Economics and Finance in Indonesia Vol. 57 (2), Page 181 – 202

Monetary Policy and Trade Balance in Indonesia: Evidence from a Structural VAR Analysis

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Abstract

This paper investigates the response of the trade balance to monetary policy innovations for the Indonesia economy using a Structural Vector Auto Regressive (SVAR) model. We estimate three type of SVAR models using monthly data, that is, for full sample (1980:M1 until 2005:M3), pre- (1980:M1-1997:M6), and post-Asian crisis periods (1998:M1-2005:M3). The aim of splitting full sample size is to take into account a sensibility of structural break due to the onset of Asian financial crisis in 1997-1998. The dynamic properties of the model are discussed and monetary policy shocks identified by means of various impulse responses and a variance decomposition.

Keywords: Monetary Policy, Exchange Rate, Trade Balance, Structural VAR, Indonesia

JEL classification: C32, E52, F41

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1. INTRODUCTION

The effects of monetary policy shocks on the trade balance and the exchange rate have long been discussed. However, theoretical models and findings of empirical studies provide controversial conclusions. The conventional model of open economy of Mundell (1963, 1964) and Flemming (1962) and the exchange rate overshooting model of Dornbusch (1976) model expect a monetary expansion leads to depreciation in the nominal exchange rate and a deterioration of the terms of trade, which results in an improvement of the trade balance, namely the expenditure-switching effect. However, another theoretical model the so-called the income-absorption effect proposed by Meade (1951) and Alexander (1952, 1959) produce a counter explanation. A monetary expansion also accelerates domestic demand, which in turns may lead to an increase in imports and a deterioration of the trade balance. Recent theoretical framework using the intertemporal model with price rigidities assumption developed by Obstfeld and Rogoff (1995) and Razin (1995) argues that a monetary expansion leads to enlarge income temporarily, so that the trade balance may improve as a consequent of consumption smoothing. On the other hand, the trade balance may degenerate if investments increase considerably caused by a fall of the real interest rate.

Empirical studies appear to be conflicting and mix at best to support the theoretical view of the effects of monetary policy shocks on the trade balance. The results of empirical investigations reviewed here employ a VAR (Vector Autoregression) model to identify the effects. Eichenbaum and Evans (1995) construct an open economy VAR model without explicitly introducing the trade sector into the model. They find that an expansionary shock to US monetary policy leads to persistent depreciations in nominal and real US exchange rates. Koray and McMillin (1999) investigate the response of the trade balance to monetary innovations for the US economy. They conclude that the immediate appreciation of the US exchange rate in response to US contractionary monetary policy shock. The deterioration in the trade balance is correlated with an appreciation of the exchange rate. The results of recent study by Kim (2001a, 2001b), applied for the US, France, Italy and the UK, confirm the existence of the income-absorption effect and the expenditure-switching effect with the first dominates in the short run and the latter dominates in the medium and long run following a monetary expansion. The result is that trade balance improves in the short run and deteriorates in the long run.

In this paper, we focus on empirical evidence of the effects of monetary policy shocks on the trade balance as well as exchange rate and output as important variables to explain the dynamic pattern of trade

balance. We investigate such effects in Indonesia for the recent period. To our knowledge, there is little empirical evidence of such effects in open economy such as Indonesia, except Sugema (2005) and Suselo et al. (2008), for instance. Sugema (2005) discusses the effects of real exchange rate depreciation and supply side shocks on trade balance in Indonesia during the episode of the Asian financial crisis in 1997-1998. Suselo et al. (2008) examines the effects of exchange rate volatility on economic growth by using trade (export and import) and investment as transmission channels. However, there is no explicit role of monetary policy shocks to elaborate trade balance behavior in their analysis.

Another major challenge of the analysis of the impact of monetary policy shocks on the trade balance is that Indonesia has been the most severely affected country in terms of economic and social costs by the Asian crisis. This will make the specification of a model a difficult task because the relevant variables of the model exhibit jump-like behavior around the crisis. Fung (2002) investigates monetary policy shocks using a VAR analysis across seven East Asian countries including Indonesia. He implements a subsample-based analysis in order to take account the importance of structural breaks owing to the Asian crisis. The same strategy is done by Hanson (2006) in analyzing the practice of monetary policy in the US before and after the Volcker disinflation periods.

The purpose of this study is to examine whether a monetary expansion improve or worsen the trade balance. We follow a more direct approach and investigate the response of the trade balance to monetary innovations within the context of a VAR model, which is based on minimal identifying restrictions and does not depend much on a specific theoretical model. Analyzing the association of the trade balance with exchange rate movements within this context requires an understanding of how monetary policy affects the economy. To achieve these objectives, we utilize newly developed models from the VAR literature on identifying monetary policy shocks for an open economy, among others, Eichenbaum and Evans (1995), Cushman and Zha (1997), Kim and Roubini (2000), and Kim (2001a, 2001b). We modify their VAR models in order to be able to document data-oriented evidence on monetary policy shocks in Indonesia. Given the regime shifts and structural break of the Asian crisis, we follow Fung (2002) to split models into a pre- and post-crisis sample-based model and to examine the sensibility of impulse responses between both models. With the same examination, we then compare both models with the model estimated using full sample size. Three types of impulse response experiments and a variance decomposition are carried out to analyze the impact of a short run monetary policy with particular attention to response of trade balance. We conclude that the trade balance improvement is found after a monetary contraction for full sample and post-crisis estimations, but the

evidence during the pre-crisis episode reveals the deterioration in trade balance. The response of other related variables are consistent with economic theory for all models, except the price puzzles exists. The variance decompositions indicates that trade balance reacts to real output shocks for full sample model and to exchange rate and the commodity price shocks for pre-crisis and post-crisis models.

The remainder of the paper is organized as follows. Section 2 presents the empirical model. We show the structural VAR modeling method and present the details of our identification scheme for monetary policy shocks. Section 3 discusses the data and its properties. Section 4 provides the empirical evidence of the dynamic effects of the monetary policy shocks on the trade balance and presents the evidence of source of fluctuations in the trade balance. Section 5 concludes.

2. THE EMPIRICAL MODEL

We set up a SVAR model for open economy based upon that of Kim and Roubini (2000) and Kim (2001a, 2001b). Their models are implicitly built up from a model of the macro-economy based upon optimizing agents developed by Sims and Zha (2006). The choice of their models is driven by the result that the structural VAR approach appears to be quite successful in explaining all the puzzles that plagued the recent literature on the effects of monetary policy in closed and open economies. To understand the model it will be useful to summarize the VAR modeling process. First, we describe the economy by the following structural form represented by a set of linear dynamic equations,

$$G(L)y_t = B_0\varepsilon_t \qquad (1)$$

where y_t is a k-vector of endogenous variables and ε_t is a vector of interpretable structural disturbances taken from the equations describing the system with a covariance matrix. B_0 is a diagonal matrix and $G(L)$ is a matrix polynomial in the lag operator L. $G(L)$ is defined as $G_0 + G^0(L)$, where $G^0(L)$ equals to $I_K - G_1L - \ldots - G_pL^p$. Equation (1) shows the structural form of the model in moving average form with $G(L)$ an infinite order lag polynomial. Although the structural model is not observable, we can express, under mild conditions of G_0 being invertible, the p-dimensional stationary autoregressive process y_t in its reduced form. Thus, the reduced form corresponding to the structural form (Equation (1)), are obtained by premultiplying with G_0^{-1}:

$$y_t = B(L)y_t + G_0^{-1}B_0\varepsilon_t = B(L)y_t + u_t \qquad (2)$$

where $B(L) = -G_0^{-1}G^0(L)$ and the relation between the reduced-form disturbance u_t and the underlying structural shocks ε_t:

$u_t = G_0^{-1}B_0\varepsilon_t$ or

$$G_0 u_t = B_0\varepsilon_t \qquad (3)$$

The relation between the variance-covariance matrices u_t and ε_t is given as follows:

$$\Sigma_u = E[u_t u_t^{'}] = G_0^{-1}B_0 E[\varepsilon_t \varepsilon_t^{'}]B_0^{'}G_0^{-1'},$$

$$\Sigma_u = G_0^{-1}B_0 I B_0^{'}G_0^{-1'} \qquad (4)$$

where $\Sigma_\varepsilon = E[\varepsilon_t \varepsilon_t^{'}]$ is specified as an identity matrix I. In any case, the structural shocks are instantaneously uncorrelated. Equation (4) can be rewritten as:

$$\Sigma_u = G_0^{-1}\Lambda G_0^{-1'} \qquad (5)$$

where Λ equals to $B_0 I B_0^{'}$ and it is a diagonal matrix. The number of parameters of the reduced-form after leaving out the parameters attached to the lagged variables is given by the number of non-redundant elements of matrix Σ_u, that is, $K(K+1)/2$. The overall number of elements of the structural form matrices G_0 and B_0 is $2K^2$. Therefore, $2K^2 - K(K+1)/2$ restrictions are required to identity full model. The G_0 in the structural VAR approach can be any structure as long as it has enough restrictions. Since the primary interest of this study is in short and medium term responses, we use contemporaneous restrictions on the G_0 matrix to identify the shocks instead of identification strategy of the Choleski decomposition pioneered by Sims (1980). We follow the identification strategy of Kim and Roubini (2000) and Kim (2001a, 2001b). We modify their identification by using trade balance instead of exports and imports and by amending the model slightly to include the exchange rate. Our identification scheme is given by Equation (6).

$$\begin{pmatrix} 1 & \alpha_{12} & 0 & 0 & \alpha_{15} & \alpha_{16} & 0 & 0 \\ \alpha_{21} & 1 & \alpha_{23} & \alpha_{24} & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & \alpha_{34} & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ \alpha_{61} & \alpha_{62} & \alpha_{63} & \alpha_{64} & \alpha_{65} & 1 & 0 & 0 \\ \alpha_{71} & \alpha_{72} & \alpha_{73} & \alpha_{74} & \alpha_{75} & \alpha_{76} & 1 & 0 \\ \alpha_{81} & \alpha_{82} & \alpha_{83} & \alpha_{84} & \alpha_{85} & \alpha_{86} & \alpha_{87} & 1 \end{pmatrix} \begin{pmatrix} u_R \\ u_M \\ u_P \\ u_Y \\ u_{FFR} \\ u_{WP} \\ u_{ER} \\ u_{TB} \end{pmatrix} =$$

$$\begin{pmatrix} \beta_{11} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & \beta_{22} & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & \beta_{33} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & \beta_{44} & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & \beta_{55} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & \beta_{66} & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & \beta_{77} & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & \beta_{88} \end{pmatrix} \begin{pmatrix} \varepsilon_R \\ \varepsilon_M \\ \varepsilon_P \\ \varepsilon_Y \\ \varepsilon_{FFR} \\ \varepsilon_{WP} \\ \varepsilon_{ER} \\ \varepsilon_{TB} \end{pmatrix}$$

(6)

where ε_R is a short-term interest rate shock and interpreted as a monetary policy shock, ε_M is a money supply shock, ε_P is a price level shock, ε_Y is an output shock, ε_{FFR} is a foreign interest rate shock, ε_{WP} is a foreign price level shock, ε_{ER} is an exchange rate shock, and ε_{TB} is a trade balance shock. The reduced form errors are denoted similarly. Note that the structural model is overidentified with 3 degrees of freedom. This leads to a testable hypothesis of the validity of the model.

It will be instructive to explain the salient features of the identification strategy. The structural model is composed of several blocks. The first two equations are money supply and money demand equations, which describe money market equilibrium. The first row in Equation (6) represents the monetary policy reaction function. It is assumed that the monetary authority is not able to contemporaneously react to output and the price level due to timing of information, that is, measures of these variables are not available at the time policy is set. However, the monetary authority can respond contemporaneously by setting the interest rate to variables that are available contemporaneously:

money, foreign interest rate, and foreign price level. The second row in Equation (6) is a money demand equation in that it assumes money demand depends contemporaneously on real income, price, and the opportunity cost of holding money, that is, the nominal interest rate.

The third and fourth rows represent the domestic goods market equilibrium. The large number of zero restrictions in these rows is consistent with a model exhibiting nominal rigidities, for example, Gordon and Leeper (1994), Sims and Zha (2006), and Dellas (2006). Firms are assumed to adjust slowly their output and price in reacting to sudden change in monetary policy (the domestic interest rates and money) or financial indicators (foreign interest rate and exchange rate). Therefore, monetary policy and financial indicators are assumed not to affect the level of real activity contemporaneously. However, domestic output is assumed to respond to changes in domestic prices immediately. The fifth row assumes that foreign interest rate is contemporaneously exogenous to all variables in the system, that is, an individual small open economy cannot influence foreign interest rate contemporaneously.

The final three rows in the model are foreign price level, exchange rate, and trade balance. These rows are contemporaneously affected by all variables in the system. Foreign price level and exchange rate equations represent financial market equilibriums. All currently available information in the system is assumed to affect these financial market equilibriums instantaneously. Justification of this assumption is that the exchange rate is a forward-looking asset price and foreign price equation is assumed to be an arbitrage equation. In line with the blocks of these equations, we do not impose any restrictions on the equation for trade balance since we are interested in how this variable responds to change in all variables in the model.

3. DATA SOURCES AND PROPERTIES

We focus on Indonesia to uncover the effects of monetary policy shocks on trade balance by employing the SVAR model developed in Section 2. We use monthly data covered from 1980:M1 until 2005:M3. Since contemporaneous restrictions are used to identify the VAR models, monthly data are more appropriate than quarterly data. It is more difficult to justify the identification assumption of no contemporaneous feedback from policy to the economy at the quarterly frequency.

The data are all from the International Financial Statistics database at the IMF. The short-term domestic interest rate (R) used to represent the stance of monetary policy is the interbank rate. Money supply (M) is nominal as listed in the monetary survey. The domestic price level and real GDP (Y) used are the consumer price index (P) and the GDP volume

index with year 2000 as a base year. The US Federal Funds rate (*FFR*) is used as the foreign interest rate, and the world commodity price is used for the foreign price level (*WP*). The nominal exchange rate (*ER*) is national currency per US dollar. The trade balance (*TB*) is defined as differences in the volumes of exports and imports. The interbank rate and the US Federal Funds rate enter the model in levels, the rest enter in logarithms. Some annual and quarterly variables are interpolated using cubic spline method[1] to obtain a complete monthly database. Those variables are real GDP and volumes of exports and imports. Figure 1 plots the (log) levels time series used in the empirical analysis. Domestic variables clearly shows structural break during the onset of the Asian crisis.

Figure 1. The underlying data

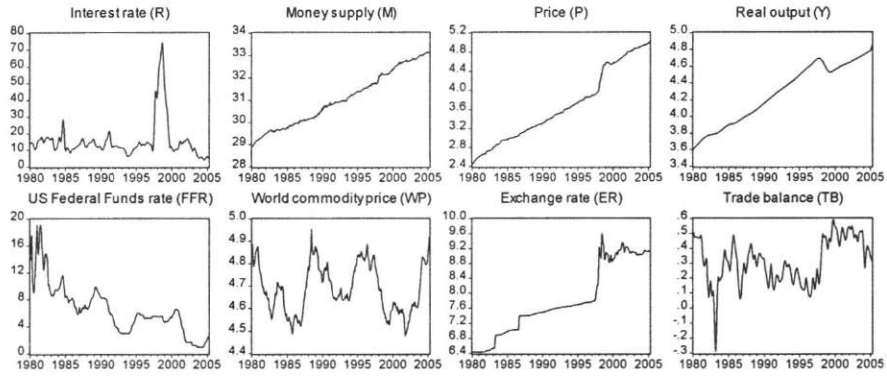

Table 1 shows the results of unit root tests, the augmented Dickey-Fuller (ADF) which tests the null that variables has a unit root against the stationary alternative. The general conclusion from the table is that all series can be treated as *I*(1). The unit root null hypothesis of the ADF test is rejected at 5% level for all sample size and series. With exceptions, real output exhibits *I*(2) in the no-trend and trend specifications for full sample and post-crisis sample. Interest rate has zero order of integration for both no-trend and trend, and exchange rate has the same order of integration for post-crisis period. In the pre-crisis episode, the outcomes of the ADF test support *I*(0) for the trade balance and price in the no-trend and trend specifications, respectively.

[1] See Boor (1978) and Fritsch and Carlson (1980) for further technical details about the cubic spline interpolation.

Table 1. Augmented Dickey–Fuller and KPSS tests results, 1980:M1–2005:M3

Variables	Augmented Dickey–Fuller statistics							
	Level				Δ			
	No trend	Lag(s)	Trend	Lag(s)	No trend	Lag(s)	Trend	Lag(s)
	Full sample 1980:M1-2005:M3							
R	-2.27	14	-2.23	14	-6.19	13	-6.19	13
M	-0.44	1	-2.03	1	-5.12	0	-5.15	0
P	-0.12	3	-2.01	3	-5.84	2	-5.85	2
Y	-1.20	3	-1.51	3	-0.99	2	-0.61	2
FFR	-2.54	13	-3.26	13	-11.49	0	-11.48	0
WP	-1.94	2	-1.87	2	-12.64	0	-12.78	0
ER	-0.02	7	-2.66	7	-5.58	6	-5.68	6
TB	-2.55	7	-3.13	7	-6.74	6	-6.75	6
Pre-crisis sample 1980:M1-1997:M6								
R	-1.94	11	-1.70	11	-5.81	10	-5.79	10
M	-0.57	2	-2.23	2	-14.08	1	-14.05	1
P	-1.87	0	-3.87	0	-8.07	3	-8.19	3
Y	-2.77	14	-1.74	14	-3.12	14	-4.46	14
FFR	-2.41	13	-2.57	13	-3.99	12	-4.16	12
WP	-2.17	1	-2.71	1	-11.16	0	-11.23	0
ER	-1.37	0	-.139	0	-14.52	0	-14.56	0
TB	-3.24	7	-3.18	7	-5.44	6	-5.48	6
Post-crisis sample 1998:M1-2005:M3								
R	-6.46	8	-8.66	8	-3.22	7	-3.70	7
M	-0.68	0	-2.36	0	-11.32	0	-11.28	0
P	-1.55	1	-3.36	1	-6.99	0	-6.66	0
Y	-2.60	2	-1.09	2	-0.13	1	-1.69	1
FFR	-1.28	1	-1.38	1	-3.80	0	-3.84	0
WP	-0.70	5	-0.67	5	-5.44	0	-4.02	0
ER	-3.40	0	-3.41	0	-8.14	0	-8.08	0
TB	-0.62	6	-1.42	6	-6.07	5	-6.43	5

Note: The number of lags for the unit root tests is based on Schwarz information criterion. At the 5% level, the ADF test critical values with trend and without trend are -3.43 and -2.87, respectively.

Table 2 presents an overview of the results of the VAR-based cointegration tests using the methodology developed by Johansen (1991) for all sample size. An intercept and a deterministic time trend are included in each long-run relation of the VECM (vector error correction model) system. Depending upon which test is used, the table shows that only a fully integrated rank can be rejected. We follow Sims et al. (1990) and estimate the models in levels since the focus of the dynamics of our system is short horizon responses. As Sims et al. (1990) note that an unrestricted VAR in levels correctly estimates the dynamics of the system taking into account whatever cointegration may be present and estimate the model in levels. In addition, Faust and Leeper (1998) emphasize that ill-specified (or arbitrarily chosen) long-run vectors (cointegrating relations) will lead to biased short-run parameters and therefore will affect impulse-response functions.

Table 2. Cointegration rank test statistics

Ho	*Ha*	Statistics	Critical value	*Ho*	*Ha*	Statistics	Critical value
LR trace				*LR max*			
Full sample 1980:M1-2005:M3							
$r = 0$	$r \geq 1$	289.04*	187.47	$r = 0$	$r = 1$	81.14*	56.71
$r \leq 1$	$r \geq 2$	207.90*	150.56	$r \leq 1$	$r = 2$	69.02*	50.60
$r \leq 2$	$r \geq 3$	138.88*	117.71	$r \leq 2$	$r = 3$	51.91*	44.50
$r \leq 3$	$r \geq 4$	86.97	88.80	$r \leq 3$	$r = 4$	35.00	38.33
$r \leq 4$	$r \geq 5$	51.97	63.88	$r \leq 4$	$r = 5$	23.11	32.12
$r \leq 5$	$r \geq 6$	28.86	42.92	$r \leq 5$	$r = 6$	17.02	25.82
$r \leq 6$	$r \geq 7$	11.84	25.87	$r \leq 6$	$r = 7$	8.54	19.39
$r \leq 7$	$r = 8$	3.29	12.52	$r \leq 7$	$r = 8$	3.29	12.52
Pre-crisis sample 1980:M1-1997:M6							
$r = 0$	$r \geq 1$	285.71*	187.47	$r = 0$	$r = 1$	77.20*	56.71
$r \leq 1$	$r \geq 2$	208.51*	150.56	$r \leq 1$	$r = 2$	58.97*	50.60
$r \leq 2$	$r \geq 3$	149.54*	117.71	$r \leq 2$	$r = 3$	46.85*	44.50
$r \leq 3$	$r \geq 4$	102.69*	88.80	$r \leq 3$	$r = 4$	41.41*	38.33
$r \leq 4$	$r \geq 5$	61.27	63.88	$r \leq 4$	$r = 5$	27.41	32.12
$r \leq 5$	$r \geq 6$	33.86	42.92	$r \leq 5$	$r = 6$	17.34	25.82
$r \leq 6$	$r \geq 7$	16.51	25.87	$r \leq 6$	$r = 7$	9.76	19.39
$r \leq 7$	$r = 8$	6.75	12.52	$r \leq 7$	$r = 8$	6.75	12.52
Post-crisis sample 1998:M1-2005:M3							
$r = 0$	$r \geq 1$	306.88*	187.47	$r = 0$	$r = 1$	106.97*	56.71
$r \leq 1$	$r \geq 2$	199.91*	150.56	$r \leq 1$	$r = 2$	70.80*	50.60
$r \leq 2$	$r \geq 3$	129.11*	117.71	$r \leq 2$	$r = 3$	40.16*	44.50
$r \leq 3$	$r \geq 4$	88.968*	88.80	$r \leq 3$	$r = 4$	30.40	38.33
$r \leq 4$	$r \geq 5$	58.56	63.88	$r \leq 4$	$r = 5$	28.58	32.12
$r \leq 5$	$r \geq 6$	29.98	42.92	$r \leq 5$	$r = 6$	13.48	25.82
$r \leq 6$	$r \geq 7$	16.50	25.87	$r \leq 6$	$r = 7$	10.05	19.39
$r \leq 7$	$r = 8$	6.45	12.52	$r \leq 7$	$r = 8$	6.45	12.52

Note: *LR trace* and *LR max* are likelihood ratio tests based on the trace and the maximum eigenvalue, respectively. The mark * denotes a significant rejection of the null hypothesis at the 5% level. r stands for number of cointegrating relations.

4. EMPIRICAL RESULTS

In this section we present the results of the three alternatively identified VAR models. We first estimate the identified VAR-model with the full sample of 303 months (1980:M1-2005:M3). Sub-sample based analysis is considered to examine the importance of structural breaks owing to the Asian crisis. Hence, we estimate the other two identified VAR-models with the pre-crisis sample of 210 months (1980:M1-1997:M6), and the post-crisis sample of 87 months (1998:M1-2005:M3). By looking at these two sub-samples, it may also be possible to shed light on whether there has been a change in the dynamic effects of monetary policy to trade balance since the 1997 Asian financial crisis.

We summarize the outcomes of the estimation exercise in two ways. First, we present the impulse response functions for the domestic interest rate shocks as a stance of monetary policy. Second, we give insight into the nature of historical shocks presenting the variance decompositions of the structural factorizations of the models. We treat all variables as endogenous variables in the SVAR models. Those variables are the money market interest rate (*R*), money balances (*M*), the domestic price level (*P*), real output (*Y*), the US Federal Fund rates (*FFR*), the world commodity price (*WP*), the nominal exchange rate (*ER*), and trade balance (*TB*). Real output and prices are seasonally adjusted. The number of lags employed in the estimation varies according to the sample period. For the full sample model and pre-crisis model estimations, 4 lags are used. We use 2 lags for the post-crisis model by bearing in mind that we have a short-time time series, we could run the risk of over-parameterization. For all models, we apply the structural identification as set out in the Section 2 and the model is estimated by maximum likelihood using the scoring algorithm of Amisano and Giannini (1997). Full seasonal dummies, an intercept and a time trend are included. The structural identification embeds overidentification of the parameters. Table 3 presents likelihood ratio tests for the validity of the over-identifying restrictions. The restrictions are not rejected at the 5% level for all models.

Table 3. Likelihood ratio tests for over-identification

	Full sample	Pre-crisis sample	Post-crisis sample
LR statistic	6.79	3.32	3.20
p-value	0.08	0.35	0.36
Degrees of freedom	3	3	3
Log-likelihood	7224.54	5478.09	1964.38

Note: The LR statistic is asymptotically distributed as a Chi-squared with degrees of freedom equal to the number of over-identifying restrictions. At the 5% level, the critical value of the LR test statistic with three degrees of freedom is 7.81. Periods covered for full sample, pre-crisis sample, and post-crisis sample are 1980:M1-2005:M3, 1980:M1-1997:M6, and 1998:M1-2005:M3, respectively.

4.1. Impulse response functions

The standard practice to compute impulse responses is to shock the system with a one standard deviation shock computed from the VAR innovations. Figure 1, 2, and 3 shows the impulse response functions of the domestic variables included in the VAR to a monetary policy shock that results in a one-standard deviation rise in the interest rate over the

full sample, pre-crisis, and post-crisis sample, respectively. The two dashed lines in each panel depict the 95% confidence bands computed using bootstrap confidence interval of Hall's percentile interval[2] with 500 replications. Given the assumption that the shocks are instantaneously uncorrelated, the corresponding impulse responses are referred to as orthogonalized impulse responses. The impulse responses are plotted over a 36 months horizon.

Following a contractionary monetary policy shock, it is expected that output, prices and money balance will all fall whilst the interest rate and the exchange rate rise. Money and the exchange rate respond to the shock more immediately and prices tend to respond more slowly. In the context of response of trade balance to monetary policy shocks, the expenditure-switching effect predicts that a monetary contraction improve the trade balance via a depreciation in the nominal exchange rate. On the other hand, the income-absorption effect expects contractionary monetary policy may reduce imports by lowering domestic demand, that is, through income absorption. These two effects move the trade balance in reverse directions. The movements of the trade balance are determined by the dominant effect. Therefore, if the income absorption effect offsets the expenditure-switching effect, then trade balance may improve.

Figure 1 shows that the general dynamic pattern of an interest rate shock in the full sample model is the following. Following the tightening policy shock, interest rates rise immediately and significantly. The interest rate increases are fairly huge revealing that Indonesia has high inflation experiences. The interest rate responses are very short-lived in which interest rates rising significantly for less than four months after the shock and then declining. These findings are consistent with the view that the liquidity effect is relatively short-lived, that is, the interest rate rises shortly subsequent to a contractionary shock. For other domestic variables, money balance falls initially and increases after eight months. Even though the world commodity prices and the US Federal Funds rate are included in the VAR system to resolve price puzzle suggested by Sims (1992) and Kim and Roubini (2000), the price puzzle exists wherein a monetary contraction is associated with a persistent increase in the price level. These findings are in the line with Fung (2002) and Lestano and Sterken (2006) who found the price puzzle as results of interest rate shocks in their VAR system for Indonesia over the sample that includes the episodes of the Asian crisis.

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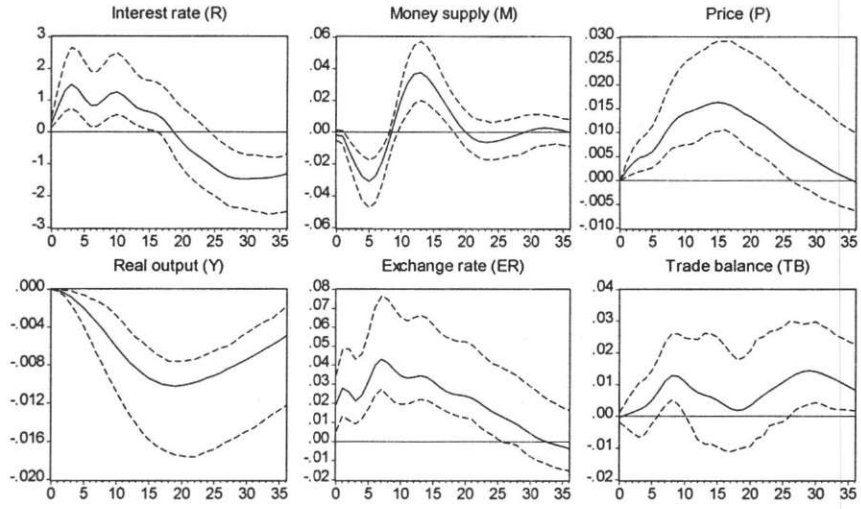
[2] See Hall (1992) and Benkwitz, et. al (2000) for a detailed exposition of Hall's bootstrap confidence interval.

Exchange rate exhibits appreciation and output falls immediately after a contractionary monetary policy shock. This decline in output is quite persistent. In the case of trade balance response to a monetary contraction, Figure 1 reveals that the improvement of trade balance is quite persistent following a monetary tightening. The income-absorption effect appears to dominate the expenditure-switching effect. An appreciation in the nominal exchange rate after a monetary tightening is much less stronger to deteriorate the trade balance compare to an improvement of the trade balance due to a fall in real output after a monetary contraction.

Figure 2. Impulse responses to monetary policy shock, 3 years: full sample, 1980:M1–2005:M3

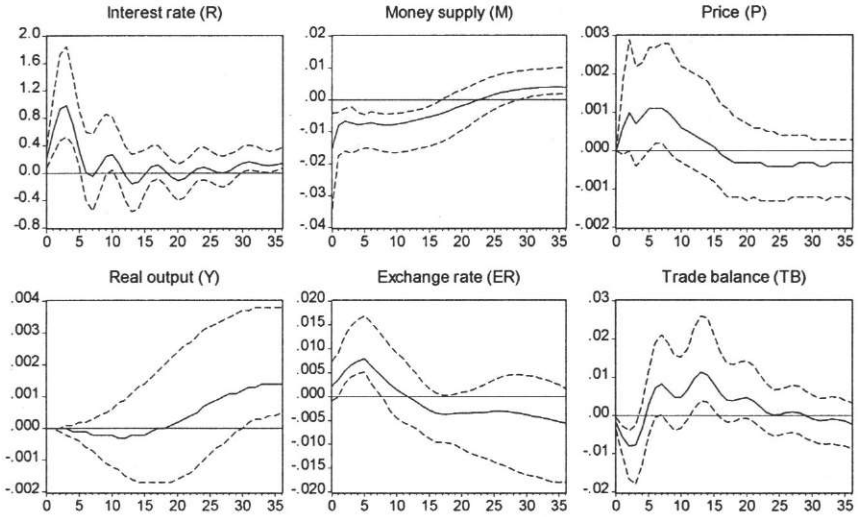


Note: Dashed lines are 95% bootstrapped confidence intervals calculated using 500 replications of Hall's percentile method.

Figure 2 and 3 depict the impulse response functions to a contractionary monetary policy shock over two sample periods: pre-crisis sample (1980:M1-1997:M6) and post-crisis sample (1998:M1-2005:M3). For both samples, the responses are in line with the predictions of economic theory: money and output fall, at the same time as the exchange rate appreciates. The response of price to a monetary tightening reveals price puzzle in which a positive response of price is more persistent in the post-crisis than the pre-crisis sample. The estimated trade balance responses are interesting. Immediately after the shock for pre-crisis periods, the trade balance decreases but the responses are not very persistent lasting for only four to five months. On the contrary, the

responses of trade balance in the post-crisis sample show the opposite direction. These facts reveal that the expenditure-switching effect dominates the income-absorption effect in the pre-crisis periods but the reverse domination is true for post-crisis periods.

Figure 3. Impulse responses to monetary policy shock, 3 years: pre-crisis sample, 1980:M1-1997:M6

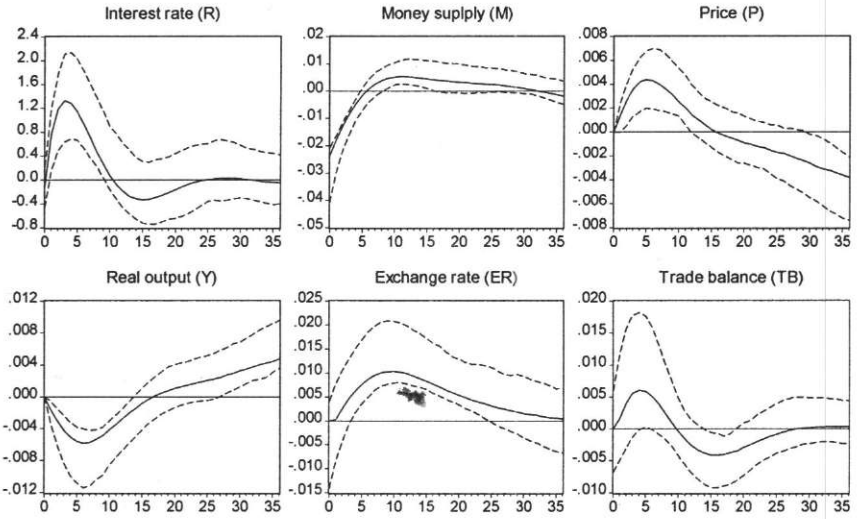


Note: Dashed lines are 95% bootstrapped confidence intervals calculated using 500 replications of Hall's percentile method.

In general, several features are common to all three estimation periods. In all models, the responses of output, money and the exchange rate are consistent with economic theory. Fung (2002) argues that in the VAR system that includes the periods of the Asian financial crisis 1997, the structural break as an outcome of the Asian financial crisis in 1997 may play a role in producing price puzzle. Therefore, splitting the full sample in order to omit the Asian crisis periods may eliminate the price puzzle. However, our results show that the price puzzle found in all VAR models is a potential problem for Indonesia economy studied here. The perverse price response after a monetary shock is more persistent in the full sample model than pre- and post-crisis models. One reason behind the price puzzle over three estimations of our VAR models is that commodity prices may provide evidence for a kind of cost channel of monetary policy viewed, for instance, by Christiano, et. al (2005), and that the nature of the cost channel may have changed over time. Another reason is that since we use monthly data, Hanson (2004) notes that the

price puzzle is generally more pronounced with monthly than lower data frequency, such as, quarterly data. In the case of full sample estimation, the Asian crisis may have distorted the outcomes in the full sample. During the crisis period, inflation rise sharply as the exchange rates in Indonesia collapsed while interest rates are raised substantially to defend the currency. The comovement of interest rates and prices in response to the huge external shock may result in the price puzzle.

Figure 4. Impulse responses to monetary policy shock, 3 years: post-crisis sample, 1998:M1-2005:M3



Note: Dashed lines are 95% bootstrapped confidence intervals calculated using 500 replications of Hall's percentile method.

Overall, the results of the dynamic pattern of trade balance following a monetary contraction show fundamentally differences between full sample, and before and after mid-1997. Table 4 summaries the results of the impact of the interest rate shock: the peak level effect and persistent effect. The first effect shows timing and point estimate in which real output, exchange rate, and trade balance reach its peak in the first 36 months following a monetary policy shock, and the latter effect exhibits the length of time the effect of the shock on the variables lasting. In the full sample estimation, the income-absorption effect outweighs the expenditure-switching effect and hence, the trade balance improves. The responses of the trade balance over the post-crisis periods are quite similar to those in the full sample. On the other hand, the reverse domination effect is happened over the pre-crisis episodes and the trade balance deteriorates.

Table 4. Summary of responses to shock in the monetary policy shock

	Level peak effect		Persistent effect	
	Timing	Point estimate	Lasting time	Sign
Full sample				
Y	20	-0.0102	>36	-
ER	8	0.0432	34	+
TB	30	0.0144	>36	+
Pre-crisis				
Y	13	-0.0003	16	-
ER	6	0.0079	13	+
TB	3	-0.0080	6	-
Post-crisis				
Y	7	-0.0068	18	-
ER	11	0.0104	>36	+
TB	5	0.0061	11	+

Note: The horizon of the level peak effect and persistent effect is within the first 36 months. *Y* represents the real output; *ER* the exchange rate; and *TB* trade balance.

For full sample estimation, the Asian crisis may play a role in the outcomes of interest rate impact to trade balance. During the crisis period, a huge rise in interest rate to defend the currency deteriorates real output substantially. The overall negative effect to real output after a monetary contraction in the full sample (lasting for more than 36 months) is far more persistent in comparison to the pre-crisis estimation (lasting for 13 months). In addition, the level peak effect of real output in the full sample reaches at the 20-th month with point estimate -0.0102 while in the pre-crisis the same variable touches its peak level at 13-th month with point estimate -0.0003. Therefore, negative real output effect in the full sample is considerably stronger to outweigh the exchange rate appreciation effect and resulted in the persistent improvement in the trade balance lasting for more than 36 months with peak level at 30-th month and point estimate 0.0144, compared to the pre-crisis estimation. The estimation results over the post-crisis sample shows that trade balance improves after a monetary shock. This result is line with the full sample estimation. The slow recovery of the Indonesia economy after the hardest hit by the Asian crisis compared to other Asian countries may be responsible to explain stronger effect of income-absorption over the expenditure-switching effect. In addition, since the sample that begins in 1998 is very short and two lags are used in the estimation of the VAR system, the results are only weakly sensible unlike the pre-crisis estimations which is considered to be relatively tranquility sample periods.

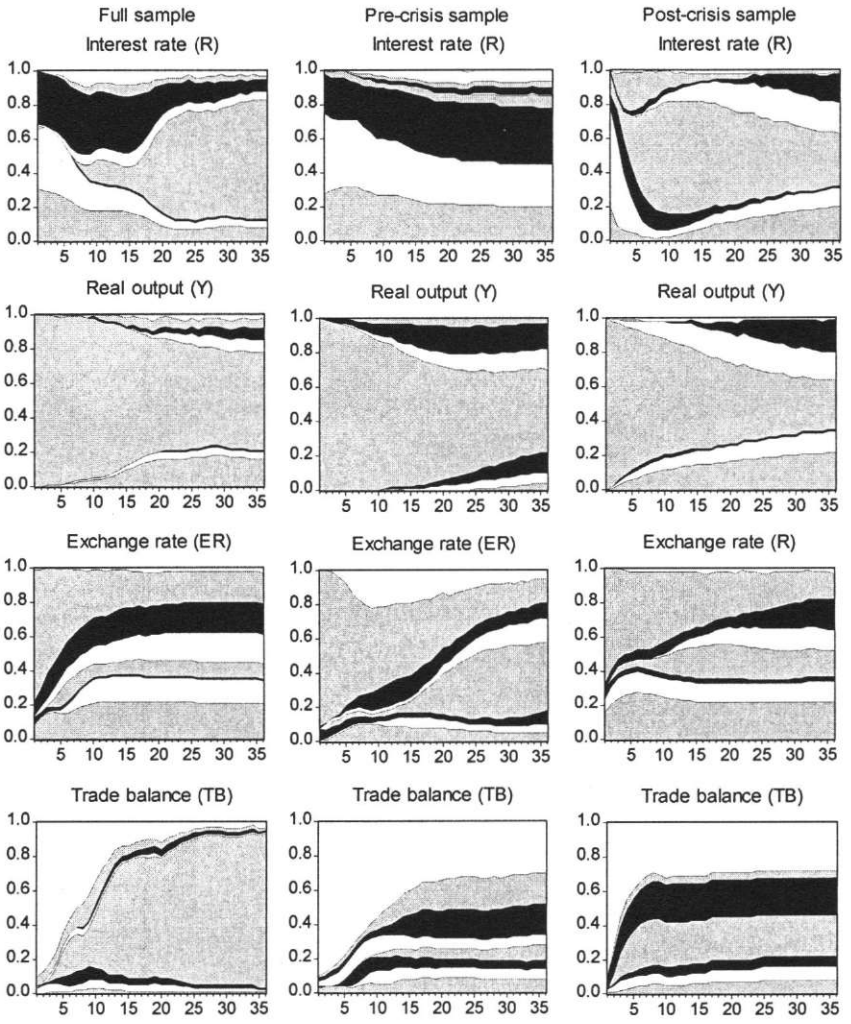
4.2. Variance decompositions

Apart from the impulse-response patterns focused on the impact of structural shocks on the economies outlined above, one can represent the variance decompositions of the models that can be used as detection for the source of variance of the endogenous variables. First, we emphasize the sources of variation in the interest rate for the model with all sample sizes, since this variance decomposition presents information on the nature of monetary policy. Next, we discuss the variance decomposition of real output, exchange rate, and trade balance. The first two decompositions are important since both variables play a role in justifying dynamic behavior of trade balance after a monetary policy shock. Figure 4 shows variance decompositions of the SVAR using full sample, pre-crisis, and post-crisis sample, respectively. The figures indicate for each of the variables the part of the forecast error variance that can be attributed, from bottom to top, interest rate, money supply, price, real output, the US Federal Funds rate, the world commodity price, exchange rate, and trade balance.

The variance decomposition for the interest rate gives us an idea of the variables in the model that the money authority reacts to most when setting interest rates. For full sample estimation, money supply and the commodity price are largely important in the beginning of the period, although as the horizon increases the other variable, real output, grows in importance. The largest proportion of variance in interest rate after three years is attributable to the real output (70%). The variance in real output and exchange rate is largely self-determined and the interest rate becomes more important as horizon increases. The variance decomposition of the trade balance is attributable to its own shocks at the beginning of periods and the attribute is taken over by the real output as horizon progresses accounted for 90% in the 36-th month.

During the pre-crisis episode, the evidence shows that the variance in the interest rate depends on money supply and price accounted for 25% and 33% at the 36-th month, respectively. Real output is largely determined by its own shocks. The variance in the exchange rate depends on the exchange rate shocks themselves lasting at 22-th month and real output becomes dominant part in explaining exchange rate variance starting at 23-th month. The contributions of its own shocks to the fluctuations in trade balance are substantial followed by the exchange rate and the commodity price. For post-crisis estimation, the story is line with the full sample model, except the variance of trade balance. The analysis of variance decomposition confirms that a large portion of trade balance variance at the first period is contributed by its own shocks. The role of the exchange rate and the commodity price become stronger as horizon increases.

Figure 5. Forecast error variance decompositions



Note: In each figure the portion of forecast error variance due to a shock, from bottom to top, interest rate, money supply, price, real output, the US Federal Funds rate, the world commodity price, exchange rate, and trade balance.

It is interesting to note that the story of the exchange rate and real output shocks to the variance decomposition of trade balance reinforce the dynamic response story of trade balance after a monetary contraction discussed in the Section 4.1. The highest portion of real output shocks to trade balance variance for full sample model reflect stronger effect of income-absorption over the expenditure-switching effect resulting in trade balance improvement. For the pre-crisis model, the same story is occurred with commodity price also plays a role in contributing trade balance variance. However, the worsening of trade balance for the pre-crisis estimation may constitute to considerably modest contribution of exchange rate shocks to trade balance. Another interesting finding is that it is important to consider commodity price in the studies on monetary policy and trade balance since the commodity price contributes quite modest in the variance of trade balance in the pre- and post-crisis model.

5. CONCLUSION

This paper examined the effects of monetary policy shocks on trade balance for Indonesia. Using monthly data on eight variables, we estimated the SVAR model for three sample size specifications–full sample, pre-crisis, and post-crisis episodes–to investigate the issue. Evidence from impulse responses for the full sample model shows that a monetary contraction worsens the trade balance. The income-absorption effect is present dominate the expenditure-switching effect. This result is in line with post-crisis estimation. However, the reverse domination effect is occurred for the pre-crisis sample and resulted in improvement of the trade balance. Impulse responses of other related variables are consistent with economic theory for all models, except we find price puzzle.

The results from the forecast error variance decomposition of the trade balance reveal that the real output is an important attribute, while the exchange rate plays a role in the trade balance variance for the pre-crisis and post-crisis models. We also found that the commodity price takes a crucial source in explaining the trade balance fluctuation.

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