中京大学経済学論叢
28号
2017年３月

Liquidity Expansion and Short-term Monetary Market in Japan [1]

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Abstract

This article investigates the long-run relationship between the current account balance at the Bank of Japan and the call rate under the non-traditional monetary policy in Japan. We estimate the effect of easing monetary policy with and without structural change models. We find that the introduction of the quantitative monetary easing policy lowers the influence on interest rate for the liquidity around 0.3.

Keywords: non-traditional monetary policy; short-term interest rate; cointegration; structural change
JEL classification: E00; E40

1. Introduction

In February, 1999, the Bank of Japan (BOJ) implemented the zero interest rate policy (ZIRP). The quantitative monetary easing policy (QMEP) was carried out afterwards in March, 2001. Under the QMEP, the policy instrument of the BOJ's monetary policy switched from the overnight uncollateralized call rate (OCR) to the current account balance (CAB) at the BOJ and the expansions of the CAB were observed by March 2006. Figure 1 plots the monthly CAB,

1 This article revises Hanabusa [Hanabusa, K., 2008. Quantitative Easing Policy and Interest Rate, *Working Paper Series*, Student-Association Graduate School of Economics, Kobe University, 227,1-8].

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OCR, and one month call rate (MCR). During the period of non-traditional monetary policy (includes the ZIRP cancellation periods), the OCR and MCR changed in approximately 0%.[3]

Under the non-traditional monetary policy, Oda and Ueda (2007) suggests that the expansion of the CAB may have been accepted by the financial market as indicating a greater willingness on the part of the BOJ to carry out zero interest rate level. Oda and Ueda (2007) argues the relation between the expansion of the CAB and the level of interest rate.

Miyao (2002), Bae, et al. (2006), and Hanabusa (2015) estimate the long-run Japanese monetary demand function in a cointegration framework and examine the liquidity traps. Miyao (2002) uses the cointegration approach with regime shift and quarterly observations for the period of 1975:1-2001:4.[4] Bae, et al. (2006) uses the nonlinear functional form and quarterly data set from 1976:1 to 2003:4. Hanabusa (2015) uses the standard cointegration approach and monthly observations for the period of 2006:8-2014:9.

This article examines the long-run relationship between the call rate and the CAB with a

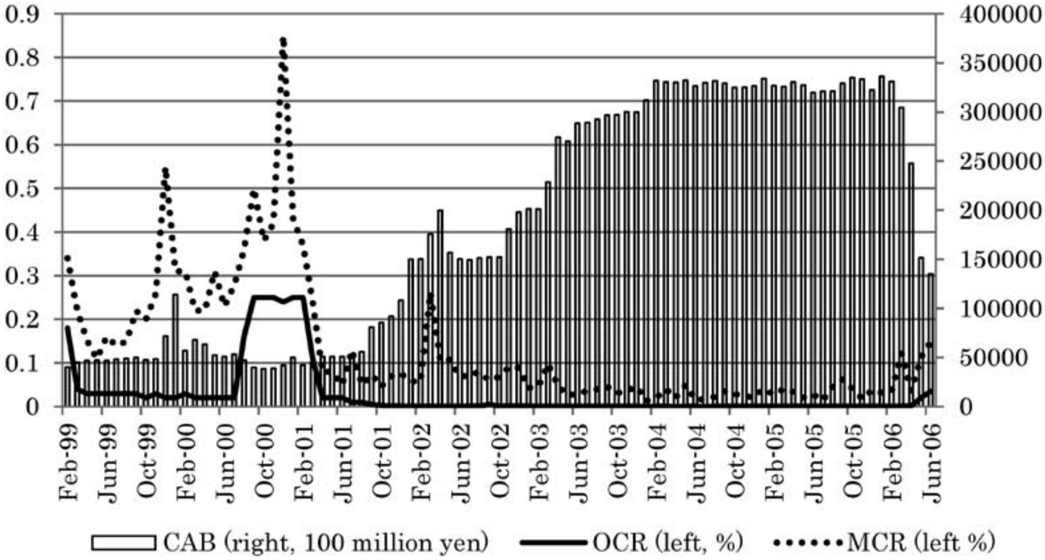

Figure 1. Current account balances and call rate

Note) CAB is current account balances at the BOJ. OCR denotes the overnight uncollateralized call rate and MCR denotes the one month call rate.

3 See Hanabusa (2010) and Ugai (2006) for recent evidence on the effectiveness of non-traditional monetary policy .

4 Hanabusa (2008) examines the relationship between the interest rate and the liquidity under the QMEP and use Japanese monthly interest rate and current account balances from 2000:8 to 2007:4.

structural change. To consider the cointegration with and without structural change, we use the technique same as Miyao (2002). This methodology is proposed by Gregory and Hansen (1996). Because we focus on the effect of easing monetary policy under the non-traditional monetary policy, we use monthly data from 1999:2 to 2006:6.

The rest of the article is organized as follows. We present the methodology and data in Section 2. Section 3 presents empirical results. Section 4 concludes.

2. Methodology and Data

We consider a single-equation model which allows for cointegration with structural change using two variables (y_t and x_t). First, we consider the standard model of cointegration with no structural change:

$$y_t = \alpha_1 + \beta_1 x_t + e_t,\ t = 1,\dots, n, \qquad (1)$$

where, y_t and x_t are integrated of order one, or, $I(1)$. If the disturbance term (e_t) is integrated of order zero, $I(0)$, then the linear combination of y_t and x_t is a cointegrating relation. Second, we use the residual-based test for cointegration with regime shift proposed by Gregory and Hansen (1996). A brief description of the methodology is given below. Before explaining it, we define the dummy variable ($D_{t\tau}$) to model structural change:

$$D_{t\tau} = \begin{cases} 0 & if \quad t \le [n\tau], \\ 1 & if \quad t > [n\tau], \end{cases}$$

where, the unknown parameter $\tau \in (0,1)$ denotes the relative timing of the change point, and [] denotes integer part. However, the test statistic is computed for each break point in the interval ([0.15n], [0.85n]). We use following the regime shift model (C/S) type:

$$y_t = \alpha_1 + \alpha_2 D_{t\tau} + \beta_1 x_t + \beta_2 D_{t\tau} x_\tau + e_t,\ t = 1,\dots, n. \qquad (2)$$

From Equation (2), α_1 represents the intercept before the shift, α_2 represents the change in the intercept at the time of the shift, β_1 represents the cointegrating slope coefficient before the regime shift, and β_2 represents the change in the slope coefficient.

We use the residual-based cointegration test (Engle-Granger type test) to test the null hypothesis of no cointegration. The candidate cointegrating relation is estimated by ordinary least squares (OLS), and a unit root test is applied to the regression errors as following equation.

$$\Delta\hat{e}_{t\tau} = \gamma_0\hat{e}_{t-1\tau} + \textstyle\sum_{i=1}^{K}\gamma_i\Delta\hat{e}_{t-i\tau} + v_t,\ t = 1,\dots n. \qquad (3)$$

The statistic ($ADF(\tau)$) is the t-statistic for the explanatory variable $\hat{e}_{t-1\tau}$, and ADF^* denotes the smallest value of $ADF(\tau)$. In this paper, y_t denotes Japanese monthly call rates (OCR or MCR) and x_t denotes the logarithm of CAB from 1999:2 to 2006:6. The data source is the homepage of Bank of Japan.

We analyze two models. Model 1 uses the logarithm of CAB as an explanation variable and uses OCR as an explained variable. Model 2 uses the logarithm of CAB as an explanation variable and uses MCR as an explained variable.

3. Empirical Results and Discussion

We examine the cointegration of y_t and x_t. Before proceeding to the cointegration tests, we must check for stationary in each variable. For this purpose, the augmented Dickey-Fuller (ADF) test is employed. The lag length of the ADF regression is selected by the Akaike information criterion (AIC). The results of unit root test are reported in Table 1. The test fails to reject the null hypothesis of a unit root for each variable in levels. On the other hand, the null hypothesis of a unit root is rejected for each variable in first differences. Thus, variables are $I(1)$.

Table 2 shows the estimated results of the standard cointegration and C/S models. The ADF statistic shows the statistic of residual-based test for cointegration. In the case of Model 1, the null hypothesis of no cointegration is accepted. However, in the case of Model 2, the null hypothesis of no cointegration is rejected. We can confirm the existence of cointegration between MCR and CAB, but can not do it between OCR and CAB. We consider a possibility of the existence of cointegration with structural change and test it. However, we find no existence of cointegration with structural change between OCR and CAB. Thus, there is no stable relation between OCR and CAB but there is the long-run relationship with structural change relation between OCR and CAB. The structural change is taking place in April, 2001.

Next, we estimate the cointegration coefficients between MCR and CAB using the dynamic OLS (DOLS) method. This method has been proposed by Saikkonen (1991), and Stock and

Watson (1993). The DOLS estimation of the standard type:

$$y_t = \alpha_1 + \beta_1 x_t + \Sigma_{j=-p}^{p} \varphi_j \Delta x_{t-j} + e_t, \; t = 1,\ldots,n . \qquad (4)$$

The empirical result is shown in Table 3. All coefficients are significant and the estimated value of α_1 and β_1 is 1.997 and −0.154, respectively. This shows the empirical result with no break. However, the BOJ decides the introduction of QMEP on March 19, 2001. In consideration of the influence of this policy, we decide to estimate the structural change. We set April, 2001 with a break point. We also find the existence of cointegration with allowing possible break in 2001:3/2001:4 between MCR and CAB in table 2. The *ADF** statistics of C/S model is −5.569. The DOLS estimation of the C/S type:

$$y_t = \alpha_1 + \alpha_2 D_{t\tau} + \beta_1 x_t + \beta_2 D_{t\tau} x_t + \Sigma_{j=-p}^{p} \varphi_j \Delta x_{t-j} + e_t, \; t = 1,\ldots, n. \qquad (5)$$

Table1: Unit root test

	OCR	lags	MCR	lags	CAB	lags
level (constant)	−1.890	8	−2.428	1	−1.524	0
1st difference (constant)	−4.832**	1	−12.734**	0	−3.697**	2

Note) **denotes significant at 1 % level, *denotes significant at 5 % level. Lag length is selected by AIC (Max lag=10).

Table2: Cointegration test

Type	ADF	lag	ADF*	lag
Model 1	−3.304	6	−4.896	5
Model 2	−4.678**	0	−5.569**	0

Note) **denotes significant at 1 % level, *denotes significant at 5 % level. The ADF statistic shows residual-based tests for cointegration, the 1 % critical value is −4.07, the 5 % critical value is −3.37 (Engle and Yoo, 1987).The ADF* statistics that are significant at the 1 % level are labeled with ** and the 5 % level are labeled with *. The 1 % critical value of the regime shift model (C/S) is −5.47 and the 5 % critical value is −4.95 (Gregory and Hansen, 1996). Lag length is selected by AIC (Max lag=6).

Table3: Estimation result

variable	DOLS(W/B)	DOLS(C/S)
α_1	1.997**	4.236**
	(0.275)	(1.567)
α_2	−	−3.584*
	−	(1.596)
β_1	−0.154**	−0.358*
	(0.022)	(0.142)
β_2	−	0.309*
	−	(0.144)
break point	−	2001:4
leads and lags	6	6

Note) **denotes significant at 1 % level, **denotes significant at 5 % level. Newy-West HAC standard errors are in parentheses.

From Table 3, the estimates of α_1, α_2, β_1, and β_2 are 4.236, −3.584, −0.358, and 0.309, respectively. All coefficients are significant. The point estimates indicate that the intercept becomes smaller by 0.652 and the slope coefficient becomes larger by -0.049. Thus the intercept coefficient is a lower value and the slope coefficient is a higher value after the break. Miyao (2002) and Bae, et al. (2006) find a stable relationship between the interest rate and the liquidity. We extend the period of data until the cancellation of zero interest rate and find the stable relationship in the same result as their previous studies. Moreover, we find that the QMEP lowers the influence on interest rate for the liquidity approximately 0.3. These empirical results imply that the QMEP causes further easing in the monetary market. Oda and Ueda (2007) also points out the effect that the QMEP reduces interest rates.

4. Conclusions

In this article, we present empirical evidence that there is the long-run relationship between the current account balance at the BOJ and the one month call rate. The relation among these variables changes after the introduction of the quantitative monetary easing policy. The purpose of this policy reduces the interest rate by supplying a large quantity of liquidity to the monetary market and is to stimulate bank lending and demand action. We find that the influence of the interest rate drop by the liquidity supply becomes small after April, 2001.

Acknowledgement

I gratefully acknowledge the financial support of Grant-in-Aid for Encouragement of Young Scientists (B) by the Japan Society for the Promotion of Science (Project no. 16K17149) and Grant-in-Aid for Scientific Research (B) by the Japan Society for the Promotion of Science (Project no. 16H03618). Needless to say, all remaining errors are mine.

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