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Abstract

This paper investigates the dynamics of oil consumption, oil price, and economic activity during the period from 1986 to 2008 in Japan. In the first step, we employ residual-based tests for cointegration with a structural change and find a stable relationship among these economic variables. Moreover, we confirm that a structural change occurred after the collapse of the bubble economy. In the second step, we find that before and after this collapse, the price elasticity of oil demand decreased approximately 0.1 and the income elasticity of oil demand decreased approximately 0.4. In the third step, we use the lag augmented-vector autoregression method, finding that oil consumption does not Granger cause economic growth before the collapse. However, oil consumption Granger causes economic growth from after it. These empirical results have implications for the environmental and economic problems that Japan will face in the future.

Keywords: Oil consumption; Oil price; Economic activities; Structural change; LA-VAR
JEL classification codes: C22; Q43

1 Introduction

In June, 2009, the Japanese government announced that by 2020 it would aim to reduce greenhouse gas (GHG) emissions by 15 percent from 2005 levels as a middle-term target. Soon after, in August 2009, the government expressed the following vision about GHG emissions. "By 2050 Japan will try to cut greenhouse gas emissions by 80 percent level as a long-term target". While these targets will contribute to reducing energy consumption, they place an additional financial burden on the industrial and household sectors.[1] Furthermore, the government must consider the feasibility of these targets given the relationship between the environmental problem and economic activity.

Under the Kyoto Protocol, which came into force in February 2005, developed countries agreed to cut their GHG emissions relative to the levels emitted in 1990. In this regard, Japan agreed to cut its emissions by 6 percent during the 2008–2012 period, principally by reducing carbon dioxide (CO_2) emissions generated by energy consumption. However, GHG emissions in fiscal year 2007 increased.[2] Moreover, Aichele and Felbermayr (2012) finds that while the commitments of the Kyoto Protocol aim to reduce the domestic CO_2 emissions of 40 countries by approximately 7 percent on average, the overall carbon footprint will not change.

[1] The government estimated the increase in the financial burden on households to achieve this goal to be 77,000 yen a year.

[2] GHG emissions in fiscal year 1990 and fiscal year 2007 were 1207.8 and 1374.3 Gg CO_2 equivalent, respectively. CO_2 emissions in fiscal year 1990 and fiscal year 2007 were 1143.2 and 1303.8 Gg CO_2 equivalent, respectively.

In recent years, the environmental problem including the issue of CO_2 emissions has become an issue that must be overcome immediately by all countries.[3] Nevertheless, economic factors might play a role in raising GHG emissions. For example, the asset price bubble economy in Japan collapsed in early 1992, resulting in a long recession. In the early 2000s, the IT bubble in the U.S. and subsequent yen depreciation also affected the Japanese economy, leading to an increase in Japanese production and exports. These factors may have been influencing factors in the rise in GHG emissions in 2007.

After the oil crisis in the 1970s, the Japanese government passed the Act on the Rational Use of Energy in 1979. This Act prescribes to take the measures required to rationalize the use of energy by factories, transportation, buildings, and machinery and equipment. However, it is necessary to allow production to increase in order to boost the domestic economy, resulting in energy consumption increases, too.

The relationship between economic activity and environmental protection has long been a hotly debated and well-researched issue. A large number of studies have investigated the relationship between oil price and economic activity in oil-importing countries (e.g., Darby, 1982; Hamilton, 1983, 1996, 2003; Burbidge and Harrison, 1984; Hooker, 1996; Bernanke et al., 1997; Lee et al., 2001; Cunado and Perez de Gracia, 2005; Jiménez-Rodríguez and Sanchez, 2005). For instance, Bernanke et al. (1997), Lee et al. (2001), and Hamilton and Herrera (2004) investigate the effect of monetary policy after oil price shocks, arguing that an increase in oil price negatively affects economic activity.

[3] Hunt and Ninomiya (2005) demonstrates that the Japanese primary energy consumption per capita expands by about 18 times between 1900 and 2000.

Many countries face problems in reconciling economic growth with improving environmental quality. This is an especially serious problem for Japan because it is the third-highest consumer of oil in the world after U.S. and China. The annual consumption of Japanese oil products is approximately 280 million kl, while 99.6 percent of its oil consumption depends on imports. Moreover, the import of oil products is approximately 234 million kl per annum in Japan.[4] Hence, although high oil consumption generates major CO2 emissions, reducing consumption in order to cut GHG emissions will have a significant influence on Japanese economic activities. Thus, the GHG emissions reduction target for Japan offers a serious challenge for the Japanese economy.

This study examines the long-run relationship and direction of its causality between oil consumption and economic activity in Japan. Before investigating the causality between these variables, we check for the existence of a stable relationship by employing the residual-based tests for cointegration with and without a structural change proposed by Gregory and Hansen (1996). Then, we estimate the price elasticity and income elasticity of oil demand. We also examine how a change in oil price or in prosperous economic conditions influences oil demand. Next, we consider the breakpoint and examine the causality by using the lag-augmented vector autoregression (LA-VAR) method developed Toda and Yamamoto (1995). The presented empirical results suggest that the energy demand function changed after the collapse of the bubble economy in Japan. This finding implies that we must examine the relationship

[4] Japan imports oil mainly from Saudi Arabia, the United Arab Emirates, Iran, and Qatar.

between oil consumption and economic growth when taking account of the structural change.

The remainder of the paper is organized as follows. In Section 2, we review previous studies of the relationship between oil demand and economic activities in Japan. We present the methodology and data for estimating the price and income elasticity of oil demand in Section 3. Section 4 examines the Granger causality between these economic variables. Section 5 concludes.

2 The relationship between oil demand and economic activities

Although no oil is produced in Japan, it remains important to the country's production activities. Therefore, Japan depends on imports for its oil resources. First, we describe the relation between oil consumption and certain economic variables. Table 1 reports the correlation coefficients between oil consumption and price level and between oil consumption and production[5], while Figures 1 and 2 present scatter diagrams that illustrate the relation between oil consumption and price level and between oil consumption and production, respectively.

A number of authors have examined the clear relationship between oil price and economic activity in Japan (Darby, 1982; Burbidge and Harrison, 1984; Jiménez-Rodríguez and Sanchez, 2005; Cunado and Perez de Gracia, 2005; Hanabusa, 2009). Darby (1982) estimates the long-run oil effect from 1957:Q1 to

[5] The sample period used in Table 1, Figure 1, and Figure 2 is 1986:1 to 2008:5.

1976:Q4 and provides a value of -0.191, suggesting that the fall in real income in this period led to an increase in the real oil price. Similarly, Burbidge and Harrison (1984) analyze the response of industrial production to a change in oil price from January 1961 to June 1982 by using a seven-variable VAR model, while Jiménez-Rodríguez and Sanchez (2005) use the same model to examine the response of GDP to an oil price shock from 1972:Q3 to 2001:Q4. Cunado and Ferez de Gracia (2005) assess the Granger causality between oil price and macroeconomic variables such as economic growth and inflation rate from 1975:Q1 to 2002:Q2. Finally, Hanabusa (2009) focuses on the level and variance of oil price and economic activity variables from July 2000 to March 2008.

In addition, authors have estimated the price elasticity and income elasticity of energy demand in Japan (Cooper, 2003; Hunt and Ninomiya, 2005; Yamaguchi, 2007), but only up until 2004. In order to bridge this gap in the literature, we analyze the price elasticity and income elasticity of energy demand until 2008 and estimate these values with a structural break.

The pioneering study of Granger causality between energy and income was carried out by Kraft and Kraft (1978). Subsequently, Erol and Yu (1987), Lee (2006), and Jinke et al. (2008) all explored the Granger causality between energy consumption and income in Japan. For example, Jinke et al. (2008) examine the causal relationship between coal consumption and GDP for 1980–2005. However, we investigate the causality between oil consumption and economic activity by using monthly data from 1986 to 2008.

[Insert Table 1 around here]

[Insert Figure 1 around here]

[Insert Figure 2 around here]

3 The price and income elasticity of oil demand

3.1 Methodology

Oil demand is expressed as:

$$Y_t = f(X_{1,t}, X_{2,t})$$

where oil demand (Y_t) is a function of oil price ($X_{1,t}$) and domestic income ($X_{2,t}$). In order to estimate the price elasticity and income elasticity of oil demand, we use the logarithmic function. The equation is as follows:

$$\ln Y_t = \beta_1 \ln X_{1,t} + \beta_2 \ln X_{2,t}$$

where β_1 denotes the price elasticity of oil demand and β_2 denotes the income elasticity of oil demand.

We consider single equation models that allow for cointegration with and without a structural change using three variables. First, we consider the standard model of cointegration with no structural change:

$$y_t = \alpha_1 + \beta_1 x_{1,t} + \beta_2 x_{2,t} + e_t, \quad t = 1, \dots, T, \quad (1)$$

where, α_1, β_1, and β_2 are unknown parameters. It is assumed that y_t, $x_{1,t}$, and $x_{2,t}$ are integrated of order one, or, *I*(1). If the disturbance term (e_t) is integrated of order zero, *I*(0), then the linear combination of y_t, $x_{1,t}$, and $x_{2,t}$ is a cointegrating relation. y_t denotes the logarithmic oil consumption, while $x_{1,t}$ and $x_{2,t}$ denote the logarithmic oil pric and Index of Industrial Production (IIP), respectively. We divide $x_{1,t}$ into $x_{1,a,t}$ or $x_{1,b,t}$. $x_{1,a,t}$ is the oil price in dollars and $x_{1,b,t}$ is that in yen. $x_{1,b,t}$ is a variable that considers the change in the exchange rate. β_1 denotes price elasticity and β_2 income elasticity. According to the conventional demand model, negative signs of β_1 and positive signs of β_2 should be expected.

Second, we employ the residual-based tests for cointegration with regime shift proposed by Gregory and Hansen (1996). We explain a brief description of the methodology. The explanation is given below. Before we explain the methodology, we define the dummy variable ($D_{t\tau}$) used to model a structural change:

$$D_{t\tau} = \begin{cases} 0 \ \ if \ t \le [\tau T] \\ 1 \ \ if \ t > [\tau T] \end{cases}$$

where the unknown parameter $\tau \in (0, 1)$ denotes the relative timing of the change point and the square brackets the integer part. The test statistic is computed for each breakpoint in the interval ([0.15T], [0.85T]). We use following the regime shift model type:

$$y_t = \alpha_{1,1} + \alpha_{1,2} D_{t\tau} + \beta_{1,1} x_{1,t} + \beta_{1,2} x_{1,t} D_{t\tau} + \beta_{2,1} x_{2,t} + \beta_{2,2} x_{2,t} D_{t\tau} + e_{\tau t}. \quad (2)$$

In Equation (2), $\alpha_{1,1}$ represents the intercept before the shift, $\alpha_{1,2}$ the change in the intercept at the time of the shift, $\beta_{1,1}$ and $\beta_{1,2}$ the cointegrating slope coefficient before the regime shift, and $\beta_{2,1}$ and $\beta_{2,2}$ the change in the slope coefficient at the time of the shift. $\beta_{1,1}$ and $\beta_{2,1}$ are the elasticity estimates before the structural change, while the size and sign of $\beta_{1,2}$ and $\beta_{2,2}$ are necessary to understand the effect of this change.

We employ the residual-based cointegration test (Engle-Granger type test) in order to test the null hypothesis of no cointegration relation among y_t, $x_{1,t}$, and $x_{2,t}$. The candidate cointegrating relation is estimated by carrying out an ordinary least squares (OLS) regression; a unit root test is conducted on the regression errors as follows:

$$\Delta\hat{e}_{t\tau} = \gamma_0\hat{e}_{t-1\tau} + \sum_{i=1}^{K}\gamma_i\Delta\hat{e}_{t-i\tau} + v_t,$$

where v_t is the disturbance term. The sign of hat (ˆ) denotes an estimated value and the *ADF*(τ) statistic is the *t*-statistic for the explanatory variable ($\hat{e}_{t-1\tau}$). ADF^* denotes the smallest value of *ADF*(τ).

3.2 Data

We use Japanese monthly oil consumption, oil price, and the IIP from 1986:1 to 2008:5. In similar previous studies, Hunt and Ninomiya (2005) use Japanese annual data series from 1987 to 2001, while Yamaguchi (2007) uses Japanese quarterly data from 1986:Q1 to 2004:Q4. However, we select a longer sample period that stops before the outbreak of the recent global financial crisis and compare the empirical result of previous research and our result.

The data source for oil consumption and oil price is the Energy Information Administration (February 2009). [6] To convert oil price into Japanese yen, we use the exchange rate provided by the Bank of Japan. [7] The data source of the IIP is the Ministry of Economy, Trade, and Industry (February 2009).[8] Note also that the data on oil consumption and the IIP are seasonally adjusted and that the logs of these data are employed for the empirical analysis. We show these data in Figures 3–5.

[Insert Figure 3 around here]

[Insert Figure 4 around here]

[Insert Figure 5 around here]

3.3 Empirical results

We analyze the cointegration of y_t, $x_{1,t}$ ($x_{1,a,t}$ or $x_{1,b,t}$), and $x_{2,t}$. Before proceeding to the cointegration tests, we must first check the stationarity of y_t, $x_{1,t}$, and $x_{2,t}$. For this purpose, the

[6] Homepage Address (Energy Information Administration): http://www.eia.doe.gov/

[7] Homepage Address (Bank of Japan): http://www.boj.or.jp/

[8] IIP mainly reflect the supply of manufacturing and utility industries. Homepage Address (Ministry of Economy, Trade and Industry): http://www.meti.go.jp/

augmented Dickey–Fuller (1979, 1981) test (ADF test) and Phillips–Perron (1988) test (PP test) are employed. The lag length of the ADF regression is selected by using the Akaike information criterion (AIC). The results of the unit root test are reported in Table 2. This test fails to reject the null hypothesis of a unit root for each variable in levels. By contrast, the null hypothesis of a unit root is rejected for each variable in first differences. Thus, these variables are confirmed to be *I(1)*.

The estimated results of the standard cointegration model and regime shift model are reported in Tables 3 and 4. While Table 4 presents the empirical results following a change in the exchange rate, Table 3 does not. The ADF statistic shows that the null hypothesis of no cointegration without the structural change is accepted. However, cointegration may exist with the structural change, while the possible breakpoint in the regime shift model is 1992:2. Regardless of the change in the exchange rate, this result is the same. This breakpoint represents the time of the collapse of the bubble economy in Japan (see also Yamaguchi (2007), who shows that the breakpoint is 1993:Q2). These empirical results imply that a long-run relationship among oil consumption, oil price, and the IIP exists with the regime shift. We can thus consider that the structural change occurred after the collapse of the bubble economy.

Next, we estimate the cointegration coefficients by using the dynamic OLS proposed by Saikkonen (1991) and Stock and Watson (1993). The dynamic OLS estimation of the regime shift model type is as follows:

$$y_t = \alpha_{1,1} + \alpha_{1,2}D_{t\tau} + \beta_{1,1}x_{1,t} + \beta_{1,2}x_{1,t}D_{t\tau} + \beta_{2,1}x_{2,t} + \beta_{2,2}x_{2,t}D_{t\tau}$$

$$+ \sum_{j=-p}^{p} \gamma_{1,j}\, \Delta x_{1,t-j} + \sum_{j=-p}^{p} \gamma_{2,j}\, \Delta x_{2,t-j} + e_{\tau t}.$$

The results are shown in Table 3. The regression model of the regime shift type with the oil price in dollars is

$$y_t = 4.968 - 0.004x_{1,a,t} + 0.787x_{2,t}, \quad \text{(1986:1-1992:1)},$$

$$y_t = 7.568 - 0.084x_{1,a,t} + 0.287x_{2,t}, \quad \text{(1992:2-2008:5)}. \qquad (4)$$

From Table 4, the regression model of the regime shift type with the oil price in Japanese yen is

$$y_t = 4.918 + 0.025x_{1,b,t} + 0.752x_{2,t}, \quad \text{(1986:1-1992:1)},$$

$$y_t = 7.876 - 0.086x_{1,b,t} + 0.311x_{2,t}, \quad \text{(1992:2-2008:5)}. \qquad (5)$$

From Equations (4) and (5), we find that the intercept coefficient is a higher value, that the slope coefficient of oil price changes from -0.004 to -0.084 or from 0.025 to -0.086 after the break, and that the slope coefficient of the IIP changes from 0.787 to 0.287 or from 0.752 to 0.311 after the break. Both price elasticity and income elasticity decrease. After February 1992, the reaction degree of oil consumption following a change in oil price (income) increases (decreases).

Cooper (2003) estimates the price elasticity of oil demand to be -0.357 during 1971–2000, [9] while Hunt and Ninomiya (2005) shows that the price elasticity and income elasticity of energy demand are -0.20 and 1.05 during 1887–2001, respectively. Similarly, Yamaguchi (2007) shows that the price

[9] The short-run price elasticity of oil demand is -0.071.

elasticity and income elasticity of energy demand are 0.043 and 1.076 during 1986:Q1–1993:Q1, respectively compared with -0.149 and 1.679 during 1993:Q2–2004:Q4. The price elasticity and income elasticity of energy demand presented herein are therefore smaller than those found in previous studies. However, the increase in the price elasticity of energy demand after the structural break is consistent with the findings of Yamaguchi (2007) (although the decrease in income elasticity after the break is not). One explanation of this result may be that the manufacturing industry aimed to reduce energy consumption, whereas other industries (e.g., the transport industry and service industry) did not. In addition, previous studies employ GDP or GNP, which include the production of the transport and service industries.

[Insert Table 2 around here]

[Insert Table 3 around here]

[Insert Table 4 around here]

4 The relationships between a change in oil consumption and economic activity

4.1 Methodology

In this subsection, we consider the LA-VAR model developed by Toda and Yamamoto (1995), which is applicable regardless of the integrated order or existence of cointegration among variables. It is assumed that an *n*-dimensional vector $\{g_t\}$ is generated by the following process:

$$g_t = \gamma_0 + \gamma_1 t + J_1 g_{t-1} + J_2 g_{t-2} +, \dots, + J_k g_{t-k} + u_t, \qquad t = 1, \dots, T, \qquad (6)$$

where *k* is the lag length, u_t is the vector of error terms with mean zero and variance-covariance matrix ($\sum_{u_t}$), and γ_0, γ_1, $J_1,\dots, J_k$ are matrices of parameters, respectively. We consider the following hypothesis:

$$H_o: f(\Phi) = 0, \qquad (7)$$

where Φ is the subset of parameters in Equation (6). To test this hypothesis, we estimate a VAR formulated in levels by using the OLS approach as follows:

$$g_t = \hat{\gamma}_0 + \hat{\gamma}_1 t + \hat{J}_1 g_{t-1} + \hat{J}_2 g_{t-2} +, \dots, + \hat{J}_p g_{t-p} + \hat{u}_t, \qquad (8)$$

where, *p*(=*k*+d_{max}) is the true lag length (*k*) plus the possible maximum integration order considered in the process (d_{max}) and $\hat{\gamma}_0$, $\hat{\gamma}_1$, $\hat{J}_1,\dots, \hat{J}_p$ are matrices of estimated parameters. However, d_{max} must not exceed length *k*. Since the true coefficients of $\hat{J}_1,\dots, \hat{J}_k$ are zero, it should be noted that the restriction (7) does not include them. Hence, Equation (8) is transformed as follows:

$$g_t = \hat{\Gamma}\tau_t + \hat{\Theta}x_t + \hat{\Psi}z_t + \hat{u}_t,$$

where $\hat{\Gamma} = (\hat{\gamma}_0,\ \hat{\gamma}_1)$, $\tau_t = (1, t)'$, $\hat{\Theta} = (\hat{J}_1,\dots,\ \hat{J}_k)$, $x_t = (g'_{t-1}, \dots, g'_{t-k})'$, $\hat{\Psi} = (\hat{J}_{k+1}, \dots, \hat{J}_p)'$, and $z_t = (g'_{t-k-1}, \dots, g'_{t-p})'$. In the matrix notation, this can be expressed as follows:

$$G' = \hat{\Gamma}T' + \hat{\Theta}X' + \hat{\Psi}Z' + \hat{U}',$$

where $T=(\tau_1,\dots,\tau_T)'$, $X=(x_1,\dots,x_T)'$, $Z=(z_1,\dots,z_T)'$, and so on. From the parameter estimates, $\hat{\Phi}=vec(\hat{\Theta})$, the Wald statistic, *W*, can be calculated as follows:

$$W=f(\hat{\Phi})'[\left(\frac{\partial f(\hat{\Phi})}{\partial\hat{\Phi}'}\right)\{\hat{\Sigma}_u\otimes(X'QX)^{-1}\}\left(\frac{\partial f(\hat{\Phi})}{\partial\hat{\Phi}'}\right)']^{-1}f(\hat{\Phi}), \qquad (9)$$

where $\hat{\Sigma}_u=T^{-1}\hat{U}'\hat{U}$, $Q=Q_\tau-Q_\tau Z(Z'Q_\tau Z)^{-1}Z'Q_\tau$, $Q_\tau=I_\tau-T(T'T)^{-1}T'$, and I_τ denotes the $T\cdot T$ identity matrix. When the null hypothesis is true, the Wald statistic has an asymptotic chi-square distribution with *m* degrees of freedom. The degrees of freedom with the Wald statistic (Equation (9)) can be expressed as follows:

$$m=rank\ (F(\cdot)), \qquad (10)$$

where $F(\cdot)=\partial f(\hat{\theta})/\partial\hat{\theta}'$ and $p\geq k+d$.[10]

4.2 Data and empirical results

In this subsection, we use the same data in order to analyze the Granger causality of y_t and $x_{2,t}$.[11] Given the time of the structural break presented in the previous subsection, we investigate the causality from 1986:1 to 1992:1 and from 1992:2 to 2008:5. The empirical results of Granger causality obtained based on the LA-VAR model are reported in Tables 4 and 5. From Equation (8), the VAR model with oil consumption, oil price and the IIP is as follows:

[10] $f(\cdot)$ is a twice continuously differentiable function with Equation (10) in the neighborhood of the true parameter.
[11] See Granger (1969).

$$y_t = \gamma_{0,1} + \gamma_{1,1}t + \sum_{q=1}^{p} J_{1,1}(q)\, y_{t-q} + \sum_{q=1}^{p} J_{1,2}(q)\, x_{1,t-q} + \sum_{q=1}^{p} J_{1,3}(q)\, x_{2,t-q} + u_{1,t},$$

$$x_{2,t} = \gamma_{0,2} + \gamma_{1,2}t + \sum_{q=1}^{p} J_{2,1}(q)\, y_{t-q} + \sum_{q=1}^{p} J_{2,2}(q)\, x_{1,t-q} + \sum_{q=1}^{p} J_{2,3}(q)\, x_{2,t-q} + u_{2,t}. \qquad (11)$$

The VAR model with oil consumption and the IIP is as follows:

$$y_t = \gamma_{0,1} + \gamma_{1,1}t + \sum_{q=1}^{p} J_{1,1}(q)\, y_{t-q} + \sum_{q=1}^{p} J_{1,3}(q)\, x_{2,t-q} + u_{1,t},$$

$$x_{2,t} = \gamma_{0,2} + \gamma_{1,2}t + \sum_{q=1}^{p} J_{2,1}(q)\, y_{t-q} + \sum_{q=1}^{p} J_{2,3}(q)\, x_{2,t-q} + u_{2,t}. \qquad (12)$$

From Equations (11) and (12), the null hypothesis is that $x_{2,t}$ does not Granger cause y_t as follows:

$$H_0{:}J_{1,3}(1) =, \ldots, = J_{1,3}(k) = 0.$$

The null hypothesis is that y_t does not Granger cause $x_{2,t}$ as follows:

$$H_0{:}J_{2,1}(1) =, \ldots, = J_{2,1}(k) = 0.$$

We choose the true lag length (*k*) by using the AIC to perform the LA-VAR analysis. The lag length before the structural break (1986:1–1992:1) is two for the three-variable model (oil consumption, oil price, and the IIP) and five for the two-variable model (oil consumption and the IIP). However, the lag length after the structural break (1992:2–2008:5) is five for both models. The length of the possible maximum integrated order (d_{max}) is one. Since the unit root test results in Table 1 indicate that d_{max} =1, we select this length. [12] The total lag length (*p*) adds *k* and 1 (*p*=*k*+1). In this case, the Wald statistic follows the chi-square distribution with k degrees of freedom.

Tables 5 and 6 show the test of Granger causality. We analyze two models by using oil price in dollars

[12] As reported by Toda and Yamamoto (1995), the length of d_{max} must not be greater than that of *k*.

and in yen (Table 5 and 6, respectively). From Table 5, we find that $x_{2,t}$ does not Granger cause y_t and that y_t does not Granger cause $x_{2,t}$ before the structural break. We thus find no Granger causality between oil consumption and the IIP, suggesting that the neutrality hypothesis holds and that Japan's oil conservation policy does not affect the IIP. [13] After the structural break, however, we see that y_t Granger causes $x_{2,t}$, while $x_{2,t}$ does not Granger cause y_t. This finding implies that oil consumption is a useful variable for predicting the economic growth rate after the collapse of the bubble economy in Japan. [14]

Table 6 shows that $x_{2,t}$ does not Granger cause y_t and that y_t does not Granger cause $x_{2,t}$ before the structural break. This empirical result is consistent with the results presented in Table 5. Therefore, we find no Granger causality between oil consumption and the IIP, implying that the neutrality hypothesis holds as before. [15] After the structural break, however, y_t Granger causes $x_{2,t}$. This result implies that oil consumption is again useful for predicting the economic growth rate after the collapse of the bubble economy.

Table 7 shows the test of Granger causality as to full sample. Both lag length of oil price in dollars and yen are five. From Table 7, we find that $x_{2,t}$ does not Granger cause y_t and that y_t does not Granger cause $x_{2,t}$. It is found that no Granger causality between oil consumption and the IIP, suggesting that the

[13] See Yu and Choi (1985) and Altinay and Karagol (2004).

[14] Erol and Yu (1987) show that energy consumption Granger causes GNP from 1950 to 1982 in Japan. Lee (2006) shows that GDP Granger causes energy consumption from 1960 to 2001 in Japan.

[15] See Yu and Choi (1985) and Altinay and Karagol (2004).

neutrality hypothesis holds from 1986:1 to 2008:5. However we find the Granger causality from the oil consumption to the IIP during 1992:2-2008:5 from Tables 5 and 6. Therefore, we provide the usefulness of considering a structural change when we investigate the causal relationship between oil consumption and IIP.

Han et al. (2004) find that an increasing energy supply results in stable economic growth in China. This finding may hold for Japan depending on the rate of oil consumption. Therefore, a change in oil consumption may play an informational role for the Japanese economy. Thus, we consider that decreasing oil consumption in order to reduce GHG emissions affects domestic economic activities.

[Insert Table 5 around here]

[Insert Table 6 around here]

[Insert Table 7 around here]

5 Conclusions

In this paper, we examined the dynamics of oil consumption, oil price, and the IIP by using monthly data over a 20-year period. The presented empirical results were then derived by applying the techniques developed by Gregory and Hansen (1996) and Toda and Yamamoto (1995). First, we found a long-run relationship among oil consumption, oil price, and the IIP with the structural change. Because the existence of the oil demand function was confirmed, we could then estimate price and income elasticity. Moreover, we showed that the shift timing among these variables was February 1992, that is when the bubble economy in Japan collapsed. Consequently, the stable relationship among oil consumption, oil price, and economic growth changed before and after the collapse of the bubble economy in Japan.

Second, we found that price elasticity decreased approximately 0.1 and income elasticity decreased approximately 0.4 before and after the structural change. Therefore, after February 1992, the decrease in oil consumption owing to the rise in oil price grew, whereas that owing to an income increase declined. Thus, the influence of the income change on oil consumption was large throughout the 20-year study period, but it decreased considerably after the collapse of the bubble economy.

Finally, we tested the Granger causality between oil consumption and the IIP by using the LA-VAR model. We found no Granger causality between oil consumption and the IIP before February 1992 but that oil consumption Granger caused the IIP thereafter. Because oil consumption is becoming an important variable for predicting economic activity, our results imply that policymakers must consider both environmental and economic problems in order to avoid both global warming and economic recession.

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Table 1: Correlation coefficient

	Oil consumption
Price level	0.723
Production	0.449

Note) Price level denotes the consumer price index and Production denotes the index of industrial production (logarithmic variables).

Table 2: Unit root test

	Variable		ADF	lag	PP
level	y_t	C	-2.521	4	-3.342*
		C/T	-1.954	4	-3.347
	$x_{1,a,t}$	C	-0.169	1	0.261
		C/T	-1.652	1	-1.841
	$x_{1,b,t}$	C	0.139	2	-0.281
		C/T	-1.607	1	-2.241
	$x_{2,t}$	C	-2.742	5	-2.018
		C/T	-3.175	5	-2.404
1st difference	y_t	C	-12.722**	3	-32.796**
		C/T	-12.959**	3	-52.860**
	$x_{1,a,t}$	C	-13.812**	0	-13.776**
		C/T	-13.857**	0	-13.855**
	$x_{1,b,t}$	C	-12.395**	1	-14.134**
		C/T	-12.514**	1	-14.236**
	$x_{2,t}$	C	-4.643**	4	-21.904**
		C/T	-4.651**	4	-21.884**

Note) y_t denotes the oil consumption, $x_{1,a,t}$ denotes the oil price (dollar indication), $x_{1,b,t}$ denotes the oil price (yen indication), and $x_{2,t}$ denotes the index of industrial production.

** denotes significant at 1% level and * denotes significant at 5% level.

The lag length is selected by AIC (Max lag=12).

Table 3: Estimation result (the oil price of the dollar indication)

Parameter	Without break	Regime shift	DOLD
$\alpha_{1,1}$	4.951**	5.220**	4.968**
	(0.519)	(0.199)	(0.226)
$\alpha_{1,2}$	-	2.315**	2.600**
	-	(0.602)	(0.716)
$\beta_{1,1}$	-0.088**	0.050*	-0.004
	(0.015)	(0.021)	(0.026)
$\beta_{1,2}$	-	-0.129**	-0.080*
	-	(0.024)	(0.028)
$\beta_{2,1}$	0.858**	0.696**	0.787**
	(0.119)	(0.048)	(0.058)
$\beta_{2,2}$	-	-0.405**	-0.499**
	-	(0.139)	(0.166)
break point	-	1992:2	1992:2
lead and lag	-	-	3
ADF	-2.537	-	-
ADF^*	-	-5.760*	-
lags	4	2	-

Note) Newy-West HAC standard errors (lag truncation=4) are in parentheses.

The ADF statistic shows residual-based tests for cointegration, the 1 % critical value is -4.35, the 5 % critical value is -3.78 (see Engle and Yoo, 1987).

The 1 % critical value of the regime shift model is -5.97 and the 5 % critical value is -5.50 (see Gregory and Hansen, 1996).

** denotes significant at 1 % level and * denotes significant at 5 % level.

The lag length of without break and regime shift model is selected by AIC (Max lag=6).

The lead and lag length of DOLS is selected by AIC (Max lag=5).

Table 4: Estimation result (the oil price of the yen indication)

Parameter	Without break	Regime shift	DOLD
$\alpha_{1,1}$	5.551**	4.811**	4.918**
	(0.361)	(0.214)	(0.177)
$\alpha_{1,2}$	-	3.087**	2.958**
	-	(0.525)	(0.561)
$\beta_{1,1}$	-0.105**	0.055*	0.025
	(0.012)	(0.017)	(0.017)
$\beta_{1,2}$	-	-0.137**	-0.111*
	-	(0.020)	(0.020)
$\beta_{2,1}$	0.849**	0.722**	0.752**
	(0.085)	(0.040)	(0.044)
$\beta_{2,2}$	-	-0.425**	-0.440**
	-	(0.124)	(0.137)
break point	-	1992:2	1992:2
lead and lag	-	-	1
ADF	-3.060	-	-
ADF^*	-	-6.190**	-
lags	4	2	-

Note) Newy-West HAC standard errors (lag truncation=4) are in parentheses.

The ADF statistic shows residual-based tests for cointegration, the 1 % critical value is -4.35, the 5 % critical value is -3.78 (see Engle and Yoo, 1987).

The 1 % critical value of the regime shift model is -5.97 and the 5 % critical value is -5.50 (see Gregory and Hansen, 1996).

** denotes significant at 1 % level and * denotes significant at 5 % level.

The lag length of without break and regime shift model is selected by AIC (Max lag=6).

The lead and lag length of DOLS is selected by AIC (Max lag=5).

Table 5: Granger causality

	1986:1-1992:1		1992:2-2008:5	
model type	three variables	two variables	three variables	two variables
lag length	k=2	k=5	k=5	k=5
$x_{2,t} \rightarrow y_t$	1.561	6.482	2.734	3.269
p-value	(0.458)	(0.262)	(0.435)	(0.659)
$y_t \rightarrow x_{2,t}$	0.108	0.679	9.645**	11.145*
p-value	(0.948)	(0.984)	(0.022)	(0.049)

Note) This table reports the result of Granger causality test with the oil price of the dollar indication.

Numbers in the table denote Wald test statistics.

y_t denotes the oil consumption and $x_{2,t}$ denotes the index of industrial production.

p-values are in parentheses.

** denotes that null hypothesis of Granger non-causality is rejected at 1 % significance level.

* denotes that null hypothesis of Granger non-causality is rejected at 5 % significance level.

Table 6: Granger causality

	1986:1-1992:1	1992:2-2008:5
model type	three variables	three variables
lag length	k=2	k=5
$x_{2,t} \rightarrow y_t$	1.840	4.989
p-value	(0.399)	(0.417)
$y_t \rightarrow x_{2,t}$	0.267	11.416*
p-value	(0.875)	(0.044)

Note) This table reports the result of Granger causality test with the oil price of the yen indication.

Numbers in the table denote Wald test statistics.

y_t denotes the oil consumption and $x_{2,t}$ denotes the index of industrial production.

p-values are in parentheses.

** denotes that null hypothesis of Granger non-causality is rejected at 1 % significance level.

* denotes that null hypothesis of Granger non-causality is rejected at 5 % significance level.

Table 7: Granger causality

	1986:1-2008:5			
	the oil price of the dollar indication		the oil price of the yen indication	
model type	three variables	two variables	three variables	two variables
lag length	k=5	k=5	k=5	—
$x_{2,t} \rightarrow y_t$	6.788	3.216	5.594	—
p-value	(0.237)	(0.667)	(0.348)	—
$y_t \rightarrow x_{2,t}$	7.603	6.812	6.597	—
p-value	(0.180)	(0.235)	(0.252)	—

Note) This table reports the result of Granger causality test with the oil price of the dollar and yen indication.

The empirical result of two variables as to the oil price of the yen indication is same as it of the dollar indication.

Numbers in the table denote Wald test statistics.

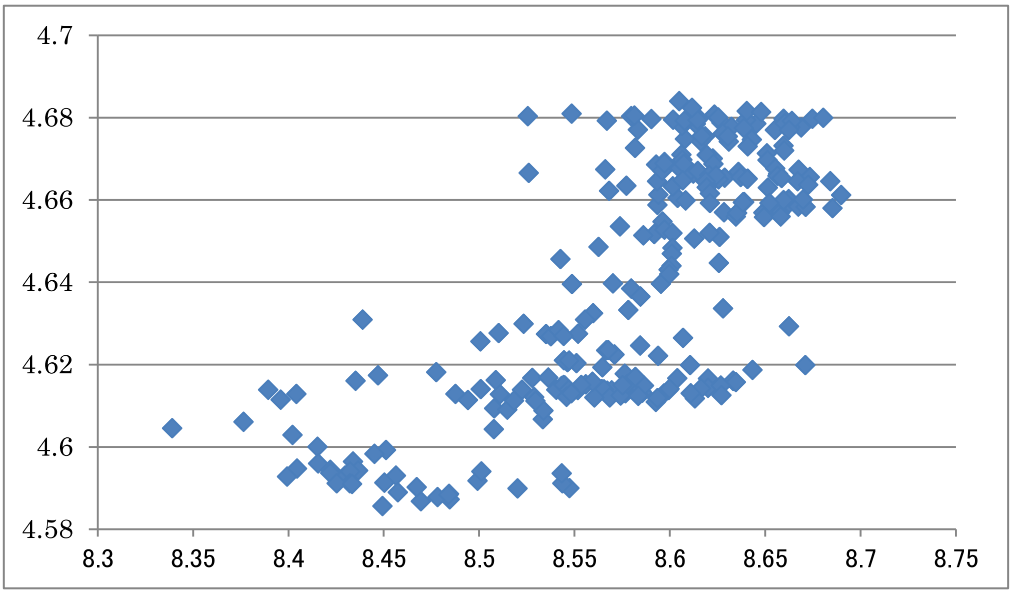
y_t denotes the oil consumption and $x_{2,t}$ denotes the index of industrial production.

p-values are in parentheses.

** denotes that null hypothesis of Granger non-causality is rejected at 1 % significance level.

* denotes that null hypothesis of Granger non-causality is rejected at 5 % significance level.

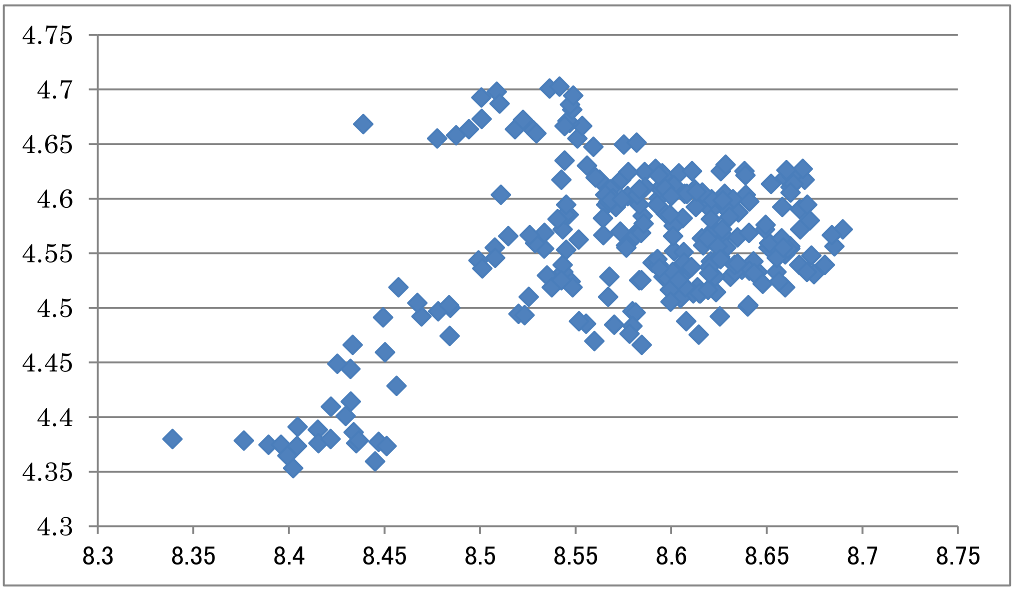
Figure 1: Scatter plot (oil consumption and price level)



Note: The horizontal axis is the oil consumption and the vertical axis is the price level.

Source: Energy Information Administration, Ministry of Economy, Trade and Industry

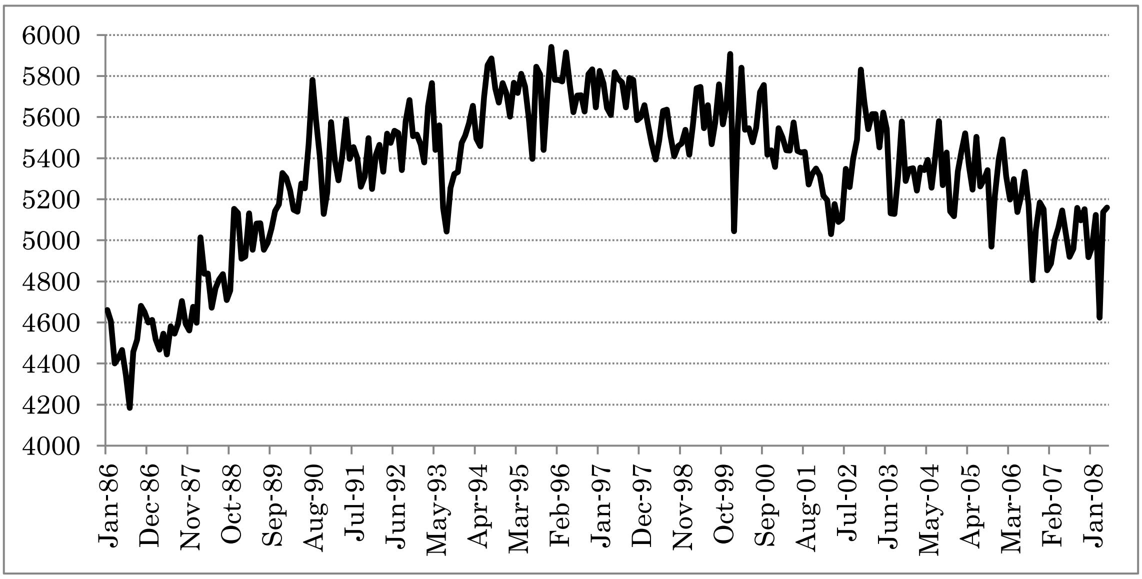
Figure 2: Scatter plot (oil consumption and economic activity)



Note: The horizontal axis is the oil consumption and the vertical axis is the economic activity.

Source: Energy Information Administration, Ministry of Internal Affairs and Communications

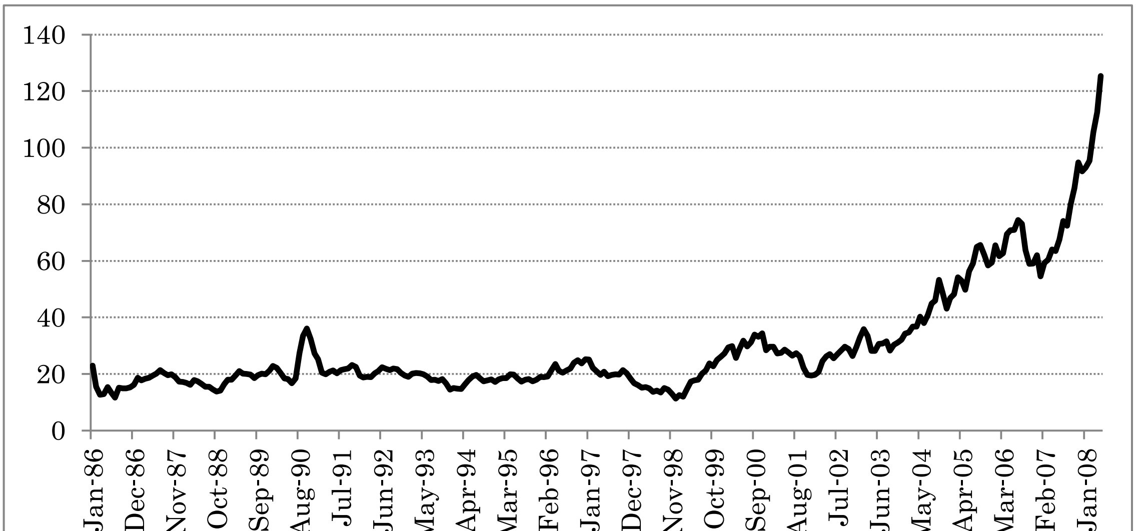
Figure 3: Oil consumption



Note: Data is seasonally adjusted (Thousand Barrels per Day).

Source: Energy Information Administration.

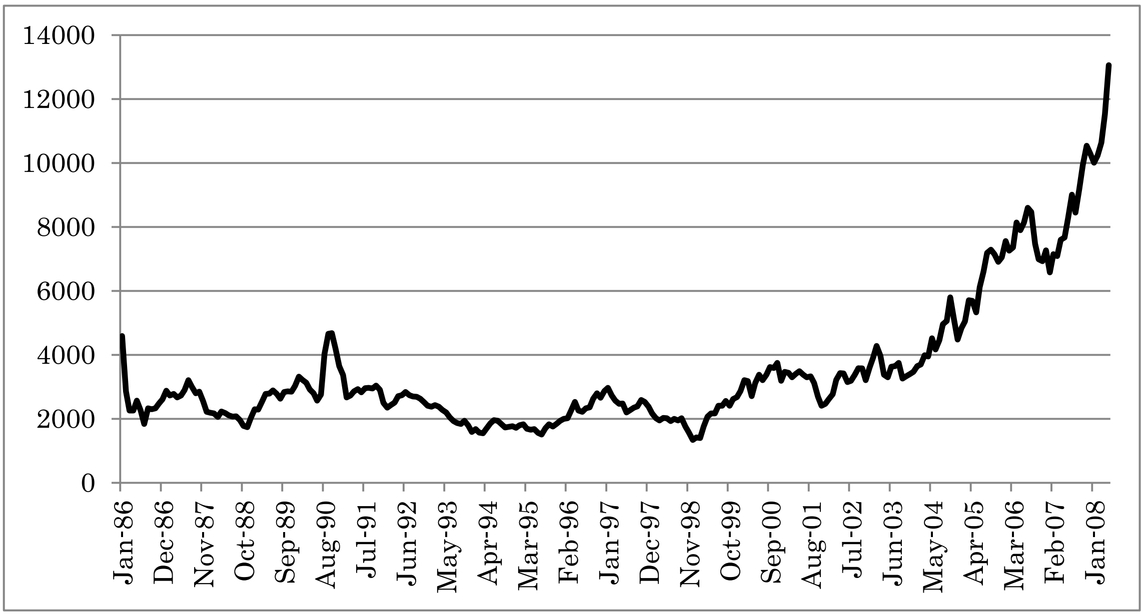
Figure 4: Oil price



Note: Cushing, OK WTI Spot Price FOB (Dollar per Barrel).

Source: Energy Information Administration.

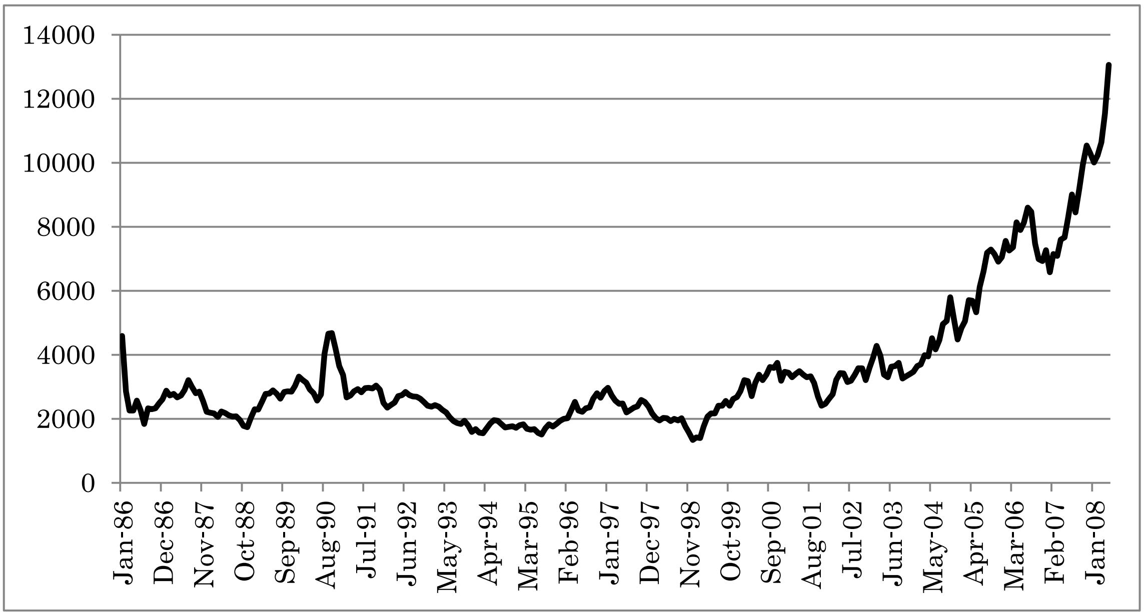
Figure 5: Oil price



Note: Cushing, OK WTI Spot Price FOB (Japanese yen per Barrel).

Source: Energy Information Administration.

Figure 6: Index of industrial production



Note: Data is seasonally adjusted.

Source: Ministry of Economy, Trade and Industry.