apply in Japan. Japan and the World Economy 18(4), 358-77.

[40] Kuttner, K.N., Posen, A.S., 2002. Fiscal policy effectiveness in Japan. Journal of the Japanese and International Economies 16(4), 536-58.

[41] Leeper, E.M., Traum, N., Walker, T.B., 2011. Clearing up the fiscal multiplier morass. NBER working paper series No. 17444.

[42] Lubik, T., Schorfheide, F., 2005. A bayesian look at new open economy macroeconomics. NBER Macroeconomics Annual 2005 20, 312-374.

[43] Mertens, K., Ravn, M.O., 2010. Measuring the impact of fiscal policy in the face of anticipation: a structural VAR approach. Economic Journal 120(544), 393-413.

[44] Miyao, R., 2002. The effects of monetary policy in Japan. Journal of Money, Credit, and Banking 34(2), 376-392.

[45] Miyao, R., 2006. Nihon keizai no hendou youin: Seisansei shock no yakuwari. BOJ working paper series No. 06-J-1. (In Japanese)

[46] Mountford, A., Uhlig, H., 2009. What are the effects of fiscal policy shock? Journal of Applied Econometrics 24(6), 960-992.

[47] Morita, H., 2013a. The effects of anticipated fiscal policy shock on macroeconomic dynamics in Japan. Global COE hi-stat discussion paper series No. 219.

[48] Morita, H., 2013b. External shocks and Japanese business cycles: Evidence from a sign-restricted VAR. Global COE hi-stat discussion paper series No. 277.

[49] Nakajima, J., 2011. Monetary policy transmission under zero interest rates: an extended time-varying parameter vector autoregression approach. The B.E. Journal of Macroeconomics 11(1), Article 32.

[50] Nakajima, J., Kasuya, Munehisa., Watanabe, T., 2011. Bayesian analysis of time-varying parameter vector autoregressive model for the Japanese economy and monetary policy. Journal of the Japanese and International Economies 25(3), 225-245.

[51] Nakazawa, M., 2002. Meimoku kinri to keizai doukou. Financial Review 66, 43-67. (In Japanese)

[52] Nakazawa, M., Onishi, S., Harada, Y., 2002. Zaisei kinyuu seisaku no kouka. Financial Review 66, 19-42. (In Japanese)

[53] Ogawa, K., Takenaka, H., Kuwana, Y., 1986. Saikin no nihon ni okeru chochiku syouhi pattern ni tuite, Financial Review 2. (In Japanese)

[54] Owyang, M.T., Ramey, V.A., Zubairy, S., 2013. Are government spending multipliers greater during periods of slack? Evidence from 20th century historical data. NBER working paper series No. 18769.

[55] Pappa, E., 2009. The effects of fiscal shocks on employment and the real wage. International Economic Review 50(1), 217-244.

[56] Peersman, G., 2005. What caused the early millennium slowdown? Evidence based on vector autoregressions. Journal of Applied Econometrics 20(2), 185-207.

[57] Ramey, V.A., 2011. Identifying government spending shocks: It's all in the timing. Quarterly Journal of Economics 126(1), 1-50.

[58] Schmit-Grohe, S., Uribe, M., 2006. Comparing two variants of Calco-type wage stickiness. NBER working paper series No. 12740.

[59] Shioji, E., 2000. Identifying monetary policy shocks in Japan. Journal of the Japanese and International Economies 14(1), 22-42.

[60] Shioji, E., Uchino, T., 2012. External shocks and Japanese business cycles: Impact of the "Great Trade Collapse" on the automobile industry. Center on Japanese Economy and Business (Columbia University) Working Paper Series No. 300.

[61] Shioji, E., Vu, T.K., Takeuchi, H., 2011. Effects of external and fiscal policy shocks in Japan: evidence from an open economy DSGE model with partial exchange rate pass-through. TCER working paper series J-4. (In Japanese)

[62] Sims, C.A., Stock, J., Watson, M., 1990. Inference in linear time series models with some unit roots. Econometrica 58(1), 113-144.

[63] Sugo, U., Ueda, K., 2008. Estimating a dynamic stochastic general equilibrium model for Japan. Journal of the Japanese and International Economies 22(4), 476-506.

[64] Tagkalakis, A., 2008. The effects of fiscal policy on consumption in recessions and expansions. Journal of Public Economics 92(5)-(6), 1486-1508.

[65] Tenhofen, J., Wolff, G.B., 2010. Does anticipation of government spending matter? the role of (non-) defense spending. Bonn econ discussion papers No.12.

[66] Uhlig, H., 2005. What are the effects of monetary policy on output? Results from an agnostic identification procedure. Journal of Monetary Economics 52(2), 381-419.

[67] Watanabe, T., Yabu, T., Ito, A., 2011. Estimation of fiscal multiplier based on institution information. Ihori, T. Fiscal Policy and Social Insurance, Economic and Social Research Institute, Cabinet Office, Government of Japan (in Japanese).