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Monetary targeting and inflation: Evidence from Indonesia's post-crisis experience

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Abstract

Using cointegration and structural vector autoregression (SVAR) techniques this paper investigates the effect of Bank Indonesia's (BI) monetary policy on inflation during the post-1997 crisis monetary-targeting period. Our analysis suggests that BI's monetary policy does not have systematic impact on the price level, apparently because of unstable money demand. Unreliable effects of BI's monetary policy are reflected in frequent and substantial deviations of the actual inflation rate from its targeted ranges.

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1. Introduction

Following the Asian financial crisis of 1997, the central banks of Korea, Indonesia, and Thailand adopted floating exchange rates, making inflation targeting its new monetary policy regime. Most economists believe that an inflation-targeting central bank conducts its monetary policy without relying on the relationship between money and inflation. More specifically, the central bank controls the expected inflation by adjusting a key policy interest rate (e.g., Charoenseang and Manakit, 2007 and Kubo, 2008).

During 2000:1–2005:6, Bank Indonesia's (BI) monetary policy regime was *de jure* inflation targeting but *de facto* monetary targeting. After the Asian crisis, BI first adopted a formal monetary-targeting regime in accordance with the International Monetary Fund (IMF) support program. After graduating from the IMF program, BI continued employing the base money as its monetary policy instrument until June 2005, despite the fact that it switched to inflation targeting in January 2000. Fane (2005) suggests that during 2000–2005, BI was not very successful in controlling inflation and was in the process of establishing a standard inflation-targeting regime similar to the one adopted in Thailand.[1]

This paper empirically investigates the effects of BI's (*de facto*) monetary targeting on inflation during the post-crisis period. Prior to the investigation, we estimate the Indonesian money demand functions by using a cointegration analysis to examine its stability on which the efficacy of monetary policy interventions depends. Subsequently, by employing a structural vector autoregression (SVAR) methodology, we estimate the impulse response functions to determine if and to what extent BI is able to control inflation. The following are the main findings of this paper: (1) the estimated Indonesian money demand functions are not stable; and (2) BI's monetary policy does not significantly affect its target variables. These results suggest that the effectiveness of BI's monetary policy is limited.

2. Indonesian money demand functions

During the period under investigation, BI missed its inflation target almost routinely (Figure 1). The poor performance of BI points to the possibility that the Indonesian money demand function is not sufficiently stable.

Some previous studies discuss the stability of Indonesia's money demand function. Bahmani-Oskooee and Rehman (2005) apply a cointegration methodology to the period 1972–2000 and judge that the Indonesian money demand functions are stable. James (2005) investigates the effects of financial liberalization on the demand for money and concludes that there was a stable money demand function for 1983–2000. However, Narayan (2007) analyzes

[1] In July 2005, BI implemented a new and enhanced inflation-targeting regime to strengthen price stability by using an interest rate adjustment as its operational target (Bank Indonesia, 2005).

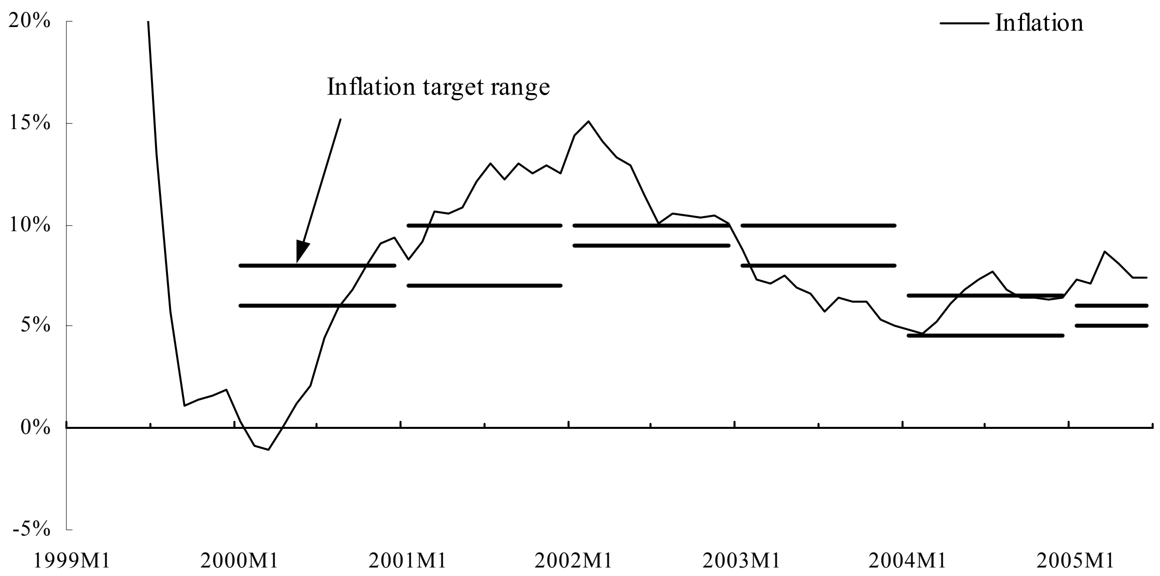

Figure 1. Inflation rates and target ranges in Indonesia

Source: Bank Indonesia

Indonesian's money demand for 1970–2005 and finds unstable money demand functions.

In this section, we use Johansen's (1995) cointegration procedure to estimate money demand in the post-crisis period. More specifically, to investigate the long-run relationship between the variables in the standard money demand function, we normalize the cointegrating vector in an estimated vector error correction model (VECM). The normalized vector can be represented as follows:

$$m_t = b_0 + b_1 r_t + b_2 y_t , \tag{1}$$

where r denotes the 3-month deposit interest rate, a variable representing the return on assets alternative to money; y is real industrial production, a proxy for real income; and m represents real money balances based on $M0$, $M1$, and $M2$ deflated by the consumer price index. In addition, we follow recent studies (e.g., Bahmani-Oskooee and Rehman, 2005) and add the exchange rate, which appears to be relevant in Indonesia:

$$m_t = b_0 + b_1 r_t + b_2 y_t + b_3 f_t . \tag{2}$$

In the above, f represents the nominal exchange rate defined as the price of a US dollar in Indonesian Rupiah. Accordingly, an increase in f denotes depreciation in the Rupiah.

In this and the next sections, all the data are monthly and refer to 1999:1-2005:6. These data were extracted from the IMF's *International Financial Statistics* (*IFS*), except for those on industrial production that were obtained from BPS-Statistics Indonesia's Web site. With the exception of the interest rate, all series are converted into natural logarithms. According to the

Table 1
Unit root test statistics

	r	y	$M0$	$M1$	$M2$	rf
ADF	-2.182 (8)	-1.391 (2)	-1.228 (1)	-2.169 (1)	-3.221 (0)**	-1.487 (2)
KPSS	0.765 (6)**	1.176 (6)**	1.006 (6)**	1.051 (6)**	0.355 (6)*	0.617 (6)**
	Δr	Δy	$\Delta M0$	$\Delta M1$	$\Delta M2$	Δrf
ADF	-4.955 (7)**	-7.825 (3)**	-13.312 (0)**	-12.464 (0)**	-6.179 (2)**	-7.527 (1)**
KPSS	0.441 (6)*	0.159 (11)	0.135 (10)	0.250 (8)	0.299 (7)	0.110 (13)

Note: The numbers in parentheses represent the optimal lag length chosen by SBIC.
** denotes the rejection of the null hypothesis of the unit roots for the *ADF* or of the stationarity for the *KPSS* at the 5% level of significance.
* denotes the rejection of the null hypothesis of the unit roots for the *ADF* or of the stationarity for the *KPSS* at the 10% level of significance.

results of unit root tests, all variables appear to be $I(1)$.[2]

First, the optimal lag length of the unrestricted vector autoregression model is set at k in each case (Table 2 and 3), by referring to the sequential modified likelihood ratio test. Hence, the lag length for the first differenced series in the VECM is k–1. Next, in order to investigate the number of cointegrating vectors, we apply Johansen's (1988) cointegration test by using trace and max-eigen statistics.[3]

Tables 2 and 3 present the estimated long-run relationships among the variables in (1) and (2).

Table 2
Long-run relationships and the money demand function: $m_t = b_0 + b_1 r_t + b_2 y_t$

m	lag (k)	Trace statistic	Max-Eigen statistic	b_0	b_1	b_2
$M0$	$k=6$	24.324	16.673	—	—	—
$M1$	$k=6$	45.149 **	28.206 **	5.660	-1.365	0.380
					(0.27)	(0.15)
					[-5.09]	[2.52]
White heteroscedasticity test (no cross term) = 79.34% Normality test (Jarque-Bera) = 5.54%						
$M2$	$k=6$	33.422 **	22.641 **	7.189	0.587	0.336
					(0.15)	(0.08)
					[3.98]	[3.97]
White heteroscedasticity test (no cross term) = 19.55% Normality test (Jarque-Bera) = 0.63%						

Note: The standard errors are in parentheses; the t-values are in square brackets.
** denotes the rejection of the null hypothesis (the cointegration rank is none) at the 5% level of significance.

[2] See Table 1.

Table 3

Long-run relationships and the money demand function: $m_t = b_0 + b_1 r_t + b_2 y_t + b_3 f_t$

m	lag (k)	Trace statistic	Max-Eigen statistic	b_0	b_1	b_2	b_3
$M0$	$k = 6$	66.375 **	32.215 **	15.811	32.628	9.456	-6.381
					(6.78)	(2.50)	(1.83)
					[4.81]	[3.78]	[-3.49]
White heteroscedasticity test (no cross term) = 64.88%				Normality test (Jarque-Bera) = 0.00%			
$M1$	$k = 7$	65.266 **	36.138 **	1.533	-2.302	0.527	0.397
					(0.35)	(0.13)	(0.09)
					[-6.60]	[4.10]	[4.37]
White heteroscedasticity test (no cross term) = 85.03%				Normality test (Jarque-Bera) = 0.04%			
$M2$	$k = 6$	85.219 **	42.557 **	5.463	-0.711	0.103	0.332
					(0.11)	(0.04)	(0.03)
					[-6.60]	[2.32]	[10.72]
White heteroscedasticity test (no cross term) = 1.86%				Normality test (Jarque-Bera) = 0.08%			

Note: The standard errors are in parentheses; the t-values are in square brackets.
** denotes the rejection of the null hypothesis (the cointegration rank is none) at the 5% level of significance.

Although at least one cointegrating vector was detected for each of these monetary aggregates, the *M0*-based equation in Table 2 appears not to have the long-run relationship between the variables. In addition, we provide residuals analysis, such as White's heteroscedasticity test and the Jarque-Bera normality test, to validate the adequacy of the trace and max-eigen statistics. The null hypotheses of no heteroscedasticity and/or that of no normality are rejected in approximately every case.[4] Moreover, in Tables 2 and 3, some of the estimated equations are not consistent with standard monetary theory. Although the real money demand is positively related to industrial production in every case, the estimated coefficient on the interest rate is statistically significant and positive for the equations for *M2* (in Table 2) and *M0* (in Table 3).[5] Thus, the Indonesian money demand function appears not to be stable. These results suggest that the money demand in Indonesia is not stable, casting doubt on the efficacy of BI's de facto monetary-targeting regime.

3. Transmission mechanism of monetary policy

To measure the dynamic responses of the Indonesian macroeconomic variables to a monetary

[3] A deterministic constant is allowed in the cointegrating space.
[4] The only exception is the *M1*-based equation in Table 2.
[5] Either a positive or negative sign for b_3 seems to be relevant because of the respective assumptions (for details, see Bahmani-Oskooee and Pourheydarian, 1990).

policy shock, we estimate a 5-variable SVAR[6] model as follows:

$$Bx_t = \mu + A_i(L)x_{t-n} + e_t, \tag{3}$$

where matrix B may be restricted by economic theory; μ is a 5-dimensional vector of constants; x_t is a 5-dimensional vector of the relevant variables in time; A_i is a 5×5 matrix of coefficients in the lag operator L; and e_t is a 5-dimensional vector of error terms with zero means and covariance matrix Σ_e.

The variables included in the SVAR model are consumer price index (p), industrial production (y), interbank call money rate (i),[7] base money aggregate (M), and nominal exchange rate expressed in terms of Indonesian Rupiah per US dollar (f). The last variable is included since the 1998 central banking law stipulates that BI pursues both price and exchange rate stability. By referring to the final predict error statistics, the optimal lag length is set at 2. As noted before, all variables other than the interest rate are expressed in natural logarithms.

We first estimate the impulse response functions using a Choleski decomposition[8] with the ordering [p, y, i, M, f] to identify the impact of monetary policy innovation. Nevertheless, some of the impulse response functions do not permit a meaningful interpretation. Thus, we pursue an alternative to the Choleski decomposition, imposing non-recursive zero restrictions[9] on the contemporaneous structural parameters.

Accordingly, (3) now takes the following specific form:

$$\begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ B_{21} & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & B_{34} & B_{35} \\ B_{41} & B_{42} & B_{43} & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} p_t \\ y_t \\ i_t \\ M_t \\ f_t \end{bmatrix} = \mu + A_i(L) \begin{bmatrix} p_{t-n} \\ y_{t-n} \\ i_{t-n} \\ M_{t-n} \\ f_{t-n} \end{bmatrix} + e_t. \tag{4}$$

The first and second rows in matrix B represent the real sector. The third row corresponds to the monetary policy rule. It is assumed that BI sets the interest rate after observing the current base monetary and the exchange rate. This assumption is consistent with BI's monetary policy instrument and objective. The fourth row corresponds to the money demand equation in (1). The fifth row assumes that the exchange rate is contemporaneously exogenous to all the variables in the model. With these restrictions, we have obtained the following five relationships:

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[6] If cointegration is present and not imposed in the estimation, there can be substantial biases in the impulse response functions (Mitchell, 2000). However, we use the level variables to analyze the short-run restrictions in accordance with the recent SVAR literature (e.g., Christiano et al., 1999) and to investigate the relationships among the contemporaneous innovations.

[7] In this section, we use the interbank call money rate—and not the 3-month deposit rate—as a suitable proxy for the monetary policy.

[8] See, for example, Sims (1980) and Christiano et al. (1999).

[9] See, for example, Sims (1986), Christiano et al. (1999), and Kim and Roubini (2000).

$$p_t = e_{pt} \tag{5}$$

$$y_t = -0.324 p_t + e_{yt} \tag{6}$$

$$i_t = 1.194 M_t + 0.050 f_t + e_{it} \tag{7}$$

$$M_t = 3.977 p_t + 0.324 y_t - 3.712 i_t + e_{Mt} \tag{8}$$

$$f_t = e_{ft} \tag{9}$$

In the above, e_{it} can be regarded as a monetary policy shock. Accordingly, (7) and (8) are the monetary policy reaction function and the money demand function, respectively. Price and exchange rate innovations in (5) and (9) are autonomous.[10] They should be predetermined as the policy objective.

Figure 2 presents the estimated responses to a contractionary monetary policy shock (a positive one-standard-deviation innovation to the interbank call money rate). The solid lines represent the point estimates of the impulse responses, while the dashed lines denote their 95% confidence interval. The horizontal axis measures the number of months following the shock. We observe that after initially rising in response to the innovation, the interbank call money rate begins to fall and returns to the initial level after 6 months. Meanwhile, the base money falls sharply and exhibits a peak decline in the first month. Industrial production reaches its lowest point after 3 months. The consumer price declines gradually and reaches the nadir in 18 months following the shock. This result implies that the effect of the monetary policy shock is not

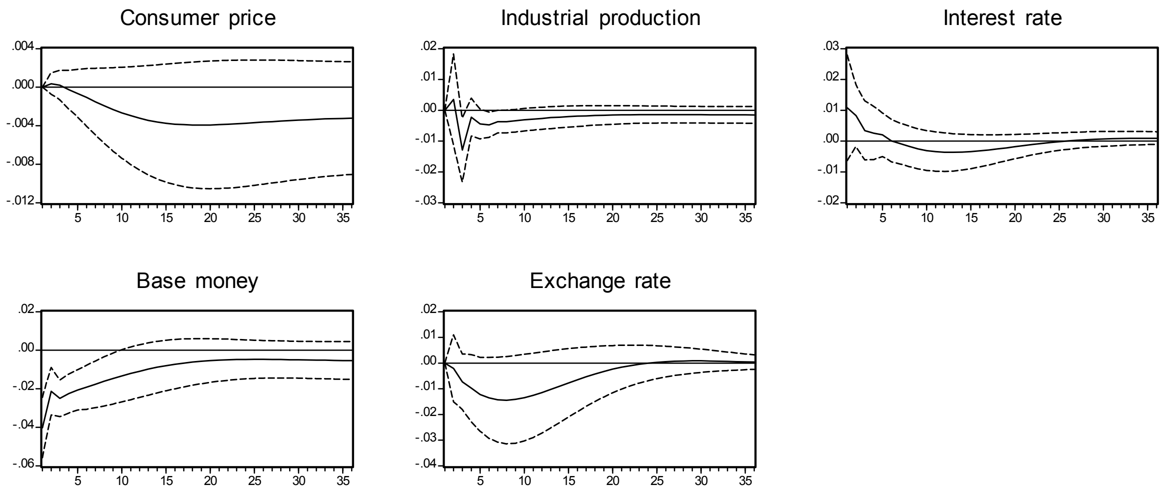

Figure 2. Impulse responses to a monetary policy shock, 1999:1–2005:6

[10] We check for the adequacy of this identification scheme by producing the correlation matrix of shocks. See Appendix A.

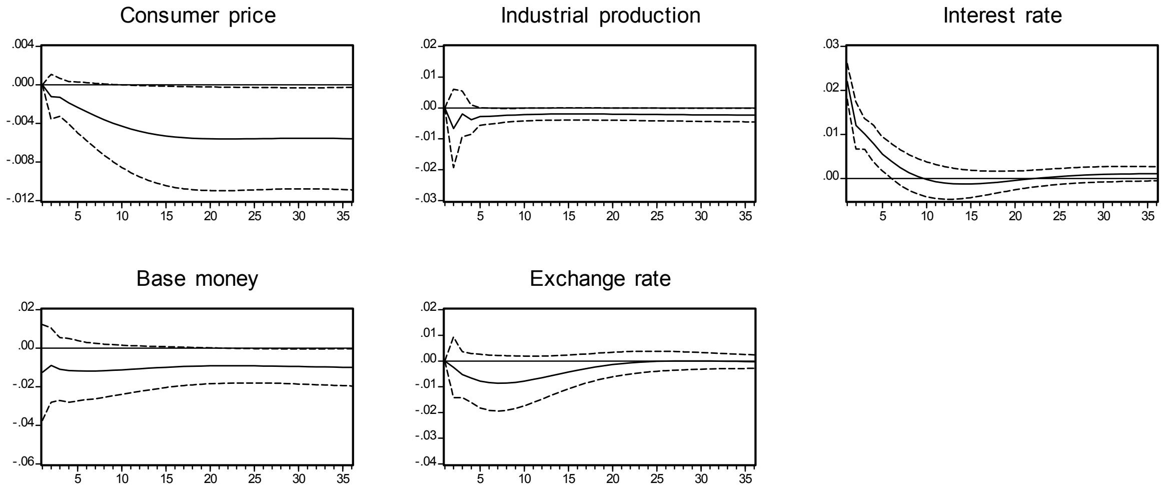

Figure 3. Impulse responses to a monetary policy shock, 1999:1–2005:12.

Note: The optimal lag length is set at 2.

entirely felt on the economy in BI's target horizon, which is reset every year. Finally, the exchange rate falls and reaches the lowest point after 8 months.

Although the previous result suggests that BI apparently has a monetary transmission mechanism to affect the policy objective variables, the declines in the consumer price and exchange rate are not statistically significant.[11] This result suggests that BI's influence on its policy objective variables is at best modest.

4. Conclusion

In this paper, we investigate the stability and consistency of the Indonesian money demand function and the effectiveness of BI's monetary policy using cointegration and SVAR techniques. The estimated money demand functions are not stable, nor does the monetary policy seem to have quantitatively significant effect on the price level. These findings suggest that BI's lever over the price level is at best limited, consistent with the deviations from the inflation target ranges observed in Indonesia almost annually.

On the basis of the empirical results, we conclude that the relationships between money and inflation as well as output in Indonesia are not sufficiently stable. Therefore, opting for a monetary-targeting regime may have been an error in judgment on BI's part.

[11] We expand the sample period for each one-month period, including the period after BI began to use the interest rate adjustment as its operational target. It is not until 2005:12 that the statistically significant effect on price is observed (see Figure 3).

Appendix A. Correlation matrix of shocks

	p	*y*	*i*	*M*	*f*
p	1.00	-0.03	0.12	0.18	0.20
y		1.00	0.22	-0.06	-0.03
i			1.00	-0.09	0.06
M				1.00	-0.02
f					1.00

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