Output Volatility Regime and Monetary Policy Rule: Evidence from Japan

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Abstract

Over the past twenty years or so, one of the most striking economic activities has been the substantial global decline in macroeconomic volatility, including real output and inflation. However, Japan has a different evolution. Employing several classical and Bayesian approaches, this paper identifies unknown structural breaks and regime switching of Japanese postwar real GDP growth volatility and finds two breaks: the first break occurred in 1974:1 toward stabilization and the second break in 1994:4 toward destabilization. This article also finds three regimes in Japan : high growth and high volatility regime – 1955:3 to 1973:2; medium growth and low volatility regime – 1975:2 to 1992:4; and low growth and high volatility regime – 1993:1 to 2001:1. The fixed exchange rate mechanisms were the main reason for the high output volatility in the first regime. Inflation targeting was the main reason for the low output volatility in the second regime. Japan's passive monetary policy caused by asset bubble bursts was the most likely explanation for the rising output volatility in the third regime. Additionally, the empirical finding supports the theory that central bank should not respond to asset price fluctuations unless they affect central bank's inflation expectation. Finally, the evidence suggests that a monetary policy that reacts directly to the exchange rate gap would not be harmful to the inflation and output fluctuations.

Keywords: Output volatility; Structural break; Regime switching; Monetary policy; Exchange rate

JEL classification: C32, E32, E58, F41

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1. Introduction

Over the past twenty years or so, one of the most striking economic activities has been the substantial global decline in macroeconomic volatility, including real output and inflation. A number of recent studies have sought to characterize this widespread decline in volatility, also called "the Great Moderation," across sectors within the U.S. as well as most of the G7 countries. Researchers, for example, Kim and Nelson (1999a), McConnell and Perez-Quiros (2000), both documented the sharply reduced volatility of U.S. real GDP growth since 1984:1. Others, such as Blanchard and Simon (2001), Stock and Watson (2003, 2005), explored other major industrial countries and found similar declines in the volatility of real output. They both documented Japan as the only country with a different evolution. Nevertheless, they didn't analyze this abnormity in more detail.

There has been a great interest in the issue of Japan's great recession in the 1990s. However, most of this literature focuses on Japanese deflation with low average GDP growth for the past decade. Rather, this paper focuses on the recent increase in observed output volatility in Japan. This could be shown in the plot of quarter-to-quarter growth rate of real GDP in the late 1990s for Japan in Figure 1. Higher volatility of output implies more volatile employment and an increase in economic uncertainty for households and firms. The rising volatility of output means that recessions have become more frequent and more severe. Hence, this paper wants to answer three questions: First, have there been one or more structural breaks in postwar Japanese real GDP toward stabilization or destabilization? Second, if so, when? Third, the nature and source of the output volatility dynamics.

This paper employs both classical and Bayesian approaches to identify the structural break for the mean and volatility of real output at an unknown change point. Empirical results suggest two breaks - the first break occurred in 1974:1 toward stabilization and the second break in 1994:4 toward destabilization for the aggregate GDP growth and many disaggregate components of GDP as well. Furthermore, this paper uses three-state Markov-switching model and find three regimes for the Japan's postwar GDP growth: high growth and high volatility regime - 1955:3 to 1973:2; medium growth and low volatility regime – 1975:2 to 1992:4; and low growth and high volatility regime – 1993:1 to 2001:1. The identification of breaks or regimes for volatility has important implications for policy decision and research modeling such as model calibration and estimation of vector autoregression.

Another goal of this research is to determine the causes of the observed opposite direction of variations for Japan. Many explanations have been proposed for the reduction of volatility in the U.S., with the three main classes of explanations. First, improved policies: specifically monetary policy, such as Clarida, Gali, and Gertler, henceforth CGG, (1998, 2000), and Bernanke (2004). Second, better practices: particularly improved management of business inventories, such as McConnell and Perez-Quiros (2000). Third, good luck: smaller and less frequent shocks hitting the economy, such as Ahmed, Levin, and Wilson (2004), and Stock and Watson (2003, 2005).

Although these three classes of explanations probably all exist at the same time, the relative importance of these explanations is still a major concern for the evaluation of policy effectiveness. What was the main source of the latest structural break in Japan? Bad policies, bad practices, or just bad luck? This paper mainly focuses on the policy hypothesis, especially monetary policy. Although enormous literature has investigated in Japanese monetary policy, few has considered about parameter stability issue, which has been assessed and overcome in this paper. Using monetary reaction function by Generalized Method of Moments (GMM) model, the research concludes that the passive monetary policy was the most likely explanation for this unwelcome structural break in output volatility. Furthermore, Andrade and Divino (2005) suggested that Japanese monetary policy has emphasized on exchange rate targeting instead of inflation rate targeting. However, CGG (1998) found evidences in favor of the inflation targeting. This paper's results support the former and against the latter.

The paper proceeds as follows: section II does the data set overview and characterizes the stationarity and linearity of postwar Japanese real GDP growth by structural stability test. Section III specifies a series of models to explore the nature and timing for both aggregate and disaggregate output fluctuations. Section IV examines the sources of volatility structural change. Section V concludes.

2. Data Overview and Stability Test

2.1. Data Overview

This paper uses quarterly Japanese real GDP data from 1955:2 to 2001:1 provided by Economic and Social Research Institute (ESRI)[1]. Table 1 reports the sample standard deviation of major macroeconomic time series for Japan by decade (1955 to 1959 are included in the 1960s; 2001:1 is included in 1990s). Each decade's standard deviation is shown relative to the full-sample standard deviation, so a value more than one means a period of relatively high volatility. Most series were less volatile in the 1980s than over the full sample while more volatile in the 1990s than over the full sample. It is worth noting that inflation (0.3) and short-term interest rate (0.62) were both less volatile in the 1990s because the liquidity trap and zero lower bound of nominal interest rate happened in Japan.

The observed changes in the volatility of output shown in Figure 1 could arise from the change in the variance of output shocks (conditional variance), changes in the dynamic process through which these shocks affect output (changes in the autoregressive coefficient; that is, conditional mean), or both. The research estimates the instantaneous variance using nonGaussian smoother based on the stochastic volatility model with heavy tails and time-varying autoregressive coefficients. Equation (1) –(3) represent the model,

[1] ESRI only provides real GDP data based on SNA63 (System National Account) from 1955:2 to 2001:1 (benchmark year 1990). ESRI also provides real GDP data SNA93 from 1980:1 to date, but the measurement components are different for SNA68 and SNA93; that is, two data sets are not consistent.

$$y_t = \mu_t + \sum_{j=1}^{4} \phi_{jt} y_{t-j} + \sigma_t e_t \tag{1}$$

$$\phi_{jt} = \phi_{jt-1} + \gamma_t v_{jt} \tag{2}$$

$$\ln \sigma_t^2 = \ln \sigma_{t-1}^2 + \omega_t \tag{3}$$

Figure 2 presents graphical evidence on the declined volatility in the 1980s and increased volatility in the 1990s for the major series shown in Table 1. The solid line is a raw estimate of the volatility of the series, that is, the absolute value of the deviation of each series from its unconditional mean. The dash line is a two-sided estimate of the instantaneous time-varying standard deviation of the series based on equation (1) to (3). The model is discussed in more detail in Appendix 1.

The results in Figure 2 reveal a more accurate picture of the dynamics in volatility. The volatility of GDP declined in the mid 1970s and rose in the mid 1990s. The volatility of consumption seems to have risen in the mid 1990s. The volatility of investment seems to have declined in mid 1960s. The volatility of change in inventories declined in mid 1970s. The volatility of government spending rose sharply in the mid 1990s. The volatility of the inflation rate declined since the mid 1970s. The volatility of short rates declined since the early 1980s. Most series for the volatility showed different dynamics from those of the U.S. or other industrial countries. Moreover, from Figure 1, we were unable to find the fewer and shorter recessions like those found in the U.S. since 1980s[2].

2.2. Stationary and Stability Tests

Our approaches focused on whether there was a parameter (structural) stability of the real output growth, which is the first-difference of the real GDP. First, this paper considered and examined the possibility of linear (deterministic) trend process of Japan real GDP. We used

[2] There is a latest recession from 2000: 4 to 2002: 1, which cannot be seen in Figure 1.

different unit root tests for real GDP and its components as well. The test statistics displayed in Table 2 all fail to reject the hypothesis of a stochastic trend of Japanese real GDP and its components (except inflation), which means that I(1) process due to frequent permanent shocks can explain the Japanese macroeconomic process better. Following Zivot and Andrews (1992) and Lumsdaine and Papell (1997), we also considered the unit root tests that allow one and two unknown-timing structural breaks respectively. Using the crash model, with an assumption of the structural change in the intercept of the Japanese real GDP trend, the minimum t statistics for one and two breaks are -2.37 and -3.15 respectively. Therefore, we still could not reject the null of I(1) process for real GDP. Consequently, GDP and its component variables were transformed to annual growth rate ($100 \times \ln(X_t / X_{t-4})$). Others were transformed to first differences ($X_t - X_{t-4}$).

Before specifying the appropriate models (nonlinear or state-space) for Japanese real GDP growth, we needed to detect the parameter stability for the linear model. We used Nyblom's L test as described in Hansen (1992). The research specified the model by using ARIMA (1, 1, 0) for log of the real GDP to test model stability.

From Table 3, we rejected both the hypothesis of the stability or homoskedasticity of intercept term (μ) and variance (σ^2) and reject the hypothesis of the joint stability of the parameters. Although we failed to reject the null for the autoregressive coefficient, this does not mean we can rule out instability in the autoregressive parameter. As Hansen (1992) mentioned, if both the autoregressive and error variance have shifted, the power of the *L* test for the autoregressive parameter is low. In summary, the result explains the possibility of heteroskedasticity of mean and variance.

3. Model Specifications for Output Growth and Volatility Regime

In this section, we will exam whether the instability of real GDP is associated with a single or multiple distinct structural breaks in the mean and volatility. If so, we then estimate the timing of the structural change in the process of the data. Alternative Markov-switching model would be presented as well.

3.1. Benchmark Method – Single Structural Break Test on Mean and Variance

We use the following model to test if there is a structural break in the conditional mean and variance proposed by Stock and Watson (2002). To test a break in the conditional variance, we let $\varepsilon_t(\tau)$ denote the errors in the autoregression in equation (4), which is the same as equation (1),

$$y_t = \mu_t + \phi(L) y_{t-1} + \varepsilon_t \tag{4}$$

where

$$\mu_t + \phi_t(L) = \begin{cases} \mu_1 + \phi_1(L), \ t \le \tau_1 \\ \mu_2 + \phi_2(L), \ t > \tau_1 \end{cases} \quad \text{and } \mathrm{E}(\varepsilon^2) = \begin{cases} \sigma_1^2, \ t \le \tau_2 \\ \sigma_2^2, \ t > \tau_2 \end{cases} \tag{5}$$

where $\phi(L)$ denotes a lag polynomial, τ_1 is the break date for conditional mean (constant and AR coefficients) and τ_2 (innovation variance) is the break for the conditional variance. Equation (5) implies conditional mean and variance might change at different dates. We use supremum of the sequence of Wald test statistics, $W_T(\pi)$, which tests the null hypothesis that the parameters are constant against the alternative that they have a single break at a fraction π through the sample. The break date is treated as an unknown priori so that the tests compute the sequence W_T for $t = t_0 + \ldots, t_1$ and then computing a supremum of the sequence. This method is called Quandt likelihood ratio (QLR[3]) proposed by Quant (1960) and also referred to sup-Wald statistic proposed by Andrews (1993).

First, the QLR statistic is used to test for a break in equation (5) in the central 70% of the whole sample. Then we test for a break in the variance at an unknown date τ_2 by computing the QLR statistic for a break in the mean of the absolute value of the residuals from the estimated AR model (4), where the autoregression allows for a break in the AR parameter at the estimated break date τ_1. For the break test in the conditional variance, under the null hypothesis that there is no break in the variance; E($\varepsilon_t^2(\tau_1)$) is constant. Under the alternative hypothesis that there is a break at date τ_2, E($\varepsilon_t^2(\tau_1)$) is σ_1^2, when $t \le \tau_2$; is σ_2^2, when $t > \tau_2$. We report the break dates at the 5% significant level and compute 67% confidence intervals by Bai (1997)[4]. The p-values associated with these statistics are computed using the approximation proposed by Hansen (1997). Although the QLR is for the single-break model, it has the power against other forms of time variation. Rejection of the no-break null by the QLR statistic is evidence of time variation, which may have single or more than one break in equation (5). In order to check another potential break, we choose the sub-sample period 1975:2 to 2001:1.

Table 4 shows the results for the full sample period 1955:3 to 2001:1. There was a structural break of conditional mean for Japanese real GDP growth in 1973:1 and a structural break of conditional variance for Japanese real GDP growth in 1974:1. The former result explains the slowdown in the average growth rates of real output widespread across major industrial countries by the first oil crisis.

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[3] We use heteroscedasticity-robust version of the QLR statistics where $W_T(\pi)$ is computed by using White (1980) heteroscedasticity-robust covariance matrix, in which the residuals were computed under the null rather than each of the alternatives for computational convenience.

[4] 95% intervals are too wide and uninformative since the break estimator has a non-normal, fat-tailed distribution. Hence, we report 67% confidence intervals rather than 95%, conventional intervals.

Table 5 reports the results for the subsample period 1975:2 to 2001:1. There is one structural break of conditional mean for Japan real GDP growth in 1991:1 since the burst of the Japanese "bubble economy" and another structural break of conditional variance for Japanese real GDP growth in 1994: 4. The former dates the beginning of Japan's persistent economic stagnation, which is also called "Great Recession."

3.2. Single Structural Break Test on Variance

McConnell and Perez-Quiros (2000) attributed the possibility of the structural break of observed residual variance for U.S. GDP growth to mean or innovation separately. To provide the robustness analysis for the structural break date that we got from the benchmark method, we estimate a break point by their method by using the innovation break test only:

$$y_t = \mu + \phi_1 y_{t-1} + \phi_2 y_{t-2} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \sigma_t^2) \tag{6}$$

where $\sigma_t^2 = \sigma_1^2$ if $t \le \tau$, and $\sigma_t^2 = \sigma_2^2$ if $t > \tau$. As reported in Table 6, we still find the same result of break date: 1974:1 for the whole sample period and 1994:4 for the subsample period.

3.3. Multiple Structural Breaks Test on Mean

Even though we found two structural breaks above, the method is based on an ad hoc choice of sample period. To do a robustness check and overcome the drawbacks from single structural break test, we use a multiple-structural-breaks test that allows one to test the null hypothesis of n breaks versus the alternative hypothesis of n+1 breaks, as proposed by Bai and Perron (1998, 2003) using least squares. We consider the linear regression with *n* breaks:

$$y_t = x_t'\beta + z_t'\delta_j + \varepsilon_t \qquad t = T_{j-1} + 1, ..., T_j \tag{7}$$

where j =1,…., n +1. y_t is the observed dependent variable at time *t*; x' $(p \times 1)$ and z' $(q \times 1)$ are vectors of covariates; β and δ_j are the vectors of coefficients; ε_t is the disturbance at time t.

We specify the model by $p = 0$ where all the coefficients are subject to change. The variance of ε_t needs not to be constant. Breaks in variance are permitted at the same dates as the breaks in means of the regression.

Table 7 reports the results. For the determination of the numbers of breaks, the $\sup F_T$ tests are all significant for n between 1 and 5. The F_T (2|1) is 11.6 and significant but F_T (3|2) is 4.08, which is not significant. Therefore, the sequential procedure using 5% significance level selects 2 breaks. Finally, the break dates on means are estimated in 1973:1 and 1991: 2, which are almost the same as the break dates found in the benchmark method.

3.4. Multivariate Structural Break Test on Variance - VAR Method

In theory, a common break date could be estimated more precisely by using multiple equations. Bai, Lumsdaine and Stock (1998) used a low-dimensional VAR to estimate common breaks across multiple series. We use their method except for keeping VAR coefficients constant. The hypothesis of no break was tested against the alternative of a common break in VAR equations by the QLR statistic, which was computed using the absolute values of the VAR residuals. We also use 67% confidence interval proposed in the same paper.

VAR components are the first difference of the logarithm of consumptions, investment, export, import, and government spending. Table 8 reports the result that there is a break date in 1975:2 for the whole sample period and another break date in 1994:3 for the subsample period. The break date results from univariate methods mentioned above are all in the confidence interval here.

3.5. The Long and Declining Trend Method on Variance

Blanchard and Simon (2001) argue that a large decline in U.S. output volatility is not a recent development but rather a steady long trend, starting in the 1950s, interrupted in the 1970s and early 1980s, with a return to trend in the late 1980s and 1990s. To investigate this possibility for Japan (an increase in output volatility since middle 1990s is a temporary deviation from the declining trend, as shown in Figure 3.A), we conduct additional specification in which the innovation variance is modeled as a time-trend linear function with a discrete jump at an unknown break date. The QLR test in equation (5) was modified as a model to include a time trend as well as the break.

$$|\varepsilon_t| = \gamma_0 + \gamma_1 t + \gamma_2 D_t(\tau) + e_t \quad (9)$$

where $D_t(\tau)$ is a binary variable that equals 1 if $t \geq \tau$ and equals 0 otherwise and e_t is an error term. We also choose the central 70% of the sample for value τ . The results are reported in the final block of Table 4. For the real GDP in the full sample, we reject the null hypothesis of no trend as well as no break. The break date is in 1994:2. Namely, we can interpret that there was a declining trend in volatility for Japanese real output prior to 1994 but a break occurred in 1994:2 towards rising volatility regime, which is shown in Figure 3.B. For the subsample, we could not reject the null hypothesis for most of the macroeconomic series because sample size is not big enough. In sum, the break model mentioned above performs better than the trend model.

3.6. Markov-Switching Model

Hamilton (1989) used a Markov-switching model to explain postwar U.S. business fluctuations, and the turning point was treated as a structural event that is inherent in the data-generating process. Parameter changes are thought to be recurrent and governed by a Markov process, which is different from the structural break test detecting nonrecurrent changes. This model can capture a particular form of nonlinear dynamics or asymmetry in the business cycle. The model allows the mean rate of GDP growth to switch between two states to show the boom with a high probability that output growth remains high if it is initially high and recession as well if output growth is initially low.

We first apply his model in the Japanese real GDP growth (y_t) for the sample 1955:3-2001:1 as follows:

$$
\begin{aligned}
&(y_t - \mu_{St}) = \phi_1(y_{t-1} - \mu_{St-1}) + ... + \phi_4(y_{t-4} - \mu_{St-4}) + \varepsilon_t \\
&e_t \sim i.i.d.N(0, \sigma^2) \\
&\mu_{St} = \mu_0(1 - S_t) + \mu_1 S_t \\
&\Pr[S_t = 1 \mid S_{t-1} = 1] = p, \quad \Pr[S_t = 0 \mid S_{t-1} = 0] = q
\end{aligned}
\tag{10}
$$

From Figure 4.A, without considering the structural change on means and variance Hamilton's model would perform too poorly to explain the recession and expansion, especially between 1970s and 1990s. To improve this situation, Kim and Nelson (1999a) proposed a modified model in which μ and/or σ^2 are subject to a one-time structural break.

However, we use an alternative model, proposed by Garcia and Perron (1996), which is three-state rather than two-state. This model allows three possible regimes affecting both the mean (μ_{St}) and variance (σ^2_{St}). We only assume that three states of means with $\mu_1 < \mu_2 < \mu_3$ to present the Japanese postwar GDP growth convergence. There is no restriction on states of variance. Furthermore, based on this model, we use the Bayesian approach by Gibbs-sampling procedure proposed by Kim and Nelson (1999b) to estimate the regime switching. It is worth noting that Bayesian approach has an advantage over classical tests for treating both model's hyperparameters and the state variable random variables. In contrast to the classical approach, which treats the state variable conditional on the estimated values of the hyperparameters, the

Bayesian approach make inference on state variables based on joint distribution of state variables and hyperparameters.

$$
\begin{aligned}
&(y_t - \mu_{St}) = \phi_1 (y_{t-1} - \mu_{St-1}) + \phi_2 (y_{t-2} - \mu_{St-2}) + \varepsilon_t \\
&\varepsilon_t \sim i.i.d.N(0, \sigma_{St}^2) \\
&\mu_{St} = \mu_1 S_{1t} + \mu_2 S_{2t} + \mu_3 S_{3t}, \quad \text{where } \mu_1 < \mu_2 < \mu_3 \\
&{\sigma_{St}}^2 = \sigma_1^2 S_{1t} + \sigma_2^2 S_{2t} + \sigma_3^2 S_{3t} \\
&\Pr[S_t = j \mid S_{t-1} = i] = p_{ij}, \quad \sum_{j=1}^{3} p_{ij} = 1, \quad i, j = 1,2,3 \quad S_{kt} = 1 \; if \; S_t = k
\end{aligned}
\tag{11}
$$

The first 2,000 draws of Gibbs-sampling are discarded, and the analysis is based on the next 10,000 draws. Table 9 reports the marginal posterior distributions of each of the model's parameters. We find three regimes as shown in Figure 4: high growth (quarterly: 2.20; annual: 8.8) and high volatility (quarterly: 1.77; annual: 7.08) regime – 1955:3 to 1973:2; medium growth (quarterly: 0.95; annual: 3.8) and low volatility (quarterly: 0.42; annual: 1.68) regime – 1975:2 to 1992:4; and low growth (quarterly: 0.10; annual: 0.4) and high volatility (quarterly: 1.90; annual: 7.6) regime – 1993:1 to 2001: 1[5] by three-state Markov-switching model. The regime-switching points are very close to the structural break points.

4. Sources of the Change in Output Volatility

There have been three major explanations proposed for the sources of the decline of U.S. output volatility since 1984. The first hypothesis is *the good macroeconomic policy*, i.e. that better monetary policy that has tamed the business cycle. If monetary policy changes, it has a direct effect on the propagation mechanism of the shocks. If agents are rational, policies changes will be incorporated into the private sector's forecasts. The second one is *the good practice*, i.e. just-in-time inventory management coming from computation and communication technology improvement, increased depth and sophistication of financial markets, deregulation in industries,

[5] Another short period of low growth and high volatility regime occurred between 1973:3 and 1975:1.

and a shift away from manufacturing toward services. The last one is *the good luck*, which simply reflects a decline in the variance of exogenous shocks hitting U.S. economy[6]. It is interesting to explore why Japan reduced their output and inflation volatility a decade earlier than U.S. (1974:1 versus 1984:1) but has had rising output volatility of late. Which hypothesis can mostly contribute to the Japan medium growth – low volatility regime from 1975:2 to 1992:4?

4.1. Bad Luck Hypothesis

Since the real exogenous shocks are difficult to identify, this paper uses VAR forecast errors (impulse) to represent the shocks. Using the reduced-form VAR method mentioned in Section 3.4 with three variables: real GDP growth, call rate, and inflation rate, we only found one common structural break date of absolute value of VAR residuals occurred in 1980:3 with confidence interval 1979:2 to 1981:4. This simply explains exogenous shocks which changed in 1980:3 have no association with the output structural breaks in 1974:1 or 1994:4. Hence, the bad luck hypothesis would not be the appropriate explanation for changing output volatility regimes. Furthermore, Hamilton (1983) argued that the stagflation period in the 1970s in the U.S. was attributed to cost-push shocks such as union wage pressures, price increases by oligopolistic firms or increase in oil prices, that is, bad luck. But we can not see that Japan's variability of inflation and output was affected by oil shock after 1975:2. It would be difficult for oil shocks to generate unstable real output and inflation for a long time without an accommodating monetary policy. However, we can not exclude the possibility of an influential role of productivity shocks on output volatility in Japanese first and third regime. This paper leaves this hypothesis open for further research.

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[6] Ferguson (2006) doubt this hypothesis because the events occurred in the late 1990s and early 2000s such as Russian default crisis, Long Term Capital Management Travails, September 11 attacks, and corporate governance scandals.

4.2. Bad Practice Hypothesis

Bad practice is also problematic to be the main cause of rising volatility since there is no obvious evidence to support this hypothesis for Japan. For example, better inventory management occurred not only in the U.S. but also in Japan as shown in Figure 2.A. And this better inventory hypothesis may play an important role for the decline in volatility of output in 1975:2 since the volatility for the change of inventory over GDP has a structural break toward stabilization in 1975:2 reported in Table 4. It is worthwhile to mention the seriousness of the banking problem in Japan, due to bad loans from the burst of the asset bubble in the early 1990s. Figure 5 shows that the M2 multiplier (M2/Monetary base) decreases sharply since 1992. We also get the structural break in mean of M2 in 1990:2 from Table 5 due to the difficulty in financing investment projects. This "credit crunch" could explain Japan's slump and its rising variance as well.

It is widely agreed that residential fixed investment has been highly volatile and procyclical. From Table 4 and 5, the evidence reports that it does not play a central role in explaining rising variability because there is no structural break of residential investment occurred in the 1990s. The services sector is less cyclically sensitive than the manufacturing sector. However, from Figure 7, we can see that the Japanese economy is shifting from manufacturing sector towards a service sector, especially in the 1990s. Consequently, this hypothesis is less persuasive to explain the rising volatility totally.

4.3. Bad Policies Hypothesis

Finally, it is not surprising to get the result: bad policy is the main source of Japan's high volatile regime since 1993:1. There has been a flood of literature investigating the sources and policy recommendations for Japan's low growth regime, but few of them have explored the association between its policy and rising volatility of output. What did Japan do that cause a rise in output volatility given relatively stable inflation (deflation)?

4.3.1. Monetary Policy Rule – Closed Economy Model

CGG (2000) estimated a forward-looking Taylor rule for the monetary policy reaction function using ex-post data and generalized method of moment (GMM). They presented evidence that US monetary policy rule changed from indeterminacy in pre-Volcker period to determinacy in Volcker-Greenspan period. Since evidence was shown two structural breaks between three regimes, we use these three regimes as subsamples to estimate Japanese monetary policy rules. These three regimes correspond to unstable, stable and unstable eras. Following CGG's method, a monetary policy that uses short-term nominal interest rate as an instrument rule affects the real economy in the short run:

$$R_t = \rho R_{t-1} + (1-\rho)(\alpha + \psi_1 E[\pi_{t+4} \mid \Omega_t] + \psi_2 E[x_{t+1} \mid \Omega_t]) + \varepsilon_t \qquad (12)$$

where R_t is the short-term nominal interest rate (call rate) set by the Bank of Japan (BOJ), ρ is an indicator of the degree of smoothing of interest rate changes, π_{t+4} [7] is the rate of inflation in one year (the percent change in the price level between period *t* and *t + 4*) measured by the four-quarter percent change of the GDP deflator, and x_{t+1} is the output gap[8] in period t + 1 (the deviation of real GDP from potential real GDP, calculated by HP-filter). E is the expectation operator, and Ω_t is the information set at the time the interest rate is determined. The instrument set includes lags of the call rate, inflation, the output gap, and the log of real exchange rate. Generally speaking, interest rate rules with ψ_1 > 1 and ψ_2 > 0 will tend to be stabilizing[9]. For example, famous Taylor rule suggests ψ_1 =1.5 (for the current inflation) and ψ_2 =0.5 (for the

[7] The central bank is assumed to have a target horizon of one year for its inflation target since the horizon roughly fits the conventional wisdom about the lag with which monetary policy affects inflation. Target horizon of 4 quarter is more realistic than one or two quarters. (CGG 2000)
[8] Although some economists point out that Japan's output gap is seriously undervalued in 1990s (Krugman 1998), it is still hard to accurately measure and predict the true productivity process. Given the limited information, we use HP filter to calculate Japan's output gap as shown in Figure 6.
[9] $\psi_1 > 1$ per se is called Taylor principle.

current output gap) proposed by Taylor (1993). On the other hand, those with $\psi_1 \leq 1$ and $\psi_2 \leq 0$ are likely to be destabilizing or, at best, accommodative of shocks to the economy.

4.3.2. Results in Closed Economy Model

Based on the results from section III, we estimate the model with three subperiods: 1961:1-1975:1, 1975:2-1992:4, and 1993:1-2001:1 using quarterly data[10]. Table 10 reports GMM estimates of the interest rate rule parameters ψ_1, ψ_2, α, and ρ. The model specification works well since the J-statistic shows that we cannot reject the overidentifying restrictions. We find significant results of ψ_1 (2.08, with *s.e.* = 0.55) in the second regime, which is stabilizing. In first and third regime, the ψ_1 are destabilizing. (-0.85, with *s.e.* = 0.46; -0.67, with *s.e.* = 0.27, respectively). The estimates of ψ_2 are not significant from zero for the first and second regimes, but it is significant in the third regime.

For robustness check, we also use alternative monthly data to estimate two subperiods[11]: 1979:1-1992:12, 1993:1-2001:12 where π_t is measured by the CPI and x_t is the output gap of industrial production. The instrument set includes 12 lags of the call rate, inflation, the output gap, the difference of the log of real exchange rate, and commodity price inflation. We find the consistent results as quarterly data. In the second regime, ψ_1 is stabilizing and significant (2.04, with *s.e.* = 0.98) and ψ_2 is insignificant (0.83, with *s.e.* = 0.53). In the third regime, ψ_1 is destabilizing and significant (0.32, with *s.e.* = 0.07) and ψ_2 is insignificant (0.02, with *s.e.* = 0.03). The estimate for ρ is high in most cases (except first regime), providing strong evidence for interest rate inertia.

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[10] The first regime begins in 1961:1 because of call rate data availability. The third regime begins in 1993:1 instead of 1994:4 because we can have a bigger sample size for the short third regime.
[11] Industrial production data is only available since 1978:1.

The results in the second regime are similar to those in CGG (1998). Using monthly data, they found ψ_1 is 2.04 (with *s.e.* = 0.19) and ψ_2 is 0.08 (with *s.e.* = 0.03) for BOJ from 1979:4 to 1994:12[12]. They suggested that BOJ appeared to have placed more weight on controlling inflation relative to output stabilization.

Thus, our estimates confirm that CGG (1998) about the monetary policy using inflation targeting that follows Taylor principle with some weight on output stabilization was the main reason for Japanese stable regime (1975:2 – 1992:4). Why would Japan apply stabilizing monetary policy in second regime but used destabilizing policy in the first and third regimes? First, it is worth noting the abandonment of Bretton Woods system of Japan in March 1973, changing from fixed-exchange-rate system to floating-exchange-rate system. It was an important institutional change and source for the structural break from high volatility regime to low volatility regime. Under the Bretton Woods system, Japan maintained a fixed nominal exchange rate of 360 yen per dollar. Monetary policy during this period was limited and passively responded to offset exchange-rate pressures (ψ_1 = -0.85, with *s.e.* = 0.46; ρ >1). From the Mundell-Fleming Model, we know that a government that fixes its currency's exchange rate loses control of the domestic money supply. Taylor (1993) also argued that the fluctuations in real output are much larger in the fixed-exchange-rate system than those in the flexible-exchange-rate system. Using a multivariate GARCH-M model, Kim (2000) found that the flexible-exchange-rate system helped the Japanese economy to absorb the foreign shocks.

Second, for the third regime, ψ_1 is passive and destabilizing (ψ_1= -0.67, with *s.e.* = 0.27). Since 1995 (see Figure 8), Japan has fallen into deflation. According to Taylor principle, BOJ has to cut the nominal interest rate or increases monetary base actively (ψ_1 >1) to mitigate this deflation gap but they did not or could not. Namely, they used too restrictive of a monetary policy,

[12] CGG (1998) mentioned that they pick April 1979 as the appropriate starting date because it is the beginning period of significant financial market deregulation. We think it is not persuasive. The appropriate starting date should be 1975 if data is available.

which is consistent with the findings in McCallum (2000) using different policy rules models. The deflationary environment let Japan fall into liquidity trap in which conventional monetary policy lost its power because the nominal interest rate was close to zero, where the quantity of money became irrelevant since money and bond became perfect substitutes. Figure 9 shows the call rate dropped to 0.25% and then cut to 0.02% since 1999[13]. Using a stochastic simulation, Reifschneider and Williams (2000) found volatility of output and employment increase significantly under the low-inflation environment with a constraint of zero bound of nominal interest rate. This is the main reason for the output instability in the third regime.

4.3.3. Monetary Policy Rule – Open Economy Model

As in Ball (1999), in a closed economy, inflation targeting and Taylor rules perform well in stabilizing both output and inflation. In an open economy, however, these policies perform poorly unless they are modified. The policy instrument should be based on both the interest rate and the real exchange rate. Svensson (2000) also mentioned that including the real exchange rate in the monetary policy rule has important consequences. The exchange rate allows an additional channel for the transmission of monetary policy. The real exchange rate will affect the aggregate demand and inflation with lags. For robustness check, we modified and estimated the open - economy monetary policy rule proposed by Taylor (2001) as follows,

$$R_t = \rho R_{t-1} + (1-\rho)(\alpha + \psi_1 E[\pi_{t+4} \mid \Omega_t] + \psi_2 E[x_{t+1} \mid \Omega_t] + \psi_3 E[e_t \mid \Omega_t]) + \varepsilon_t \qquad (13)$$

where R_t π_{t+4}, x_{t+1}, are variables as mentioned in (7) and e_t is the real exchange rate gap[14] (the deviation from its HP-trend; an increase in e_t is a real appreciation). The model is similar to that

[13] Even though conventional monetary policy is ineffective in the liquidity trap, Japan still can benefit by expanding monetary base and using open market purchases for long-term government bonds and/or foreign exchanges (Svensson 2004).

[14] In Obstefel and Rogoff (1995), they pointed out that Japan, with an unusually high differential between productive growth in its tradable- and nontradable-goods, will typically experience a rise in the relative price of nontraded goods and an appreciation of its real exchange rate. Any attempt to use PPP as a guide to monetary policy intervention must allow for productivity-based PPP deviations.

in CGG (1998). The difference is that we use HP-trend to get real exchange rate gap rather than linear-trend they used. The reason for different forecast horizons is because of the realistic lags of the different channels for the transmission of monetary policy (Svensson 2000): The direct exchange rate channel to the inflation (through import prices) has the shortest lag, the aggregate demand channel to the output gap has an intermediate lag, and the aggregate demand and expectations channels on inflation have the longest lag. We use quarterly data to estimate the forward-looking model by two subperiods: 1975:2-1992:4, 1993:1-2001:1 and monthly data for two subperiods: 1979:1-1992:12, 1993:1-2001:12 where π_t is measured by the CPI and x_t is the output gap of industrial production. The instrument set includes four lags of the call rate, inflation, the output gap, and the difference of real exchange rate.

4.3.4. Results in Open Economy Model

Table 11 presents the results for the open-economy model. Surprisingly, in the second (stable) regime, ψ_1 is 0.54 (with *s.e.* = 0.14), that is, BOJ did not raise the nominal interest rate sufficiently (ψ_1>1) to increase the real rate when inflation moves above its long-run target as reported in CGG (1998). Including real exchange rate data, they found ψ_1 was 1.92 (with *s.e.* = 0.11) for BOJ in the period from 1979:4 to 1994:12[15]. One potential explanation would be that they used inappropriate real exchange data (deviation from linear trend) while real exchange rate is I(1) process as shown in Table 2. Using cointegration analysis and impulse response functions for the same period data as in CGG (1998), Andrade and Divino (2005) found BOJ has tried to stabilize exchange rate and the interest rate is counter-cyclical to the exchange rate and the coefficient of inflation is smaller than 1. In the second regime, BOJ responded actively via exchange rate channel (as an indicator of future inflation expectation) rather than inflation

[15] CGG (1998) mentioned that they pick April 1979 as the appropriate starting date because it is the beginning period of significant financial market deregulation. We think it is not persuasive. The appropriate starting date should be 1975 if data is available.

channel directly because the exchange rate channel gives the central bank a possibility to stabilize inflation sufficiently.

As argued by Obstfeld and Rogoff (1995), an appreciation of the real exchange rate accompanied by slow output growth would call on the central bank to lower the short-term interest rate to relax the monetary policy. Nonetheless, Japan did not follow this rule of thumb in the third regime. As shown in Table 10 and 11, either quarterly or monthly data, Japan has all the significant negative policy coefficients in the second regime but positive coefficients in the third regime. This implies that BOJ used relatively passive and destabilizing monetary policy rule in the third regime because of $\psi_1 < 1$ and $\psi_3 > 0$. This destabilizing interest-rate response to real exchange rate also caused rising volatility in real exchange rate (structural break occurred in 1992:4 toward destabilization) shown in Table 4 and 5 and Figure 2.B. And more volatile real exchange rate caused larger output fluctuations through IS curve. This would be the reason for the output instability in the third regime in the open-economy model.

4.3.5. Reasons for Using Bad Monetary Policy

The results above lead to the following important question: why did BOJ persistently use clearly inferior monetary policy resulting in high output volatility in the 1990s? Three possible answers for this question are the following.

First, after the asset price bubbled in the late 1980s, BOJ was more concerned about asset price so they were reluctant to use expansionary policy. Bernanke and Gertler (1999) argued that it is neither necessary nor desirable for monetary policy to respond to asset prices cycle, except to the extent that they could help to forecast inflationary or deflationary pressures. Did BOJ conduct excessive monetary easing to generate the asset bubble in the second regime, in particular, 1985-1989? From the results of our open economy model, it is persuasive that BOJ tried to stabilize real exchange rate gap rather than the explanation that BOJ was accommodating asset boom. As in Ueda (1997) using structural VAR, given the reasonable aggregate price level, BOJ's monetary policy could not have been conducted better to avoid the asset bubbles and bursts in the late 1980s and the early 1990s caused by private autonomous optimism.

Second, policymakers treated deflation as a good idea since they thought it came from a financial system reform and technological progress rather than a deflationary spiral. Third, they were probably waiting for the Pigou effect to come in play. If prices fell enough, expenditure would increase because of rising wealth. The truth is that deflation is coming from falling aggregate demand or IS curve with falling output. And this mistake is closely related to the output gap mismeasurement hypothesis proposed by Orphanides (2004). In the 1970s, U.S. monetary policymakers overestimated the negative output gap and thus continued using expansionary monetary policy, which led to high volatility of inflation and output. In the 1990s, Japan underestimated the output gap and thus continued using passive/contractionary monetary policy, which then led to deflation as shown in Figure 8.

4.3.6. Fiscal Policy

Based on the Ricardian equivalence argument, the aging population and huge rise in public debt in Japan more or less limited the effectiveness of fiscal policy. The surveys shows discretionary fiscal policy still has the effects on Japan such as tax cuts and government spending increases (Kuttner and Posen 2002). With a growth recovery in 1996, the recession seemed to be over, and Ministry of Finance (MOF) decided that it was time to clean up the budget deficit, which reached 4% of GDP with the aging population in prospect. Therefore MOF implemented so-called expansionary fiscal contractions by raising the value-added tax in April 1997, with rising government spending at the same time. This ill-timed tax increases drove Japanese economy down into an even deeper recession. As shown in Table 5, evidence shows the structural breaks in volatility toward destabilization which occurred in 1997:2 for both government spending and nonresidential investment. More research is needed to find the association between Japanese recent fiscal policy and rising volatility in the past decade. It will be interesting to see how much of the Japanese output volatility is attributed to the monetary policy by removing fiscal effects.

5. Conclusion

This paper employs several approaches to identify an unknown structural break and regime switching of postwar Japanese real GDP growth. This paper has found two breaks: the first break occurred in 1974:1 toward stabilization and the second break in 1994:4 toward destabilization for the aggregate GDP growth and many disaggregate components of GDP as well. Furthermore, we find three regimes as follows, the first regime: high growth and high volatility regime – 1955:3 to 1973:2; the second regime: medium growth and low volatility regime – 1975:2 to 1992:4; and the third regime: low growth and high volatility regime – 1993:1 to 2001:1 by three-state Markov-switching model.

The fixed exchange rate mechanisms were the main reason for the high output volatility in the Japanese first regime. This conclusion supports Obstfeld and Rogoff (1995) and CGG (1998). In the second regime, the monetary policy which follows inflation targeting (raise or lower nominal interest rate sufficiently when expected inflation moves from its long-run target) with little weight on output stabilization was the main reason for the low output volatility. This finding is consistent with CGG (1998, 2000) and supports the view that central bank is undesirable to respond strongly to the output gap. Furthermore, the empirical finding supports the theory (Bernanke and Gertler 1999) that central bank should not respond to asset price fluctuations unless they affect central bank's inflation expectation.

In the open-economy framework, the flexible exchange-rate-targeting (negative nominal interest rate response to the real exchange rate appreciation) instead of Taylor rule or inflation targeting would be a more appropriate explanation for successfully stabilizing monetary policy in the second regime in Japan. This result suggests that a monetary policy rule that reacts directly to the exchange rate gap would not be harmful to the inflation and output fluctuations[16]. This evidence favors empirical result in Andrade and Divino (2005), and theoretical models in Ball (1999) and Svensson (2000) over CGG (1998) and Taylor (2001).

Aside from the inconsistent fiscal policy in Japan, the passive monetary policy caused by asset bubble bursts, especially positive nominal interest rate response to the real exchange rate appreciation was the main reason for the rising output volatility in the third regime. Even though BOJ can not lower the interest rate when the nominal interest rate are bounded at zero, Japan still can let their currency depreciate as suggested in Svensson (2003) to reach the negative coefficient (ψ_3) as that in the second stable regime.

[16] Taylor (2001) argued that reacting to the exchange rate (not exchange rate gap/deviation) might not improve the macroeconomic performance.

Table 1. Sample Standard Deviations, by Decade, of Annual Growth Rates or Changes of Major Japan Macroeconomic Time Series

Series	Standard Deviation	Standard Deviation, relative to Whole Period			
	1955-2001	1955-1969	1970-1979	1980-1989	1990-2001
GDP	0.038	1.25	0.82	0.49	1.13
Consumption	0.034	1.22	0.83	0.64	1.09
Investment	0.120	1.37	0.90	0.48	0.89
Fixed investment	0.100	1.30	0.90	0.49	1.00
Nonresidential	0.117	1.38	0.86	0.42	0.92
Residential	0.118	1.02	0.91	0.87	1.15
Δ Inventory/GDP	0.006	1.06	1.35	0.61	0.82
Exports	0.076	1.09	1.12	0.97	0.78
Imports	0.110	1.26	0.93	0.92	0.73
Government Spending	0.051	0.95	1.16	0.79	1.08
Inflation rate	3.753	1.07	1.61	0.50	0.30
Real exchange rate	12.597	-	0.88	1.01	1.05
Short interest rate	2.199	0.96	1.26	1.09	0.62
10-year bond rate	0.842	-	1.07	0.80	1.13

1. Most series are annual growth rates, which are the first difference of the logarithm of the original series, except for the change in inventories as a fraction of GDP, which is the level of series; inflation rate, real exchange rate, short term interest rate, and long term interest rate (10 year government bond yield) are the first difference of the series.
2. All growth rates are four-quarter growth rate, which is $100 \times \ln (Y_t / Y_{t-4})$. And standard deviation is the absolute value of the deviation of each series from its mean.
3. Inflation rate is the four-quarter change in the annual inflation rate measured from GDP deflator.
4. Most of the data are from Japan ESRI, website: http://www.esri.cao.go.jp/en/sna/qe011-68/gdemenue68.html.
5. Real exchange rate is the index of weighted average of the yen's real exchange rates versus 15major currencies (26 countries) which are calculated from exchange rates and price indexes of the respective countries. The data are available since1973: 1 from Bank of Japan, website: http://www2.boj.or.jp/en/dlong/stat/stat2.htm#03 "ehrate.csv" file.
6. Short-term interest rate is call rate (Collateralized overnight interest rate), which is available since 1960: 1. Short term interest rate data are from Bank of Japan, website: http://www.boj.or.jp/en/stat/stat_f.htm "cdab0720.csv"file.
7. 10-years government bond yield data is available since 1972: 1. Short-term interest rate data are from Bank of Japan, website: http://www.boj.or.jp/en/stat/stat_f.htm "cdab0740.csv" file.

Table 2. Unit Root Test Statistic of Japan Major Macroeconomic Time Series 1955:3 to 2001:1

		ADF				Phillips-Perron				DF-GLS			KPSS		
Variables	equation	Test stat	Lag lenth	P-value	Result	Test stat	Band width	P-value	Result	Test stat	Lag lenth	Result	Test stat	Band width	Result
Real GDP	Intercept +trend	-0.48	0	0.98	I(1)	-0.55	7	0.98	I(1)	-0.21	4	I(1)	0.42	10	I(1)
Consumption	Intercept +trend	-0.47	1	0.98	I(1)	-0.63	1	0.98	I(1)	0.26	3	I(1)	0.43	10	I(1)
Investment	Intercept +trend	-2.26	3	0.45	I(1)	-2.46	6	0.35	I(1)	-0.71	3	I(1)	0.36	10	I(1)
Export	Intercept +trend	-0.58	1	0.98	I(1)	-0.49	8	0.98	I(1)	0.01	1	I(1)	0.44	10	I(1)
Import	Intercept +trend	-2.33	2	0.41	I(1)	-2.38	4	0.39	I(1)	-0.99	2	I(1)	0.37	10	I(1)
Government Spending	Intercept +trend	0.53	1	0.99	I(1)	0.45	4	0.99	I(1)	0.10	0	I(1)	0.42	10	I(1)
Call rate	Intercept	-2.20	1	0.21	I(1)	-2.02	4	0.28	I(1)	-1.66	1	I(1)	1.08	10	I(1)
Real exchange rate	Intercept	-1.55	0	0.51	I(1)	-1.79	4	0.39	I(1)	-1.24	0	I(1)	0.93	9	I(1)
Inflation rate	Intercept	-3.08	3	0.03	I(0)	-3.54	4	0.01	I(0)	-2.56	3	I(0)	0.94	10	I(1)

1. ADF is the augmented Dickey-Fuller test of a unit root against no unit root by Dickey and Fuller (1979). Lag length is automatic based on SIC, the maximum lag is 13. Test critical value is -3.43 at 5% significant level. P-value is from Mackinnon (1996).
2. For Phillips –Perron test, Bandwith is based on Newey-West using Bartlett kernel.
3. DF-GLS is a Dickey-Fuller test based on GLS-detrended series by Elliott, Rothenberg, and Stock. (1996).
4. KPSS test is based on the null hypothesis of stationary process proposed by Kwiatkowski, Phillips, Schmidt, and Shin (1992).

Table 3. **Nyblom's L Test for Stability of Japan Real GDP Growth 1955:3 to 2001:1**

Specification: $y_t = \mu + \phi y_{t-1} + \varepsilon_t, \;\; E(\varepsilon_t^2) = \sigma^2$			
	Estimate	L_c	CV (5 percent)
μ	0.94 (0.15)	3.02	0.48
ϕ	0.26 (0.09)	0.48	0.48
σ^2	1.67 (0.28)	0.67	0.48
Joint L_c		3.64	1.01

Notes: Nyblom’s L test from Hansen (1992). y is real GDP growth. Standard errors are in parentheses. L_c is the test statistic for a break point in each of the coefficients in the first column. CV (5 percent) is the 5 – percent critical value for the null hypothesis of no break from Hansen.

Table 4. Benchmark Structural Break Tests for Japanese Macroeconomic Variables 1955:3 to 2001:1

	Conditional Mean			Conditional Variance: Break Only			Conditional Variance: Trend and Break		
Series	p-value	Break date	Confidence interval	p-value	Break date	confidence interval	p-value trend	p-value break	Break date
Real GDP	0.00	1973:1	1972:1-1973:3	0.01	**1974:1**	1973:2-1977:3	0.00	0.00	1994:2
Consumption	0.00	1973:4	1973:2-1974:2	0.26	-	-	0.39	0.13	-
Investment	0.00	1970:1	1969:3-1970:3	0.00	**1966:2**	1965:4-1967:2	0.23	0.00	1963:1
Fixed investment	0.01	1990:4	1990:2-1991:2	0.00	**1969:3**	1968:4-1971:2	0.00	0.08	-
Nonresidential	0.60	-	-	0.00	**1969:2**	1968:4-1970:4	0.83	0.00	1969:2
Residential	0.00	1973:4	1973:2-1974:2	0.62	-	-	0.20	0.32	-
Inventory/GDP	0.00	1974:3	1974:1-1975:1	0.00	**1975:2**	1975:1-1977:3	0.64	0.00	1975:2
Exports	0.00	1978:1	1977:3-1978:3	0.01	**1978:3**	1977:4-1982:3	0.97	0.60	-
Imports	0.00	1972:1	1971:3-1972:3	0.01	**1973:1**	1971:4-1976:3	0.02	0.67	-
Government Spending	0.00	1973:3	1973:1-1974:1	0.30	-	-	0.07	0.02	1978:2
Inflation rate	0.00	1969:3	1969:1-1970:1	0.00	**1963:3**	1963:2-1964:3	0.05	0.20	-
Real exchange rate	0.37	-	-	0.03	**1992:4**	1987:2-1994:2	0.99	0.64	-
Nominal exchange rate	0.00	1986:3	1986:1-1987:1	0.00	**1995:1**	1991:4-1995:3	0.35	0.19	-
Short interest rate	0.49	-	-	0.00	**1982:2**	1982:1-1984:4	0.00	0.00	1973:2
10-year bond rate	0.02	1990:4	1990:2-1991:2	0.25	-	-	0.08	0.04	1995:3
Monetary base	0.01	1996:2	1995:4-1996:4	0.31	-	-	0.82	0.38	-
M2+CDs	0.00	1990:2	1989:4-1990:4	0.03	**1980:3**	1980:1-1983:2	0.00	0.22	-

Table 5. Benchmark Structural Break Tests for Japanese Macroeconomic Variables 1975:2 to 2001:1

	Conditional Mean			Conditional Variance: Break Only			Conditional Variance: Trend and Break		
Series	p-value	Break date	Confidence interval	p-value	Break date	confidence interval	p-value trend	p-value break	Break date
Real GDP	0.00	1991:1	1990:3-1991:3	0.00	**1994:4**	1992:2-1995:3	0.96	0.25	-
Consumption	0.00	1997:1	1996:3-1997:3	0.05	**1994:4**	1989:4-1996:1	0.95	0.89	-
Investment	0.00	1995:4	1995:2-1996:2	0.99	-	-	0.62	0.82	-
Fixed investment	0.00	1990:4	1990:2-1991:2	0.43	-	-	0.33	0.53	-
Nonresidential	0.02	1997:1	1996:3-1997:3	0.00	**1997:2**	1995:4-1998:1	0.87	0.26	-
Residential	0.04	1979:3	1979:1-1980:1	0.47	-	-	0.86	0.75	-
Inventory/GDP	0.00	1997:2	1996:4-1997:4	0.05	**1993:3**	1992:3-1996:1	0.23	0.01	1993:3
Exports	0.69	-	-	0.80	-	-	0.85	0.83	-
Imports	0.01	1980:2	1979:4-1980:4	0.20	-	-	0.27	0.06	-
Government Spending	0.00	1996:2	1995:4-1996:4	0.01	**1997:2**	1994:3-1998:1	0.00	0.04	1988:2
Inflation rate	0.35	-	-	0.13	-	-	0.98	0.67	-
Real exchange rate	0.39	-	-	0.05	**1992:4**	1987:1-1994:2	0.97	0.61	-
Nominal exchange rate	0.00	1986:3	1986:1-1987:1	0.00	**1995:1**	1990:1-1995:3	0.63	0.72	-
Short interest rate	0.17	-	-	0.00	**1993:4**	1993:3-1995:2	0.00	0.44	-
10-year bond rate	0.02	1990:4	1990:2-1991:2	0.25	-	-	0.08	0.04	1995:3
Monetary base	0.00	1997:2	1996:4-1997:4	0.03	**1997:2**	1987:1-1997:3	0.99	0.31	-
M2+CDs	0.00	1990:2	1989:4-1990:4	0.51	-	-	0.11	0.29	-

Table 6. McConnell and Perez-Quiros' Tests for Structural Change on Variances

1955:3-2001:1	1975:2 -2001:1
Sup W_T	Sup W_T
15.20	19.65
Estimated break date	Estimated break date
1974:1	**1994:4**

Table 7. Bai and Perron's Tests for Multiple Structural Changes on Means
1955:3 to 2001:1

Sup F_T (1)	Sup F_T (2)	Sup F_T (3)	Sup F_T (4)	Sup F_T (5)
94.66*	53.75*	36.17*	26.01*	22.32*
Sup F_T (2|1)	Sup F_T (3|2)	Sup F_T (4|3)		
11.6*	4.08	0.53		

Number of Breaks Selected	Estimates with Two Breaks	
2	$\hat{T}_1$	$\hat{T}_2$
	1973:1	**1991:2**
	(1972:1-1974:1)	(1983:3-1994:2)

1. All tests allow for heteroskedasticity and autocorrelation in the disturbances, AR(1) prewhitening, and 15 percent of sample trimmed.
2. * represents significance at the 5% level.
3. We use a 5% size for the sequential test sup F (*n*+1|*n*).
4. In parentheses are the 95% confidence intervals for $\hat{T}_i$.

Table 8. Bai, Lumsdaine and Stock's VAR Structural Change on Variances

1955:3-2001:1	1975:2 -2001:1
QLR p-value **0.00**	QLR p-value **0.00**
Estimated break date	Estimated break date
1975:2	**1994:3**
67% confidence interval	67% confidence interval
1974:1 – 1976:3	**1993:3 – 1995:3**

Table 9. Bayesian Gibbs-sampling approach to a three-state Markov-switching mean-variance model of quarterly real GDP growth 1955:3 to 2001:1

parameter	Posterior Mean	Standard deviation	median
ϕ_1	-0.0669	0.1021	-0.0670
ϕ_2	0.1126	0.1002	0.1093
σ_1^2 (1995:1-2001:1)	1.8978	0.8126	1.7146
σ_2^2 (1975:2-1994:4)	0.4228	0.0904	0.4092
σ_3^2 (1955:3-1974:1)	1.7689	0.3505	1.7137
μ_1 (1995:1-2001:1)	0.1037	0.2642	0.1479
μ_2 (1975:2-1994:4)	0.9526	0.1035	0.9585
μ_3 (1955:3-1974:1)	2.1963	0.1956	2.2026

Table 10. Estimates for Japanese Closed Economy Monetary Policy Rule

	ψ_1	ψ_2	α	ρ
$R_t = \rho R_{t-1} + (1-\rho)(\alpha + \psi_1 E[\pi_{t+4}] + \psi_2 E[x_{t+1}]) + \varepsilon_t$				
Quarterly Data 1961:1 – 2001:1 Output: Real GDP, Inflation: GDP deflator				
1961:1- 1975:1	-0.85	0.67	11.85*	1.16*
	(0.46)	(0.50)	(2.96)	(0.06)
1975:2- 1992:4	**2.08***	-0.14	1.34*	0.82*
	(0.55)	(0.60)	(1.20)	(0.05)
1993:1- 2001:1	**-0.67***	**0.38***	-0.01	0.97*
	(0.27)	**(0.18)**	(0.27)	(0.01)
$R_t = \rho R_{t-1} + (1-\rho)(\alpha + \psi_1 E[\pi_{t+12}] + \psi_2 E[x_t]) + \varepsilon_t$				
Monthly Data 1978:1 – 2001:12 Output: Industrial Production, Inflation: CPI				
1979:1- 1992:12	**2.04***	0.83	3.88*	0.98*
	(0.98)	(0.53)	(1.53)	(0.01)
1993:1- 2001:12	**0.32***	0.02	0.18*	0.99*
	(0.07)	(0.03)	(0.06)	(0.00)

Table 11. Estimates for Japanese Open Economy Monetary Policy Rule

	ψ_1	ψ_2	ψ_3	α	ρ
$R_t = \rho R_{t-1} + (1-\rho)(\alpha + \psi_1 E[\pi_{t+4}] + \psi_2 E[x_{t+1}] + \psi_3 E[e_t]) + \varepsilon_t$					
Quarterly Data 1961:1 – 2001:1 Output: Real GDP, Inflation: GDP deflator					
1975:2- 1992:4	0.17	**0.62***	**-0.17***	5.15*	0.77*
	(0.11)	**(0.25)**	**(0.04)**	(0.29)	(0.03)
1993:1- 2001:1	**0.83***	**-0.29***	**0.10***	1.32*	0.92*
	(0.11)	**(0.06)**	**(0.02)**	(0.10)	(0.02)
$R_t = \rho R_{t-1} + (1-\rho)(\alpha + \psi_1 E[\pi_{t+12}] + \psi_2 E[x_t] + \psi_3 E[e_t]) + \varepsilon_t$					
Monthly Data 1978:1 – 2001:12 Output: Industrial Production, Inflation: CPI					
1979:1- 1992:12	**0.54***	0.13	**-0.31***	4.63*	0.94*
	(0.14)	(0.14)	**(0.06)**	(0.46)	(0.01)
1993:1- 2001:12	**0.70***	-0.02	**0.09***	0.07	0.99*
	(0.16)	(0.04)	**(0.03)**	(0.13)	(0.00)

1. For the quarterly data model, we all can not reject the null hypothesis for test of overidentifying restrictions (the validity of instruments). The instruments are 1, lagged value (t-1, t-2, t-3, t-4) of the call rate, inflation, the output gap, the difference of exchange rate, and the commodity price inflation (Japanese wholesale price index).
2. For the monthly data model, we all can not reject the null hypothesis for test of overidentifying restrictions (the validity of instruments). The instruments are 1, lagged value (t-1, t-2, t-3, t-4, t-5, t-6, t-9, t-12) of the call rate, inflation, the output gap, the difference of exchange rate, and the commodity price inflation (Japanese wholesale price index).
3. The numbers in the parentheses mean standard errors.

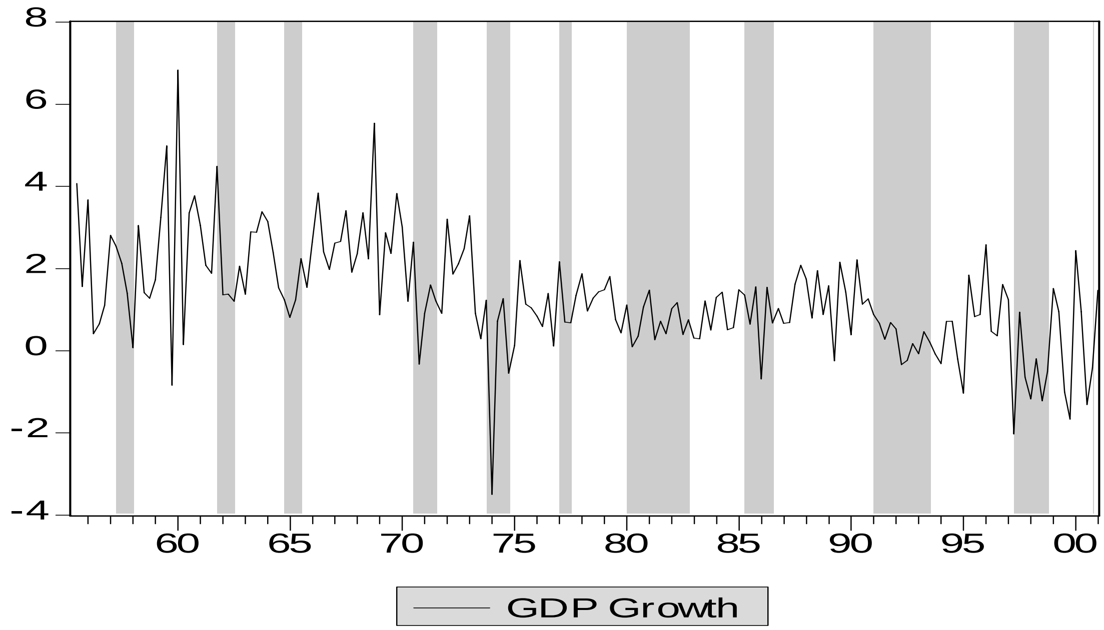

Figure 1. A. Quarter–To–Quarter Growth Rates in Japan real GDP from 1955:3 to 2001:1.

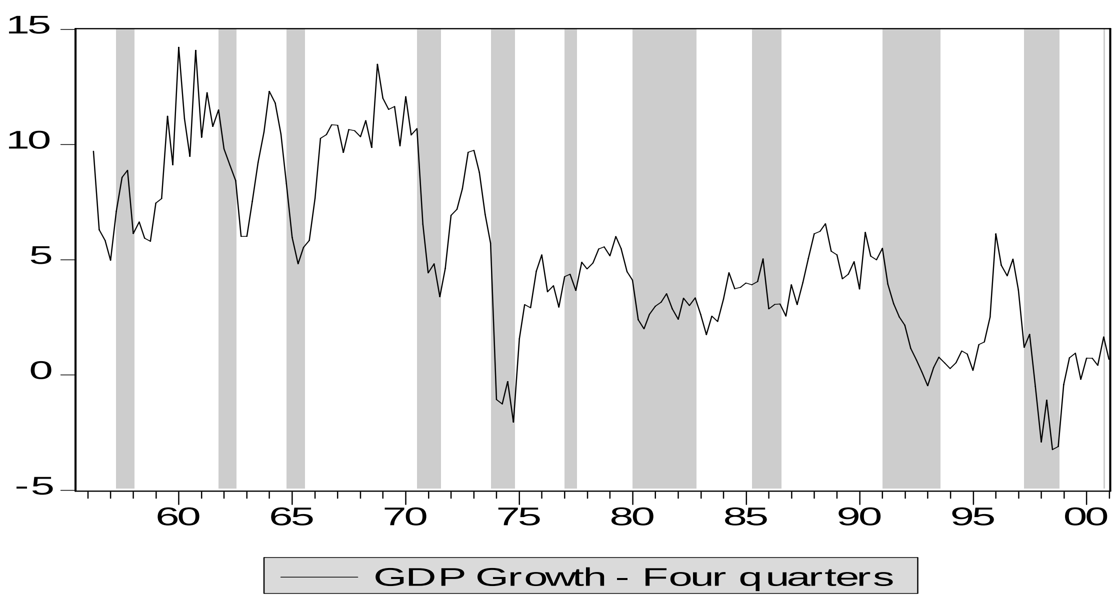

Figure 1. B. Four–Quarter Growth Rates in Japan real GDP from 1955:3 to 2001:1

Notes: Shaded area represents recessions (from peak to through of business cycle) in Japan. Reference dates are from ESRI (Official report for Japan business cycle), website: http://www.esri.cao.go.jp/en/stat/di/041112rdates.html.

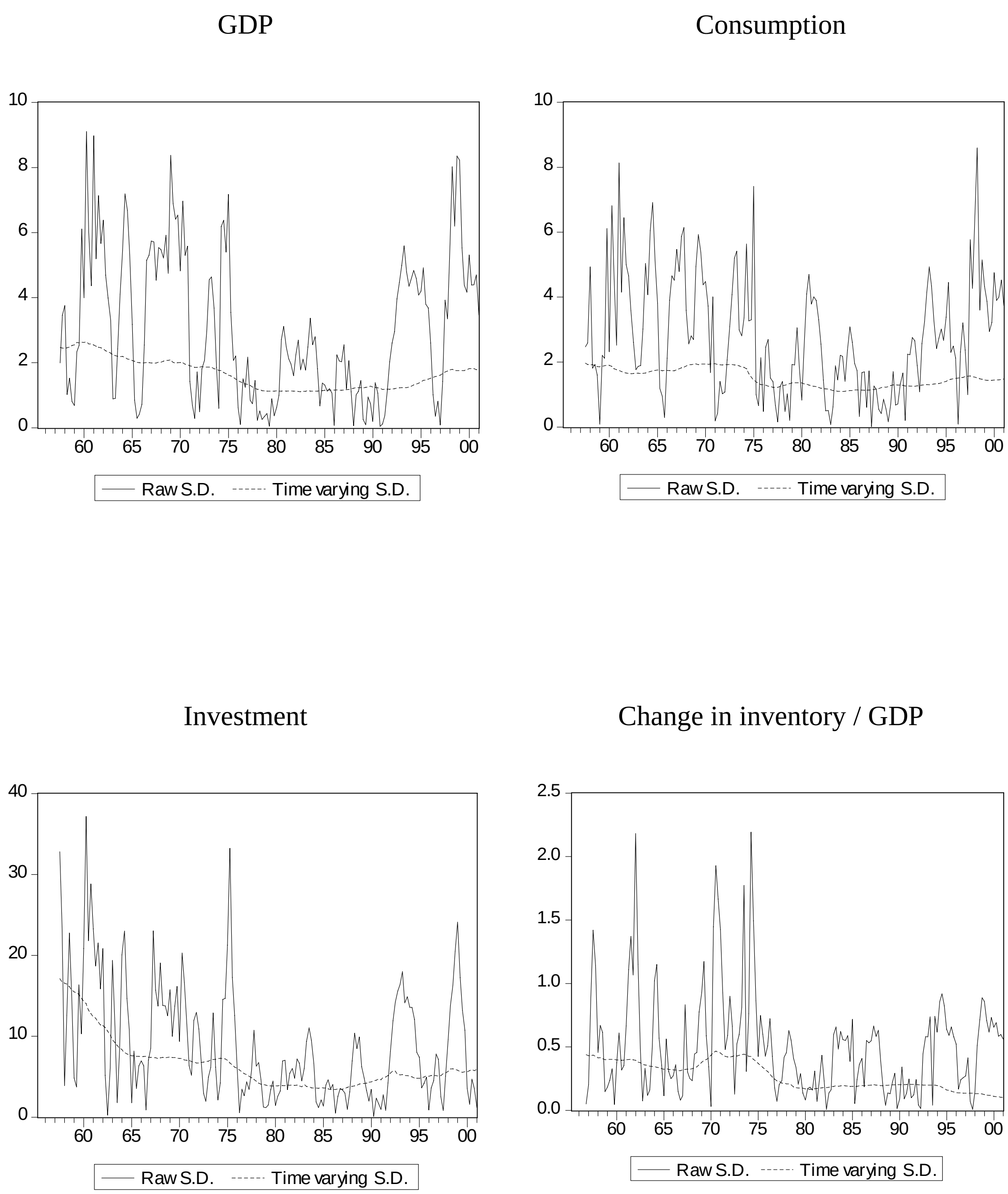

Figure 2. A. Time Varying Standard Deviations

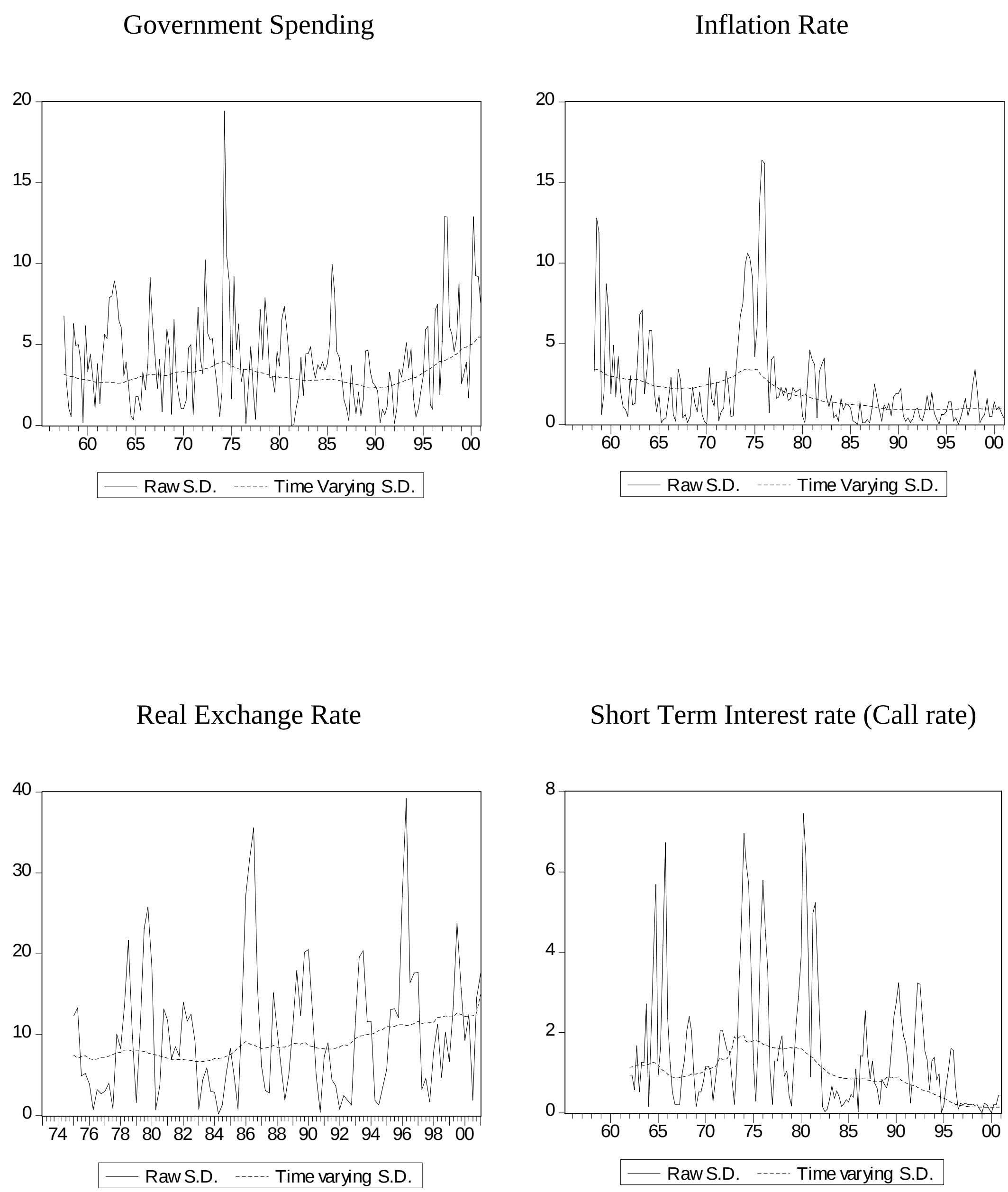

Figure 2. B. Time Varying Standard Deviations

Break Hypothesis Trend Hypothesis

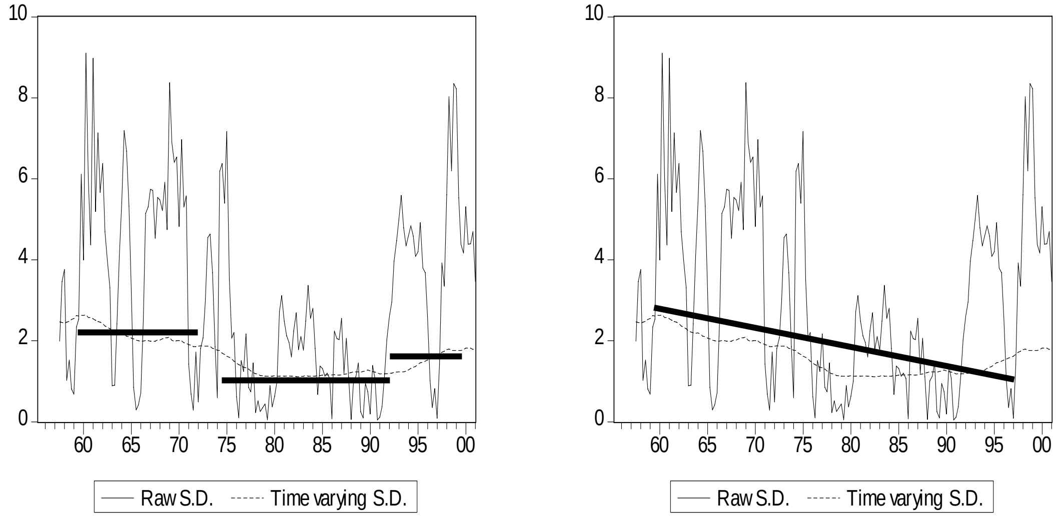

Figure 3. A. Conditional Variance: Trend and Break Test

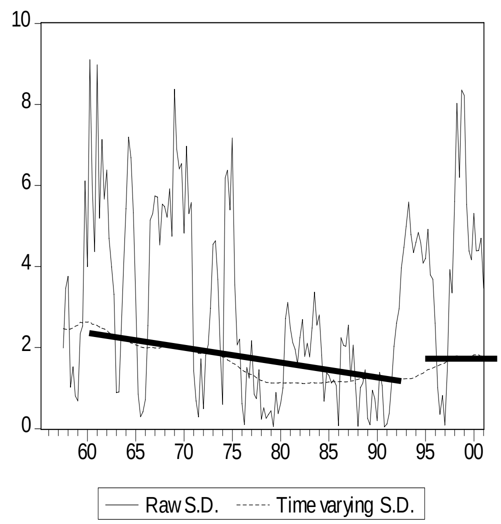

Figure 3. B. Conditional Variance: Trend and Break Test
Japan Real GDP 1955:3 to 2001:1

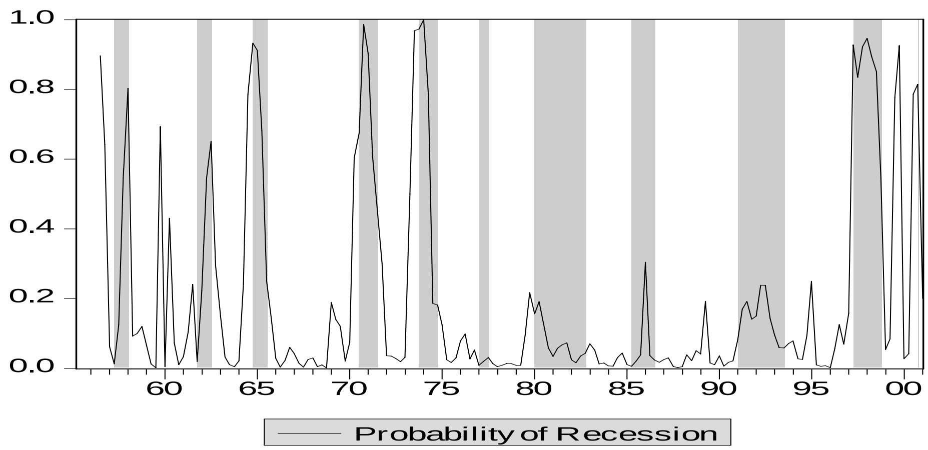

Figure 4. A. Hamilton's Two-State Markov-Switching Mean Model
Probability of recession

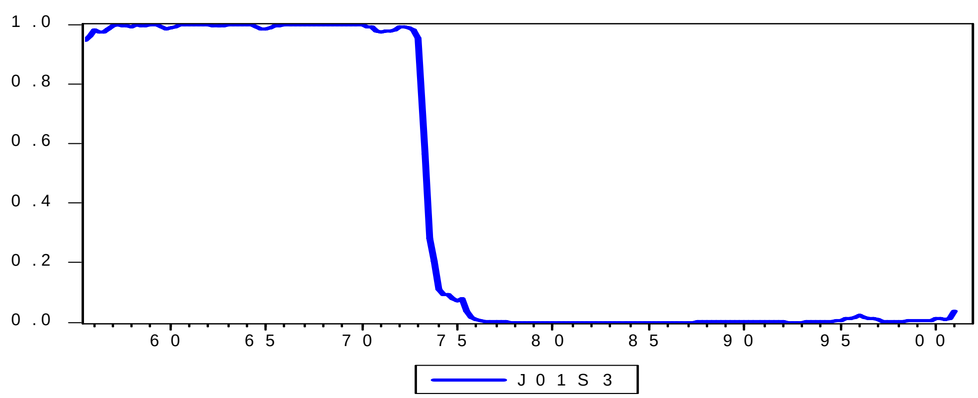

Figure 4. B. Three-State Markov-Switching Mean-Variance Model
- Probability of High Growth - High Volatility Regime

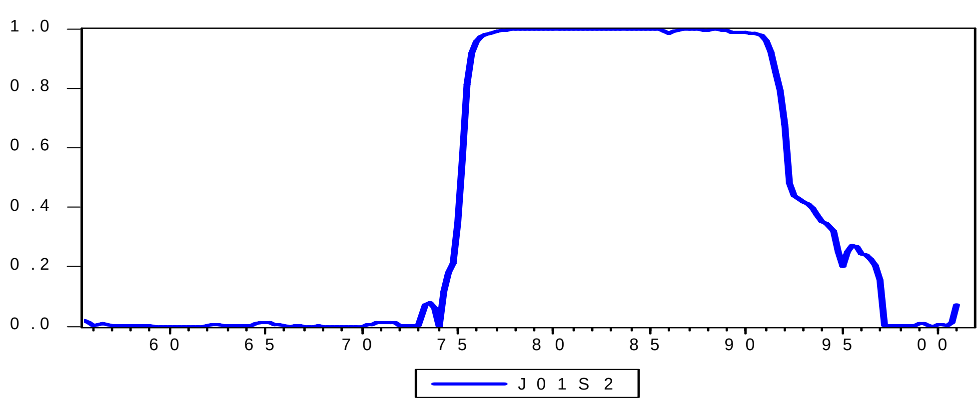

Figure 4. C. Three-State Markov-Switching Mean-Variance Model
- Probability of Medium Growth - Low Volatility Regime

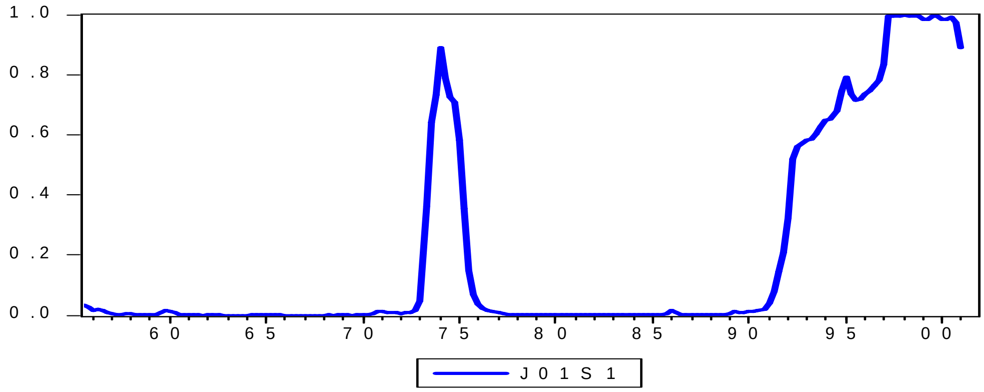

Figure 4. D. Three-State Markov-Switching Mean-Variance Model - Probability of Low Growth - High Volatility Regime

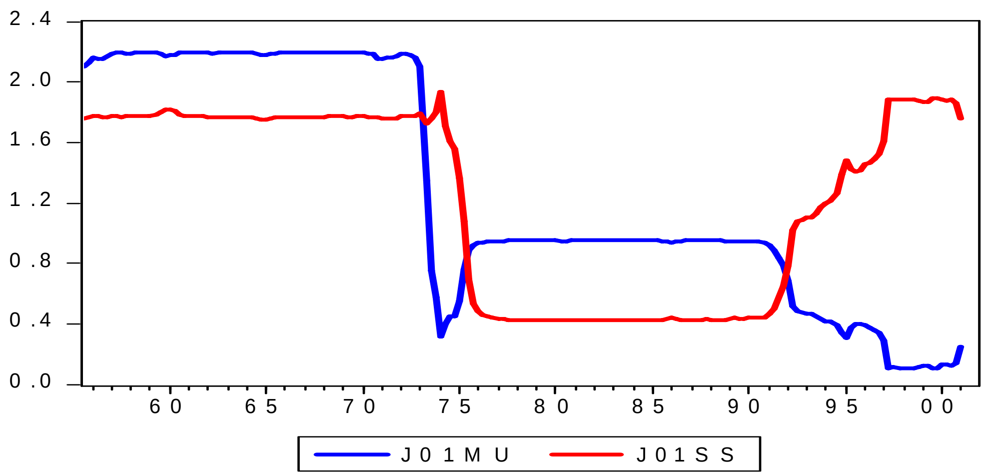

Figure 4. E. Three-State Markov-Switching Mean-Variance Model - Quarterly GDP Growth Mean and Variance

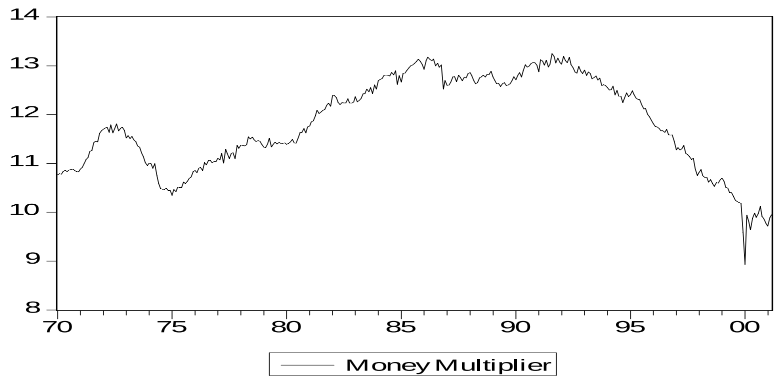

Figure 5. M2 Multiplier, 1970 to 2001

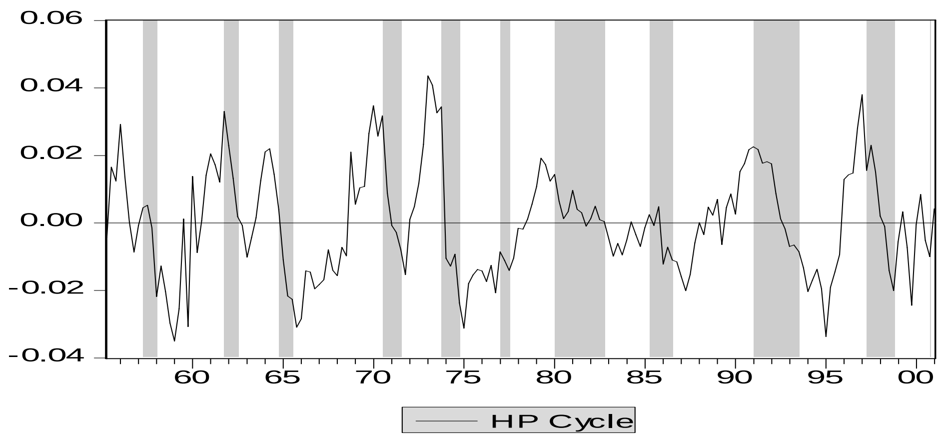

Figure 6. Output Gap computed by HP filter, 1955 to 2001

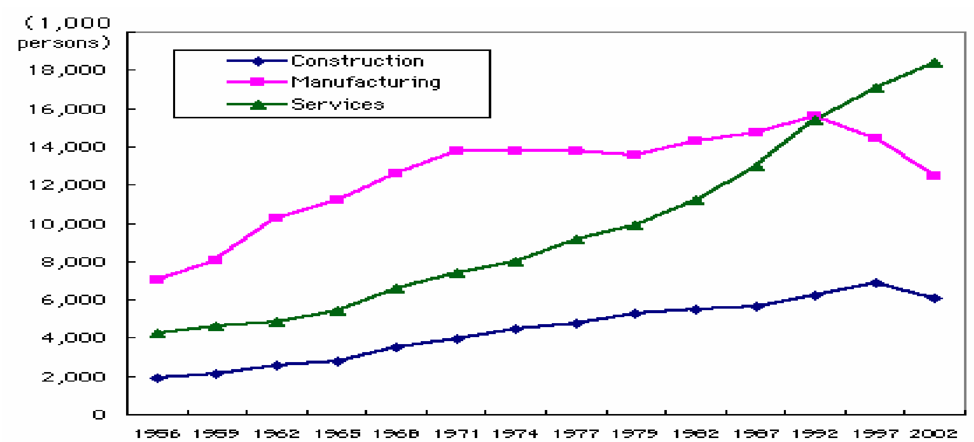

Figure 7. Trend of number of persons engaged in work by major industry

Data source: Japan Statistics Bureau -The 2002 Employment Status Survey

http://www.stat.go.jp/english/data/shugyou/2002/kakuhou/youyaku.htm#1

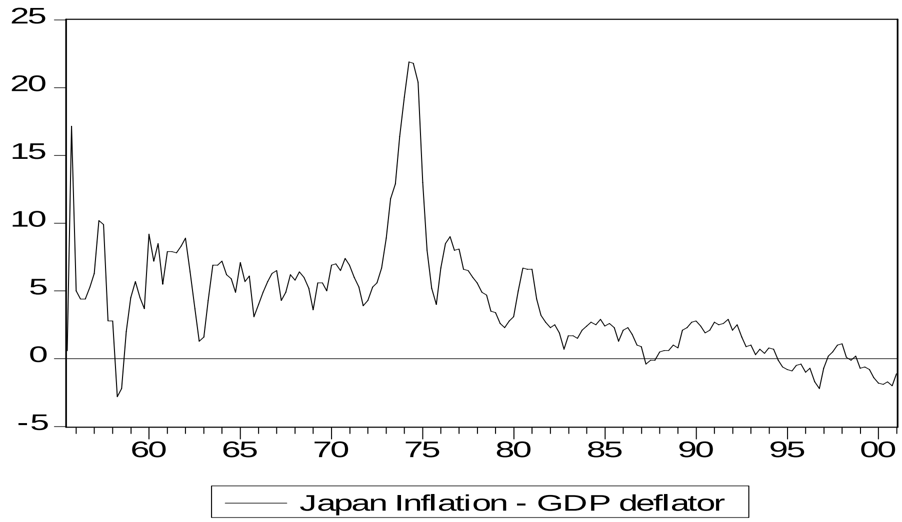

Figure 8. Japan Inflation rate – GDP Deflator 1955 to 2001

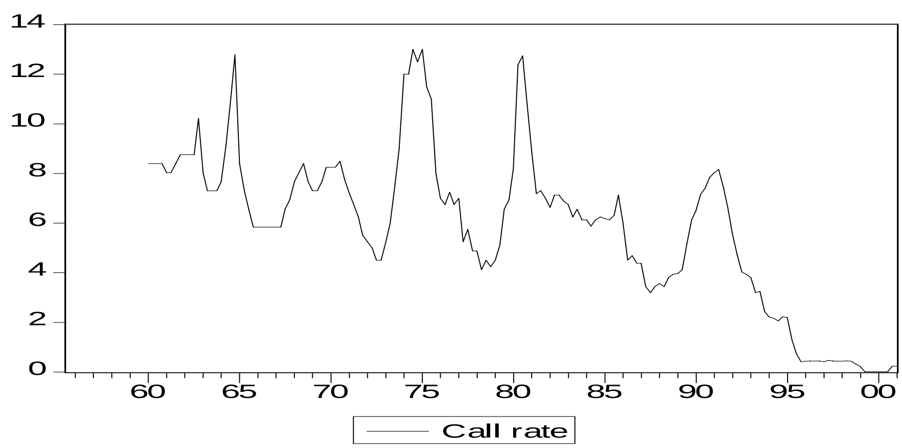

Figure 9. Japanese call rate 1960 to 2001

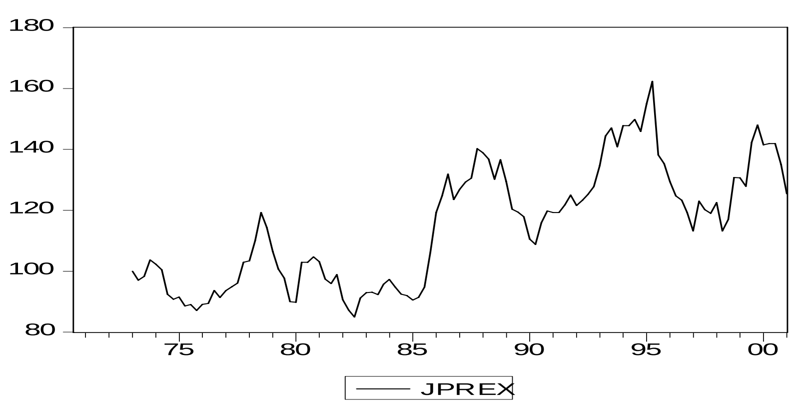

Figure 10. Japanese real exchange rate 1973 to 2001

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Appendix 1

We use stochastic volatility model following geometric random walks described by Stock and Watson (2002) to estimate the smoothed instantaneous standard deviations (σ_t) with random-walk time-varying autoregressive coefficients (ϕ_{jt}) as shown in Figure 2.

$$y_t = \mu_t + \sum_{j=1}^{4} \phi_{jt} y_{t-j} + \sigma_t e_t$$
$$\phi_{jt} = \phi_{jt-1} + \gamma_t v_{jt}$$
$$\ln \sigma_t^2 = \ln \sigma_{t-1}^2 + \omega_t$$

where $e_t, v_{1t}, ... v_{4t}$ are $i.i.d.N(0,1)$. And ω_t is $N(0, \omega_1^2)$ with probability p and $N(0, \omega_2^2)$ with probability *1-p*. The model shows the output shock is the sum of two underlying shocks. One is from a normal distribution (e_t); the other is from a mixture of normal distribution with smaller variance (ω_1^2) with probability *p* and larger variance (ω_2^2) with probability 1-*p* to catch fat-tailed disturbancees. The series y_t is standardized before the computations, and we set, scale factor, γ_t = 7 / T, which is consistent with Stock and Watson's previous estimates of parameter drift in autoregressions. We set ω_1^2 = 0.04, ω_2^2 = 0.2 and p = 0.95. This nonGaussian smoother for the time-varying parameters is computed using Markov Chain Monte Carlo (MCMC) method. Assume Y denotes $y_1, ... y_T$; A denotes { ϕ_{jt}, where j = 1, …4, t = 1, …T}; and S denotes $\sigma_1, ... \sigma_T$. The MCMC iterates between the three conditional distributions of Y| A, S; of A| Y, S; and S| A, Y. The first and second distributions are normal but the third one is nonnormal which is computed by a mixture of normal distribution. We use the log chi-square distribution to match the first four moments for the mixture means and variances. Initial conditions were set by a flat prior and a diffuse conjugate prior was used for the parameter values.