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Inflation expectations in Japan: Forecast revision and forecast trend

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Abstract

Inflation expectation is one of the most essential components of monetary policy decisions. Upon examining Japanese inflation expectations between 1991:Q4 and 2025:Q1, we propose a modified empirical model that includes a forecast trend term in addition to the forecast revision term. We found that the forecast trend term affects the forecast errors. The full sample results indicate that people in Japan form non-rational expectations with information rigidity. However, this holds only in the recent episode of inflation following the post-COVID period. During the zero-inflation periods, people formed full-information rational expectations. In addition, we find evidence that consumption tax hikes affect the forecast errors, partly due to the uncertainty about future implementation. In both periods, the possibility of deviating from rational expectations cannot be ruled out.

Keywords: Forecast Trend, Inflation Expectations, Information Rigidity, Japanese Inflation.
JEL Classification codes: C53, D83, D84, E31, E37

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1 Introduction

Japan had experienced low or zero inflation for more than two decades despite the unprecedented monetary policy easing (see Figure 1). Only after COVID-19 did we observe inflation in Japan rise gradually. The BOJ shifted from a zero-interest-rate policy in March 2024, raising the policy rate to 0.5 percent in January 2025 and to 0.75 percent in December 2025. Once the BOJ confirms a higher, well-anchored expected inflation rate, it is likely to raise the policy rate further. Understanding how inflation expectations evolve is paramount in Japan.

In this study, we investigate how expected inflation is formed in Japan: Is inflation expectation rational, and is information fully updated at all times? To answer the question, we apply the approach of Coibion and Gorodnichenko (2015) to the Japanese inflation expectation dataset recently constructed by Osada and Nakazawa (2024). We estimate the forecast-error regression over the period from 1991:Q4 to 2025:Q1.

This study is not the first to apply Coibion and Gorodnichenko (2015) to Japanese data. The main objective of Inatsugu et al. (2019) was to apply the approach of Coibion and Gorodnichenko (2015) to the Japanese dataset; however, they used an entirely different econometric approach to the BOJ's Tankan data. We apply the forecast-error model in Coibion and Gorodnichenko (2015) to the Japanese inflation expectation dataset. In doing so, a crucial difference in the data framework exists between the US Survey of Professional Forecasters (SPF) dataset and the Japanese dataset. To address the dataset problem, we develop an alternative framework to test the rational expectation under information rigidity.

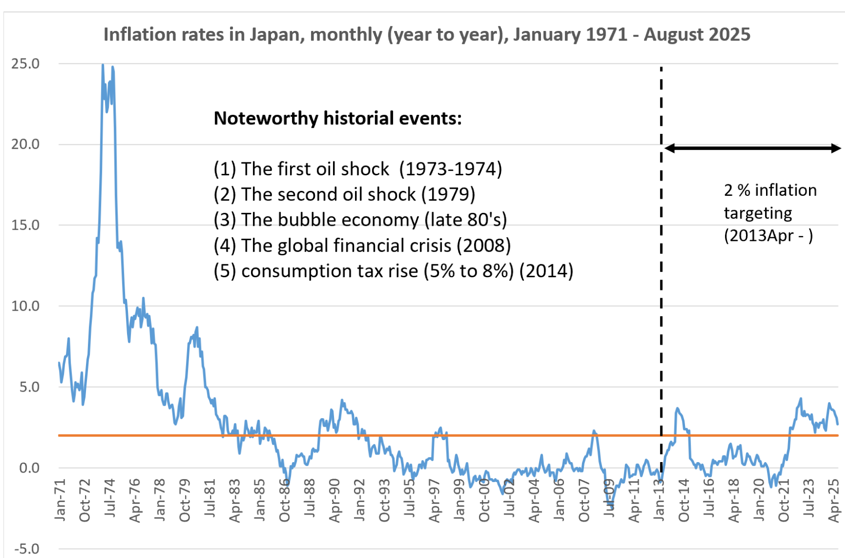

Figure 1: Japan's inflation from January 1971 to August 2025, year on year

The data source: the Consumer Price Index, monthly inflation rates year on year, the Statistics Bureau of Japan.

In particular, we propose an alternative forecast-revision term in the forecast-error regression model of Coibion and Gorodnichenko (2015). The alternative forecast revision differs from the original forecast revision in the deviation between the target forecast periods. The modified forecast revision allows the forecasted periods to differ between the past forecast and the current forecast. This seemingly contradictory modified forecast revision, as shown by decomposition, equals the sum of the original forecast revision and the forecast trend.

We find that the coefficients on the forecast revision, whether in the original or modified form, are positive and statistically significant regardless of the specifications. Based on this evidence, we conclude that people in Japan form rational expectations about inflation. Due to the information rigidity, the forecasts are under-reactive to new information. Moreover, we have evidence that this underreaction is observed only in the recent episode of inflationary economics following the post-COVID period. During the zero-inflation period, there was no evidence of information rigidity. We estimated the subsample only up to 2019:Q4 and found that forecast revisions are not statistically significant. Moreover, the time-varying parameter estimates provided by Müller and Petalas (2010) and Inoue et al. (2025) indicate that the forecast revision coefficients become positive and statistically significant only after 2022.

These results are robust to alternative specifications and econometric approaches. (i) The expected value of forecast errors is zero. Therefore, we re-estimated the forecast errors on forecast revisions by imposing a zero constant. (ii) The forecast revision and the error term might be correlated. We re-estimated the model by instrumental variable estimation. (iii) To validate the applicability of the proposed model, we apply it to the US inflation survey dataset and obtain similar results. (iv) In addition, we present a theoretical model and the necessary assumptions that yield the modified model.

In addition, we introduced three consumption tax hike dummies in the empirical model to control for special events in April 1997, April 2013, and October 2019. The null hypothesis of the rational expectation model is that consumption tax hike dummies do not affect forecast errors. Backed by the media coverage of future tax hikes a year earlier, there was uncertainty about the actual implementation of the tax increase. In the forecast-error regression model, we find the tax-hike dummies statistically significant.

Our findings have important implications for Japan's monetary policy. As the subsample analysis and time-varying parameter models suggest, we have evidence that inflation expectation formation in Japan may have shifted. Judging the expected inflation based on experience during the low- and zero-inflation period before 2020 can be misleading. Given elevated uncertainty due to geopolitical risk in the recent period, inflation expectations in Japan are formed in a noisy information environment.

The rest of the paper is structured as follows. The next section reviews the inflation-expectation model under information rigidity. Section 3 describes the dataset on Japanese expectations and discusses the crucial differences across forecast horizons. Section 4 discusses the modified version of forecast revision, partly necessitated by the lack of shorter forecast horizons in the Japanese data. We introduce a forecast-trend term in an empirical model, in addition to the standard forecast-revision term. Section 5 discusses the empirical results. Section 6 investigates the possible shift between the unconventional monetary policy regime and the recent inflation period. Section 7 discusses the empirical results regarding the forecast trend, and the last section concludes.

2 Literature review: Inflation expectations with information rigidity

We first review the two information rigidity models of Coibion and Gorodnichenko (2015). The first model is a sticky-information model, in which only a portion of economic agents can update their information. Agents update their information sets each period with probability $(1-\lambda)$, and have rational expectations. Averaging across agents' inflation forecasts yields the following relationship between forecast errors and forecast revisions.

$$\pi_{t+h} - \bar{E}_t[\pi_{t+h}] = \frac{\lambda}{1-\lambda}(\bar{E}_t[\pi_{t+h}] - \bar{E}_{t-1}[\pi_{t+h}]) \tag{1}$$

The second model is a noisy-information model, in which agents receive noisy information and update their forecasts via the Kalman filter, which balances current information and past forecasts, with the weight G placed on new information. Averaging across agents yields the following relationship between forecast errors and forecast revisions.

$$\pi_{t+h} - \bar{E}_t[\pi_{t+h}] = \frac{1-G}{G}(\bar{E}_t[\pi_{t+h}] - \bar{E}_{t-1}[\pi_{t+h}]) + v_{t+h,t} \tag{2}$$

From both models, as equation (11) in Coibion and Gorodnichenko (2015), they propose the following simple empirical model, see the discussion in Angeletos et al. (2020).

$$\begin{aligned} Forecast\ error_t &= \alpha + K_{CG} \cdot Forecast\ revision_t + u_t \\ \pi_{t+3} - \bar{E}_t[\pi_{t+3}] &= \alpha + K_{CG}(\bar{E}_t[\pi_{t+3}] - \bar{E}_{t-1}[\pi_{t+3}]) + u_t \end{aligned} \tag{3}$$

The key parameter in this model is K_{CG}. If agents form rational expectations without information frictions and without noisy information, K_{CG} must be equal to zero. With information rigidity, if agents are forming rational expectations, K_{CG} should be greater than zero. In this case, agents under-react to new information, i.e., make a too-low forecast when updating with an upward revision. In addition, the model imposes that no other control variables have explanatory power, as it does not include any other variables. If some variables are statistically significant, it implies that agents are not following rational expectations.[1]

[1] To be precise, it only implies that either expectations are formed non-rationally or the underlying assumptions and models are not correct or both. A modified theoretical model with rational expectation assumption may allow these variables to have explanatory power in the forecast-error regression.

The estimated coefficients of ‘forecast revision’ are always positive and statistically significant in Coibion and Gorodnichenko (2015), indicating the result is consistent with the information rigidity model. With control variables of inflation, T-bill rates, and oil price changes, additional control variables were not statistically significant, and they imply that full-information rational expectations (FIRE) were not violated. “In the case of unemployment, however, there is additional predictive power even after controlling for forecast revisions, although the coefficient on the unemployment rate is cut by approximately 40 percent. This finding suggests that deviations from FIRE may exist above and beyond those captured by simple models of information rigidities, and further exploration of these deviations is a fruitful avenue for future research.” Coibion and Gorodnichenko (2015)(p.2655).

2.1 Underreaction and overreaction

Coibion and Gorodnichenko (2015) assumed that each of the heterogeneous agents acts rationally with information rigidity. At the averaging level, or at the consensus level, the empirical results on inflation forecasts showed that people respond less to information, i.e., the realized inflation is higher than the expectation when the expectation is revised. Following the seminal work of Coibion and Gorodnichenko (2015), many studies report that at the individual level, households, firm managers, or professional forecasters overreact, see Angeletos et al. (2020). Following the diagnostic expectations in Bordalo et al. (2018), Bordalo et al. (2020) reconciled with their diagnostic expectation model the evidence that individual forecasters overreact to news, while consensus forecasts underreact. Proposition 2 in Bordalo et al. (2020) proves this.

3 Inflation expectation data in Japan

Osada and Nakazawa (2024) constructed the CIE, using the six measures of inflation expectations: two measures for households, one measure for firms, and three measures for experts. In particular, for the two household measures, they use ”The Opinion Survey on the General Public’s Views and Behavior” by the BOJ, which comprises two separate responses: one qualitative and the other quantitative. The firm measure is Tankan by the BOJ. Three measures for experts are ’Consensus Forecasts’, ’QUICK’, and inflation swap rates.[2]

The CIE index is based on the principal component and forecasting power, constructed by Osada and Nakazawa (2024). They introduce a time-series model and use estimation to interpolate values for forecast horizons that were not observed. The forecast horizons of the CIE comprise annual incremental forecasts extending from one to ten years ahead.

[2]As survey-based inflation forecasts, we have two sources: the Consumer Confidence Survey by the CAO and the Opinion Survey by the BOJ. The CAO survey data is conducted monthly, and the forecast horizon is one year. The BOJ survey is conducted quarterly, with forecast horizons of one and five years.

The CIE forecasts are updated every quarter. We use the CIE forecast data for the period from 1991:Q4 to 2025:Q1.

3.1 The crucial difference between the SPF and the CIE index

The forecast frequencies for the SPF and the CIE are both quarterly; however, the forecast horizons differ critically between the two datasets. The CIE has no short-horizon forecasts, whereas the SPF has forecasts for one, two, three, and four quarters ahead. The forecast horizons of the CIE range from one to ten years ahead.

This crucial difference in forecast horizons makes it difficult to closely follow the empirical model of Coibion and Gorodnichenko (2015). In the next section, we discuss how we choose between two options. One is to follow the empirical model of Coibion and Gorodnichenko (2015) precisely, with forecast revisions occurring within one year rather than within one quarter. The other option is to modify the empirical model of Coibion and Gorodnichenko (2015) to keep the forecast revision within one quarter.

4 Modified empirical model

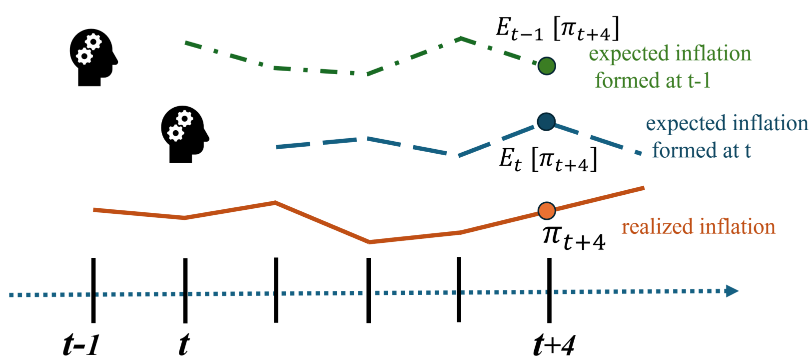

Figure 2: The original Coibion and Gorodnichenko model

Note: An increment of time is a quarter. The forecast horizon of the original model is three quarters. In this figure, we use a four-quarter-ahead forecast to be consistent with our empirical models.

For the sake of comparison, we restate the original model in Coibion and Gorodnichenko (2015) in equation (3), by changing the forecast horizon to be four instead of three quarters. This slight change does not alter the qualitative implications of the model. Equation (4) shows this slightly modified version, and Figure 2 illustrates the model structure graphically.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+4}]) + u_t \tag{4}$$

With the CIE dataset having forecast horizons available at only annual frequency, i.e., every four quarters, to perfectly follow the model of Coibion and Gorodnichenko (2015) in equation (4), we need to replace the second term in the forecast revision with $\bar{E}_{t-4}[\pi_{t+4}]$.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + u_t \tag{5}$$

This model in equation (5) differs from equation (4) only in that the previous forecast period is four quarters ago, i.e., one year instead of one quarter. This difference in forecast revision does not affect the empirical results or their implications. Note that, even in the original CG model in equation (3), the two forecast horizons in the forecast-revision term differ, i.e., four-quarter and five-quarter ahead forecasts. In equation (5), the difference in forecast horizons is much greater; The forecast horizon in the first term is one year, whereas the forecast horizon in the second term is two years.

We then propose further modifying an empirical model for the CIE dataset. A proposed model adjusts forecast horizons to be at the same length in the forecast revision as shown in equation (6). However, the forecast revision in this modified model (6) is fundamentally different from those in the original models (3) and (4). The forecasted inflation on the right-hand side has different targets, i.e., π_{t+4} and π_t, as shown visually in figure 3.[3]

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]) + u_t \tag{6}$$

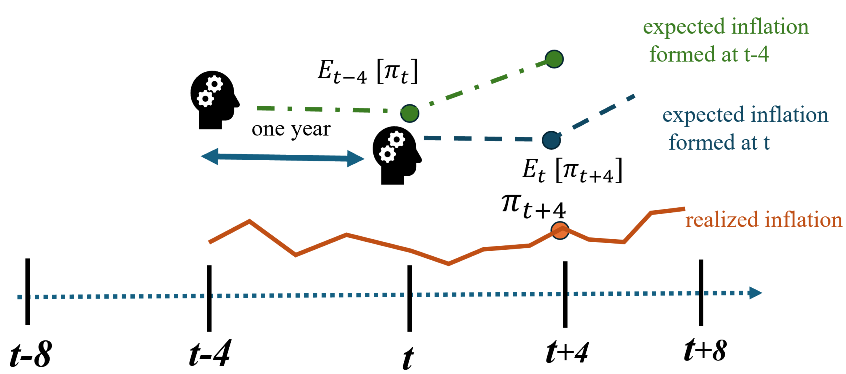

Figure 3: The modified Coibion and Gorodnichenko model

Note: An increment of time is a quarter. The CIE forecast horizon is every four quarters. For the inflation forecast at t+4, we use a four-quarter-ahead forecast made in t, and a eight-quarter-ahead forecast made in t-4.

How can we reconcile these differences between equation (5) and equation (6)? We can show in the following subsection that the modified model in equation (6) can be

[3]As it will be clear later in section 4.3, Coibion and Gorodnichenko (2015) also suggests a similar empirical model to equation (6).

interpreted as adding a new term 'forecast trend' to the original model of Coibion and Gorodnichenko (2015).

4.1 The link between the CG model and the modified model

We start by restating equation (6) below.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]) + u_t$$

Then, we add and subtract the same term, $\bar{E}_{t-4}[\pi_{t+4}]$ in the right-hand side of the equation, and rearranging yields the following equation.

$$\begin{aligned}\pi_{t+4} - \bar{E}_t[\pi_{t+4}] &= \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + K'_{CG}(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]) + u_t \\ &= \alpha + K_1(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + K_2(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]) + u_t \\ Forecast\ error_t &= \alpha + K_1 \ \cdot Forecast\ revision_t \ + K_2 \ \cdot Forcast\ trend_t + u_t \end{aligned} \tag{7}$$

Equation (7) decomposed the modified forecast revision term in equation (6) into two independent terms. The first term on the right-hand side of the equation is the CG term in equation (5). The second term is the difference between an eight-quarter-ahead forecast and a four-quarter-ahead forecast, both forecast at the same quarter, $t-4$. We coined the latter term as 'forecast trend'.

Figure 4 shows the time series of forecast errors, forecast revision, and forecast trend consistent with the decomposition equation (7).[4] With a simple rational expectation hypothesis, the expected forecast error is zero. With a finite sample, this implies that, on average, forecast errors should be zero. In Figure 4, the forecast errors seem to satisfy the hypothesis, fluctuating around zero. However, we frequently observe large forecast errors around the global financial crisis and the consumption tax hikes.

The forecast revision, shown in the dotted line, is less volatile and persistent. It is generally negative in the 90s and the early 00s. The forecast trend in the dashed line takes low but positive values in most of the sample. This implies that consumers and firms in Japan have a higher inflation expectation for the more distant future.[5]

4.2 The decomposed model and the test of equal coefficients

In the previous sections, we started from the modified model in equation (6), and we showed that this modified model can be decomposed into two terms. In this section, we start from the original model in equation (5). This model is based on a theoretical

[4]In appendix figures, we show the time series of each component in forecast revisions, modified forecast revision, and various forecast trends.

[5]Here and elsewhere, we do not always explicitly list all agents that consist of the forecast index. The forecast index is constructed by combining the forecasts of consumers, firms, and professional forecasters such as economists working in financial institutions (see section 3).

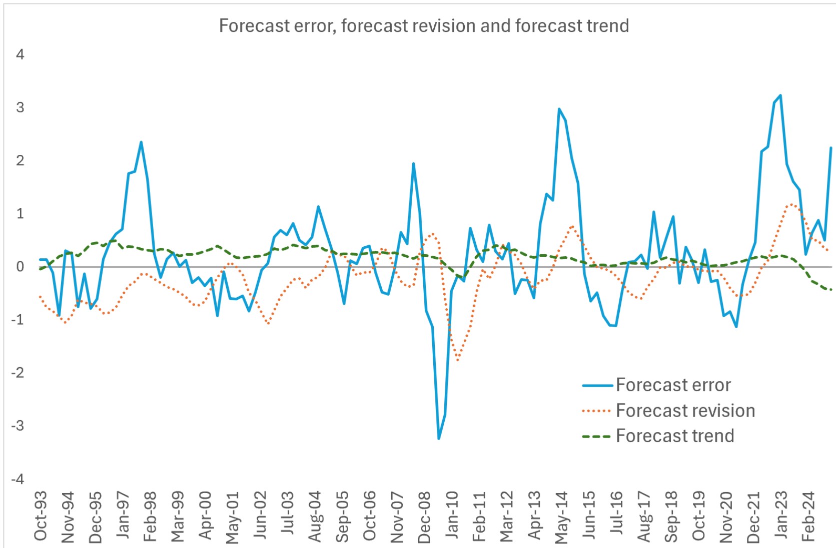

Figure 4: Forecast errors, forecast revisions, and forecast trend

Note: The figure follows the specification of equation 5. For the data points in 2020:Q4, the realized inflation is that of 2020:Q4, the four-quarter-ahead forecast is made in 1999:Q4, the eight-quarter-ahead forecast is made in 1998:Q4, and we also have the 1999:Q4 forecast made in 1998:Q4.

framework with information rigidity and includes only a forecast-revision term on the right-hand side. Any other control variables have no explanatory power if the rational expectation assumption holds. In Coibion and Gorodnichenko (2015), except for one variable, i.e., unemployment rate, control variables were not statistically significant.

In a similar manner, we can introduce forecast trend as a control variable to equation (5). From this point of view, there is no a priori justification for assuming that the coefficients on forecast revision and forecast trend are equal. To capture this argument, the second and last equations specify different coefficients, K_1 and K_2, for the forecast revision and the forecast trend, respectively. Having explained that, these two coefficients can be equal if the specification in equation (6) is correct. On the other hand, equation (6) is misspecified if K_1 and K_2 are not equal.[6] We can test whether the specification in equation (6) is correct by testing the null hypothesis of equal coefficients of K_1 and K_2 by the classical Wald F-test.

We have two-step tests of the decomposed model. The first step is to test the hypothesis of the original CG forecast revision using control variables. The null hypothesis is $H_0 : K_2 = 0$ against the alternative $H_1 : K_2 \neq 0$. A rejection of the null hypothesis implies that the underlying inflation expectation is consistent with our model but not with the CG model. In section 7.3 and Appendix D, we show that the modified forecast revision model is consistent with a variant of a noisy information model. The second test is to test

[6] For example, if $K_1 > K_2$, the additional term, $-(K_1 - K_2) \cdot forecast\ trend_t$ appears on the right-hand side of equation (6).

the hypothesis of the modified forecast revision model introduced in this study. The null hypothesis is that $H_3 : K_1 = K_2$, given $K_2 \neq 0$. A rejection of the null hypothesis implies that neither the original CG nor the modified empirical model in this study is sufficient to describe the underlying formation of expectations.

4.3 Modified forecast revisions with different revision spans

Using the decomposition technique, an alternative empirical model for the CIE dataset is proposed to reflect forecast revisions within one quarter as in equation (8). However, the forecast-revision term in this modified model is fundamentally different from that in the original model. The forecasted inflation on the right-hand side has a one-quarter difference, i.e., π_{t+4} and π_{t+3}, as shown visually in figure 5.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]) + u_t \tag{8}$$

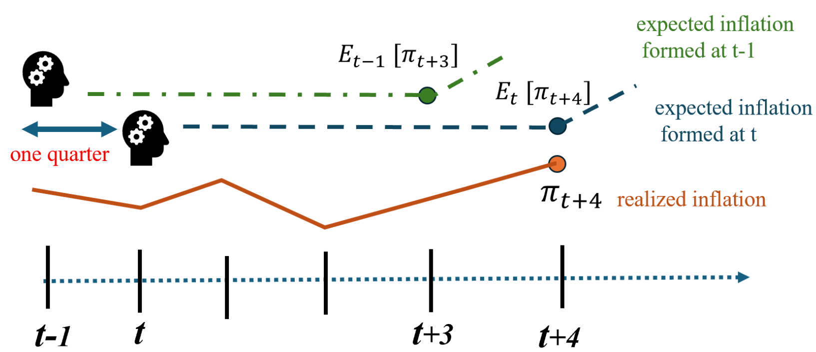

Figure 5: The alternative modified Coibion and Gorodnichenko model

Note: An increment of time is a quarter. The CIE forecast horizon is every four quarters. In this framework, the forecast horizon is four quarters. However, there is a quarter gap between the two inflation forecasts.

To be precise with alternative specifications, we define the triplets that determine the forecast revision in the forecast-error regression models, including that of Coibion and Gorodnichenko (2015). The triplet, $FR(h, r, g)$, for the forecast revision that contains the forecast horizon (h), the forecast revision span (r), and the gap in forecast horizon (g). The original forecast revision term in Coibion and Gorodnichenko (2015), as shown in equation (3), is $FR(3, 1, 0)$. The forecast horizon is three-quarters ahead, the forecast was revised after a quarter, and there is no gap in forecast horizons between the two forecasts. Similarly, the four quarters ahead forecast version of Coibion and Gorodnichenko (2015) in equation (4) is $FR(4, 1, 0)$. Equation (5), obtained by applying it directly to the CIE dataset, has $FR(4, 4, 0)$. For these three specifications, there is no gap between the forecast horizons in the forecast revision term.

Now, for equation (8), the triplet is $FR(4,1,1)$. In the modified forecast revision term, there is a one-quarter difference in the targeted forecast quarters. For equation (6), the triplet is $FR(4,4,4)$.

By understanding the structure of these triplets, we can define the forecast error regression models analogously for $FR(4,3,3)$, $FR(4,2,2)$, and $FR(4,1,1)$. The forecast error regression models with $FR(4,3,3), FR(4,2,2)$, and $FR(4,1,1)$ are the following, respectively. Note that the last equation, i.e., equation (11), is exactly the same as equation (8).

The $FR(4,3,3)$ model has a four-quarter-ahead forecast horizon, a three-quarter revision span, and a three-quarter forecast gap.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]) + u_t \tag{9}$$

The $FR(4,2,2)$ model has a four-quarter-ahead forecast horizon, a two-quarter revision span, and a two-quarter forecast gap.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-2}[\pi_{t+2}]) + u_t \tag{10}$$

The $FR(4,1,1)$ model has a four-quarter-ahead forecast horizon, a one-quarter revision span, and a one-quarter forecast gap.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K'_{CG}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]) + u_t \tag{11}$$

It is important to note that equation (11) is equivalent to equation (18) in Coibion and Gorodnichenko (2015) when they apply their empirical model to the Michigan Surveys of Consumers and the extracted measures of market expectations. They had to modify their empirical model for the same reason as in our case: The data structure was different from that of the SPF. In addition, they discuss that the forecast revision is correlated with the error term, and they estimate the model by IV. We will address this issue when we estimate our models in section 5.1.

The decomposed versions are similarly shown as follows, respectively. For $FR(4,3,3)$ model,

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K_1(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + K_2(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]) + u_t \tag{12}$$

For $FR(4,2,2)$ model,

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K_1(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + K_2(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-2}[\pi_{t+2}]) + u_t \tag{13}$$

For $FR(4,1,1)$ model,

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha + K_1(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + K_2(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]) + u_t \tag{14}$$

5 The empirical results

5.1 The comparison between the CG and the modified forecast revisions

For the Japanese inflation forecast, the CG model results in Table 1 show that it is consistent with the information rigidity model. Column (i) shows the results for the forecast revision that is consistent with Coibion and Gorodnichenko (2015) but with a different span between two forecasts. In Coibion and Gorodnichenko (2015), the forecast revision is measured between forecasts one quarter apart. Instead, in this study, the forecast revision is updated after four quarters. The estimated coefficient is positive and statistically significant at the one percent level. It rejects the null hypothesis of the FIRE that the forecast error is independent of the forecast revision.

In column (ii), apart from the original model of Coibion and Gorodnichenko (2015), we introduced the modified forecast revision. Interestingly, the coefficient of the modified model is similar to that of the CG model. The coefficient is positive and statistically significant. By decomposing the modified forecast revision, we obtain the original forecast revision in Coibion and Gorodnichenko (2015) and the forecast trend term. Column (iii) shows that the forecast trend is not statistically significant, whereas the original forecast revision is statistically significant at the one percent level. These results support the original CG model over the modified forecast-revision model, although the Wald test does not reject the equality between K_1 and K_2.

As to rational expectations, there should not be any variables that have explanatory power for forecast errors. That also includes a constant term. However, the constant terms are statistically significant in Table 1. As in Coibion and Gorodnichenko (2015), we re-estimated the models by imposing zero constant terms and find that forecast revisions remain statistically significant. The estimation results are shown in Appendix Table A.1

As discussed in section 4.3, the modified forecast revision terms are correlated with the error terms, and they should be estimated with instrumental variable estimation or two-stage least squares. The estimated results are shown in Appendix Table A.2. In columns (i) and (ii) of Table A.2, the estimated results of IV estimation with a half-year percentage change in WTI price as an instrumental variable are shown.[7] The estimates remain statistically significant. However, compared with the OLS estimates of 0.606 and 0.614, the IV estimates are substantially larger, at 1.297 and 1.446, respectively. The F-statistics in the first stage reject the null hypothesis of no explanatory power, see Appendix Table A.3.

[7]As instrumental variable candidates, we considered the percentage changes of WTI price and the BOJ's Corporate Goods Price Index (CGPI) in alternative periods and lags. The oil price is applied in Coibion and Gorodnichenko (2015), and CGPI is shown to be closely related to the CPI in Sasaki et al. (2022). In a case of multiple instrumental variables, we also checked Wooldridge's overidentification tests. As a result, the percentage change in the oil price over half a year is selected. The coefficients of forecast revisions often become insignificant in instrumental variable estimation when the consumption tax hike dummies are included. The consumption tax hike dummies will be discussed in section 5.3.

Table 1: CG forecast revision and the modified forecast revision

	(i) eq.(5) $FR(4,4,0)$	(ii) eq.(6) $FR(4,4,4)$	(iii) eq.(7) $FR(4,4,4)$	(iv) eq.(5) $FR(4,4,0)$	(v) eq.(6) $FR(4,4,4)$	(vi) eq.(7) $FR(4,4,4)$
CG forecast revision	0.606***		0.622***	0.451**		0.454**
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]$	(0.181)		(0.186)	(0.191)		(0.192)
Modified forecast revision		0.614***			0.430**	
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]$		(0.182)			(0.186)	
Forecast trend (4Qs)			0.500			0.111
			(0.513)			(0.496)
dummy 3% to 5%				1.756***	1.684***	1.738***
				(0.157)	(0.167)	(0.171)
dummy 5% to 8%				1.849***	1.878***	1.849***
				(0.376)	(0.365)	(0.377)
dummy 8% to 10%				-0.333*	-0.268*	-0.317*
				(0.168)	(0.159)	(0.189)
Constant	0.373***	0.260***	0.283*	0.241**	0.156*	0.221
	(0.107)	(0.086)	(0.146)	(0.112)	(0.088)	(0.154)
Observations	126	126	126	126	126	126
Adj. R-squared	0.085	0.091	0.084	0.256	0.254	0.250
Wald			0.06			0.44

Note: The dependent variable is the forecast errors. Robust standard errors are in parentheses. ***, **, * represent one, five, and ten percent significance levels. Wald indicates the F-value for testing equal coefficients of forecast revision and forecast trend, and it follows F(1, 120). The triplet, $FR(h, r, g)$, for the forecast revision that contains the forecast horizon (h), the forecast revision span (r), and the gap in forecast horizon (g).

5.2 Alternative forecast revision spans

In columns (i) through (iii) in Table 2, the estimated coefficients of modified forecast revision are all statistically significant at the one percent level. Regardless of the specifications of the modified forecast-revision terms, the empirical results support the rational-expectations model with information rigidity. Interestingly, the degree of impact becomes larger as the revision span, and the forecast gap becomes shorter.

In the decomposition specifications in columns (iv) through (vi), 'Forecast trends' are statistically significant regardless of the trend span. Again, these results may show that the inflation forecasts in Japan are not rational. However, in section 7.3 and Appendix D, we show that the modified forecast revision or forecast revision plus forecast trend are consistent with a variant of the noisy information model. More importantly, the estimated coefficients of the CG forecast revisions and the modified forecast revisions are quite similar. The estimated coefficients for the US inflation forecasts in Coibion and Gorodnichenko (2015) range from 1.06 to 1.20. This is close to the estimated coefficients of 0.92 in column (i) and 1.38 in column (ii).

Regarding the equality of coefficients, K_1 and K_2, F-values for the Wald test cannot reject the null of $K_1 = K_2$, for columns (iv) through (vi). These results support that the modified forecast revision models introduced in this study fit better in the Japanese inflation expectation than the original CG model.

Again, we addressed the issue of endogeneity by applying the IV estimation, and the estimated results are shown in columns (iii) through (v) in Table A.2. The F-tests in the first-stage regressions show that using a half-year percentage change in WTI price as the instrumental variable is adequate.

5.3 The effect of consumption tax hikes

On examining Japanese inflation, we should consider the episodes of consumption tax hikes as well as the introduction of the tax. Obviously, inflation increases after a rise in the consumption tax. Shoji (2022) investigated the effect of the consumption tax hike in 2014 on firms' behavior to raise prices. The consumption tax was first introduced in Japan in April 1989. Our sample period covers from 1991:Q4 to 2025:Q1. Therefore, the introduction of the consumption tax does not need to be considered in this study. However, the consumption tax was raised on three occasions later: from three to five percent in April 1997, from five to eight percent in April 2013, and from eight to ten percent in October 2019.

We constructed consumption hike dummies that take the value of one in the rising quarter and in the following three quarters. We need to note that inflation is defined as the percentage increase on a year-on-year basis. The rational expectation model with the information rigidity framework of Coibion and Gorodnichenko (2015) clearly indicates that no other control variables affect the forecast errors. However, these episodes occur as temporary events, not as continuous variables. Therefore, we should understand that the

Table 2: Alternative forecast revision spans

	(i) eq.(9) FR(4,3,3)	(ii) eq.(10) FR(4,2,2)	(iii) eq.(11) FR(4.1,1)	(iv) eq.(12) FR(4,3,3)	(v) eq.(13) FR(4,2,2)	(vi) eq.(14) FR(4.1,1)
Modified forecast revision	0.924***					
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]$	(0.228)					
Modified forecast revision		1.375***				
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-2}[\pi_{t+2}]$		(0.335)				
Modified forecast revision			2.528***			
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]$			(0.548)			
CG Forecast revision				0.938***	1.392***	2.460***
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]$				(0.238)	(0.348)	(0.574)
Forecast trend (3Qs)				1.260***		
				(0.471)		
Forecast trend (2Qs)					1.479***	
					(0.489)	
Forecast trend (1Q)						2.335***
						(0.663)
Constant	0.259***	0.258***	0.256***	0.199	0.242**	0.280***
	(0.084)	(0.082)	(0.080)	(0.128)	(0.112)	(0.104)
Observations	126	126	126	126	126	126
Adj. R-squared	0.145	0.182	0.208	0.144	0.176	0.204
Wald				0.75	0.11	0.35

Note: The dependent variable is the forecast errors. Robust standard errors are in parentheses. ***, **, * represent one, five, and ten percent significance levels. Wald indicates the F-value for testing equal coefficients of forecast revision and forecast trend, and it follows F(1, 120). The triplet, $FR(h, r, g)$, for the forecast revision that contains the forecast horizon (h), the forecast revision span (r), and the gap in forecast horizon (g).

null hypothesis of the rational expectation model with information rigidity is tested only on these event periods. If the null hypothesis is rejected, we interpret it as a temporal deviation from the rational expectation under information rigidity.

From columns (iv) to (vi) in Table 1, two earlier consumption tax hike dummies are statistically significant at the one percent level. These results hint that Japanese inflation may follow non-rational expectations. Interestingly, the fitness of regression has more than doubled. The results are qualitatively the same as in columns (i) through (iii). For the decomposed regressions, the estimated results are presented in Appendix Table A.4. For columns (i) through (iii), two of the three consumption tax hike dummies are statistically significant at the one percent level. This invalidates the null hypothesis that no other control variables affect forecast errors, indicating a deviation from the rational expectation hypothesis on these events.

5.3.1 Why the consumption tax hike pre-announcement is not orthogonal to forecast errors?

One crucial issue remains: whether Japanese consumers/voters correctly anticipated the consumption tax hikes in quarter one year earlier. It is essential to note that raising the consumption tax rate was a political issue between the majority political party and competing parties. On the two occasions, the planned increase of the consumption tax was postponed.

Forecast errors on inflation observed in the month of the consumption tax increase should be orthogonal to the forecast revision made one year earlier if the tax increase is correctly anticipated. However, the forecast error can be positive in a substantial size if consumers are taken by surprise after making a forecast. Can we resort to the external sources of information to assess how confident consumers were that a tax hike would occur at the time of making a forecast? In this study, we suggest using media coverage to represent the degree of tax-hike anticipation, following the works of Baker et al. (2016) and Chahrour et al. (2025).[8]

The Nikkei Telecom allows keyword search across eight newspapers, three news bulletins, and press releases from private firms. In this study, we selected the Nikkei newspaper's morning and evening issues. Note that the evening issue is not a revised version of the morning issue. In Japan, newspapers are published twice daily. We used Japanese keywords related to 'consumption tax increase' for an entire month and counted the number of articles containing either keyword.[9] The records of article counts are recorded as a screenshot by one researcher, and another researcher checked the screenshots for the

[8]Chahrour et al. (2025) examines the effect of news information on the formation of inflation expectations. They use survey questions that ask a consumer respondent whether she heard any good/bad news about business conditions and inflation. Appendix Figure A1 in their study compares the newspaper coverage of inflation and inflation expectations.

[9]The following Japanese Keywords are used in a union boolean operation. These are 'Shouhizeiritsu Hikiage' and 'Shoubizeiritsu jump.'

integrity of using the correct keywords, the correct date span, and whether the correct number of articles is recorded.

In Figure 6, the number of newspaper articles containing the relevant phrase of 'consumption tax increase' appearing in the corresponding months is depicted. For the three consumption tax hikes in April 1997, April 2014, and October 2019, the monthly counts of articles for the previous 18 months are shown.

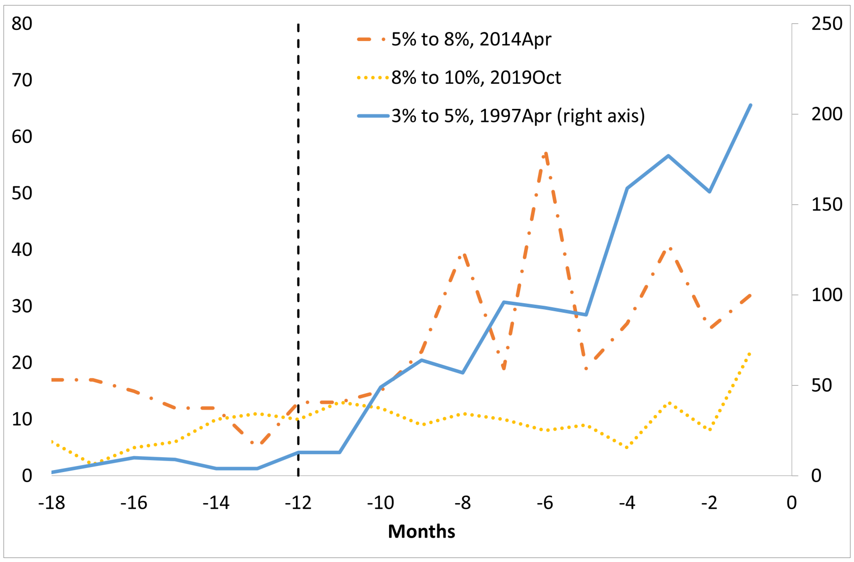

Figure 6: Consumption tax hike anticipations

Note: The horizontal axis indicates the months before the consumption tax hikes in April 1997, April 2014, and October 2019. The vertical dotted line indicates one year before the implementation of a tax hike. The solid line represents the number of articles per month before the implementation of the tax hike in April 1997, as indicated on the right vertical scale. The dash-dotted line represents the tax hike in April 2014, and the dotted line represents the tax hike in October 2019, both on the left vertical scale.

The media coverage of the 1997 consumption tax hike, shown as a solid line, was relatively small in comparison to the later coverage in a few months of April 1997. There were only thirteen articles in April 1996; however, there were 159, 177, 157, and 205 in December 1996, January 1997, February 1997, and March 1997, respectively. The anticipation of the 1997 consumption tax hike by consumers one year earlier is far from certain. The ruling coalition in Japan was still contemplating raising the consumption tax to the planned level of five percent on April 1, 1997.[10]

The corresponding number of articles was much more modest in the 2014 consumption tax hike. The monthly counts never exceeded 60. The consumption tax hike in 2014 was also confusing because it was planned to raise the tax to eight percent in April 2014 and again to ten percent in October 2015. In hindsight, the latter rise was postponed.[11]

[10]The Nikkei Newspaper, April 20, 1996.

[11]Takahashi and Takayama (2025) investigates the effects of the 2014 tax hike episode on economic ac-

The media was the calmest about the 2019 consumption tax hike. The monthly counts never exceeded 30. This is partly due to the fact that raising the rate to ten percent was first planned four years earlier, and it was postponed twice.[12] The government was moving forward with adjusting other related frameworks, such as the medical payment system and the exception clauses, to maintain the current eight percent rate. Only on the last tax hike, the dummy was not statistically significant in specifications (i) through (vi) of Table A.4 at the five percent level.

To conclude this subsection, we note that the media coverage in the three episodes tends to increase as the time approaches the day of the tax increase. As stated earlier, forecast errors on inflation observed in the month of the consumption tax increase should be orthogonal to the forecast revision made one year earlier if the tax increase is correctly anticipated. However, we found that the consumption tax dummies are statistically significant, possibly due to uncertainty still prevailing one year earlier. On the contrary, the last tax hike did not affect the inflation forecast error, possibly due to the reduced uncertainty ironically brought by two postponements.

6 During the zero inflation period

Both inflation and inflation expectations in Japan have been peculiar over the last three decades; they became positive only recently in the post-COVID period. After the collapse of the bubble economy, symbolized by the plummeting stock market prices and real estate prices in the early 1990s, Japan has faced low or zero inflation and accommodative zero interest rates by the BOJ.[13] Only after the higher inflation in the rest of the world, driven partly by the rebounding demand and the bottleneck of global supplies, did Japan start experiencing domestic inflation in 2022. Given this background, it is a natural question to ask whether inflation expectations in Japan have gone through a structural break in the recent period.

We re-estimated the models using only the subsample up to 2019:Q4. The estimated results are shown in Table 3. In all six specifications, namely, the original CG forecast revisions in columns (i) and (ii), the modified forecast revisions in (iii) and (iv), and the decomposed framework in (v) and (vi), both the original and modified forecast revisions are not statistically significant. Under the zero-inflation regime in Japan, people formed their forecasts in accordance with rational expectations under full information. This evidence is convincing because people rationally expected the future inflation to be around zero percent, and the realized inflation later confirmed those expectations during the two-decade-long zero-inflation regime. As a result, forecast errors and forecast revisions were

tivities. They found that information about future tax hikes affects only future inflation and consumption, not other economic activities.

[12] In November 2014, the first scheduled date of October 2015 was postponed to April 2017. It was once again announced in June 2016 that raising the consumption tax to ten percent was rescheduled to October 2019.

[13] See Aoki and Ueda (2025), for the survey on the effects of unconventional monetary policy in Japan.

statistically independent. [14]

The forecast trends and the consumption tax hike dummies are statistically significant. Note that the null of $K_1 = K_2$ is rejected in columns (v) and (vi), indicating that the modified forecast-revision model of this study is not adequate, refer to the discussion in section 4.2. This evidence warrants reconsidering the information rigidity model in Coibion and Gorodnichenko (2015) and further investigating whether tax hikes were anticipated in the quarter one year earlier.[15]

Table 3: Under the zero inflation regime

	(i) eq.(5) $FR(4,4,0)$	(ii) eq.(5) $FR(4,4,0)$	(iii) eq.(6) $FR(4,4,4)$	(iv) eq.(6) $FR(4,4,4)$	(v) eq.(7) $FR(4,4,4)$	(vi) eq.(7) $FR(4,4,4)$
CG forecast revision	0.219	-0.162			0.165	-0.243
	(0.225)	(0.179)			(0.241)	(0.179)
Modified forecast revision			0.294	-0.039		
			(0.191)	(0.154)		
Forecast trend (4Qs)					1.393**	1.445***
					(0.597)	(0.521)
dummy 3% to 5%		1.919***		1.915***		1.729***
		(0.165)		(0.167)		(0.173)
dummy 5% to 8%		2.496***		2.386***		2.655***
		(0.341)		(0.331)		(0.334)
dummy 8% to 10%		-0.239**		-0.277***		-0.020
		(0.100)		(0.083)		(0.146)
Constant	0.206	-0.062	0.160*	-0.016	-0.117	-0.404**
	(0.131)	(0.106)	(0.093)	(0.081)	(0.222)	(0.187)
Observations	105	105	105	105	105	105
Adj. R-squared	0.003	0.371	0.016	0.366	0.034	0.407
Wald					2.83*	7.50***

Note: The dependent variable is the forecast errors. Robust standard errors are in parentheses. ***, **, * represent one, five, and ten percent significance levels. Wald indicates the F-value for testing equal coefficients of forecast revision and forecast trend, and it follows F(1, 99). The triplet, $FR(h, r, g)$, for the forecast revision that contains the forecast horizon (h), the forecast revision span (r), and the gap in forecast horizon (g).

[14] However, we cannot rule out the possibility that people formed rational expectations under noisy information during the zero inflation period. Due to zero inflation and limited variation in this period, even with sticky information or noisy information, forecast errors seemed independent of forecast revisions. In terms of the models, λ close to zero in equation (1) or G close to one in equation (2) might have been at play in Japan.

[15] A variant of the noisy information model proposed in section 7.3 and Appendix D suggests a possibility of misspecifications in Table 3.

6.1 Time-varying parameter estimates

So far, we have split the sample into the zero-inflation (zero-interest-rate) period and the post-COVID period, the latter of which is associated with higher inflation. Although this particular structural break point may be acceptable to most economists, there is no statistical ground to support this break. In this section, we follow the methodology proposed in Inoue et al. (2024), Inoue et al. (2025), and Müller and Petalas (2010) to capture the smooth dynamics of parameter changes of forecast revision effect on the forecast errors.[16]

Müller and Petalas (2010) suggested the efficient estimation of the parameter path by minimizing weighted average risks (WAR). Inoue et al. (2024) extends the method to local projections and vector autoregressive models. Inoue et al. (2025) provides examples of applications and a description of the codes, applicable to the OLS, IV, VAR, and local projection estimations.

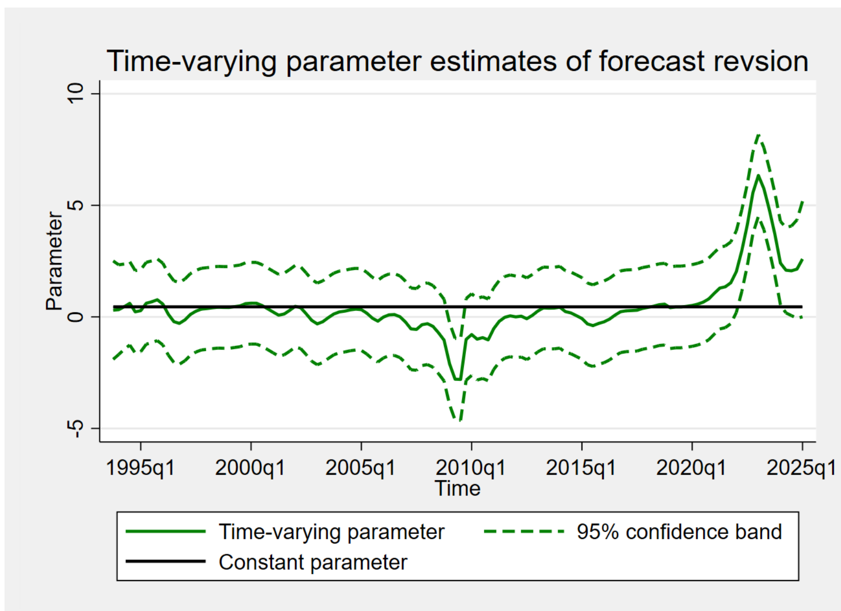

Figure 7: Time-varying parameter estimates of forecast revisions, FR(4,4,0)

Figure 7 shows the time-varying parameter of forecast revision on forecast error in the model of equation (5). The figure confirms that forecast revision has no effect on forecast errors in the 90s, 00s, and 10s, except in the wake of the global financial shock. By contrast, the forecast revision effect is statistically significant in the post-COVID period, thus confirming the earlier results. Specifically, the peak occurs in the first quarter of 2023, the point estimate reaching 6.34. The positive effect of forecast revision is statistically significant between 2022:Q1 and 2025:Q1, except for the last quarter in 2024. Time-

[16] Alternatively, one can adopt methodologies for inferring the break dates, given the number of breaks.

varying parameter estimates for other modified forecast-revision models are presented in Appendix C. They all demonstrate qualitatively the same results.

6.2 The shifts in the BOJ monetary policy

The sample in this research spans three decades, and the economic and social environments in which Japanese consumers and firms formed inflation expectations have evolved through numerous significant episodes. In this subsection, we focus on the Bank of Japan's monetary policy and discuss the empirical results from the previous sections regarding the evolution of BOJ monetary policy.

The sample begins in the first quarter of 1991, i.e., with the collapse of Japan's bubble economy. The policy rate reached six percent in August 1990, following five consecutive rate increases to tame overheating in the stock and real estate markets. After observing the stock price plummeting, the BOJ lowered the policy rate to 5.5 percent in July 1991, and subsequent cuts brought it to 0.5 percent in September 1995. Figure 8 shows Japan's policy rate at the end of the month from January 1990 to September 2025.

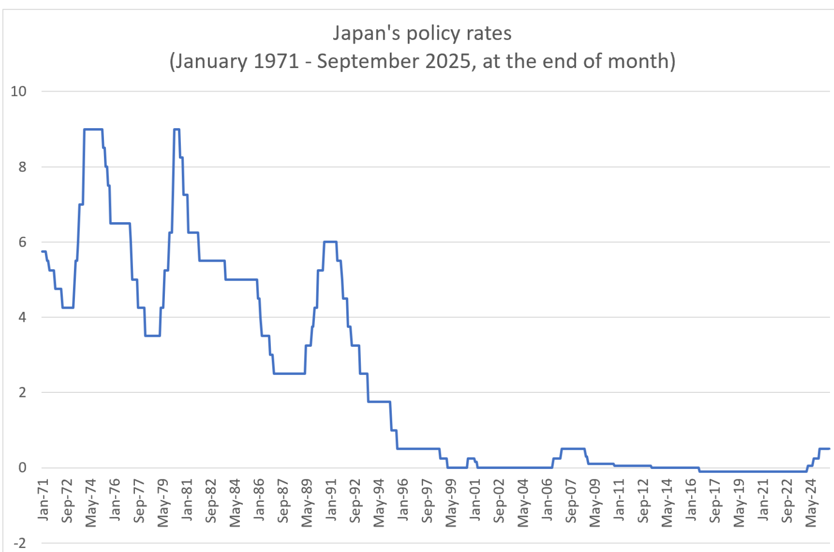

Figure 8: Japan's policy rate from January 1971 to August 2025

Note: The data source is the Central Bank Policy Rates, BIS (2025).

Meanwhile, the Bank of Japan Act was amended to provide the Bank with independence from the government and accountability for its monetary policy conduct. Mr. Masaru Hayami was the first governor appointed by the new BOJ Act in March 1989. Governor Hayami lowered the policy rate to 0.25 percent in September 1998 and introduced the so-called zero interest policy in February 1999. During his five-year term, which ended in March 2003, he also introduced quantitative easing as a monetary policy. The next governor was Toshihiko Fukui, who served from March 2003 to March 2008. During his term, the BOJ terminated the zero-interest policy and raised the policy rate to 0.5 percent.

Governor Masaaki Shirakawa, serving in office from April 2008 to March 2013, faced the turmoil of the global financial crisis and the devastating shock of the 2011 great earthquake. He quickly reintroduced the zero-interest policy and the quantitative easing. Following Governor Shirakawa, Mr. Haruhiko Kuroda began his governorship by introducing an inflation target and massive quantitative easing at a level unmatched by previous policy. Since March 2013, the BOJ's unconventional monetary policy has differed markedly in its operations.

However, the time-varying parameter estimates of the forecast revision effect in Figure 7 indicate that evolving monetary policies over the last three decades did not affect the formation of inflation expectations in Japan. At least from the perspective of the forecast error regressions in this study, we find no evidence that either monetary policy before March 2013 or after March 2013 influenced the formation of inflation forecasts in Japan.

7 Discussion

7.1 What is the forecast trend?

We found that the forecast trends, for j = 1, 2, 3, $\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4-j}]$ introduced in this study, have explanatory power in the forecast error regression.[17] However, we have not thoroughly discussed what the forecast trend represents. The forecast trend defined so far in this study is tricky because two inflation forecast terms have different forecast targets, $t+4$ and $t+4-j$, and different forecasting periods at $t-4$ and $t-j$.

The concept of trend shock discussed in Mertens (2016) has some similarity with the forecast revision. Trend inflation in Mertens (2016) is identified as a forecast of inflation at the infinite horizon conditional on an information set in t, $E(\pi_{t+\infty}|\boldsymbol{\Omega_t}) = \tau_t$. The trend shock is obtained as $E_t[\pi_{t+\infty}] - E_{t-1}[\pi_{t+\infty}]$, by differencing the trend inflation. The forecast's target date is the same, i.e., the infinite horizon, but the information set is different. This is a variant of forecast revision in Coibion and Gorodnichenko (2015), with an infinite forecast horizon.

Instead of discussing the current form of the forecast trend, let us further decompose it into two terms.

$$\begin{aligned}\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4-j}] \\ = (\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4}]) + (\bar{E}_{t-j}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4-j}])\end{aligned} \tag{15}$$

[17]The forecast trend for j = 4 does not need further decomposition introduced in this subsection. It should be noted that it is not statistically significant in Table 1.

The two terms in the first parenthesis on the right-hand side of equation (15), $\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4}]$, have the same forecast target at $t+4$, but with different information sets. This first part is another variant of the CG forecast revision.[18] Therefore, we do not repeat the discussion about the CG forecast revision to understand how we should interpret the first part.

The two terms in the second parenthesis in equation (15), $\bar{E}_{t-j}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4-j}]$, now have the same information set at $t-j$, but with different forecast horizons. This is a more natural form of forecast trend that indicates whether a forecaster expects higher or lower inflation at a longer horizon.

To conclude, our definition of the forecast trend, $\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-j}[\pi_{t+4-j}]$, is the sum of a variant of forecast revision in Coibion and Gorodnichenko (2015) and a change in inflation forecast for a longer horizon. We can rely on this representation in equation (15) to interpret the forecast trend.

7.2 A long-lasting positive forecast trend

The forecast trend in Figure 4 is positive in general throughout the sample period, except during the global financial crisis and after the post-pandemic. Therefore, even in the low-inflation regime of the 00s and 10s, people in Japan expected higher inflation over longer horizons. This observation is somewhat surprising because people continuously anticipated a rise in inflation in the distant future, only to discover later that they were wrong. They continued to make the same mistakes repeatedly.

In a different context in international finance, a similar phenomenon was observed in the late 70's as the Peso problem in Mexico, see Lizondo (1983). After the collapse of the Bretton Woods system, Mexico continued to fix the Mexican Peso to the US dollar at 0.08. If this fixed exchange rate is assured with 100 percent, the exchange rate in the forward market should also be at 0.08; however, the forward transactions continued to be exchanged at a lower price for the Peso by about two percent. This episode was shown as an example against the rational expectation hypothesis. In September 1986, the Peso was revalued to 0.05. However, Krasker (1980) argues that testing market efficiency, rational expectations, and forward market requires considering events that cause a large change with a small probability, and that traditional tests need to be reevaluated. He applies his testing methodology to the Mark/Pound forward market during the hyperinflation in 1921-23. Therefore, the long-lasting positive forecast trend in Japan remains consistent with rational expectations, given a very low probability that Japan may experience very high inflation.

[18]This representation has similarity with the forecast smoothing model discussed in Coibion and Gorodnichenko (2015). In the forecasting smoothing model, two forecast revision terms appear.

7.3 Is the modified forecast model based on a theory?

The empirical model of Coibion and Gorodnichenko (2015) is restated as equation (3) in the literature review section. The theoretical models of sticky information in equation (1) and of noisy information in equation (2) are based on some specific underlying assumptions. Some modifications in assumptions can lead to alternative forms of the empirical model. For example, Shintani and Ueda (2023) proposes an empirical forecast error model with four forecast-revision terms by combining sticky-information and noisy-information models. In addition, as discussed in section 4.3, forecast revision and the error term are correlated in equation (3). To address this issue, we implemented IV estimations in the Appendix. Here, we explicitly derive the empirical model so that the error term consists solely of the rational-expectations error.

Table 4: Noisy information model

	(i) Table 1 (iii)	(ii) Noisy inf.	(iii) Table 1 (vi)	(iv) Noisy inf.
Forecast revision	0.622***	0.703***	0.454**	0.547***
	(0.186)	(0.192)	(0.193)	(0.201)
Forecast trend (4Qs)	0.500	0.218	0.111	-0.189
	(0.513)	(0.513)	(0.496)	(0.496)
correction term		-0.323***		-0.315***
$\pi_t - \bar{E}_{t-4}[\pi_{t+4}]$		(0.112)		(0.111)
dummy 3% to 5%			1.738***	1.835***
			(0.171)	(0.173)
dummy 5% to 8%			1.849***	1.701***
			(0.377)	(0.326)
dummy 8% to 10%			-0.317*	-0.343**
			(0.189)	(0.155)
Constant	0.283	0.419***	0.221	0.363**
	(0.146)	(0.152)	(0.154)	(0.161)
Observations	126	126	126	126
Adj. R-squared	0.084	0.140	0.250	0.305

Note: The dependent variable is the forecast errors. Columns (i) and (iii) repeat, for the comparison, the estimated results shown in Table 1. Columns (ii) and (iv) add the correction term variable to be consistent with the noisy information model in the appendix. Robust standard errors are in parentheses. ***, **, * represent one, five, and ten percent significance levels.

The empirical specification used in this study, which deviates from the original model of Coibion and Gorodnichenko (2015), was necessitated by the lack of short-horizon inflation forecasts in the Japanese data. We showed by decomposition that our modified model, in fact, adds an additional forecast trend term to the original model. Up to this point, we interpreted the rejection of the forecast trend as evidence that the data conforms to the rational expectations hypothesis. Conversely, we rejected the information rigidity models of Coibion and Gorodnichenko (2015) when we found the forecast trend statistically significant, refer to section 4.2 for the null hypotheses in this study. However, one question arises. Is the modified empirical specification consistent with any information rigidity model?

In Appendix D, we show how a variant of the noisy information model can be transformed to include the forecast trend term introduced in this study. It also shows that an additional term, $\pi_t - \bar{E}_{t-4}[\pi_{t+4}]$, is necessary to comply with the noisy information model. Therefore, our empirical model with the correction term can be shown as follows:

$$\begin{aligned}\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \alpha &+ K_1(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) \\ &+ K_2(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]) + K_3(\pi_t - \bar{E}_{t-4}[\pi_{t+4}]) + u_t\end{aligned} \tag{16}$$

The estimated results are shown in Table 4. The first and third columns repeat the estimated results of columns (iii) and (vi) in Table 1 for comparison. The second column of Table 4 shows the estimated results for equation (16). The statistical significance of the forecast revision and forecast trend remains unchanged. The correction term, $\pi_t - \bar{E}_{t-4}[\pi_{t+4}]$, is negative and statistically significant at the one percent level. The fitness of regression increases substantially from 0.08 to 0.14. The specification in the fourth column, with tax-hike dummies, shows a similar result. With this additional result, we may conclude that inflation forecasts in Japan are formed under rational expectations with noisy information.

7.4 Application to the US SPF data

Is the modified forecast revision in this study only applicable to the Japanese inflation forecasts? We apply the model to the US Survey of Professional Forecasters (SPF) data to equations (5) through (7). For the four-quarter-ahead forecast, we used 'CIP6' from the SPF mean level dataset. For the eight-quarter-ahead forecast, we used 'CIPC' from the SPF mean level dataset. Given the availability of two-year-ahead forecasts, the sample period begins in the third quarter of 2005. The data sample spans from 2005:Q3 to 2025:Q3. Quarterly realized US inflation is constructed from the FRED monthly CPI.[19] First, we averaged the monthly CPI over three months to construct a corresponding quarterly CPI. Then, a percentage change is calculated over four quarters.

[19]Consumer Price Index for All Urban Consumers: All Items in U.S. City Average, Index 1982-1984=100, Monthly, Not Seasonally Adjusted.

Table 5: Application of the modified model to the US SPF data

	(i) eq.(5)	(ii) eq.(6)	(iii) eq.(7)
CG forecast revision	1.424*		1.479*
	(0.795)		(0.827)
Modified forecast revision		1.213*	
		(0.718)	
Forecast trend (4Qs)			0.555
			(0.844)
Constant	0.355	0.261	0.322
	(0.248)	(0.223)	(0.237)
Observations	73	73	73
Adj. R-squared	0.042	0.039	0.045

Note: Robust standard errors are in parentheses. ***, **, * denote significance at the 1, 5, and 10 percent levels, respectively.

The estimated results are shown in Table 5. In the first column, the forecast revision term, consistent with Coibion and Gorodnichenko (2015) but with a longer span between two forecasts, is positive and statistically significant. The estimated coefficient of 1.424 is 19 percentage points higher than the estimate reported in Coibion and Gorodnichenko (2015). This difference arises from differences in the sample period and the span between the two forecasts. In column (ii), the estimated coefficient for the modified forecast revision is positive and statistically significant at the ten percent level. These results provide support for our modified forecast revision in this study. Our framework works not only for Japanese forecasts but also for US forecasts. Column (iii) provides the estimates for the decomposed regression. Similar to the results in Table 1 for Japan, the forecast trend with a four-quarter span is not statistically significant for the US either.

8 Conclusion

In this study, we estimated inflation forecast errors in Japan using the Coibion and Gorodnichenko (2015) framework for the period from 1991:Q1 to 2025:Q1. We devise an alternative model to suit the data framework in Japan. Necessitated by the data unavailability in forecast horizons in the Japanese dataset, we introduced the modified form of 'forecast revision', which is closely related to the original 'forecast revision' of Coibion and Gorodnichenko (2015). We find that this new 'forecast revision' term is useful in investigating inflation forecasts. On the surface, the alternative model fundamentally deviates from the original model of Coibion and Gorodnichenko (2015), but we show that they are closely connected in the decomposed representation, with only an additional term for 'forecast trend'.

The finding of both $K_{CG} > 0$ and $K'_{CG} > 0$ rejects full-information rational expectations (FIRE) in Japan. More precisely, they are consistent with the rational expectations model with information rigidity. When the consensus forecast needs to be revised upward for the inflation expectation, they always underreact to the new information.

In the decomposition model, we find that the positive coefficient on the forecast trend, $K_2 > 0$, indicates that the forecast is too low when the forecast trend is rising. This contradicts the rational expectation hypothesis under full information. We conclude from this evidence that people in Japan forecast inflation under information rigidity and form their expectations that do not deviate from the rational expectations.

Finally, under the zero-inflation regime in Japan, people formed their forecasts in accordance with rational expectations, and there was no information frictions. This evidence is convincing because people rationally expected the future inflation to be around zero percent during the two-decade-long zero-inflation regime, and the realized inflation later confirmed those expectations. Both the subsample analysis and time-varying parameter estimates revealed that forecast errors and forecast revisions were statistically independent during the zero-inflation regime.

Regarding the implications for monetary policy decision-making, it is essential to acknowledge that inflation expectation formation in Japan has undergone a structural shift, particularly in the post-pandemic period. Moreover, based on our estimation results, consistent with those from the US dataset, we found that people in Japan underreact to new information. Regarding the BOJ's forward guidance policy, this implies that inflation forecasts will likely be lower than the level the BOJ aims to guide. This research result suggests that the BOJ reassess the impact of the forward guidance policy on the inflation forecasts.

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Appendix:

A The robustness checks

Table A.1: Estimation results by restricting no constant term

	(i) Table 1 (i)	(ii) Table 1 (ii)	(iii) Table 1 (iii)	(iv) No-constant (i)	(v) No-constant (ii)	(vi) No-constant (iii)
CG forecast revision	0.606***		0.622***	0.371**		0.565***
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]$	(0.170)		(0.171)	(0.161)		(0.190)
Modified forecast revision		0.614***			0.617***	
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]$		(0.167)			(0.198)	
Forecast trend Q4			0.500			1.317***
			(0.534)			(0.409)
Constant	0.373***	0.260***	0.283**			
	(0.092)	(0.086)	(0.134)			
Observations	126	126	126	126	126	126
Adj. R-squared	0.085	0.091	0.084	0.029	0.086	0.110

Note: The dependent variable is the forecast errors. Standard errors are in parentheses. ***, **, * denote significance at the 1%, 5%, and 10% levels, respectively.

Table A.2: Two-stage least squares regressions

	(i) eq.(5) $FR(4,4,0)$	(ii) eq.(6) $FR(4,4,4)$	(iii) eq.(9) $FR(4,3,3)$	(iv) eq.(10) $FR(4,2,2)$	(v) eq.(11) $FR(4,1,1)$
CG forecast revision	1.297*				
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]$	(0.751)				
Modified forecast revision		1.446*			
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]$		(0.845)			
Modified forecast revision			1.732*		
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]$			(0.941)		
Modified forecast revision				2.643*	
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-2}[\pi_{t+2}]$				(1.547)	
Modified forecast revision					5.330**
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]$					(2.594)
Constant	0.501***	0.258***	0.257***	0.255***	0.251***
	(0.158)	(0.094)	(0.089)	(0.089)	(0.093)
Observations	126	126	126	126	126

Note: The regression is estimated by the two-stage least squares with a half-year percentage change in WTI price as an instrumental variable. The dependent variable is the forecast error. Robust standard errors are in parentheses. ***, **, * denote significance at the 1, 5, and 10 percent levels, respectively. The triplet, $FR(h, r, g)$, for the forecast revision that contains the forecast horizon (h), the forecast revision span (r), and the gap in forecast horizon (g).

Table A.3: First-stage regressions

Model	Robust F(1,124)	Prob > F
CG forecast revision $\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]$	26.135	0.00
Modified forecast revision $\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]$	19.114	0.00
Modified forecast revision $\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]$	20.816	0.00
Modified forecast revision $\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-2}[\pi_{t+2}]$	13.011	0.00
Modified forecast revision $\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]$	6.572	0.01

Note: The table reports robust F-statistics from the first-stage regressions.

Table A.4: Alternative forecast revision spans

	(i) eq.(9) FR(4,3,3)	(ii) eq.(10) FR(4,2,2)	(iii) eq.(11) FR(4.1,1)	(iv) eq.(12) FR(4,3,3)	(v) eq.(13) FR(4,2,2)	(vi) eq.(14) FR(4.1,1)
Modified forecast revision	0.708***					
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]$	(0.223)					
Modified forecast revision		1.107***				
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-2}[\pi_{t+2}]$		(0.327)				
Modified forecast revision			2.173***			
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-1}[\pi_{t+3}]$			(0.542)			
CG Forecast revision				0.718***	1.128***	2.169***
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]$				(0.230)	(0.336)	(0.558)
Forecast trend (3Qs)				1.004**		
				(0.441)		
Forecast trend (2Qs)					1.263***	
					(0.465)	
Forecast trend (1Q)						2.161***
						(0.641)
dummy 3% to 5%	1.650***	1.629***	1.640***	1.604***	1.609***	1.641***
	(0.159)	(0.193)	(0.188)	(0.157)	(0.199)	(0.185)
dummy 5% to 8%	1.763***	1.725***	1.807***	1.802***	1.763***	1.803***
	(0.337)	(0.268)	(0.236)	(0.335)	(0.274)	(0.277)
dummy 8% to 10%	-0.259	-0.256*	-0.256	-0.214	-0.235	-0.257
	(0.160)	(0.150)	(0.160)	(0.185)	(0.166)	(0.172)
Constant	0.159*	0.161*	0.156*	0.105	0.134	0.158
	(0.085)	(0.083)	(0.080)	(0.134)	(0.118)	(0.109)
Observations	126	126	126	126	126	126
Adj. R-squared	0.293	0.326	0.366	0.292	0.322	0.360
Wald				0.59	0.25	0.00

Note: The dependent variable is the forecast errors. Robust standard errors are in parentheses. ***, **, * represent one, five, and ten percent significance levels. Wald indicates the F-value for testing equal coefficients of forecast revision and forecast trend, and it follows F(1, 120). The triplet, $FR(h, r, g)$, for the forecast revision that contains the forecast horizon (h), the forecast revision span (r), and the gap in forecast horizon (g).

B Forecast revisions and modified forecast revisions

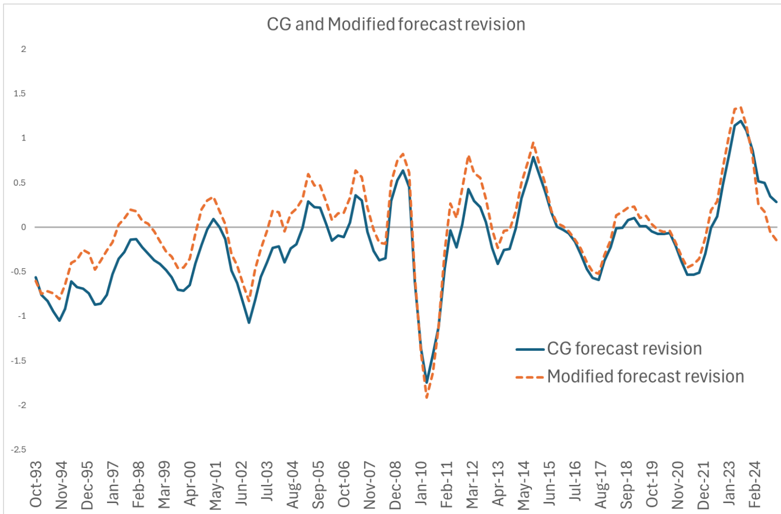

Figure B.1: Forecast revisions: FR(4,4,0) and FR(4,4,4)

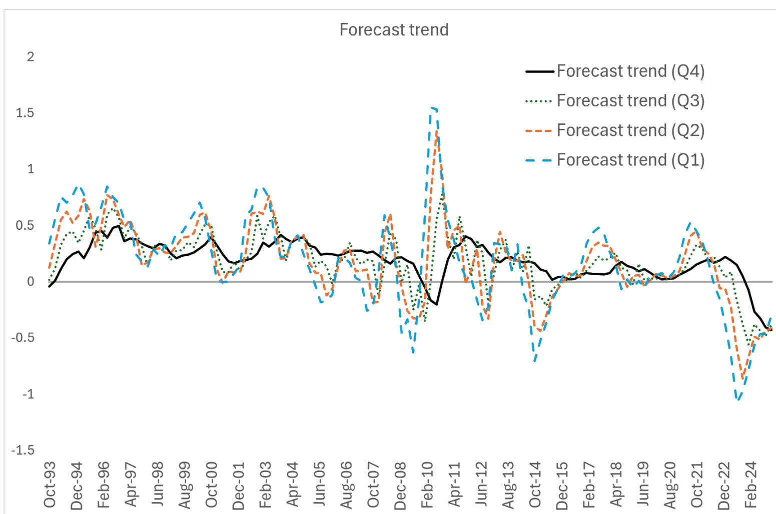

Figure B.2: Alternative forecast trends

C Time-varying parameter estimates

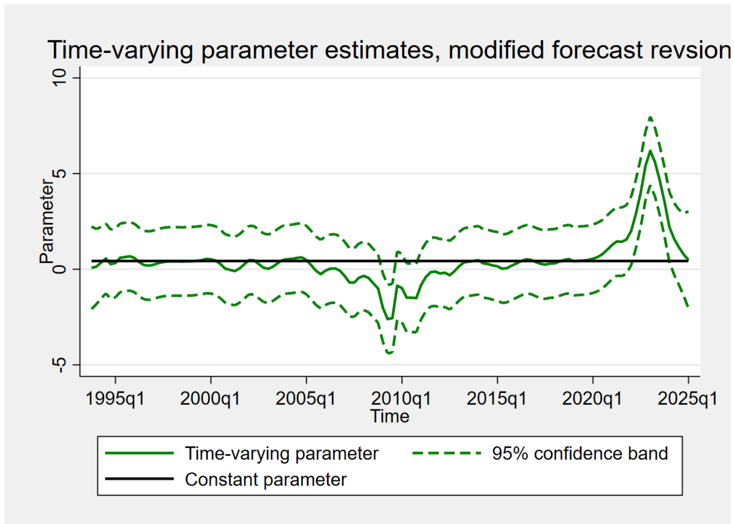

Figure C.1: Forecast revisions, FR(4,4,4)

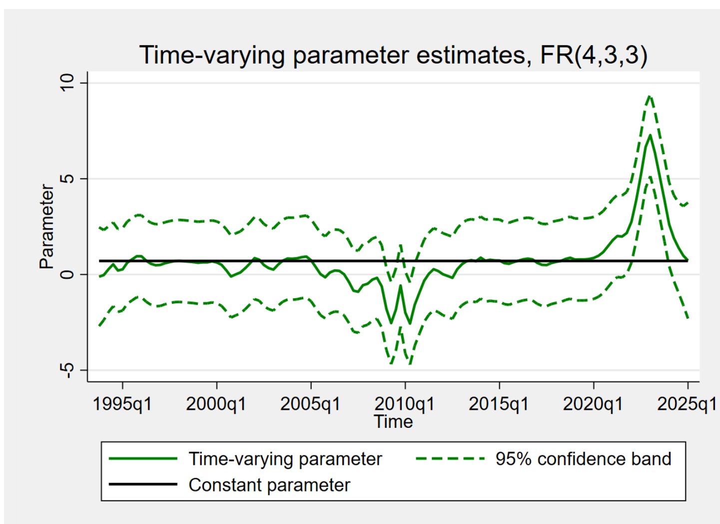

Figure C.2: Forecast revisions, FR(4,3,3)

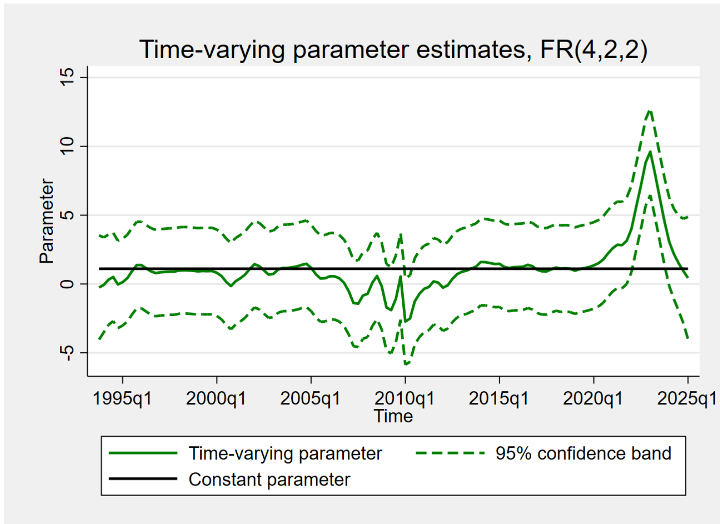

Figure C.3: Forecast revisions, FR(4,2,2)

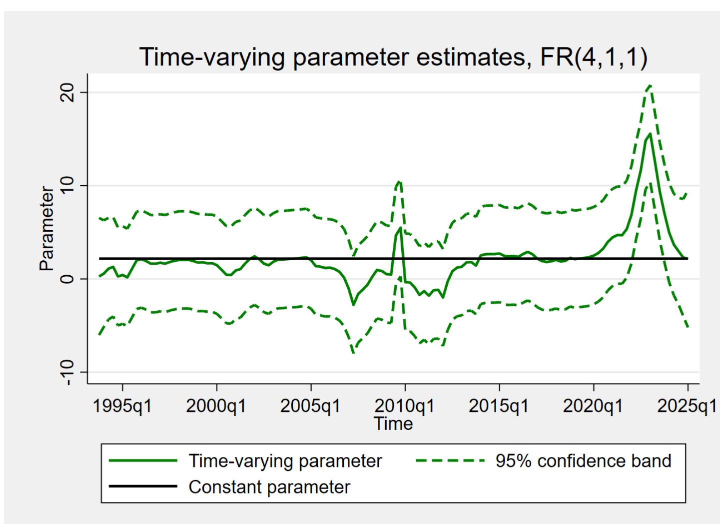

Figure C.4: Forecast revisions, FR(4,1,1)

D Noisy information model

In this appendix, we follow the modeling of Coibion and Gorodnichenko (2015) and Shintani and Ueda (2023) and make necessary changes to fit our modified model in equations (7) and (16).

We assume that the true inflation follows AR(n) process. The observed inflation consists of the true inflation and the noise. This system can be represented in the state-space framework. The state transition equation is as follows.

$$\Pi_t \equiv \begin{bmatrix} \pi_t \\ \vdots \\ \pi_{t-n+1} \end{bmatrix} = \begin{bmatrix} \rho_1 & \cdots & \cdots & \rho_n \\ 1 & 0 & 0 & 0 \\ \ddots & \ddots & \ddots & \vdots \\ 0 & \ddots & 1 & 0 \end{bmatrix} \Pi_{t-1} + G\omega_t = F\Pi_{t-1} + G\omega_t \tag{D.1}$$

where $G = [1\,0\ldots0]'$, $\omega_t \sim iidN(0, \sigma_\omega^2)$ and $cov(G\omega_t) = \sigma_\omega^2 GG'$. The observation equation is assumed to follow

$$z_{it} = H\Pi_t + v_{it} \tag{D.2}$$

where $H = [1\,0\cdots0]$, $v_{it} \sim iidN(0, R)$ and we assume $E(\omega_t v_{it}) = 0$.

Denote the forecast as predicted state estimate $\Pi_{t|t-1}(i) = F\Pi_{t-1|t-1}$ and the forecast covariance as $\Sigma^\Pi_{t|t-1} = F\Sigma^\Pi_{t-1|t-1}F' + \sigma_\omega^2 GG'$. Denoting the Kalman gain as $K = \Sigma^\Pi_{t|t-1}H'S^{-1}$ where innovation covariance $S = R + H\Sigma^\Pi_{t|t-1}H'$. With the Kalman gain, the forecast covariance is shown as $\Sigma^\Pi_{t|t-1} = F\Sigma^\Pi_{t|t-1} - KH\Sigma^\Pi_{t|t-1}F' + \sigma_\omega^2 GG'$.

The updating equation for true inflation is as follows,

$$\Pi_{t|t}(i) = \Pi_{t|t-1}(i) + K(z_{it} - z_{t|t-1}(i)) = \Pi_{t|t-1}(i) + K(H\Pi_t + v_{it} - H\Pi_{t|t-1}(i)). \tag{D.3}$$

After taking averages across agents, we note that

$$\pi_{t+h} - \pi_{t+h|t} = H(\Pi_{t+h} - \Pi_{t+h|t}) = HF^h(\Pi_t - \Pi_{t|t-1}(i)) + RE\ error \tag{D.4}$$

After rearranging equations (D.3) and (D.4), see the online appendix of Coibion and Gorodnichenko (2015) in detail, we obtain the h-period ahead inflation forecast error as follows.

$$\begin{aligned} \pi_{t+h} - \pi_{t+h|t} = {} & \beta_{11}(\pi_{t+h|t} - \pi_{t+h|t-1}) + \beta_{12}(\pi_{t+h-1|t} - \pi_{t+h-1|t-1}) \\ & \cdots + \beta_{1n}(\pi_{t+h-(n-1)|t} - \pi_{t+h-(n-1)|t-1}) + RE\ error \end{aligned} \tag{D.5}$$

The above equation is the same specification as equation (B.5) in Coibion and Gorodnichenko (2015). Now, we assume h=1, n=1, and four quarters in this paper as one period. The notation for the averaged expectation over agents is replaced by $\bar{E}$ in the above equation. Then, equation (D.5) is reduced as follows.

$$\pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \beta_{11}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) + \beta_{12}(\bar{E}_t[\pi_t] - \bar{E}_{t-4}[\pi_t]) + RE\ error \quad \text{(D.6)}$$

The above equation indicates that the forecast error is related to two forecast revision terms when agents have rational expectations under the noisy information environment, and the true inflation follows an AR(1) process. Now, we add and subtract $\beta_{12}\bar{E}_{t-4}[\pi_{t+4}]$ to equation (D.6).

$$\begin{aligned} \pi_{t+4} - \bar{E}_t[\pi_{t+4}] = \beta_{11}(\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-4}[\pi_{t+4}]) \\ + \beta_{12}(\bar{E}_{t-4}[\pi_{t+4}] - \bar{E}_{t-4}[\pi_t]) \\ + \beta_{12}(\pi_t - \bar{E}_{t-4}[\pi_{t+4}]) + RE\ error \end{aligned} \quad \text{(D.7)}$$

The first two terms on the right-hand side are the same as equation (7). Adding a constant and relabeling coefficient parameters yields equation (16)