Inflation expectations in Japan: Forecast revision and forecast trend

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Abstract

Inflation expectations are 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. The forecast trend term affects the forecast errors. In addition, we find evidence that consumption tax hikes affected the forecast errors, partly due to the uncertainty of actual implementation in the future. 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 rational expectations without information rigidity, but with the possibility of deviating from rational expectations.

Keywords: Forecast trend, Inflation expectations, Information rigidity, Japanese inflation.

JEL Classification codes: C53, D83, D84, E13, E31, E37

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1 Introduction

Japan had experienced low or zero inflation for two decades despite the unprecedented monetary policy easing. Only after COVID-19 did we observe inflation in Japan rise gradually. The BOJ shifted from a zero-interest-rate policy in March 2024. In the July 2025 Monetary Policy Meeting, the BOJ decided not to increase the policy rate further. Yet, with a higher expected inflation rate, the BOJ is likely to raise the policy rate.

In this study, we investigate how expected inflation in Japan is formed. Is inflation expectation rational, and is information fully updated at all times? To provide an answer to 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 for the period between 1991:Q4 and 2025:Q1.

This study is not the first study to apply the Coibion and Gorodnichenko (2015) to the 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 Coibion and Gorodnichenko (2015) forecast error model to the Japanese inflation expectation dataset. In doing so, a crucial difference in the data framework exists between the US SPF dataset and the Japanese dataset. To circumvent the dataset problem, we devise an alternative framework to test the rational expectation with information rigidity.

In particular, in this study, we propose an alternative forecast revision term in the forecast error regression model of Coibion and Gorodnichenko (2015). The alternative forecast revision term differs from the original forecast revision term in the deviation between the target periods for forecasts. The modified forecast revision term allows the forecasted periods to differ between the past forecast and the current forecast. This seemingly contradictory modified forecast revision is shown through decomposition to be equal to the sum of the original forecast revision and the forecast trend.

We find that the coefficients of the forecast revision, whether in the original form or the modified form, are positive and statistically significant regardless of the specifications. From this evidence, we conclude that people in Japan, in general, form rational expectations about inflation. However, due to the information rigidity, the forecasts are under-reactive to new information. Moreover, we have evidence that this underreaction can only be observed in the recent episode of inflationary economics after 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), show that the forecast revision coefficients become positive and statistically significant only after 2022.

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. However, the null hypothesis of the rational expectation model with information rigidity is that no consumption tax hike dummies affect forecast errors. Backed by the limited 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 dummies statistically significant.

The rest of the paper is structured as follows. The next section reviews the inflation expectation model when information is rigid. Section 3 describes the dataset for the Japanese expectations and discusses the crucial differences in 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, 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 will lead to the following relationship between the forecast errors and the 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 that balances the current information and the past forecasts, with the weight, G, placed on new information. Averaging across agents will lead to the following relationship between the forecast errors and the 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, Coibion and Gorodnichenko (2015) proposes the following simple empirical model.

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

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 are under-reacting to new information, i.e., making a too low forecast when updating with an upward revision. In addition, the model imposes that no other control variables should have any explanatory power because the model does not include any other variables. If some variables are found to be statistically significant, it implies that agents are not following rational expectations.

The estimated coefficients of 'forecast revision' are always positive and statistically significant, 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 for inflation forecasts showed that people respond less to information, i.e., the realized inflation is higher than the expectation when the expectation is revised. After the seminal work of Coibion and Gorodnichenko (2015), many studies reported that at the individual levels, households, firm managers, or professional forecasters, people overreact. 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

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.

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 consists of two separate responses; one is qualitative and the other is quantitative. The firm measure is Tankan by the BOJ. Three measures for experts are 'Consensus Forecasts', 'QUICK', and inflation swap rates.

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 available. The forecast horizons of the CIE consist of annual incremental forecasts, extending from one to ten years ahead. 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 SPE and the CIE index

The forecast frequencies of the SPE and the CIE are both quarterly; however, the forecast horizons are crucially different between the two datasets. The CIE does not have any short-horizon forecasts, whereas the SPE has one, two, three, and four months ahead forecasts every quarter. The forecast horizons of the CIE are one through 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 perfectly follow the empirical model of Coibion and Gorodnichenko (2015), with forecast revision within one year instead of 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

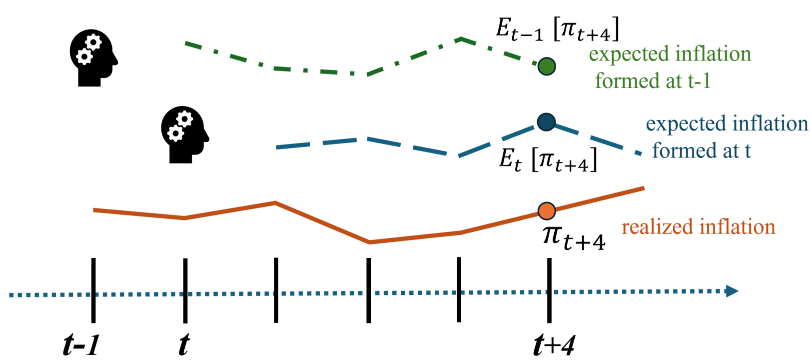

Figure 1: The original Coibion and Gorodnichenko model

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 1 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) only differs from equation (4) in the previous forecast period being 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. It should be noted that even in the original CG model in equation (3), the two forecast horizons in the forecast revision term are different. 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.

Then, we propose to further modify an empirical model for the CIE dataset. A proposed model adjusts forecast horizons to be at the same length in the forecast revision term as shown in equation (6). However, the forecast revision term in this modified model (6) is fundamentally different from the original model (3) or 4. The forecasted inflations on the right-hand side have different targets, i.e., π_{t+4} and π_t, as shown visually in figure 2.

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

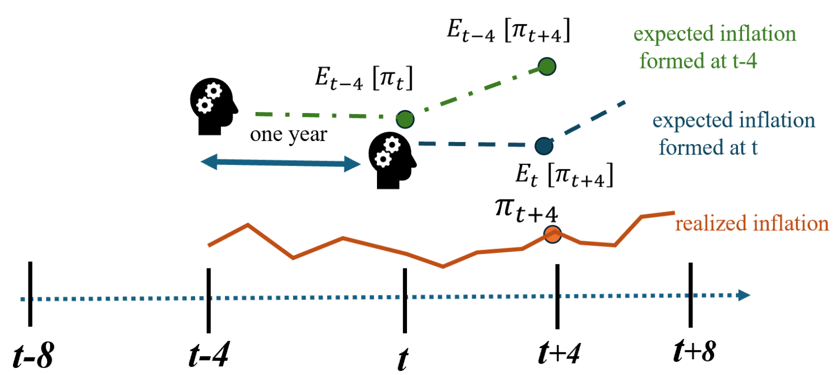

Figure 2: The modified Coibion and Gorodnichenko model

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 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 in an eight-quarter ahead forecast minus a four-quarter ahead forecast, both forecasted at the same quarter,$t-4$. We coined the latter term as 'forecast trend'.

Figure 3 shows the time series of forecast errors, forecast revision, and forecast trend consistent with the decomposition equation (7).[1]

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 model with information rigidity and only includes 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 ot 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 the coefficients of forecast revision and forecast trend to be equal. To capture this argument, the second and last equations specify different coefficients, K_1 and K_2, for forecast revision and forecast

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

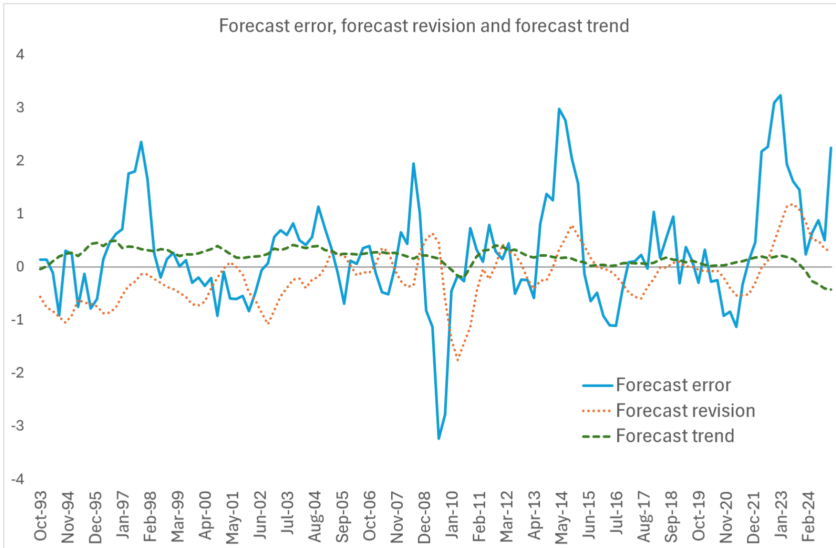

Figure 3: Forecast errors, forecast revisions, and forecast trend

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.

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.[2] 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 alternative interpretations of the decomposed model. One interpretation is that we test the hypothesis of the original CG forecast revision with control variables. The null hypothesis is $K_2 = 0$. The other interpretation is that we test the hypothesis of the modified forecast revision model introduced in this study. The null hypothesis is that $K_1 = K_2$ and $K_2 \neq 0$.

4.3 Modified forecast revisions with different revision spans

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

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

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

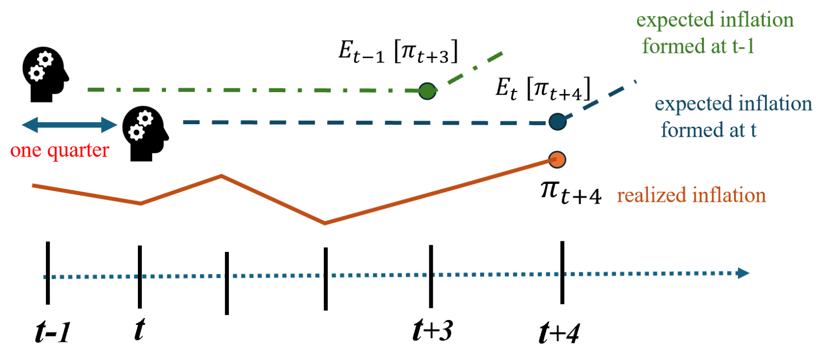

Figure 4: The alternative modified Coibion and Gorodnichenko model

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 one quarter difference in the targeted quarters for the forecast. 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 similarly 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).

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

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

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

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

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

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

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) rejects the null hypothesis of the FIRE that the forecast error is independent of the forecast revision. In column (ii), the coefficient of the modified model is similar to that of the CG model.

Column (iii) shows that the forecast trend is not statistically significant, whereas the 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 cannot reject the equality of K_1 and K_2.

5.2 The effect of consumption tax hikes

On examining Japanese inflation, we should consider the episodes of the 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,

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 Q4			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 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 the consumption hike dummies that take the value of one in the rising quarter and for the consecutive three quarters. We need to note that inflation is defined as a percentage increase on a year-on-year basis. However, 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. Therefore, the null hypothesis of the rational expectation model with information rigidity is that no consumption dummies affect forecast errors.

From columns (iv) to (vi) in Table 1, two earlier consumption tax hike dummies are statistically significant at the one percent level. These results support that Japanese inflation follows non-rational expectations. Interestingly, the fitness of regression has more than doubled. The results are qualitatively the same as in columns (i) through (iii).

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

One crucial issue remains: whether the consumption tax hikes were correctly anticipated by Japanese consumers/voters in the quarter one year earlier. It is important to be reminded that raising the consumption tax rate had been a political issue between the majority political party and the competing political parties. The planned increase of the consumption tax was once 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 measure how sure consumers anticipated a tax hike at the time of making a forecast? In this study, we suggest using the media coverage to represent the degree of tax hike anticipation, following the works of Baker et al. (2016) and Yoshida (2025).

In Figure 5, the number of newspaper articles containing the relevant phrase of 'consumption tax increase' appearing in the corresponding months is depicted.[3] On the three occasions of the consumption tax hike in April 1997, April 2014, and October 2019, the monthly counts of articles up to the previous 18 months are shown.

[3]The detailed description of the procedure is discussed in Appendix A.

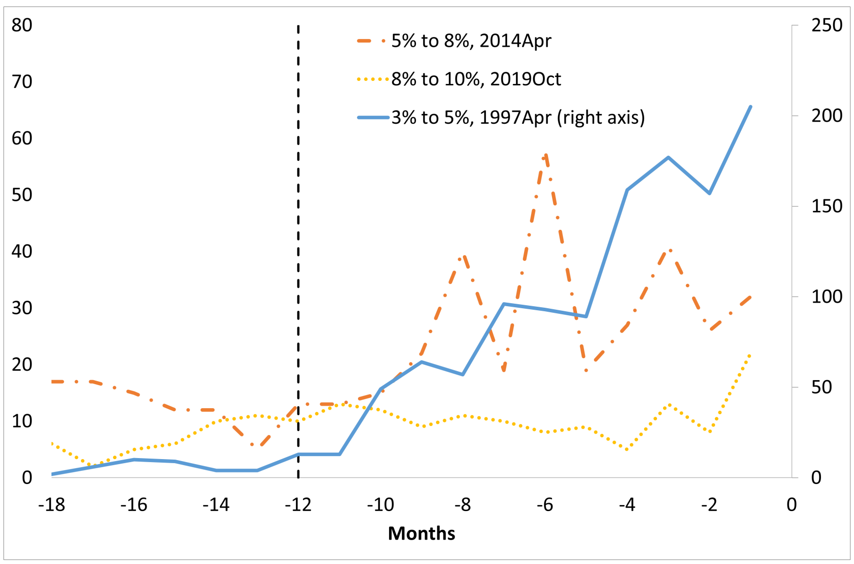

Figure 5: Consumption tax hike anticipations

Note: The horizontal axis indicates the months prior to 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 prior to 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 consumption tax hike in 1997 by consumers one year earlier is far from certain. The ruling coalition of Japan was still contemplating raising the consumption tax to the planned level of five percent on April 1, 1997.[4]

The corresponding number of articles was much 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.[5]

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. The government was moving ahead with adjusting other related frameworks, such as the medical payment system and the exception clauses, to maintain the current eight percent rate.

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.

5.3 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 expectation with information rigidity. Interestingly, the degree of impact becomes larger as the revision span and the forecast gap become shorter. 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.

In the decomposition specifications in columns (iv) through (vi), 'Forecast trends' are statistically significant regardless of the trend span. Again, these results show that the inflation forecasts in Japan are not rational. More importantly, the estimated coefficients of the CG forecast revisions and the modified forecast revisions are quite similar. The estimated coefficients for the US consumption forecasts in Coibion and Gorodnichenko

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

[5] Takahashi and Takayama (2025) investigates the effects of the 2014 tax hike episode on economic activities. They found that future tax hike information only affects future inflation and consumption but not other economic activities.

(2015) range from 1.06 to 1.20. This is close to the estimated coefficients of 1.13 in column (v).

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 over the original CG model.

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	0.708***					
$\bar{E}_t[\pi_{t+4}] - \bar{E}_{t-3}[\pi_{t+1}]$	(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 Q3				1.004**		
				(0.441)		
Forecast trend Q2					1.263***	
					(0.465)	
Forecast trend Q1						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).

5.4 During the zero inflation period

Both inflation and inflation expectations in Japan have been peculiar over the last three decades; they only became positive 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. 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 with the subsample of only 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. 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. Forecast errors and forecast revisions were statistically independent.

The forecast trends and the consumption tax hike dummies are statistically significant. This evidence warrants us to reconsider the information rigidity model of Coibion and Gorodnichenko (2015) and further investigate whether tax hikes were anticipated in quarter one year earlier. Note that the null of $K_1 = K_2$ is rejected in columns (v) and (vi).

Table 3: Under the zero inflation regime

	(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) $FR(4,4,4)$
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 Q4					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).

5.4.1 Time-varying parameter estimates

So far, we have split the sample into the zero-inflation or zero-interest rate period and the post-COVID period, the latter of which is associated with a non-zero inflation period. Though this particular structural break point may be agreeable to the majority of 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.[6]

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.

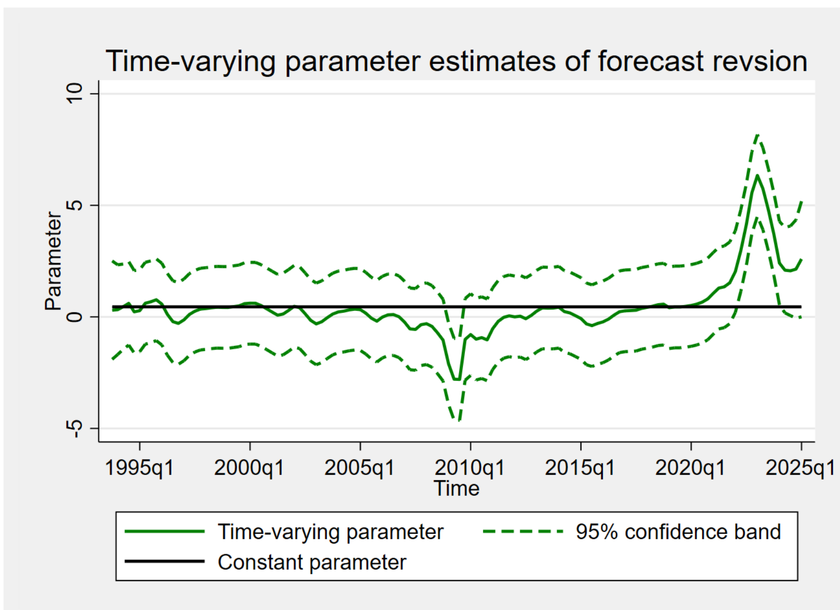

Figure 6: Forecast revisions, FR(4,4,0)

Figure 6 shows the time-varying parameter of forecast revision on forecast error in the model in 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. It is clear and 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-varying parameter estimates for

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

other modified forecast revision models are shown in Appendix C. They all demonstrate qualitatively the same results.

6 Conclusion

In this study, we estimated inflation forecast error in Japan via the Coibion and Gorodnichenko (2015) framework for the period between 1991:Q1 and 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, only with an additional term of '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 the coefficient of forecast trend, $K_2 > 0$, indicating that the forecast is too low when the forecast trend is moving upward. This contradicts the rational expectation hypothesis. We conclude from this evidence that people in Japan forecast inflation under information rigidity and form their expectations that 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. This study demonstrates that the monetary authority should also consider the forecast trend. The forecast will be too low when the forecast trend is moving upward.

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Appendix A. The expectations of consumption tax hikes

The Nikkei Telecom allows keyword search in eight different newspapers, three news bulletins, and press releases by 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, the newspapers are released twice per day. The keywords in Japanese relating to 'consumption tax increase' are used for an entire month, and we counted the number of articles containing either of these keywords.[7] The records of article counts are recorded as a screenshot by one researcher, and another researcher checked the screenshots for the integrity of using the correct keywords, the correct date span, and whether the correct number of articles is recorded.

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

Appendix B. Forecast revisions and modified forecast revisions

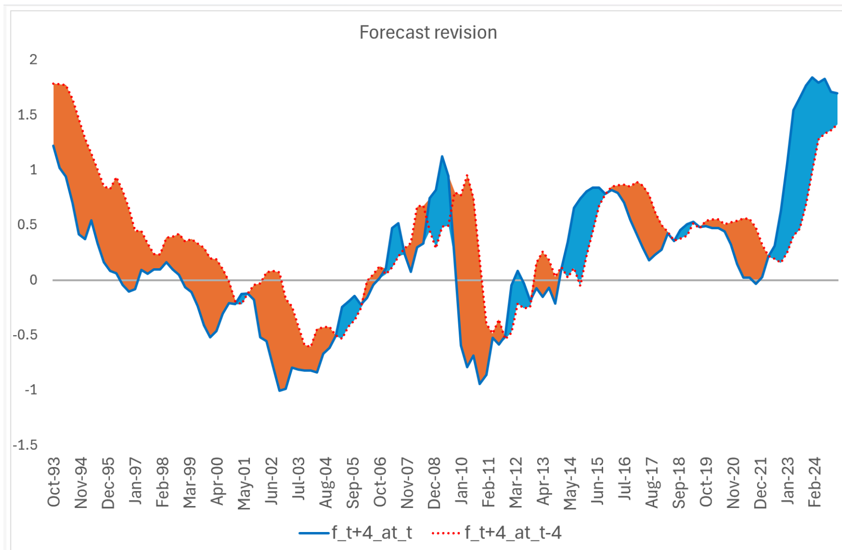

Figure B.1: Forecast revisions, FR(4,4,0)

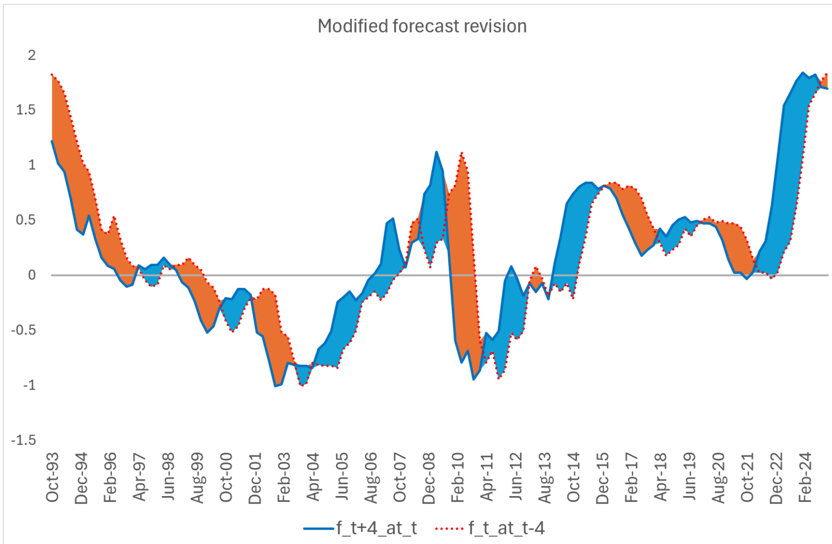

Figure B.2: Modified forecast revisions, FR(4,4,4)

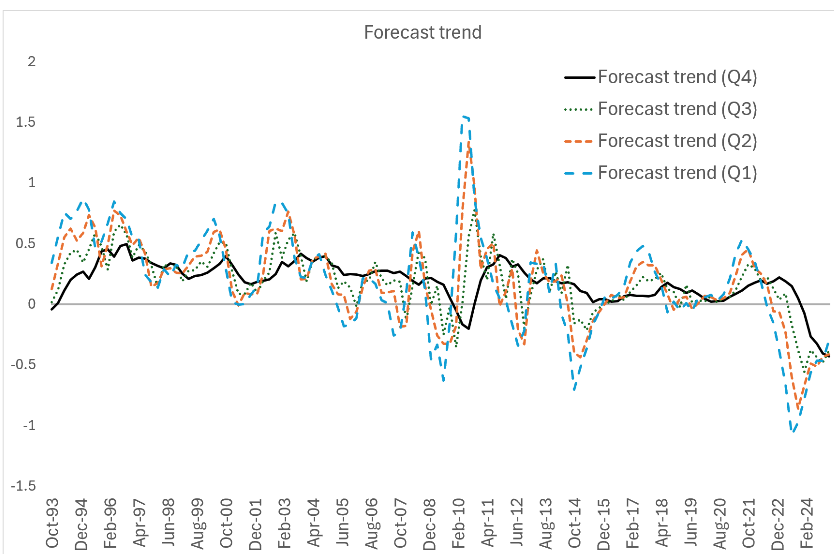

Figure B.3: Alternative forecast trends

Appendix C. Time-varying parameter estimates

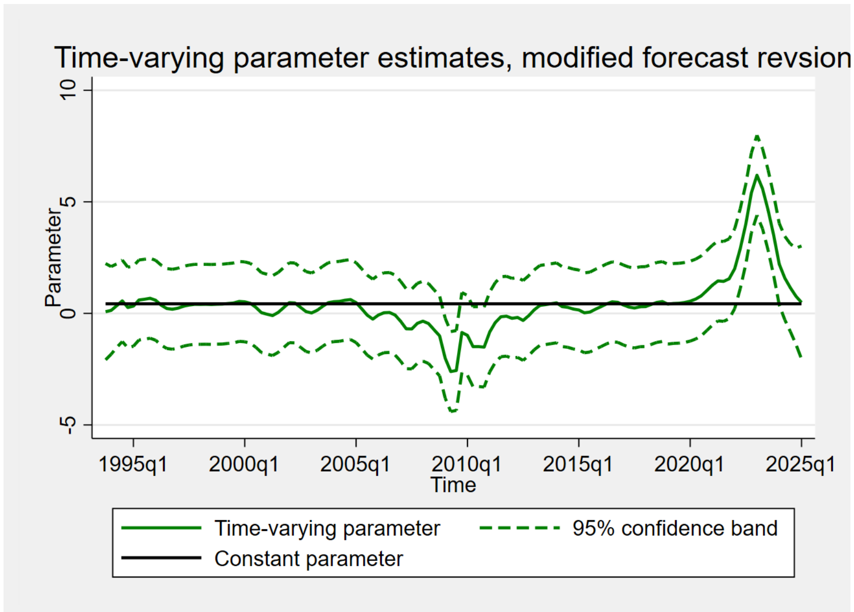

Figure C.1: Forecast revisions, FR(4,4,4)

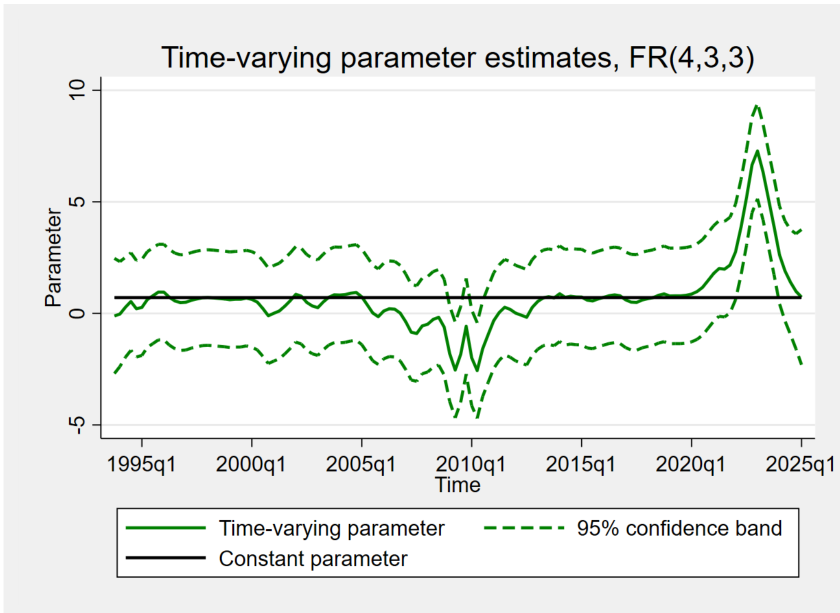

Figure C.2: Forecast revisions, FR(4,3,3)

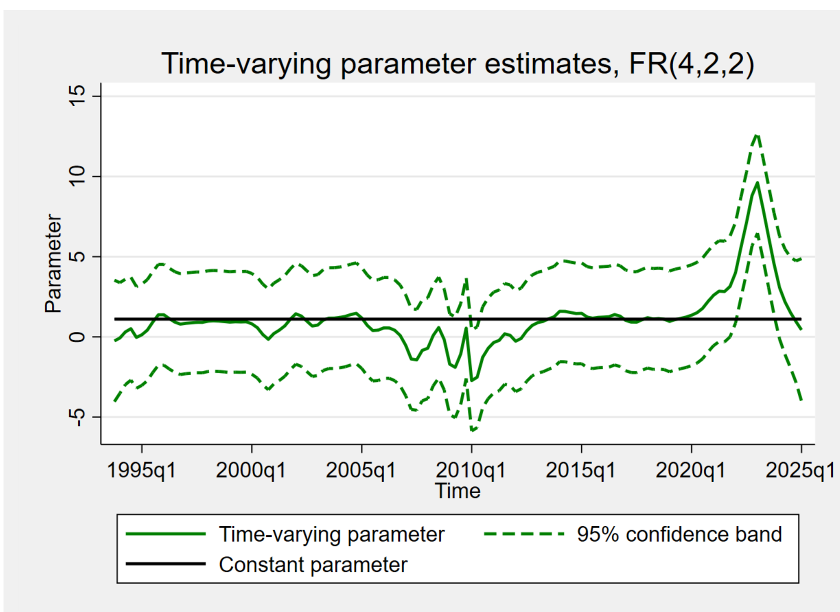

Figure C.3: Forecast revisions, FR(4,2,2)

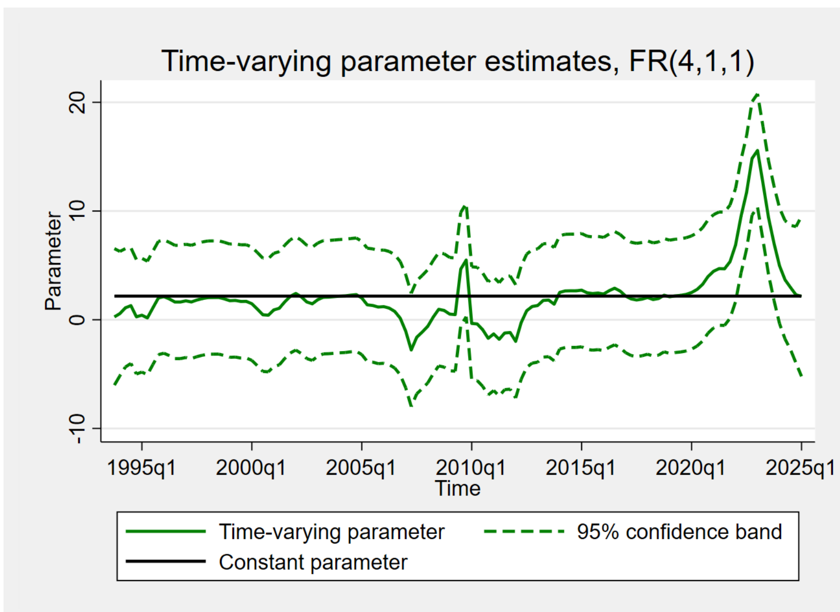

Figure C.4: Forecast revisions, FR(4,1,1)