



The Japanese Taylor rule estimated using censored quantile regressions

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Abstract This paper conducts quantile regressions and obtains detailed estimates of monetary policy rules in Japan using a sample that includes recent periods of zero interest rates. Taking into account censoring and endogeneity, we compute censored quantile instrumental variable estimators and compare them with estimates from uncensored quantile regressions. The estimation results indicate that not accounting for censoring of interest rates tends to result in downwardly biased estimates. Moreover, our censored quantile regressions lead to relatively flat coefficients of inflation and insignificant coefficients of the output gap over the conditional interest rate distribution, suggesting that monetary policy in Japan may be well described by a linear rule.

Keywords Quantile regression · Censoring · Japan · Taylor rule · Zero lower bound

JEL Classification C21 · C26 · E52 · E58

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1 Introduction

Japan's monetary policy was used to be considered peculiar with the nominal interest rate hitting the zero lower bound (ZLB). Japan's policy is no longer unusual. Many countries, including the USA, have adopted such a policy since the recent financial crisis of 2008. Motivated by Japan's relatively long periods of zero interest, we obtain detailed estimates of monetary policy in Japan, using quantile regressions and taking into account possible bias due to the ZLB. In addition, we find that estimates tend to be biased downward if censoring of interest rates is not considered.

We consider an estimable monetary policy reaction function in which short-term interest rates systematically respond to the inflation rate and the output gap, that is, a version of the [Taylor \(1993\)](#) rule. To the best of our knowledge, the present paper is the first to conduct censored quantile regressions of the Japanese Taylor rule.

In our view, observations from the recent zero interest period provide useful information for estimating the reaction function. Although the nominal interest rate in Japan has remained low compared with earlier periods since the first implementation of the zero interest policy in February 1999, the interest rate has not necessarily remained at the ZLB at all times. In particular, the interest rate was approximately 0.5% between early 2007 and late 2008.

One of the most appealing features of the quantile regression is its ability to estimate quantile-specific effects that describe the impact of covariates along the tails of the interest rate distribution. On the other hand, conventional estimation methods¹ provide estimates of the coefficients of the reaction functions at the conditional mean and, accordingly, lack information about the policy reactions over the conditional distribution of nominal interest rates.

We take the interest rate data as censored because of the ZLB and apply censored quantile regressions. In addition, we address endogeneity in our estimation. Specifically, we compute the censored quantile instrumental variable (CQIV) estimators developed by [Chernozhukov et al. \(2015\)](#), who build on the control function approach addressing endogenous regressors and the [Powell \(1986\)](#) censored quantile regressions. Using this CQIV methodology with the control function approach, we systematically estimate how the Bank of Japan's reaction varies at different quantiles of the conditional interest rate distribution. In addition, this methodology can carry out uncensored and censored quantile regressions, controlling for endogeneity in a unified framework. Hereafter, we refer to an estimator from the uncensored version of the CQIV methodology as an uncensored quantile instrumental variable (UQIV) estimator.² To obtain detailed estimates, we apply the CQIV methods as well as the following uncensored quantile regressions: a preliminary quantile regression that ignores endogeneity, a two-stage quantile regression (2SQR), and the UQIV methods.

¹ Among studies using such conventional methods, [Clarida et al. \(1998\)](#) estimate the monetary policy reaction functions for six countries—France, Germany, Italy, Japan, the UK, and the USA—and [Bernanke and Gertler \(1999\)](#) study those of the USA and Japan. See also [Kuttner and Posen \(2004\)](#) for estimations of the Japanese policy reaction function.

² The UQIV methods are implemented by setting an arbitrarily small left-censoring point under the CQIV methodology.

We obtain the following results. Our estimates of the inflation coefficient from CQIV are relatively flat over the conditional quantiles. On the other hand, those from uncensored quantile regressions tend to be relatively low at lower quantiles and tend to increase and approach CQIV estimates moving toward the upper tail of the conditional interest rate distribution. In addition, the coefficient of the output gap from 2SQR is negative and significant, whereas that from CQIV is mostly insignificant. Accordingly, without taking censoring into account, estimates tend to be biased downward. Moreover, given such flat estimates of the inflation coefficient together with insignificant estimates of the output-gap coefficient, Japan's monetary policy may be well described in a linear policy rule. That is, deviations from monetary policy responses from the conditional mean are not systematically related to inflation or to the output gap.

We merge two strands of the literature by estimating monetary policy rules using censored quantile regressions. First, several works estimate the Taylor rule using quantile regressions, such as [Chevapatrakul et al. \(2009\)](#), who study the USA and Japan, and [Wolters \(2012\)](#), who focuses on the USA. Note that [Chevapatrakul et al. \(2009\)](#) and [Wolters \(2012\)](#) do not need to account for censoring since these studies conduct estimations using a sample that ends prior to the period of zero interest rates.³ To address endogeneity, [Chevapatrakul et al. \(2009\)](#) apply 2SQR, while [Wolters \(2012\)](#) applies an instrumental variable quantile regression (IVQR).⁴ Building upon these works that adopt uncensored quantile regressions to estimate monetary policy rules, we take into account the ZLB in our estimation, since we include the periods of zero interest rates in our sample.

Second, there are works that adopt Tobit-type censored regressions to estimate monetary policy rules, dealing with the ZLB. Such studies include [Kuttner and Posen \(2004\)](#), [Kim and Mizen \(2010\)](#), and [Kiesel and Wolters \(2014\)](#). We complement these works by obtaining quantile-specific estimates.

This paper proceeds as follows. Section 2 discusses the data. Section 3 describes our estimation methodology, and Sect. 4 presents the empirical results. Finally, Sect. 5 sets forth our conclusions.

2 Data

The policy instrument is the overnight call rate, and we use monthly data available from the Bank of Japan's Web site. Our sample spans from July 1985 through September 2014.⁵ The monthly Japanese data for the consumer price index (CPI) for all items less fresh food are collected from the Ministry of Internal Affairs and Communications. Using these CPI data, we construct 1-year-ahead ex post core inflation⁶ rates and,

³ Although [Chevapatrakul et al. \(2009\)](#) present data up to 2005 for Japan, their estimation uses a sample that ends before February 1999, to study the monetary policy when the ZLB is approached.

⁴ The 2SQR methodology is based on the fitted value approach, which extends [Amemiya \(1982\)](#), whereas the IVQR method builds on the instrumental variable approach suggested by [Chernozhukov and Hansen \(2005, 2006, 2008\)](#).

⁵ The call rate data are available from July 1985 at the Bank of Japan's web site.

⁶ In Japan, inflation of CPI less fresh food is called core inflation, and it is the main inflation series considered by the Bank of Japan.

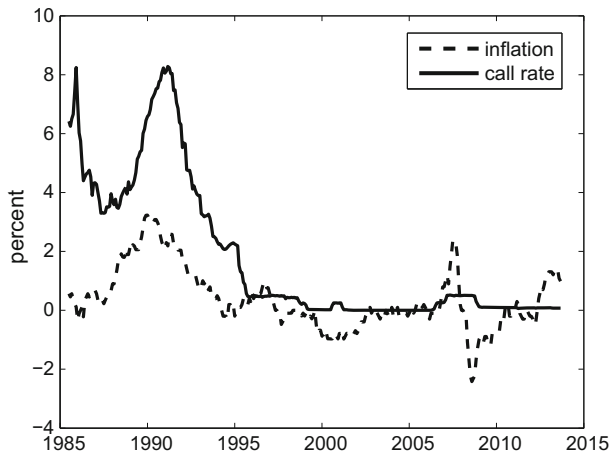


Fig. 1 Core inflation and nominal interest rates

accordingly, our core inflation data span from July 1985 through September 2013.⁷ We use industrial production indices to measure output, following [Clarida et al. \(1998\)](#), [Bernanke and Gertler \(1999\)](#), and [Chevapatrakul et al. \(2009\)](#). The data are available at a monthly frequency and are collected from the Ministry of Economy, Trade and Industry. Output-gap data are constructed using the Hodrick–Prescott (HP) filter with smoothing parameter 14400, as in [Chevapatrakul et al. \(2009\)](#).

Figure 1 plots the uncollateralized overnight call rate and the core inflation rate (in percent) from July 1985 through September 2013. In the sample period, the call rate peaked in March 1991 (8.28%). Before then, cycles are evident in the call rate. Starting at 6.41% in July 1985, the call rate exceeded 8% in December 1985 (8.24%) and decreased to 3.3% in May 1987. The call rate continued to rise until it peaked in March 1991. Since then, it has shown a downward trend and approached zero as the Japanese economy remained in recession.

Japan initiated its zero interest rate policy in February 1999. In August 2000, the zero interest rate policy was terminated and the call rate jumped to 0.16% from 0.02% in the preceding month. However, the call rate hit the ZLB again after implementation of quantitative easing in March 2001. In July 2006, when the zero interest policy was terminated, the nominal interest rate began to rise and stayed above the ZLB. The call rate was 0.16% in July 2006 and approximately 0.5% in most months of 2007 and 2008. In December 2008, the Bank of Japan changed its target for the call rate to 0.1%. In October 2010, this target was changed to “0 to 0.1%.”⁸ In April 2013, the Bank of Japan introduced “quantitative and qualitative monetary easing” and announced that it would “conduct money market operations so that the monetary base will increase at an annual pace of about 60–70 trillion yen.”

⁷ Following the Bank of Japan, the effects of consumption taxes have been controlled for.

⁸ For example, in December 2012, the uncollateralized call rate was 0.082%.

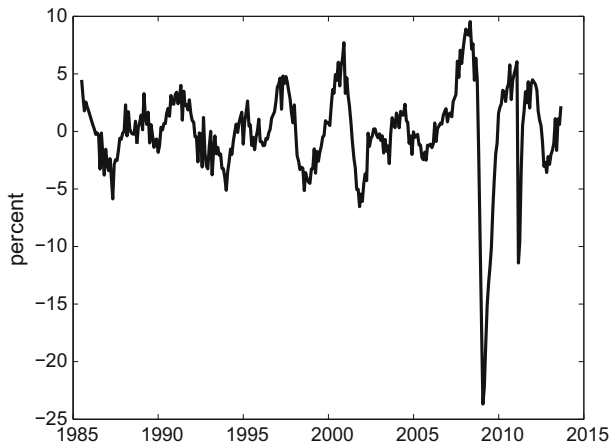


Fig. 2 Output gap

Japan has been experiencing deflation or very low inflation since the mid-1990s. Although the core inflation rate rose substantially above zero in late 2006 and in 2007,⁹ the Japanese core inflation rate stayed below or close to zero after 2008. Due to initiation of “quantitative and qualitative monetary easing” in April 2013, the ex post core inflation series began to rise from April 2012.

The output gap is plotted in Fig. 2, as percent deviation of industrial production indices from their HP trend. Compared with the output gap, interest rate moves gradually and suggests an interest smoothing term in the estimation equation. However, as we discuss below, we abstract from interest smoothing in our estimation due to technical problems.

3 Estimation methodology

Following [Chevapatrakul et al. \(2009\)](#), we consider the following Taylor rule:

$$i_t = \omega + \alpha\pi_{t+n} + \kappa y_t + \epsilon_t, \quad (1)$$

where i_t , π_{t+n} , and y_t are the nominal interest rate, n -period forward-looking core inflation rate, and output gap, respectively. We abstract from an interest rate smoothing term in the specification of the monetary policy rule to avoid convergence problems as discussed by [Chevapatrakul et al. \(2009\)](#). Besides, when we consider censored quantile regressions, the regressor (the interest rate smoothing term) would also be censored as well as the dependent variable; it remains unclear how to propose a valid algorithm to estimate quantile autoregression with endogeneity.¹⁰

⁹ The core inflation of late 2006 and 2007 is explained largely by energy prices. If the CPI for all items less food and energy is used to construct inflation rates (called core–core inflation in Japan), the resulting series shows low or negative inflation during that period.

¹⁰ See [Choi and Portnoy \(2016\)](#).

Conventional estimation methods [e.g., ordinary least squares (OLS), two-stage least squares (2SLS), and two-step efficient generalized method of moments (GMM)] estimate parameters (ω, α, κ) at the conditional mean of nominal interest rates. In addition to these conventional methods, we implement uncensored quantile regressions (a preliminary quantile regression without considering endogeneity, 2SQR, and UQIV). Moreover, taking the ZLB into account, we treat the nominal interest rate data as censored and compute the CQIV estimators as well.

Let τ be a real number in $(0, 1)$. Quantile regressions can estimate policy rules at any τ th conditional quantile of the nominal interest rate. For a given τ , we estimate the corresponding policy coefficients ω_τ, α_τ , and κ_τ . These parameters are estimated by

$$\min_{\omega_\tau, \alpha_\tau, \kappa_\tau} \sum_{t=1}^T \rho_\tau (i_t - \omega_\tau - \alpha_\tau \pi_{t+n} - \kappa_\tau y_t), \tag{2}$$

where $\rho_\tau(u) = (\tau - \mathbb{I}(u \leq 0)) u$. Note that $\mathbb{I}(u \leq 0)$ is an indicator function such that

$$\mathbb{I}(u \leq 0) = \begin{cases} 1 & \text{if } u \leq 0, \\ 0 & \text{if } u > 0. \end{cases}$$

As for 2SQR, we follow the methodology suggested by [Chevapatrakul et al. \(2009\)](#), in which (in the first stage) the forward-looking inflation rate and the output gap are regressed by OLS on appropriately chosen instruments and (in the second stage) the quantile regression is carried out using least-squares-fitted values from the first stage as the explanatory variable.

The aforementioned estimators do not take censoring into account. Given the ZLB, a version of the [Powell \(1986\)](#) censored quantile regression objective function,

$$\min_{\omega_\tau, \alpha_\tau, \kappa_\tau} \sum_{t=1}^T \rho_\tau (i_t - \max\{\omega_\tau - \alpha_\tau \pi_{t+n} - \kappa_\tau y_t, \underline{i}\}), \tag{3}$$

is considered, where \underline{i} is the left-censoring point and $\underline{i} = 0$ in our estimation. Avoiding strong parametric assumptions, the Powell-type procedure addresses censoring semi-parametrically through the equivariance property of quantile functions to monotone transformation. To further control for endogeneity, we utilize a control function approach that adds a variable to the non-separable quantile regression with nonadditive error terms such that conditioning on this variable, regressors and error terms become independent. This control variable (function) is unobservable and needs to be estimated in a nonadditive, nonparametric first stage. In light of the control function approach, controlling for both censoring and endogeneity, the CQIV estimator of [Chernozhukov et al. \(2015\)](#) is obtained by

$$\min_{\omega_\tau, \alpha_\tau, \kappa_\tau, \beta_\tau} \sum_{t=1}^T \mathcal{T}_\tau \rho_\tau (i_t - \max\{\omega_\tau - \alpha_\tau \pi_{t+n} - \kappa_\tau y_t - \hat{v}'_t \beta_\tau, \underline{i}\}), \tag{4}$$

where \mathcal{T} is an exogenous trimming indicator and \hat{v} is a vector of estimated control variables, given chosen instruments.¹¹ The uncensored case (i.e., UQIV) is covered by making \underline{i} arbitrarily small. A computationally attractive approximation to this estimator, the associated asymptotic inference, and the control variable estimation are demonstrated in great detail by Chernozhukov et al. (2015). We obtain confidence intervals (CIs) through a weighted bootstrap procedure that is also advocated by Chernozhukov et al. (2015), instead of carrying out a complex analytical inference on the parameter of interest.

4 Empirical results

We estimate the Japanese Taylor rule using the methods described in Sect. 3. To address the difference between the results from uncensored and censored quantile regressions, we describe the results from uncensored quantile regressions first, then report the results from censored quantile regressions.

4.1 Uncensored quantile regressions

We first run a preliminary quantile regression together with OLS. Implicitly, in this estimation, we assume no endogeneity issue, that is, the core inflation rate and the output gap are independent of the error term. However, as Chevapatrakul et al. (2009) claim, the inflation rate is forward-looking and likely to be correlated with the error term. In addition, considering the standard IS equation, where the output gap reacts to the interest rate, an endogeneity problem is likely to exist between the interest rate and the output gap as well. Later, we conduct 2SQR and UQIV estimations to address endogeneity.

The conditional mean estimate of α is 1.622 (0.090), where the number in parentheses indicates the standard error. The OLS estimate of the output-gap coefficient κ is -0.001 (0.023). These estimates imply that Japan tended to respond to core inflation and there is no significant response to the output gap. This tendency is consistent with previous studies, including Clarida et al. (1998), Bernanke and Gertler (1999), and Kuttner and Posen (2004).

Figure 3 illustrates the estimation results from the preliminary quantile regression. The left and right panels present the estimates of coefficients of core inflation and output gap (i.e., α and κ), respectively. Each panel plots the quantile regression estimates¹² (bold solid line) together with their 95 % CIs (shaded area).

The estimated core inflation coefficient tends to be small and often insignificant at lower quantiles of the interest rate distribution. This coefficient tends to increase moving toward the upper tail of the interest rate distribution. The estimated coefficient

¹¹ The control variables here are estimated through the quantile-regression-based procedure constructed by Chernozhukov et al. (2015), which is valid in a nonadditive quantile regression setup. The specific construction of the trimming indicator, \mathcal{T} , is found in their paper.

¹² In our quantile regressions (both uncensored and censored), we estimate the coefficients at the following quantiles of the conditional interest rate distribution: 0.05–0.95 with a grid of 0.05.

of the output gap is negative and/or insignificant over the conditional quantiles, except for the extreme upper tail, where this coefficient is positive and significant.

We now conduct 2SQR. In addition, 2SLS and the two-step efficient GMM are implemented. For the robustness analysis, we consider the following instruments. For inflation, we construct $\widehat{inflation}$, which is the fitted value of π_{t+n} regressing on a constant, $\pi_{t-1}, \dots, \pi_{t-3}, y_{t-1}, \dots, y_{t-3}$, and i_{t-1}, \dots, i_{t-3} , and $\widehat{inflation}$, which is the fitted value of π_{t+n} regressing on a constant, $\pi_{t-1}, \dots, \pi_{t-12}, y_{t-1}, \dots, y_{t-12}$, and i_{t-1}, \dots, i_{t-12} . Similarly, we define $\widehat{output\ gap}$ and $\widehat{outputgap}$ for the output gap. A GMM Hansen J test is used to validate overidentifying restrictions and to determine the lag order for $\widehat{inflation}$ and $\widehat{output\ gap}$. Construction of $\widehat{inflation}$ and $\widehat{output\ gap}$ is motivated by the works of Clarida et al. (2000), Orphanides (2001), and Wolters (2012), who use lagged variables up to four quarters as instruments.

Figure 4 presents the results from 2SQR with $\widehat{inflation}$ and $\widehat{output\ gap}$ used as the fitted variables obtained from the first-stage regression. As in Fig. 3, the left and right panels illustrate the estimates of α and κ , respectively. The solid line represents the 2SQR estimates, with the shaded area indicating the 95% CIs.

The coefficient of the core inflation rate is higher than in the case in which no endogeneity is assumed. The 2SLS and two-step efficient GMM estimates are, respectively, 2.680 (0.121) and 2.680 (0.321), and significant. The 2SQR estimates of α at low quantiles are substantially higher than those from the preliminary quantile regression presented in Fig. 3. Still, the 2SQR estimates of α tend to increase moving toward the upper tail.

On the other hand, the 2SLS and two-step efficient GMM coefficients of κ are, respectively, -0.051 (0.027) and -0.051 (0.044). The 2SQR coefficients are

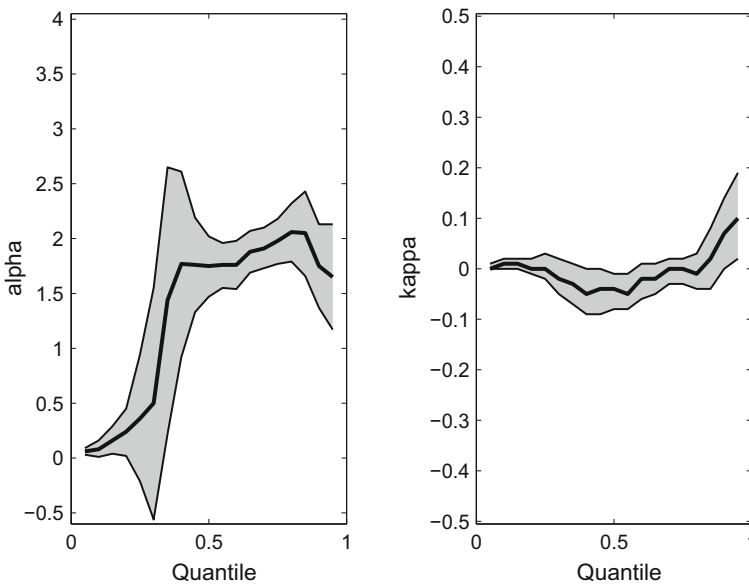


Fig. 3 Preliminary quantile regression (uncensored)

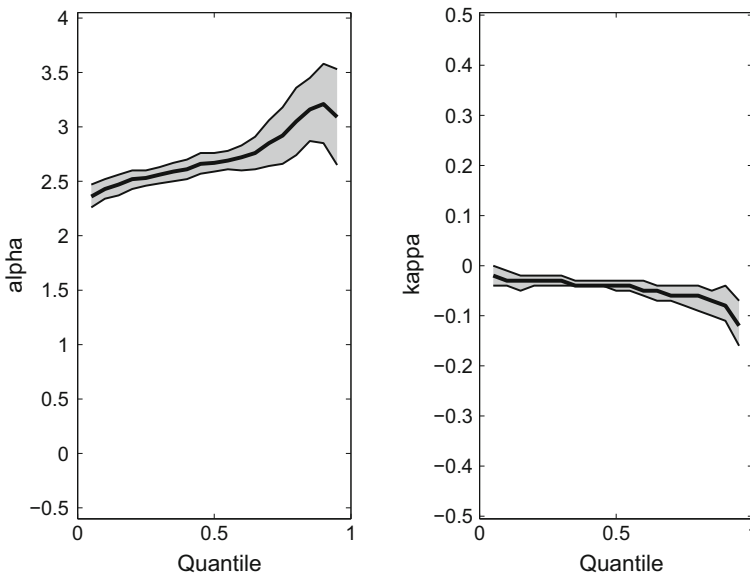


Fig. 4 2SQR (uncensored), $\widehat{\text{inflation}}$ and $\widehat{\text{output gap}}$

significantly below zero across the conditional quantiles. These estimates are counter-intuitive, since the implied policy rule lowers the interest rate when the output gap is positive. These estimates are likely to be biased. In fact, according to [Chevapatrakul et al. \(2009\)](#), who estimate monetary policy rules for Japan prior to zero interest rates (together with 90% CIs) using 2SQR, the coefficient of the output gap in the upper tail of the interest rate distribution is significantly positive, whereas the coefficient is otherwise insignificant.

A similar pattern is observed in [Fig. 5](#), which depicts the estimation result using $\widehat{\text{inflation}}$ and $\widehat{\text{output gap}}$. As in [Fig. 4](#), the estimates of α tend to increase moving toward the upper quantiles, and the estimates of κ are significantly negative except for lower quantiles. In addition, the 2SLS and two-step efficient GMM estimates of α are, respectively, 2.585 (0.114) and 2.585 (0.296). Those of κ are, respectively, -0.034 (0.026) and -0.048 (0.041).

To obtain estimates from another econometric methodology controlling for endogeneity (but not censoring), we also implement UQIV estimations. Estimations using 2SQR and UQIV provide a benchmark for comparison with the results from CQIV estimations reported later, such that we can examine the effect of censoring. The instrumental variables are the same as those used in our 2SQR estimations. Due to technical difficulties, the CQIV method allows only for one endogenous regressor in the algorithm of [Chernozhukov et al. \(2015\)](#).¹³ Therefore, we control for the endogenous core

¹³ The same limitation arises in identification and estimation of average structural effect in a triangular simultaneous equations model without additivity. The setups considered in the literature on control variables restrict heterogeneity: the residual in the structural first stage is one-dimensional, cf. [Imbens and Newey \(2009\)](#).

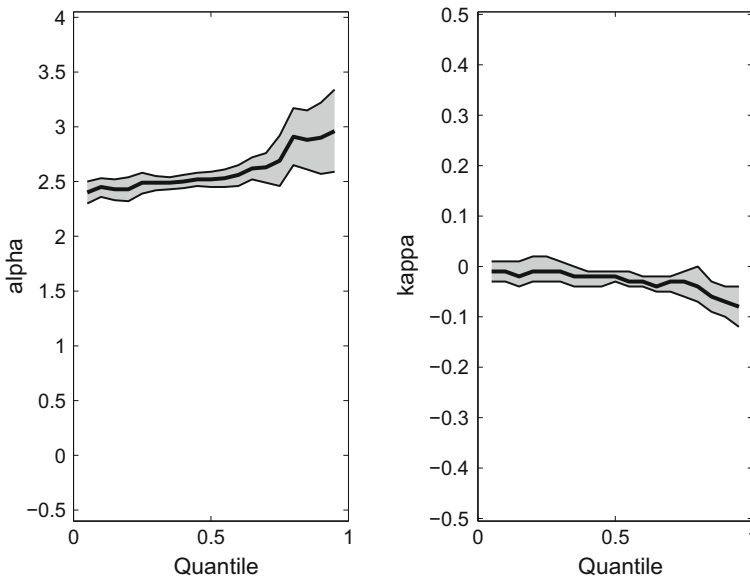


Fig. 5 2SQR (uncensored), $\widehat{inflation}$ and $\widehat{outputgap}$

inflation regressor via these authors' control function approach, whereas we replace output-gap variable with the fitted values.

Figures 6 and 7 illustrate the results from UQIV using, respectively, $(\widehat{inflation}, \widehat{output\ gap})$ and $(inflation, \widehat{output\ gap})$. Each figure reports the estimates of the coefficients of core inflation and the output gap (i.e., α and κ , respectively) together with their 95% CIs at different quantiles.¹⁴

The UQIV estimates of α in Fig. 6 present an increasing pattern moving toward the upper tail of the interest rate distribution. On the other hand, those in Fig. 7 are relatively flat. The estimates of κ are insignificant except for a few quantiles in Fig. 6. Note that even at the quantiles where the estimates are significant (and negative), the upper bounds of the CIs are very close to zero, compared with the 2SQR results.

4.2 Censored quantile regressions

We compute CQIV estimates using the projected instruments constructed in the previous section. Figures 8 and 9 depict the estimation results using $(\widehat{inflation}, \widehat{output\ gap})$ and $(inflation, \widehat{output\ gap})$, respectively. The coefficients of the output gap are insignificant except for the upper quantiles in Fig. 8, whereas they are insignificant across all quantiles in Fig. 9. In fact, accounting for censoring does not change the estimation results of κ very much compared with UQIV (Figs. 6, 7); however, it does so sub-

¹⁴ For UQIV and CQIV, the CIs for the parameter of interest are constructed through a weighted bootstrap with 400 repetitions.

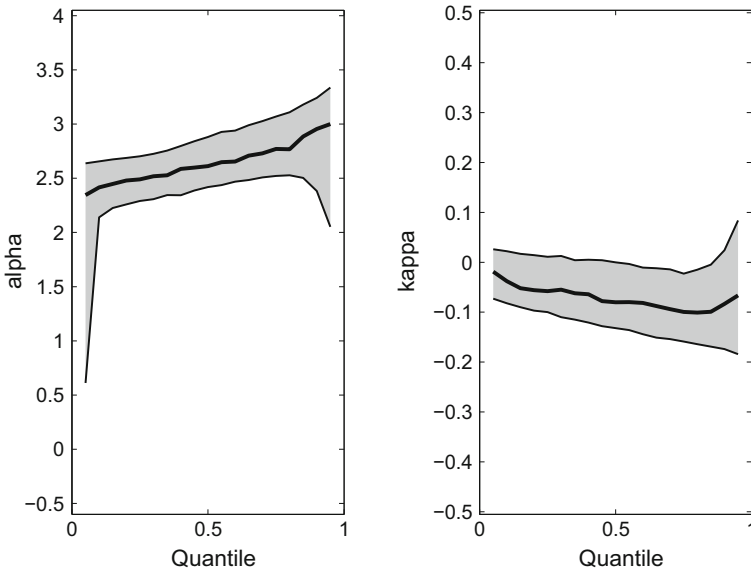


Fig. 6 UQIV: $\widetilde{\text{inflation}}$ and $\widetilde{\text{output gap}}$

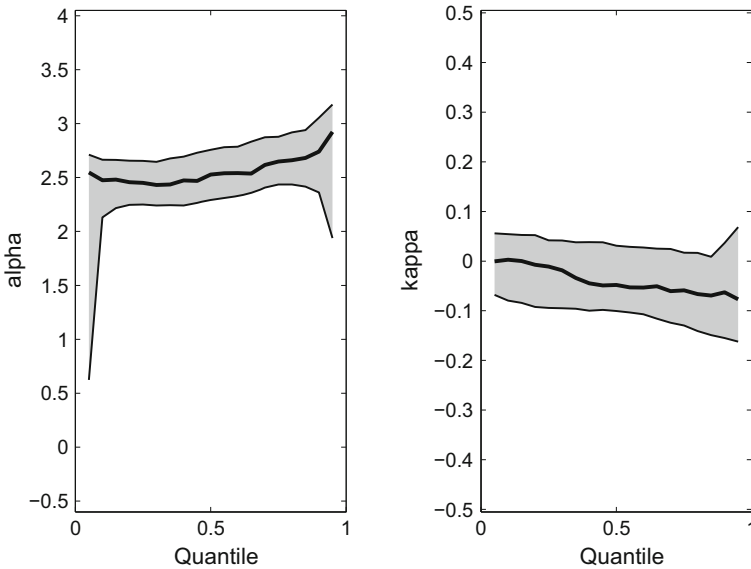


Fig. 7 UQIV: $\widetilde{\text{inflation}}$ and $\widetilde{\text{output gap}}$

stantially compared with the 2SQR results (Figs. 4, 5). The estimates of κ from CQIV (and UQIV) seem more reasonable than those from 2SQR, where the estimates are negative and significant at many quantiles.

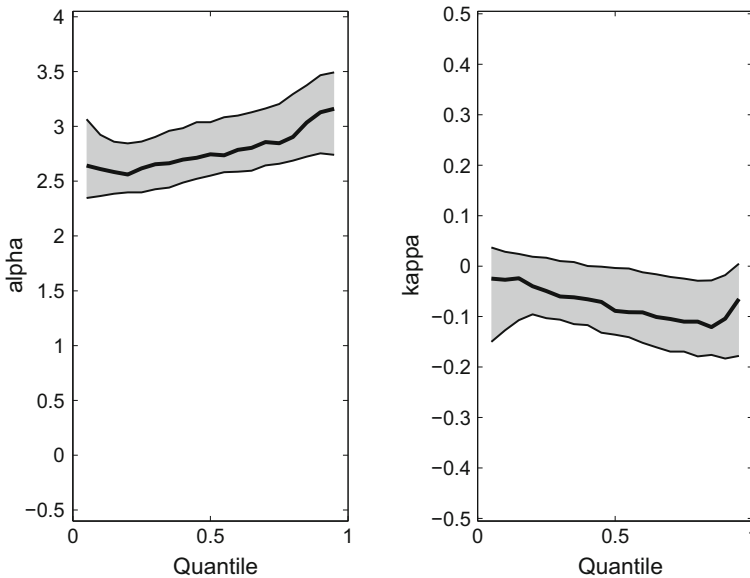


Fig. 8 CQIV: $\widehat{\text{inflation}}$ and $\widehat{\text{output gap}}$

As for the coefficient of inflation (α), the CQIV results illustrated in both Figs. 8 and 9 are relatively flat over the conditional interest rate distribution. On the other hand, those from uncensored quantile regressions (except for Fig. 7) present an increasing pattern and approach to the CQIV estimates moving toward the upper tail of the conditional distribution. Given this observation and significantly negative estimates of κ from 2SQR, estimates from uncensored quantile regressions tend to be biased downward.¹⁵

Results from CQIV, namely, relatively flat estimates of α and insignificant estimates of κ , suggest that deviations of monetary policy responses from those implied by estimates at the conditional mean are not systematically related to core inflation or the output gap, but are purely discretionary shocks (or possibly variables not captured by the Taylor rule). For example, if the interest rate was set lower than estimates at the conditional mean, this could not be explained by aggressive reaction to inflation or the output gap. In other words, monetary policy in Japan seems to be quite well described by a linear monetary policy rule. This is in contrast to the results for the USA of [Chevapatrakul et al. \(2009\)](#) and [Wolters \(2012\)](#), who find systematic patterns in the inflation and output-gap coefficients over the range of conditional quantiles.¹⁶

¹⁵ This result may appear to contrast with [Kim and Mizen \(2010\)](#), who point out that ignoring the ZLB on interest rates biases the estimates upward. Note that, however, their estimation results are derived from a Tobit-type model which relies on parametric distributional assumptions. [Chung and Goldberger \(1984\)](#) show that under considerably weak distributional assumptions (even without normality), whether an upward or downward bias of the estimate occurs depends on the covariance between latent and actually observed variables, the variance of the latent variable, and the probability of the event that the latent variable is larger than the censoring point.

¹⁶ We thank a referee for raising the points addressed in this paragraph.

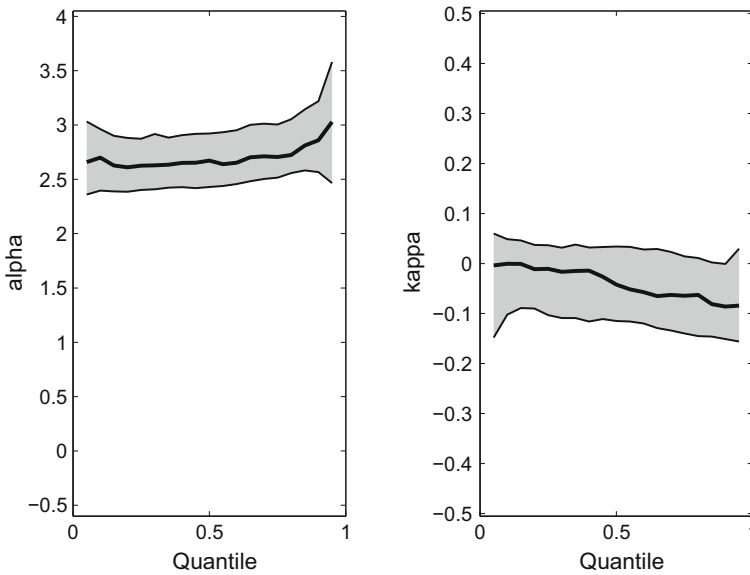


Fig. 9 CQIV: $\widetilde{\text{inflation}}$ and $\widetilde{\text{output gap}}$

Some studies point out, theoretically, that a central bank should adopt an aggressive monetary policy when the ZLB is approached (see, e.g., [Orphanides and Wieland 2000](#); [Jung et al. 2005](#); [Kato and Nishiyama 2005](#); [Adam and Billi 2006](#)). Given our estimates, such aggressiveness does not appear to be observed in actual policy rules. In this respect, our CQIV results appear to complement those of [Chevapatrakul et al. \(2009\)](#), who observe no tendency to increase responsiveness to inflation as the ZLB approaches using Japanese data prior to the initiation of the zero interest policy. Note that [Chevapatrakul et al. \(2009\)](#) control for endogeneity, but do not need to take censoring into account because of their sample period. In addition, [Chevapatrakul et al. \(2009\)](#) obtain relatively flat estimates of the inflation coefficient for a wide range of the conditional quantiles. We also find relatively flat estimates of the inflation coefficient using a sample that includes the zero interest period, by controlling for endogeneity and censoring.

However, it is worth noting that our CQIV results do not lead to a clear statement about the responsiveness of monetary policy at the ZLB. This is because the estimated parameters address monetary policy responses to the latent interest rate and not the actual one, conditional on the (ex post) inflation rate and the output gap in the current period. To see this, consider the following simple relation between the latent interest rate i_t^* and the actual interest rate i_t : $i_t = \max\{i_t^*, 0\}$, where $i_t^* = \omega + \alpha\pi_{t+n} + \kappa y_t + \epsilon_t$. Then, for simplicity, consider an estimation at the conditional mean. In this case, the inflation coefficient α corresponds to $\partial \mathbb{E}_t[i_t^* | \pi_{t+n}, y_t] / \partial \pi_{t+n}$.

By conducting quantile regressions to estimate the Taylor rule for Japan using a sample that includes recent observations, we find that estimates tend to be biased downward if censoring is not taken into consideration. In addition, we obtain estimates of α that are relatively flat and of κ that are insignificant over the conditional interest

rate distribution, suggesting that monetary policy in Japan may be described by a linear policy rule. These results complement the findings from the literature that estimates monetary policy rules using data prior to the zero interest rate policy.

5 Conclusion

We obtain detailed estimates of monetary policy in Japan using a sample that includes recent periods hitting the ZLB. Taking into account endogeneity and censoring, we apply the CQIV method developed by Chernozhukov et al. (2015). In addition, we compare the CQIV results with those from uncensored quantile regressions to address possible estimation bias due to the ZLB.

We find that estimates from uncensored quantile regressions tend to be biased downward. Estimates of the inflation coefficient from CQIV are relatively flat over the conditional quantiles, whereas those from uncensored quantile regressions tend to be increasing and tend to approach CQIV estimates moving toward the upper tail of the conditional distribution. In addition, the estimated coefficients of the output gap from 2SQR are negative and significant, whereas those from CQIV are mostly insignificant. Given such flat estimates of the inflation coefficient and insignificant estimates of the output-gap coefficient, Japan's monetary policy rule may be well described by a linear policy rule.

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