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Trend Inflation in the Japanese pre-2000s: A Markov-Switching DSGE Estimation *

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Abstract

In Japan, the inflation rate declined to near-zero, whereas the monetary policy faced a zero lower bound (ZLB) in the 1990s. We examine whether trend inflation had fallen to near-zero prior to the ZLB. To achieve this, we estimate Japanese pre-2000 trend inflation developing a Markov-switching New Keynesian dynamic stochastic general equilibrium (DSGE) model in which non-zero trend inflation is explicitly incorporated as a Markov chain state. Our estimation results indicate that (i) the trend inflation remained broadly stable at 3.0–3.5 percent from the 1960s to the late 1970s prior to the second Global Oil Crisis. (ii) Then, over time, trend inflation gradually declined to nearly 1.0 percent toward the Plaza Accord in 1985. (iii) Up until 1997 when the ZLB was hit, trend inflation hovered well above zero, mostly near 1.0 percent.

JEL Classification: E31, E52, C54

Keywords: Trend Inflation, Markov-Switching Dynamic Stochastic General Equilibrium Model, Deflation, Monetary Policy, Lost Decade of Japan.

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1 Introduction

Japan's *lost decade* – a protracted stagnation of the economy since the 1990s – has attracted considerable interest of economists. A seminal work in the literature is Hayashi and Prescott (2002), who argued that real and supply-side factors of the economy, specifically a slowdown in technological progress and reduction in labor hours, can broadly account for low GDP growth rates during the periods. In contrast, Nishizaki, Sekine and Ueno (2014) and many others detected multiple factors that caused the economic stagnation over the decade, including declines in inflation expectations.[1] The debate on the major driving factors that gave rise to Japan's slow growth since the 1990s is still ongoing.

Looking at the nominal aspect of the economy apart from the real factors, deflation began in the late 1990s in Japan. Figure 1 presents monthly year-on-year changes in the consumer price index (CPI). On average, CPI inflation was 4.3 percent before 1960, 9.1 percent in the 1970s and fell to nearly one percent, remaining low from 1980 to 2005. In the 2000s, CPI inflation hovered around zero. In the meantime, nominal interest rates declined, hitting the zero lower bound (ZLB) in the late 1990s.

Against this backdrop, this study aims to contribute to the literature by examining whether trend inflation has fallen to near zero *prior to* the ZLB. Although the data reveal the possible declines in trend inflation that pre-date the 2000s, the timing and rate of such declines require further exploration. To this end, we focus on pre-2000s macroeconomic time series data, including high inflation eras since the 1960s, particularly the 1970s, during which CPI inflation reached 25 percent in the wake of Global Oil Crises.

We specifically estimate a small-scale New Keynesian (NK) model using pre-2000s Japanese data. However, with higher than 5 percent average inflation, linearizing the model around the zero-inflation steady-state can elicit significant estimation bias, as argued by Cogley and Sbordone (2008) and Ascari and Sbordone (2014, hereafter denoted as AS 2014). Because our study covers transition periods from high- to low-trend inflation eras, we estimate a generalized New Keynesian (GNK) model that explicitly allows non-zero trend inflation as a Markov-switching unobserved state.

The main results reveal that trend inflation remained broadly stable at 3.0–3.5 percent from 1958 to 1977. Then, after the first Global Oil Crisis, trend inflation started to decline, nearing 1.0 percent by the Plaza Accord in 1985. The results are consistent with the narra-

[1] See Kaihatsu and Kurozumi (2014) for similar arguments based on an estimated structural macroeconomic model with financial market frictions. On the other hand, Kato and Nishiyama (2005) highlighted the role of monetary policy in the 1990s.

tives which suggest that the power shifted from the government to the Bank of Japan (BoJ) where more hawkish officials took office in 1974. From 1985 to 1997, the trend inflation hovered well above zero, mostly at near 1.0 percent, except for the short bout of higher inflation during the asset price "bubble" period. Thus, little evidence was found that trend inflation fell to near zero percent or even to the negative prior to 1996 when the ZLB was hit in Japan.[2]

Separately, our results detect regime switches regarding monetary policy stance and supply-shock volatility. The estimated regime-switching monetary policy rule suggests that the monetary policy stance has become "hawkish" in the sense that the short-term policy rate responded more to inflation since the 1980s than in the earlier periods. Meanwhile, the results reveal that the supply-shock volatility surged at the time of the two global oil crises in 1973 and 1979, while it remained low and stable during the entire non-crisis periods.

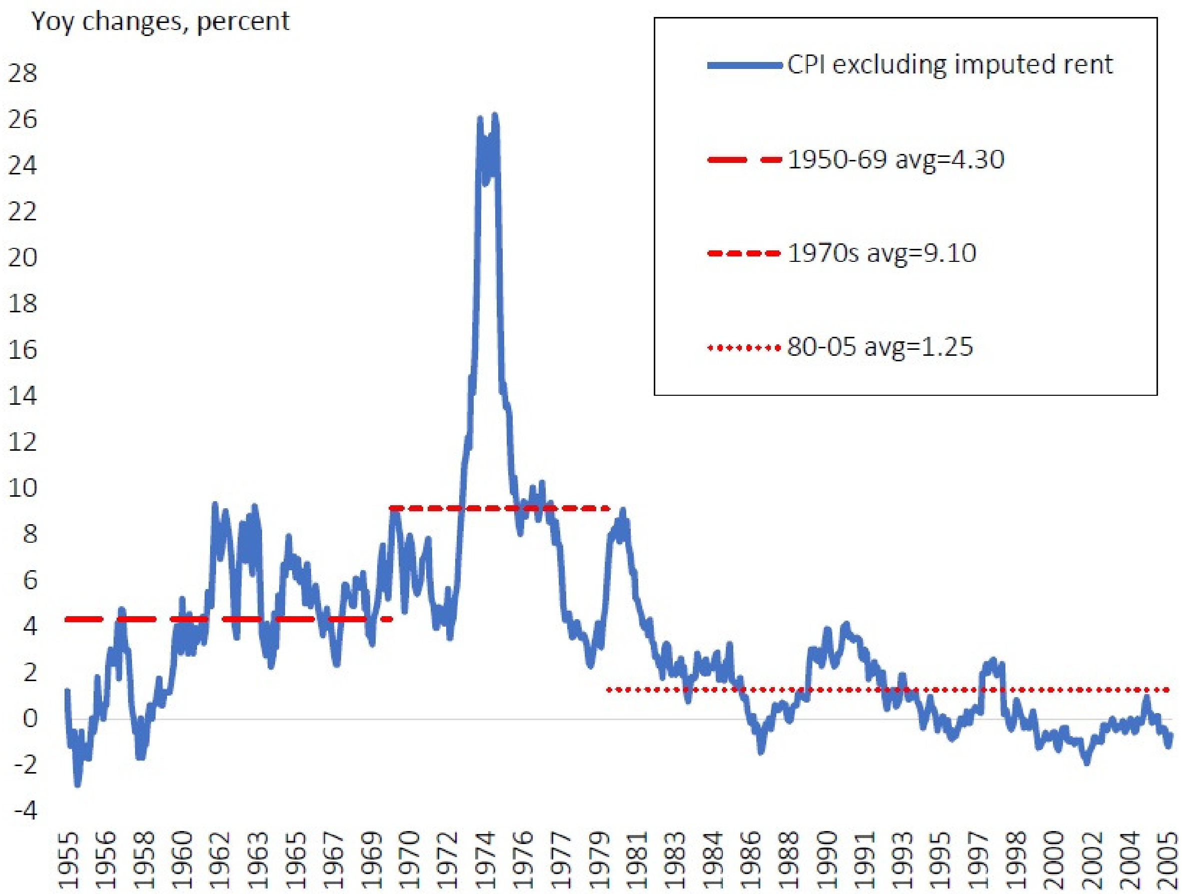
This study is related to three strands of literature. First, because we estimate a GNK dynamic stochastic general equilibrium (DSGE) model using data in which trend inflation is likely to be significantly above zero, we apply the non-zero trend inflation model proposed by AS (2014) rather than the standard NK models.[3] Second, regarding the estimation procedure, we apply a Markov-switching rational expectations framework that has been developed in the literature on Bayesian inference for structural macroeconomic models. Specifically, we follow the framework introduced by Maih (2014) and later demonstrated by Bjørnland, Larsen, and Maih (2018, hereafter denoted as BLM 2018). Third, many studies, including Sugo and Ueda (2008), Aruoba, Cuba-Borda, and Schorfheide (2018), Abe, Fueki and Kaihatsu (2019), Hirose (2020), and Iiboshi, Shintani, and Ueda (2022), estimate NK-DSGE models using Japanese data. In contrast to ours, however, most of the earlier studies relied on non-regime-switching DSGE models.[4] The most closely related studies are Aruoba, Cuba-Borda and Schorfheide (2018) and Abe, Fueki and Kaihatsu (2019), both of which allow their models to regime-switch, but they do not consider the non-zero trend inflation discussed by AS (2014) and the current study.

[2]Our data sample period ends at 1997. In 1996, the effective lower bound of nominal interest rate was hit and then the zero interest rate policy was officially introduced in February 1999.

[3]See Woodford (1999) and Clarida, Gali, and Gertler (1999), among many others, for the standard NK models

[4]See also, Ichiue, Kurozumi and Sunakawa (2013), Kaihatsu and Kurozumi (2014) and Fueki et al. (2016). All these DSGE estimations use the data starting later than 1980, thus excluding the high inflation periods.

Figure 1: CPI Inflation in Japan



Note: Monthly year-over-year changes in Consumer Price Index, excluding imputed rent.
Source: Statistics Bureau of Japan

The remainder of this paper is organized into five sections. Section 2 presents the GNK model applied to this study. Section 3 establishes the estimation procedure used. Section 4 details the results of the analyses and Section 5 discusses the interpretations of the main estimation results relying on narratives. Section 6 concludes.

2 A New Keynesian model with non-zero trend inflation

As depicted in Figure 1, as average inflation was quite high – far above 2 percent – before the 1970s, we adopt a small-scale GNK-DSGE model developed by AS (2014), which explicitly incorporates non-zero trend inflation $\bar{\pi}_t$. Formally, trend inflation is defined in terms of the infinite horizon forecast as follows:

$$\bar{\pi}_t = \lim_{j \to \infty} E_t \pi_{t+j}.$$

In the context of our GNK-DSGE model, trend inflation is the steady state inflation under the rational expectations equilibrium.

Our model is a log-linearized version of the full model developed by AS (2014). We first articulate the supply-side of the economy by presenting a generalized Phillips curve. The four equations of the generalized Phillips curve are derived from the log-linearized optimal conditions of the Calvo-price setting firms and the standard resource constraint. The model's set-up is presented in the Appendix.

2.1 Aggregate supply side: GNK Phillips curve

A standard small-scale NK model comprises three equations, the NK Phillips curve, the consumption–output Euler equation, and a monetary policy rule. In our model, however, the generalized Phillips curve is expressed by four separate equations. We let P_t be the price index of final goods and define $\pi_t = P_t / P_{t-1}$ as gross inflation rate and $\bar{\pi}_t$ as gross trend inflation rate. In contrast to early studies, we allow $\bar{\pi}_t$ to vary over time, such that $\bar{\pi}_t = \bar{\pi}\left(\mathcal{S}_t^{\bar{\pi}}\right)$ where the Markov-switching state $\mathcal{S}_t^{\bar{\pi}}$ will be defined later combined with other unobserved states. Further, y_t denotes output which is equal to consumption in this model, and z_t^{AS} indicates the aggregate supply shock. In the remainder of the paper, the variables with tilde denote log-deviations from (Markov-switching) steady-state values of the variable, i.e., $\tilde{x}_t = ln(x_t/\bar{x}_t)$. The four equations of the GNK Phillips curves are as follows:

$$\tilde{\pi}_t = \frac{1-\Gamma_a(\bar{\pi}_t)}{\Gamma_a(\bar{\pi}_t)}(\tilde{\psi}_t - \tilde{\phi}_t) + \rho\tilde{\pi}_{t-1}, \tag{1}$$

$$\begin{aligned}\tilde{\psi}_t &= \{1-\beta\Gamma_b(\bar{\pi}_t)\}\left\{\chi\tilde{s}_t + (1+\chi)\left(\tilde{y}_t - z_t^{AS}\right)\right\} \\ &\quad +\beta\Gamma_b(\bar{\pi}_t)\left(E_t\tilde{\psi}_{t+1} + \varepsilon E_t\tilde{\pi}_{t+1} - \rho\varepsilon\tilde{\pi}_t\right), \end{aligned} \tag{2}$$

$$\begin{aligned}\tilde{\phi}_t &= \{1-\beta\Gamma_a(\bar{\pi}_t)\}(1-\sigma)\tilde{y}_t \\ &\quad +\beta\Gamma_a(\bar{\pi}_t)\left\{E_t\tilde{\phi}_{t+1} + (\varepsilon-1)E_t\tilde{\pi}_{t+1} + \rho(1-\varepsilon)\tilde{\pi}_t\right\}, \end{aligned} \tag{3}$$

$$\tilde{s}_t = \Gamma_b(\bar{\pi}_t)\tilde{s}_{t-1} + \frac{\varepsilon\bar{\pi}_t\Gamma_a(\bar{\pi}_t)}{1-\Gamma_a(\bar{\pi}_t)}\tilde{\pi}_t - \varepsilon\rho\Gamma_b(\bar{\pi}_t)\tilde{\pi}_{t-1}, \tag{4}$$

where $\tilde{\psi}_t$ and $\tilde{\phi}_t$ represent auxiliary variables inherited from the first-order condition of the Calvo-price setting firms. In (4), $\tilde{s}_t$ denotes the degree of the real distortion stemming from price dispersion. AS (2014) demonstrated that (i) this additional endogenous variable $\tilde{s}_t$ shows up only in the "generalized" NK Phillips curve and (ii) plays an important role in creating richer dynamics if the trend inflation is away from zero.

Further, β, σ, and χ indicate the discount factor, the degree of risk aversion, and inverse labor supply elasticity, respectively. $\varepsilon > 1$ denotes the elasticity of substitution among intermediate goods. $\rho \in [0, 1)$ represents the degree of inflation indexation by non-optimizing price-setting firms. In the model, $\Gamma_a(\bar{\pi}_t) \equiv \theta\bar{\pi}_t^{(\varepsilon-1)(1-\rho)}$ and $\Gamma_b(\bar{\pi}_t) \equiv \theta\bar{\pi}_t^{\varepsilon(1-\rho)}$ wherein θ is a Calvo-parameter that represents the fraction of firms that cannot change prices in any given period. The model is linear in terms of endogenous variables while the parameters are nonlinear functions of trend inflation. By setting $\bar{\pi}_t = \bar{\pi} = 1$ with $\rho = \chi = 0$ and $\sigma = 1$ in (1)–(4), $\tilde{\psi}_t$, $\tilde{\phi}_t$ and $\tilde{s}_t$ will be eliminated from the simultaneous equations, resulting in a standard NKPC, which is $\tilde{\pi}_t = \beta E_t\tilde{\pi}_{t+1} + \theta^{-1}(1-\beta\theta)(1-\theta)\left(\tilde{y}_t - z_t^{AS}\right)$.

2.2 Aggregate demand side and structural shocks

The model closes with a standard consumption-output Euler equation and a monetary policy rule, such that

$$\tilde{y}_t = E_t\tilde{y}_{t+1} - \sigma^{-1}(\tilde{\imath}_t - E_t\tilde{\pi}_{t+1}) + z_t^{IS}, \tag{5}$$

$$\tilde{\imath}_t = \rho^{MP}\tilde{\imath}_{t-1} + \left(1-\rho^{MP}\right)(\alpha_y\tilde{y}_t + \alpha_{\pi,t}\tilde{\pi}_t) + e_t^{MP}, \tag{6}$$

where $\tilde{\imath}_t$ denotes nominal interest rate and α_π, α_y, and $\rho^{MP} \in [0,1)$ represent monetary policy response parameters. Following many early works, we allow Taylor rule's coefficient α_π to switch such that $\alpha_{\pi,t} = \alpha_\pi \left(\mathcal{S}_t^{MP}\right)$. The Markov-switching state $\mathcal{S}_t^{MP}$ is defined together with other states in subsection 2.3.

The model contains three structural shocks; the aggregate supply shock z_t^{AS}, the IS shock z_t^{IS}, and the monetary policy shock e_t^{MP}. While monetary policy shock e_t^{MP} is distributed as $N(0, \sigma_{MP}^2)$, the other two follow stationary first-order autoregressive processes such that,

$$z_t^x = \rho^x z_{t-1}^x + e_t^x,\ e_t^x \sim N(0, \sigma_x^2),$$

where $\rho^x \in [0,1)$ and $x \in \{AS, IS\}$. The aggregate supply (AS) shock contains shocks that affects the supply-side of the economy, such as technology shocks and various markup shocks, including oil price shocks. IS shock could be a composite of preference shock, fiscal policy shock, and shocks to the natural rate of interest.

2.3 Markov-switching states

In our model, we stipulate four independent Markov chain processes with the following Markov-switching states,

$$\begin{aligned} \mathcal{S}_t^{\bar{\pi}} &\in \{\text{High}, \text{Mid}, \text{Low}\}, &(7)\\ \mathcal{S}_t^{MP} &\in \{\text{Hawkish}, \text{Dovish}\}, &(8)\\ \mathcal{S}_t^{dy^*} &\in \{\text{High growth}, \text{Low growth}\}, &(9)\\ \mathcal{S}_t^{AS} &\in \{\text{High volatility}, \text{Low volatility}\}. &(10)\end{aligned}$$

First, we allow trend inflation $\bar{\pi}_t$ to switch according to a Markov chain process with $\mathcal{S}_t^{\bar{\pi}}$ taking three states, $\{\text{High}, \text{Mid}, \text{Low}\}$. Second, we allow two monetary policy states given by (8). We define a "hawkish" state as the periods during which the BoJ responded more sensitively to inflation. Specifically, Taylor rule's coefficient α_π in (6) follows a Markov chain process, and $\mathcal{S}_t^{MP}$ takes a high (i.e. hawkish) or low value (i.e. dovish). Third, we also allow the steady-state real per-capita GDP growth denoted by dy_t^* to follow a Markov chain process with $\mathcal{S}_{t,}^{dy^*}$ as given by (9). We explicitly consider $\mathcal{S}_t^{dy^*}$ because our data cover the 1950s–1960s, the post-war Japanese "economic miracle" era.[5] However, $\mathcal{S}_t^{dy^*}$ does not affect

[5] See Patrick and Rosovsky (1976) for discussions on the Japan's rapid growth in the 1960s.

any parameters in the model presented in this section, but in dy_t^*, which appears only in the observation equation presented later.

We also allow the volatility of the supply shock σ_{AS}^2 to switch according to a Markov chain process with $\mathcal{S}_{t,}^{AS}$ as given by (10). In our small-scale NK model, σ_{AS}^2 includes shocks arising from global oil markets. BLM (2018) emphasized the importance of considering the role of oil price volatility in accounting for business cycles. In the same spirit, we allow $\sigma_{AS}^2 = \sigma_{AS}^2(\mathcal{S}_t^{AS})$ to Markov-switch because (i) our sample periods include the two global oil crises experienced in the 1970s, and (ii) Japan's economy was heavily dependent on imported crude oil during this period.

3 Estimation methodology

3.1 Markov-switching rational expectations framework

The methodology employed in our study operates within a versatile Markov-Switching DSGE (MS DSGE) framework whose generic representation is given as follows:

$$E_t \sum p_{\mathcal{S}_t,\mathcal{S}_{t+1}} \mathbf{d}_{\mathcal{S}_t} \left(\mathbf{x}_{t+1}\left(\mathcal{S}_{t+1}\right), \mathbf{x}_t\left(\mathcal{S}_t\right), \mathbf{x}_{t-1}, \mathbf{e}_t\right) = 0, \tag{11}$$

where, $\mathbf{d}_{\mathcal{S}_t}$ is an $n_d \times 1$ vector of functions with arguments including $\mathbf{x}_{t+1}\left(\mathcal{S}_{t+1}\right)$, $\mathbf{x}_t\left(\mathcal{S}_t\right)$, $\mathbf{x}_{t-1}$, and $\mathbf{e}_t$.

In our log-linearized model, $\mathbf{d}_{\mathcal{S}_t}$ is expressed as a set of linear functions. $\mathcal{S}_t$ denotes the regime at time t, $\mathbf{x}_t$ represents an $n_x \times 1$ vector of endogenous variables, and $\mathbf{e}_t$ is an $n\varepsilon \times 1$ vector of Gaussian shocks with $\mathbf{e}_t \sim N(0, \mathbf{I}_{n_\varepsilon})$. $p_{\mathcal{S}_t,\mathcal{S}_{t+1}}$ signifies the transition probability from regime $\mathcal{S}_t$ at period t to $\mathcal{S}_{t+1}$ at $t+1$, with the constraint that $\sum_{\mathcal{S}_{t+1}=1}^{h} p_{\mathcal{S}_t,\mathcal{S}_{t+1}} = 1$. For our primary model, we adopt $h = 24$, reflecting three trend inflation states (high, mid and low), two Taylor coefficient states (hawkish and dovish), two potential GDP growth states (high growth and low growth), and two aggregate supply shock volatility states (high volatility and low volatility), amounting to a total of $3 \times 2 \times 2 \times 2 = 24$ possible regimes.

The true minimum-state-variable solution to (11) takes the following form

$$\mathbf{x}_t\left(\mathcal{S}_t\right) = \mathbf{T}_{\mathcal{S}_t}\left(\mathbf{x}_{t-1}, \mathbf{e}_t\right)$$

where $\mathbf{T}_{\mathcal{S}_t}$ is an unknown function.

Due to the complexity of the problem, analytical solutions are generally unattainable, even when $\mathbf{d}_{\mathcal{S}_t}$ is linear in its arguments. To tackle this, we employ perturbation techniques for Markov-Switching DSGE models (See e.g Maih 2015 or Chang, Maih and Tan 2021). The perturbation technique delivers an approximated solution of the form

$$\mathbf{x}_t\left(\mathcal{S}_t\right) \approx \bar{x}(\mathcal{S}_t) + \mathcal{H}_{\mathcal{S}_t}\mathbf{x}_{t-1} + \mathcal{G}_{\mathcal{S}_t}\mathbf{e}_t.$$

The stability of our MS DSGE models is determined using the concept of mean-square stability (MSS), following methodologies outlined by Farmer, Waggoner, and Zha (2011) and others.[6]

3.2 Data and Bayesian inference

Our dataset is in quarterly frequency, consisting of real per-capita GDP (y_t), CPI, excluding imputed rent, (P_t), and the official discount rate (logR_t). The real GDP and CPI, excluding imputed rent, are derived from 68SNA and from CPI (with 2015 as the base year index) databases. The official discount rate is taken from the Bank of Japan's Time Series Data Search website. The per-capita real GDP is computed as the real GDP divided by the total labor force. The log differences of the real per-capita GDP and CPI are used for the estimation. All data, including the official discount rate, are presented in annualized rates.

Our estimation framework includes the following observation equations summarized as follows:

$$\begin{bmatrix} 400\Delta\log y_t \\ 400\Delta\log P_t \\ 100\log R_t \end{bmatrix} = \begin{bmatrix} dy_t^* \\ \pi_t^* \\ \pi_t^* + r_t^* \end{bmatrix} + \begin{bmatrix} 4\left(\tilde{y}_t - \tilde{y}_{t-1} + z_t^{AS}\right) \\ 4\tilde{\pi}_t \\ 4\tilde{\imath}_t \end{bmatrix} + \begin{bmatrix} \eta_t^y \\ \eta_t^\pi \\ 0 \end{bmatrix}, \tag{12}$$

where $r_t^* = 400\left(1/\beta - 1\right) + \sigma^{-1}dy_t^*$ and η_t^{obs} is a measurement error distributed as $N(0, Var(\eta_t^{obs}))$, and $obs \in \{y, \pi\}$. As noted in the previous section, we assume that the steady-state real per-capita GDP growth rate $dy_t^* = dy^*(\mathcal{S}_t^{dy^*})$ and trend inflation $\bar{\pi}_t = \bar{\pi}\left(\mathcal{S}_t^{\bar{\pi}}\right)$ switch according to different Markov chain processes given by (7) and (9). For notational convenience, we redefine the *net* annual trend inflation rate as $\pi_t^* = 100 \times (\bar{\pi}_t - 1)$. Accordingly, we will report $\pi_{high}^* = 100 \times \{\bar{\pi}\left(\mathcal{S}_t^{\bar{\pi}} = \text{High}\right) - 1\}$, $\pi_{mid}^* = 100 \times \{\bar{\pi}\left(\mathcal{S}_t^{\bar{\pi}} = \text{Mid}\right) - 1\}$ and $\pi_{low}^* = 100 \times \{\bar{\pi}\left(\mathcal{S}_t^{\bar{\pi}} = \text{Low}\right) - 1\}$ in line with data counterpart $400\Delta\log P_t$ in (12). Similarly, $4\tilde{\imath}_t = 100\log(R_t/\bar{R}_t)$ connects the nominal interest rates in the model and in the

[6]Maih's (2015) perturbation method, along with the stability criterion, is implemented using the RISE toolbox for Matlab, accessible at https://github.com/jmaih/RISE_toolbox/

data.[7]

The sample period of our dataset is from 1958Q2 to 1997Q1. Because our interest lies in the transition of the Japanese economy from high- to low-trend inflation eras, including the 1970s when inflation was historically high, assessing this period is essential. The end of the sample period is chosen for the following three reasons. First, our key question is whether Japan's trend inflation declined before the ZLB. Nominal short-term interest rates, including the official discount rate and the overnight call rate, were cut to 0.5 percent in 1996, which were then considered the effective lower bound (ELB). Moreover, including data beyond 1997 can seriously distort the estimation results due to the nonlinearity arising from ZLB or ELB.[8] The second reason arises from the hike of the consumption tax rate from three percent to five percent in April 1997. Reflecting this exogenous shock, CPI inflation reveals a blip in 1997Q2 as depicted in Figure 1, providing a ground for excluding the data after 1997Q2. Third, because we are estimating a regime-switching model, ensuring consistent and continuous time series data is critically important. One advantage of our dataset is that all variables remain continuous on the same official basis, and neither artificial discontinuation nor connection of different time series is found, such as revision of the base-year, sampling method, and data definitions. In particular, longitudinal time series of real GDP covering the 1950s to the 2010s do not exist due to base year changes and other statistical revisions.

Before discussing the priors of the parameters to be estimated, we calibrate two parameters to avoid identification issues. We calibrate the discount factor $\beta = 0.999$ based on the sample medians of ex-post real interest rate and output growth rate. We set the inverse labor supply elasticity $\chi = 2$ following Hirose (2020). Table 1 summarizes the prior distributions of the parameters. Most of the priors for the structural parameters $(\varepsilon, \theta, \sigma, \rho)$ are taken from Hirose (2020). The main parameters of our interest are trend inflation, π^*_{high} π^*_{mid} and π^*_{low} of which priors are set to 5.0, 3.0, and 0.5 percent, respectively. In contrast to gamma distributions for π^*_{high} and π^*_{mid}, normal distribution is assumed for π^*_{low} in order not to over-restrict the domain of the lower bound of trend inflation. Assuming a normal distribution for π^*_{low} flexibly allows that negative trend inflation, if it is the case, can be estimated.

Regarding the priors for the Taylor rule parameters, we set $\alpha_y = 0.5$, α_π (Hawkish) $= 1.5$ and α_π (Dovish) $= 1.0$. Finally, considering the sample sub-period averages, we set dy^*(High growth) $= 8$ and dy^*(Low growth) $= 2$ percents, respectively.

[7] See A.4 in the appendix for greater details.

[8] A number of early studies, such as Aruoba, Cuba-Borda, and Schorfheide (2018), Inoue and Okimoto (2008), and Hirose (2020) argue that the ZLB/ELB gave rise to a structural break in Japan in the late 1990s.

We compute the posteriors of the parameters by combining the likelihood function of the model with the priors. Exploiting the linear-Gaussian nature of the model, the likelihood function is evaluated based on the Kalman filter. In the procedure, 10,000 draws from the posterior distribution are generated by the Metropolis-Hastings algorithm. The convergence of posterior distributions has been checked based on trace plots and Gelman et al. (2004) statistics.

4 Results

4.1 Parameter estimates

Table 2 summarizes the parameter estimation results. In the table, our main interest lies in the estimates of π^*_{high}, π^*_{mid}, and π^*_{low}. The posterior modes of high, mid, and low trend inflation states are 3.4, 1.9, and 0.7 percent, respectively. The upper 2.5 percentile of π^*_{mid} is 2.3 percent, which is lower than the mode of π^*_{high}. Likewise, the lower 2.5 percentile of π^*_{mid} is 1.1 percent which is higher than the mode of π^*_{low}. Therefore, three states of trend inflation are reasonably identified.

Other important Markov-switching parameters include the Taylor coefficients on inflation rate $\alpha_\pi(\mathcal{S}_t^{MP})$, the potential GDP growth rate $dy^*(\mathcal{S}_t^{dy^*})$ and the supply-shock volatility $\sigma^2_{AS}(\mathcal{S}_t^{AS})$. When monetary policy stance is hawkish, the mode of Taylor coefficient is estimated at 2.2, whereas it is at 0.6 under the dovish state. The higher potential GDP growth rate is 7.5 percent compared compared with 2.8 percent in the lower growth state. The supply-shock volatility in terms of standard deviation $\sigma_{AS}(\mathcal{S}_t^{AS})$ in the high volatility state is 1.0 while that in the low volatility state is 0.1, each of which is translated into annual 3.9 and 0.3 percent, respectively.

The remainders are non-switching parameters and their 95% credible intervals are also reasonably narrow. Although the outright comparison is not appropriate because of differing sample periods, most of the posterior modes and means of the parameters are broadly similar to those in early studies using Japanese data. Relatively, the Calvo parameter θ and the inflation indexation ρ are estimated at lower values compared with early studies. The lower θ implies more flexible price adjustments. This is consistent with the higher average and volatility of actual inflation in our pre-2000 sample periods as shown in Figure 1. While the inflation indexation has been a controversial "structural" parameter in the literature, AS (2014) claim that applying the GNK model can eliminate estimation bias and yield lower

Table 1: Calibration and Prior Distributions of Parameters

Parameter	Definition	Dist.	Mean	S.D.
ε	Elasticity of substitution among intermediate goods	G	8.00	1.50
θ	Calvo-Yun parameter	B	0.66	0.10
σ	Inverse intertemporal elasticity of substitution	G	3.00	1.00
α_y	MP reaction coefficient of output	G	0.50	0.30
ρ^{IS}	Persistence of IS shock	B	0.80	0.10
ρ^{AS}	Persistence of AS shock	B	0.80	0.10
ρ^{MP}	Interest rate smoothing	B	0.80	0.10
ρ	Inflation indexation	B	0.50	0.20
σ_{IS}	S.D. of IS shock (Q%)	IG	0.125	0.125
σ_{MP}	S.D. of MP shock (Q%)	IG	0.125	0.125
π^*_{high}	Trend inflation when $S_t^{\overline{\pi}}$ = High	G	5.00	1.00
π^*_{mid}	Trend inflation when $S_t^{\overline{\pi}}$ = Mid	G	3.00	1.00
π^*_{low}	Trend inflation when $S_t^{\overline{\pi}}$ = Low	N	0.50	1.00
$\alpha_\pi(S_t^{MP} = \text{Hawkish})$	Taylor coefficient when S_t^{MP} = Hawkish	G	1.50	0.50
$\alpha_\pi(S_t^{MP} = \text{Dovish})$	Taylor coefficient when S_t^{MP} = Dovish	G	1.00	0.30
$dy^*(S_t^{dy*} = \text{High growth})$	Potential GDP growth when S_t^{dy*} = High growth	G	8.00	2.00
$dy^*(S_t^{dy*} = \text{Low growth})$	Potential GDP growth when S_t^{dy*} = Low growth	G	2.00	1.00
$\sigma_{AS}(S_t^{oil} = \text{High volatility})$	S.D. of tech. shock when S_t^{oil} = High volatility	IG	1.00	1.00
$\sigma_{AS}(S_t^{oil} = \text{Low volatility})$	S.D. of tech. shock when S_t^{oil} = Low volatility	IG	0.25	0.25
p_{High,Low}	Transition prob. from High $\bar{\pi}$ to Low $\bar{\pi}$	Calib.	0	–
p_{High,Mid}	Transition prob. from High $\bar{\pi}$ to Mid $\bar{\pi}$	B	0.05	0.025
p_{Mid,High}	Transition prob. from Mid $\bar{\pi}$ to High $\bar{\pi}$	B	0.05	0.025
p_{Mid,Low}	Transition prob. from Mid $\bar{\pi}$ to Low $\bar{\pi}$	B	0.05	0.025
p_{Low,Mid}	Transition prob. from Low $\bar{\pi}$ to Mid $\bar{\pi}$	B	0.05	0.025
p_{Low,High}	Transition prob. from Low $\bar{\pi}$ to High $\bar{\pi}$	Calib.	0	–
p_{Dovish,Hawkish}	Transition prob. from Dovish to Hawkish	B	0.20	0.10
p_{Hawkish,Dovish}	Transition prob. from Hawkish to Dovish	B	0.20	0.10
p_{High growth,Low growth}	Transition prob. from High growth to Low growth	B	0.20	0.10
p_{Low growth,High growth}	Transition prob. from Low growth to High growth	B	0.20	0.10
p_{Low vol.,High vol.}	Transition prob. from Low vol. to High vol.	B	0.025	0.05
p_{High vol.,Low vol.}	Transition prob. from High vol. to Low vol.	B	0.20	0.10
$stderr\_\eta^y$	S.D. of measurement error for GDP growth	IG	4.00	2.00
$stderr\_\eta^\pi$	S.D. of measurement error for inflation	IG	0.50	0.25
β	Quarterly discount factor	Calib.	0.999	–
χ	Inverse labor supply elasticity	Calib.	2.00	–

NOTE: MP stands for monetary policy, S.D. stands for standard deviation, and Calib. stands for calibrated parameter. B, G, IG, and N stand for Beta, Gamma, inverse Gamma, and normal distribution, respectively. The unit of trend inflation and potential GDP growth rate is in annual percentage.

estimates for ρ.

4.2 Smoothed state probabilities

Another important output of our estimation is the smoothed state probabilities. The shaded areas in Figure 2 indicate recession periods identified by the Economic and Social Research Institute (ESRI) of the Japanese government. The first panel in the figure presents the probability for being in π^*_{mid} state. From 1958 to 1977, π^*_{high} state was likely to be dominant. In 1977, as indicated in the first panel, the probability for being in π^*_{high} sharply increased. Then, it continued declining gradually reaching zero in 1986. In 1986 onwards, π^*_{low} remains dominant until 1997, the end of sample period, except for a short bout in the late phase of the "bubble" period of 1989–1991. During the period, probability for π^*_{mid} rose, but the hike was short-lived as indicated by the second panel.

The fourth panel of Figure 2 presents the probability for being in the hawkish monetary policy state. The figure indicates that monetary policy was dovish until 1977. In 1978–80, slightly later than the start of the decline in the trend inflation, the once and for all change in BoJ's monetary policy stance from dovish to hawkish took place. In the following section, we will discuss (i) the start of the decline in trend inflation in 1977 and (ii) change in the policy stance in 1978–80, relying on some narratives. The sixth panel shows the probability for being in the high supply shock volatility state. In the figure, two spikes reaching 100 percent clearly identify the well known first and second global oil crisis episodes: the 1973–74 OAPEC embargo and the 1978 Iranian revolution. This result suggests that the two recessions in the Japanese 1970s were precipitated by the elevated global oil price volatility.

4.3 Point estimate of the trend inflation

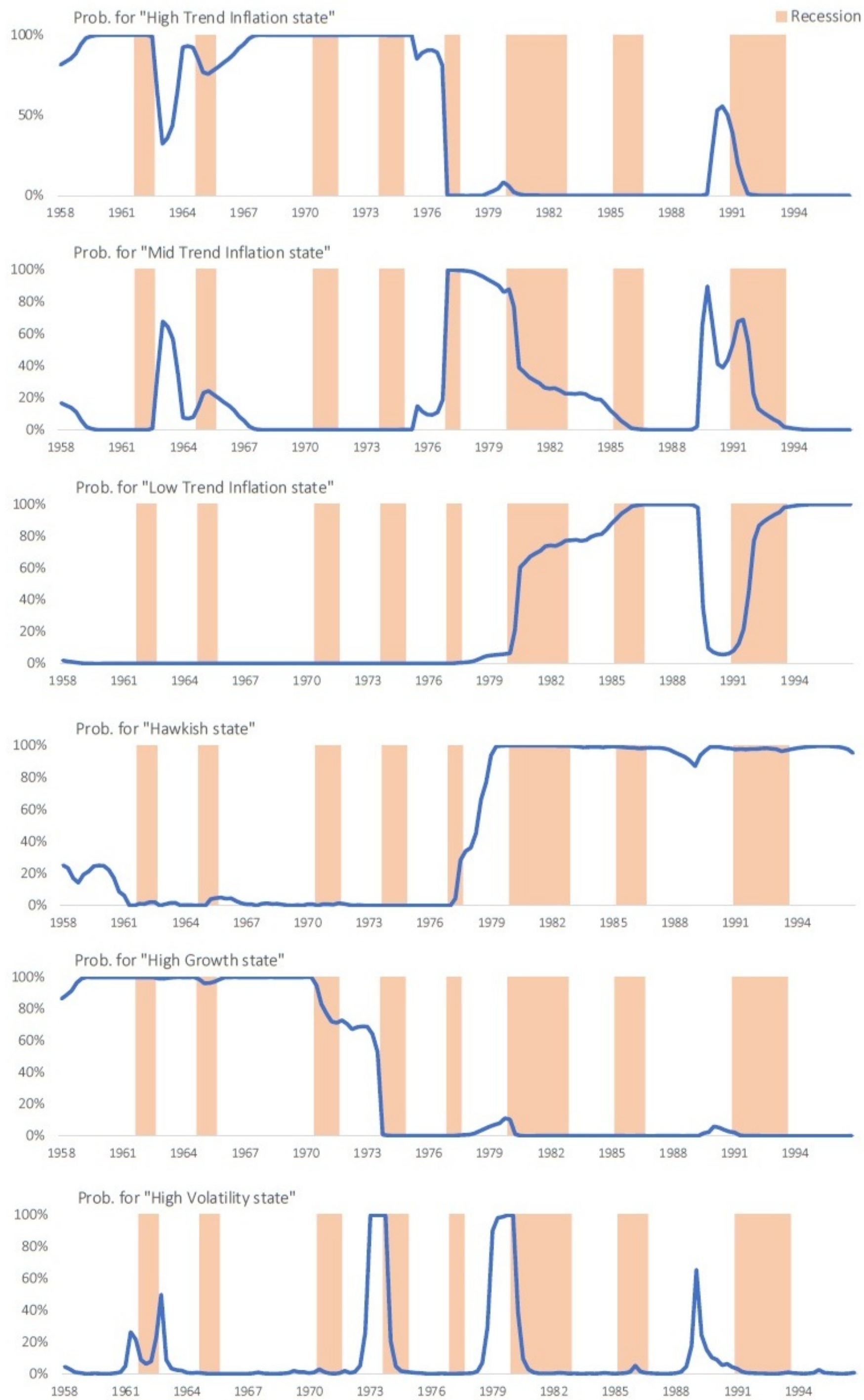
By combining the posterior modes for π^*_{high}, π^*_{mid}, and π^*_{low} and the smoothed probabilities for each state, the point estimate for the expected rate of trend inflation can be calculated over time. Figure 3 depicts the point estimate of the trend inflation. The figure reveals that the Japanese trend inflation was mostly dominated by π^*_{high}, staying at 3.0–3.5 percent from 1958 to 1977. Even well before the first oil crisis in 1973, trend inflation was already as high as 3.5 percent. As already mentioned, in 1977, trend inflation started to fall sharply followed by gradual declines until the mid-1980s. The downward trend continued through the second oil crisis when actual inflation rose to 10 percent. From 1986 onward, trend inflation hovered well above zero, mostly near 1 percent, until 1997. During the bubble

Table 2: Posterior Estimation Results

Parameter	Posteriors				
	Mode	Mean	Median	2.5%	97.5%
ε	8.048	7.014	6.767	5.052	9.996
θ	0.421	0.457	0.458	0.393	0.533
σ	5.346	4.485	4.522	3.227	5.828
α_y	0.671	0.579	0.526	0.207	1.266
ρ^{IS}	0.931	0.900	0.902	0.855	0.934
ρ^{AS}	0.916	0.907	0.907	0.884	0.933
ρ^{MP}	0.898	0.884	0.886	0.846	0.911
ρ	0.188	0.280	0.267	0.143	0.462
σ_{IS}	0.016	0.027	0.027	0.016	0.040
σ_{MP}	0.099	0.092	0.092	0.077	0.107
π^*_{high}	3.396	3.483	3.501	2.736	4.017
π^*_{mid}	1.923	1.778	1.825	1.090	2.261
π^*_{low}	0.708	0.710	0.764	-0.103	1.287
$\alpha_\pi(S_t^{MP} = \text{Hawkish})$	2.231	1.972	1.965	1.495	2.512
$\alpha_\pi(S_t^{MP} = \text{Dovish})$	0.574	0.447	0.450	0.242	0.638
$dy^*(S_t^{dy*} = \text{High growth})$	7.527	7.754	7.721	6.626	9.128
$dy^*(S_t^{dy*} = \text{Low growth})$	2.819	3.073	3.086	2.345	3.692
$\sigma_{AS}(S_t^{oil} = \text{High volatility})$	0.978	1.112	1.093	0.606	1.821
$\sigma_{AS}(S_t^{oil} = \text{Low volatility})$	0.072	0.120	0.114	0.058	0.218
$p\_\{\text{High,Mid}\}$	0.028	0.026	0.023	0.008	0.059
$p\_\{\text{Mid,High}\}$	0.050	0.039	0.035	0.012	0.083
$p\_\{\text{Mid,Low}\}$	0.035	0.038	0.036	0.011	0.085
$p\_\{\text{Low,Mid}\}$	0.028	0.028	0.026	0.009	0.055
$p\_\{\text{Dovish,Hawkish}\}$	0.069	0.053	0.048	0.016	0.111
$p\_\{\text{Hawkish,Dovish}\}$	0.049	0.052	0.047	0.015	0.103
$p\_\{\text{High growth,Low growth}\}$	0.040	0.042	0.038	0.009	0.090
$p\_\{\text{Low growth,High growth}\}$	0.016	0.024	0.023	0.007	0.050
$p\_\{\text{Low vol.,High vol.}\}$	0.006	0.025	0.020	0.003	0.077
$p\_\{\text{High vol.,Low vol.}\}$	0.296	0.198	0.190	0.071	0.371
$stderr\_\eta^y$	4.115	4.167	4.162	3.756	4.655
$stderr\_\eta^\pi$	0.315	0.541	0.490	0.220	1.060

NOTE: Posteriors are based on 10,000 draws of Metropolis-Hastings algorithm. Quantiles of posterior distributions are 2.5% and 97.5% for each parameter.

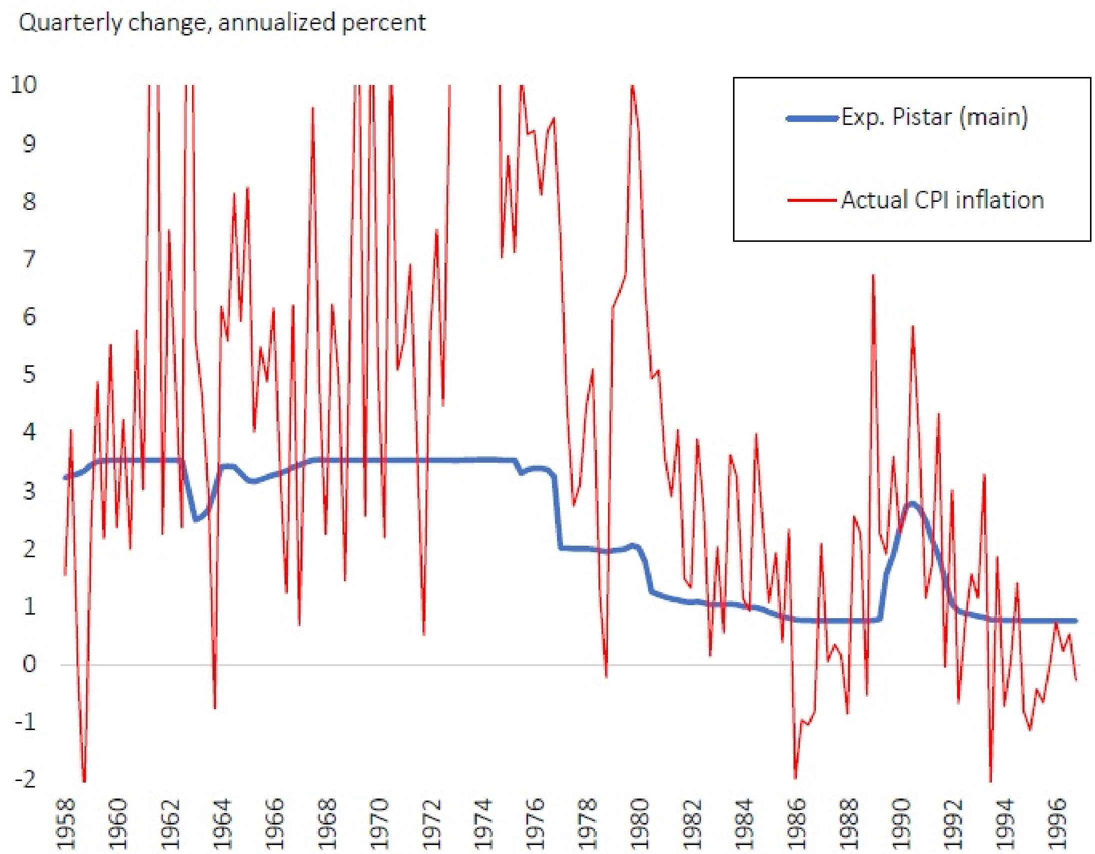
Figure 2: Smoothed State Probabilities



Note: The shaded areas correspond to the dated ESRI recession.

period, the point estimate rose higher than 2 percent, but the hike was short-lived. In our results, little evidence is found that trend inflation fell anywhere close to zero percent. As presented in Table 2, the probability is less than 2.5 percent for trend inflation being lower than -0.1 percent in the 1990s .

Figure 3: Trend Inflation



Note: The blue bold line is the point estimates of the expected trend inflation. The thin red line is the actual CPI inflation. Both are seasonally adjusted quarterly changes in price levels expressed as annualized percentage rates.

4.4 Role of time-varying volatility

In the model, $\mathcal{S}_t^{AS}$ is the only Markov chain process which affects the 'second moment' and hence, one may argue that the Markov chain is of the least importance compared to others. Here, we examine whether the estimation results considerably change if the aggregate supply shock is assumed to be homoskedastic. Figure 4 compares the point estimates of the trend inflation across the two specifications. The dotted line indicates the trend inflation in the case that σ^2_{AS} is non-switching – i.e., constant. Trend inflation moderately increases during the first and second global oil crises. This result can be interpreted as follows. Because σ^2_{AS} is assumed to be constant, the periods of oil crises are identified as the states where the levels of trend inflation are elevated, instead of high volatility period. This result under the constant volatility specification suffers from mis-identification of 'first moment' estimates, including changes in trend inflation, and shocks to the 'second moment,' – the volatility of the aggregate supply shock in our model. The nature of the two estimates is different and the two global oil crises need to be identified as a time-varying second moment.[9] Our overall assessment is that allowing σ^2_{AS} to follow a Markov chain process gives a better fit of the model to the data, which is similar to the U.S. context as demonstrated by BLM (2018).

5 Discussion

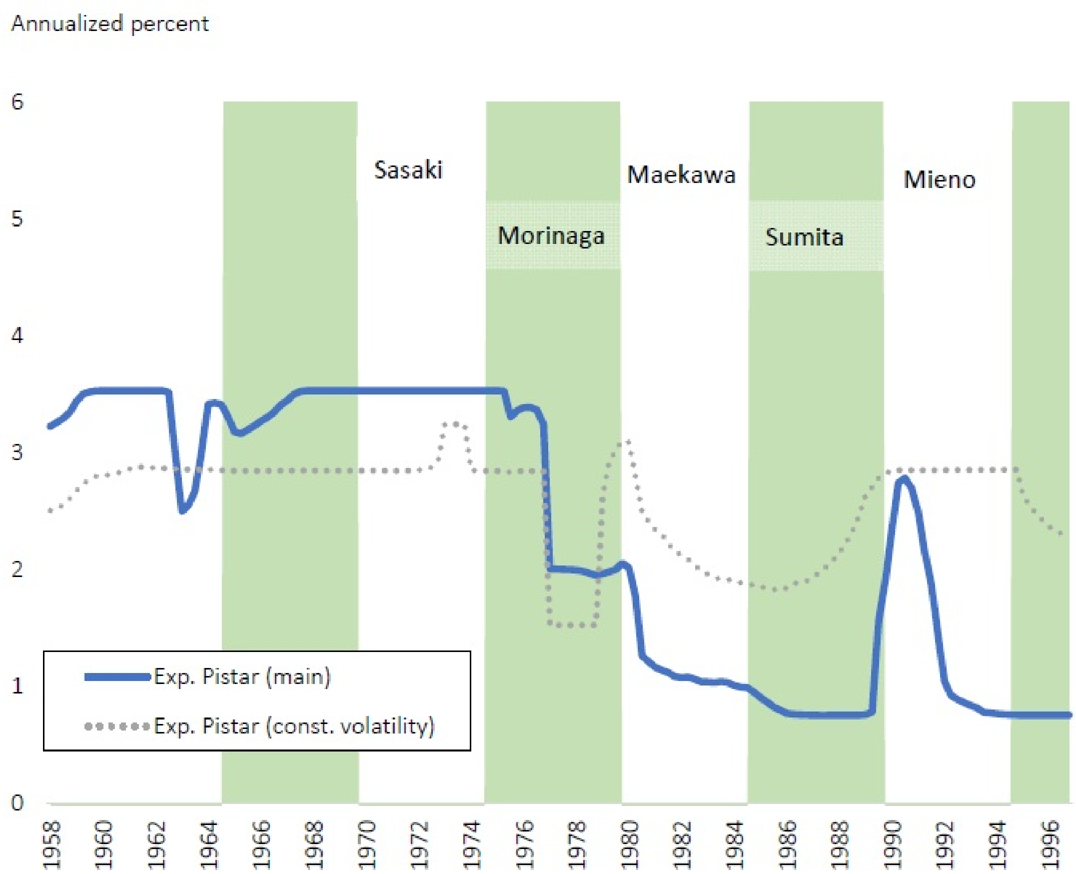
5.1 Two percent trend inflation as a vantage point

As the estimated π^*_{mid} is quite close to 2 percent, (i.e., 1.9 percent as the posterior mode in Table 2) the smoothed state probability for being in a π^*_{mid} state can serve as a framework for assessing the monetary policy stances. (Figure 5) If we assume that BoJ was pursuing 2 percent inflation target prior to the 2000s, which is neither a claim nor an estimate, then the smoothed state probability of π^*_{mid} close to 1 suggests that the monetary policy stance is 'neutral' or just about right. Conversely, if the probability of being at π^*_{mid} is close to zero, BoJ's stance is judged as either being 'too tight' or 'too accommodative' to be in line with the 2 percent trend inflation.

Our sample period, which is up to 1997, fully covers the five-year terms of five BoJ governors –Tadashi Sasaki (1969-1974), Tei-ichiro Morinaga (1974-1979), Haruo Maekawa (1979-1984), Satoshi Sumita (1984-1989), and Yasushi Mieno (1989-1994), –each of whom played different roles in terms of Japan's disinflation transition from 1977 to 1997. We will discuss the assessment of the BoJ's policy rate cuts and hikes led by each governor from our vantage point of the 2 percent trend inflation as depicted in Figure 5.

[9] It could however be noted that, in the case of constant σ^2_{AS}, π^*_{low} is estimated even higher than one percent.

Figure 4: Trend Inflation under an Alternative Specification



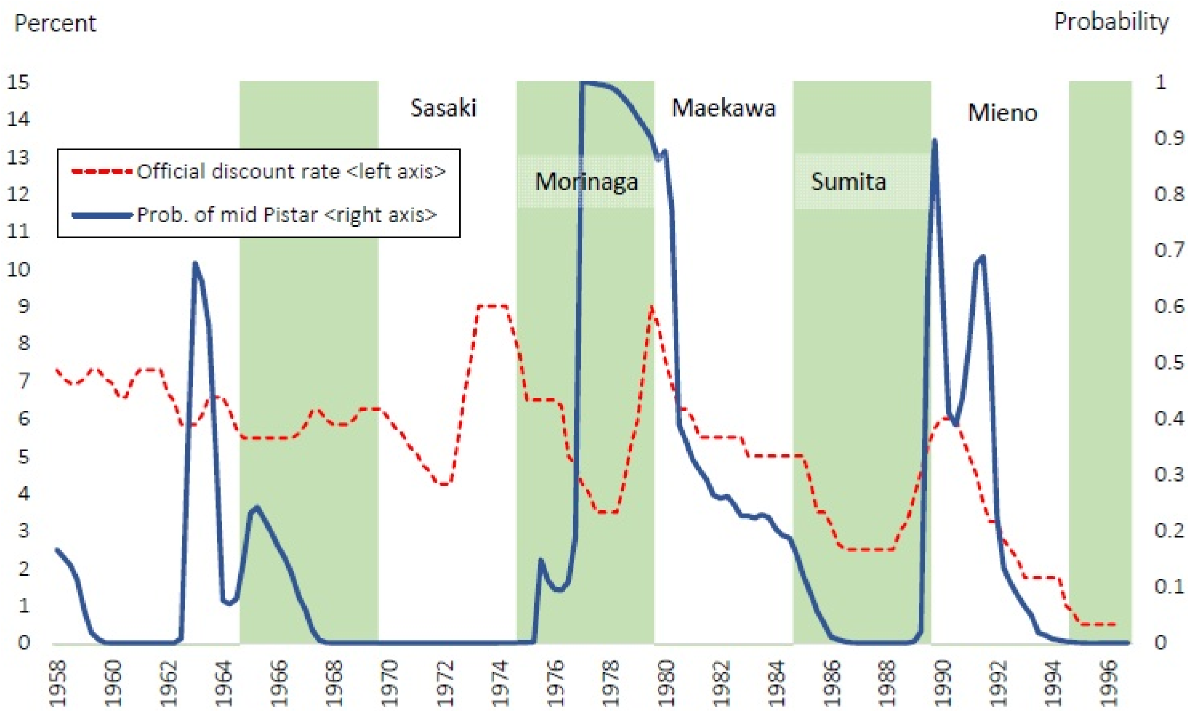
Note: The blue bold line is the point estimates of the expected trend inflation in the main model. The dotted gray line is that in the constant volatility model.

5.2 Disinflation transition and the five Governors of BoJ

The estimated π^*_{high} at 3.5 percent, which prevailed until 1977, may have reflected the power balance between the less independent BoJ led by Sasaki and the pro-fiscal expansionary administration under Prime Minister Kakuei Tanaka. In the run-up to the first Global Oil Crisis, i.e., the embargo on oil exports by OAPEC in October 1973, BoJ allowed money supply to grow annually by 25 percent.[10] (Figure 6) The BoJ's highly accommodative stance may have been the result of the influence of Tanaka's administration, which was then promoting nation-wide infrastructure projects and, thus, seeking monetary supports from the

[10]For instance, Sasaki testified that 20 percent cash growth rate is 'never too high' when he was summoned to the Diet in 1974. See the official minutes of Finance Committee, House of Counsilors (1974).

Figure 5: ‘Scorecard’ and Policy Rate



Note: The red dotted line indicates the official discount rate in annualized percentage rates (left axis). The blue line indicates the smoothed probability of trend inflation being at mid-state (right axis).

central bank.

After Sasaki stepped down in 1974, Morinaga took office as BoJ governor under the Miki administration, which was relatively up for fiscal consolidation. Both of the major regime switches, – (1) the sharp decline in trend inflation, followed by a disinflation transition phase until the mid-1980s, and (2) the rise in the response parameter in the Taylor rule on inflation, –took place almost simultaneously during Morinaga’s term (1974-1979).

Maekawa, the deputy of Morinaga, was appointed as the successor to Morinaga in 1979. When he took office, he inherited Morinaga’s hawkish stance, continuing the rate hikes in response to the second Global Oil Crisis. Figure 5 confirms that Maekawa’s initial rate hikes are mostly consistent with 2 percent trend inflation while the score, i.e., the probability of the 2 percent trend inflation, quickly dropped and kept decreasing toward the end of his term. Although the policy rate was mostly being cut throughout his term, our assessment reveals that Maekawa’s policy stance would have been too tight if the BoJ had been pursuing the 2 percent trend inflation.

In September 1985, or ten months after Sumita took office, together with then Finance Minister Noboru Takeshita, they participated in the meeting at the Plaza Hotel in New York City. At the meeting, the finance ministers and central bank governors of the U.S., France, Germany, Japan, and the U.K. reached an agreement that the US dollar was overvalued and that they were ready for coordinated intervention in foreign exchange markets to stimulate depreciation of the US dollar against major currencies. This is the historic Plaza Accord after which the yen significantly appreciated against the US dollar. Subsequently, to revive the economy from the "strong-yen recession," Sumita cut the policy rate to 2.5 percent, –a historically lowest level at that time, –and kept the rate there for more than two years until May 1989. Existing studies have argued that during Sumita's 'accommodative' policy periods, expectations for a low interest rate continued for an extended period, thereby fostering and fueling the asset price bubble in later years.[11] Others claim that the intention of the BoJ or the Ministry of Finance to keep the rate at a low level was to prevent the yen from appreciating and to be consistent with the international coordination following the Plaza Accord.[12] While we do not further discuss Sumita's monetary easing in relation to the bubble or foreign exchange rate, Figure 5 indicates that Sumita's policy stance was mostly too tight, rather than being accommodative with the "ultra-low" rate level, in light of the 2 percent trend inflation.

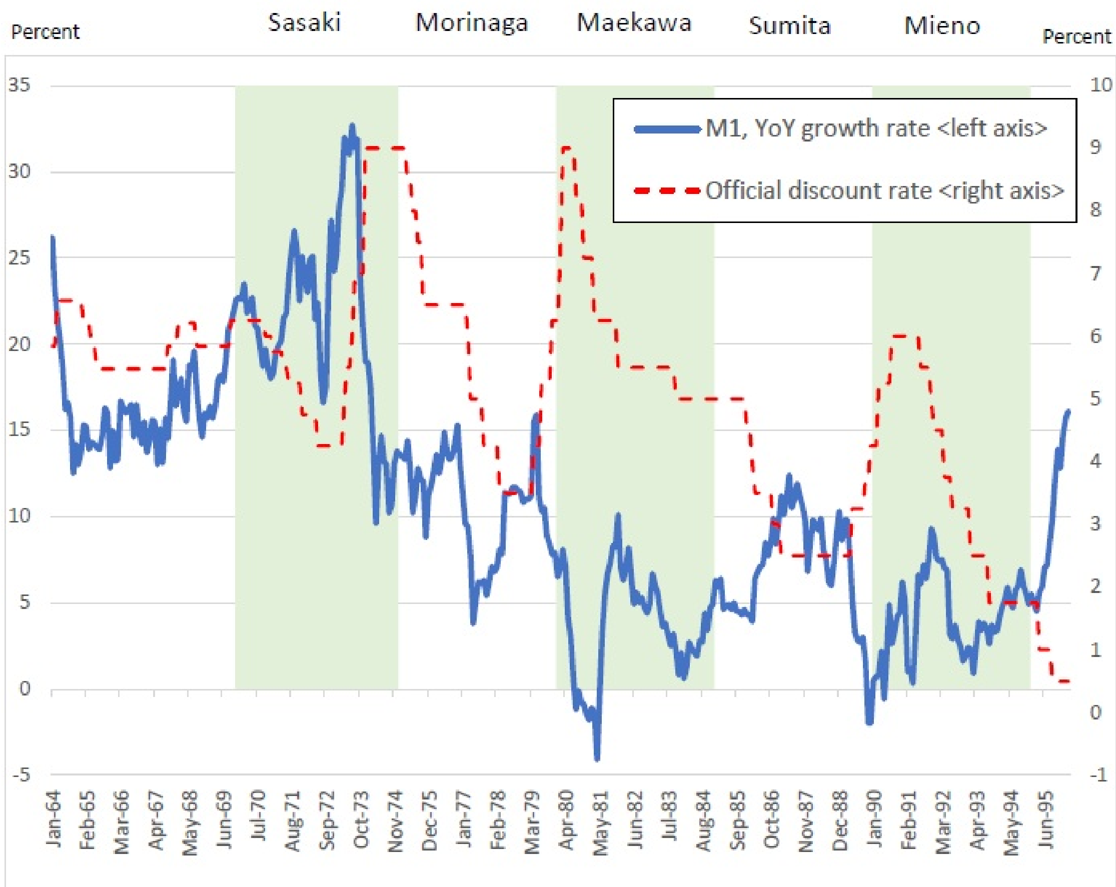
Mieno, the former deputy of Sumita, took office in 1989 and raised the policy rate reaching as high as 6 percent allegedly to combat the incipient asset price bubble. In 1991, he turned to a rate-cut cycle to mitigate the adverse shocks arising from the burst of the asset price bubble. The policy rate declined markedly to 2 percent toward the end of his term. Figure 3 depicts that trend inflation jumped beyond 2 percent during the early days of Mieno's term and then quickly started declining and kept declining over the rest of his term, reaching a π^*_{low} state in 1992. Since 1992, the low trend inflation state was dominant until the end of the sample period. From our vantage point of 2 percent trend inflation, Mieno's policy stance was too tight, and his policy easing was consistent with the 1 percent trend inflation rather than the 2 percent trend inflation.[13]

[11] See Okina, Shirakawa and Shiratsuka (2000) for example.

[12] See Arioka, Ito, and Kosai (2000) and Ito (2015).

[13] Similar assessment can be found in Ito and Mishkin (2004) and Kato and Nishiyama (2005). Ito and Mishkin (2004) note that Mieno's monetary easing responses as being 'too little, too late.'

Figure 6: Money Supply Growth



Note: Blue line indicates the monthly year-over-year growth rate of M1 (left axis). Red line indicates the official disount rate in annualized percentage (right axis).

6 Concluding remarks

Our estimation results broadly indicate that, in Japan, trend inflation was unlikely to decline to near- or even below-zero level prior to 1996 when the ZLB was hit. Japan's deflation began around 2000 and continued for protracted periods. Hence, the results imply that deflation did not trigger the ZLB as a natural outcome. In the meantime, we do not overstate the implications of our results. As noted, our estimates indicate that ZLB precedes deflation, while it remains yet to be explored whether ZLB caused the deflation or the other way around. In the run-up to the ZLB arrival, it can be a case that the low trend inflation rate, albeit well above zero, may have constrained the Bank of Japan's capacity to rate-cut in response to adverse IS shocks. Further studies, possibly relying on some counterfactual simulations, are awaited to inspect the chain of causality.

Further cautions and limitations need to be borne in mind in interpreting our results. First, the estimation is not based on real-time data but on historical data. As of 1997, that is, at the end of our sample period, policymakers were observing real-time data which were later revised. Estimation using real-time data may elicit more nuanced implications. Second, our estimation is based on a small-scale model. Medium-scale models that include more variables, particularly financial sector and asset market variables, will provide richer information regarding trend inflation and possible state transitions. These are also remaining issues to be explored in future studies.

References

[1] Abe, Nobuhiro, Takuji Fueki and Sohei Kaihatsu, 2019, "Estimating a Markov Switching DSGE Model with Macroeconomic Policy Interaction," Bank of Japan Working Paper Series 19-E-3, Bank of Japan.

[2] Aruoba, S. Borağan, Pablo Cuba-Borda, and Frank Schorfheide, 2018, "Macroeconomic Dynamics Near the ZLB: A Tale of Two Countries," *Review of Economic Studies*, 85(1), pages 87–118.

[3] Ascari, Guido, and Argia M. Sbordone, 2014, "The Macroeconomics of Trend Inflation," *Journal of Economic Literature*, 52(3), pages 679–739.

[4] Bjørnland, Hilde C., Vegard H. Larsen, and Junior Maih, 2018, "Oil and Macroeconomic (In)Stability," *American Economic Journal*, 10(4), pages 128–151.

[5] Chang, Yoonsoon, Junior Maih, and Fei Tan, 2021, "Origins of Monetary Policy Shifts: A New Approach to Regime Switching in DSGE Models," *Journal of Economic Dynamics and Control*, 133, issue C, number S0165188921001706.

[6] Cogley, Timothy and Argia M. Sbordone, 2008, "Trend Inflation, Indexation, and Inflation Persistence in the New Keynesian Phillips Curve," *American Economic Review*, 98(5), pages 2101–2126.

[7] Clarida, Richard, Jordi Galí, and Mark Gertler, 1999, "The Science of Monetary Policy: A New Keynesian Perspective," *Journal of Economic Literature*, 37(4), pages 1661–1707.

[8] Collins, Susan M. and Barry Bosworth, 1996, "Economic Growth in East Asia: Accumulation versus Assimilation," *Brookings Papers on Economic Activity*, 27(2), pages 135–204.

[9] Farmer, Roger, E., Daniel F. Waggoner, and Tao Zha, 2011, "Minimal State Variable Solutions to Markov-Switching Rational Expectations Models," *Journal of Economic Dynamics and Control*, 35(12), pages 2150–2166.

[10] Fueki, Takuji, Ichiro Fukunaga, Hibiki Ichiue, and Toyoichiro Shirota, 2016, "Measuring Potential Growth with an Estimated DSGE Model of Japan's Economy," *International Journal of Central Banking*, 12(1), pages 1–32.

[11] Gelman A., J. B. Carlin, H. S. Stern, and D. B. Rubin, 2004, *Bayesian Data Analysis*, London, Chapman & Hall.

[12] Hayashi, Fumio and Edward C. Prescott, 2002, "The 1990s in Japan: A Lost Decade," *Review of Economic Dynamics*, 5(1), pages 206–235.

[13] Hirose, Yasuo, 2020, "An Estimated DSGE Model with A Deflation Steady State," *Macroeconomic Dynamics*, 24(5), pages 1151–1185.

[14] House of Councillors, The National Diet of Japan, 1974, Minutes of the 9th Finance Committee. https://kokkai.ndl.go.jp/txt/107214629X00919740307/48

[15] Ichiue, Hibiki, Takushi Kurozumi, and Takeki Sunakawa, 2013, "Inflation Dynamics and Labor Market Specifications: A Bayesian Dynamic Stochastic General Equilibrium Approach for Japan's Economy," *Economic Inquiry*, 51(1), pages 273-287.

[16] Iiboshi, Hirokuni, Mototsugu Shintani, and Kozo Ueda, 2022, "Estimating a Nonlinear New Keynesian Model with the Zero Lower Bound for Japan," *Journal of Money, Credit and Banking*, 54(6), pages 1637-1671.

[17] Inoue, Tomoo and Tatsuyoshi Okimoto, 2008, "Were There Structural Breaks in the Effects of Japanese Monetary Policy? Re-Evaluating Policy Effects of the Lost Decade," *Journal of the Japanese and International Economies*, 22(3) pages 320–342.

[18] Ito, Takatoshi and Frederic S. Mishkin, 2004, "Two Decades of Japanese Monetary Policy and the Deflation Problem," NBER Working Papers 10878, National Bureau of Economic Research, Inc.

[19] Ito, Takatoshi, 2015, "The Plaza Agreement and Japan: Reflection on the 30th year Anniversary," Rice University's Baker Institute for Public Policy Working Paper.

[20] Kaihatsu, Sohei and Takushi Kurozumi, 2014, "What Caused Japan's Great Stagnation in the 1990s? Evidence from an Estimated DSGE Model," *Journal of the Japanese and International Economies*, 34(C), pages 217–235.

[21] Kato, Ryo and Shin-Ichi Nishiyama, 2005, "Optimal Monetary Policy When Interest Rates Are Bounded at Zero," *Journal of Economic Dynamics and Control*, 29(1-2), pages 97–133.

[22] Kosai, Yutaka, Osamu Ito, and Ritsuko Arioka, 2000, "Bubble-ki no Kin'yu Seisaku to Sono Hansei" (Re-examination of Monetary Policy in Bubble Period), IMES Discussion Paper, No. 2000-J-27, Institute for Monetary and Economic Studies, Bank of Japan, 2000 (in Japanese).

[23] Maih, Junior, 2014, "Efficient Perturbation Methods for Solving Regime-Switching DSGE Models," Working Papers No 10/2014, Centre for Applied Macro- and Petroleum economics (CAMP), BI Norwegian Business School.

[24] Nishizaki, Kenji, Toshitaka Sekine, and Yoichi Ueno, 2014, "Chronic Deflation in Japan," *Asian Economic Policy Review*, 9(1), pages 20–39.

[25] Okina, Kunio, Masaaki Shirakawa, and Shigenori Shiratsuka, 2000, "Asset Price Bubble and Monetary Policy: Japan's Experience in the Late 1980s and the Lessons," IMES Discussion Paper, No. 2000-E-12, Institute for Monetary and Economic Studies, Bank of Japan.

[26] Patrick, Hugh T. and Henry Rosovsky, 1976, *Asia's New Giant: How the Japanese Economy Works*, Washington DC, Brookings Institution.

[27] Sugo, Tomohiro and Kozo Ueda, 2008, "Estimating a Dynamic Stochastic General Equilibrium Model for Japan," *Journal of the Japanese and International Economies*, 22(4), pages 476–502.

[28] Woodford, Michael, 1999, *Interest and Prices: Foundations of a Theory of Monetary Policy*, Princeton NJ, Princeton University Press.

A Appendix: Set-up of the model with non-zero trend inflation

Our model is a version of the generalized New Keynesian DSGE model developed by Ascari and Sbordone (2014). This appendix provides the setup of our model.

A.1 Households

A representative household exhibits a utility function that is separable in consumption C_t and labor supply N_t:

$$U(C_t, N_t) = \frac{C_t^{1-\sigma}}{1-\sigma} - v\frac{N_t^{1+\chi}}{1+\chi},$$

where v is a constant parameter, with the period-by-period budget constraint as follows:

$$P_t C_t + R_t^{-1} B_t = W_t N_t + D_t + B_{t-1}, \tag{13}$$

where B_t, R_t, W_t and D_t denote one-period bond holdings and its (gross) interest rate, nominal wage and distributed dividend, respectively. The utility maximization yields the following first-order conditions,

$$\beta E_t R_t \left(\frac{P_t}{P_{t+1}}\right)\left(\frac{C_{t+1}}{C_t}\right)^{-\sigma} = 1, \tag{14}$$

$$v N_t^{\chi} C_t^{\sigma} = \frac{W_t}{P_t}. \tag{15}$$

The consumption-output Euler equation corresponds to (14), whereas (15) eliminates real wages in firms' first-order conditions.

A.2 Production

An intermediate goods producer i, has a linear production function in which labor is the only input:

$$Y_{i,t} = A_t N_{i,t},$$

where A_t denotes productivity that follows a stationary stochastic process. Then, the aggregate labor demand is as follows:

$$N_t = \int_0^1 N_{i,t} di = \underbrace{\int_0^1 \left(\frac{P_{i,t}}{P_t}\right)^{-\varepsilon} di}_{\equiv s_t} \frac{Y_t}{A_t} = \frac{s_t Y_t}{A_t}, \tag{16}$$

where s_t denotes the price dispersion arising from the Calvo pricing.

In the economy, the final good producer aggregates intermediate goods $Y_{i,t}$ according to the followings:

$$Y_t = \left[\int_0^1 Y_{i,t}^{\frac{\varepsilon-1}{\varepsilon}} di\right]^{\frac{\varepsilon}{\varepsilon-1}},$$

where ε represents the elasticity of substitution among intermediate goods.

A.3 Firms' pricing

In each period, a fraction $1-\theta$ of firms re-optimize their prices denoted as $P^*_{i,t}$. The rest of the firms index their prices to the previous period's inflation rate such that $P_{i,t} = \pi_t^\rho P_{i,t-1}$ where $\pi_t = P_t/P_{t-1}$ and ρ represents the degree of indexation. The profit maximization problem for the firms is given as follows:

$$\max_{P^*_{i,t}} : E_t \sum_{j=0}^{\infty} \mathfrak{D}_{t,t+j} \theta^j \left(\frac{P^*_{i,t} \Pi^\rho_{t-1,t+j-1}}{P_{t+j}} Y_{i,t} - \frac{W_{t+j}}{P_{t+j}} \frac{Y_{i,t+j}}{A_{t+j}}\right),$$

subject to the demand constraint,

$$Y_{i,t+j} = \left(\frac{P_{i,t} \Pi^\rho_{t-1,t+j-1}}{P_{t+j}}\right)^{-\varepsilon} Y_{t+j},$$

where $\mathfrak{D}_{t,t+j}$ is a stochastic discount factor and $\Pi_{t,t+j}$ indicates cumulative inflation from period t to $t+j$ such that,

$$\Pi_{t,t+j} = \frac{P_{t+1}}{P_t} \frac{P_{t+2}}{P_{t+1}} \times \cdots \times \frac{P_{t+j}}{P_{t+j-1}},$$

for $t \geqq 1$.

Let $p^*_{i,t} = P^*_{i,t}/P_t$ and $w_t = W_t/P_t$. Then, the first-order condition for the firms' price-setting can be written as follows:

$$p_{i,t}^{*} = \frac{\varepsilon}{\varepsilon - 1}\frac{\psi_t}{\phi_t}, \tag{17}$$

where

$$\begin{aligned}
\psi_t &\equiv E_t \sum_{j=0}^{\infty} (\theta\beta)^j \, \frac{Y_{t+j}^{1-\sigma} w_t}{A_t} \left(\frac{\Pi_{t-1,t+j-1}^{\rho}}{\Pi_{t+j}} \right)^{-\varepsilon}, \\
\phi_t &\equiv E_t \sum_{j=0}^{\infty} (\theta\beta)^j \, Y_{t+j}^{1-\sigma} \left(\frac{\Pi_{t-1,t+j-1}^{\rho}}{\Pi_{t+j}} \right)^{1-\varepsilon}.
\end{aligned}$$

Combining (17) with (15) and s_t defined in (16) results in (1), (2), and (3).

A.4 Monetary policy

Recall that the (gross) official discount rate denoted by R_t in (12) is the data counterpart of the one-period risk-free interest rate in (13). Let $4\tilde{\imath}_t = 100\log(R_t/\bar{R}_t)$. Then, in the model, monetary policy follows a standard Taylor rule given as follows:

$$\frac{R_t}{\bar{R}} = \left(\frac{R_{t-1}}{\bar{R}}\right)^{\rho^{MP}} \left[\left(\frac{\pi_t}{\overline{\pi}_t}\right)^{\alpha_{\pi,t}} \left(\frac{Y_t}{Y_t^{*}}\right)^{\alpha_y} \right]^{1-\rho^{MP}} exp(e_t^{MP}),$$

which corresponds to (6).