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Base Money and Exchange Rate: Sources of Inflation in Indonesia during the Post-1997 Financial Crisis

Reza Y. Siregar
University of Adelaide, Australia

Gulasekaran Rajaguru
Bond University, Australia

Abstract

The magnitude of the rise in inflation rate in Indonesia during the height of the 1997 financial crisis was among the sharpest that the East Asian economies has ever witnessed in the recent decades. This paper empirically tests the monetary hypotheses of inflation and compares and contrasts the sources of price changes during the pre- and post-1997 financial crisis. We find a high explanatory power of the monetary model for the post-crisis period, but not for the pre-crisis. The high volatilities of the local currency and the unprecedented rapid growth rate of base money during the post-crisis are found to be the two key monetary determinants of the inflation in the country.

- **JEL Classification:** E41, E51, E52, E58, F41
- **Key words:** Inflation, Base Money, Expected Depreciation, Exchange Rate Policy, East Asian Financial Crisis

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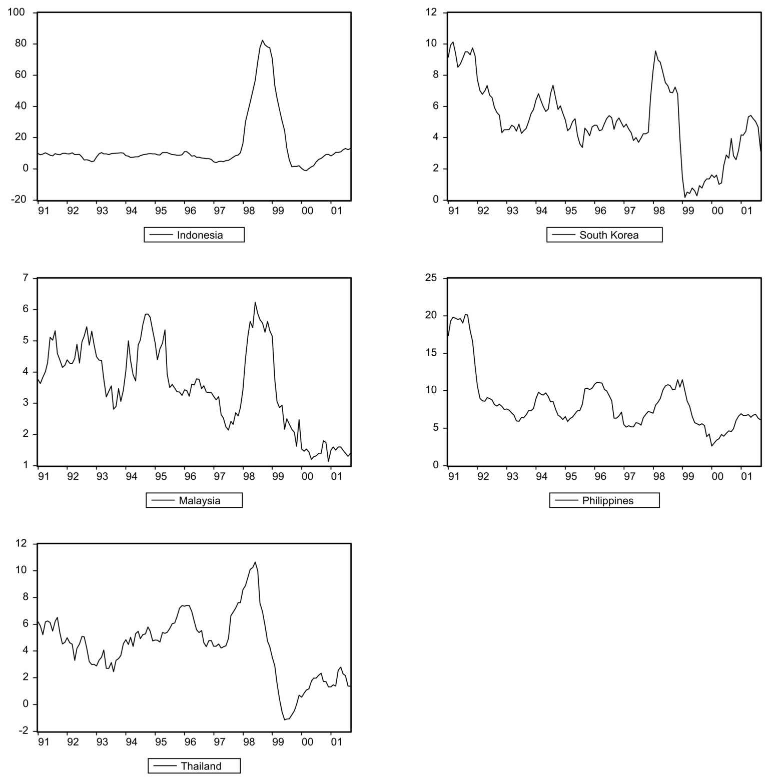
*Corresponding address: Reza Y. Siregar, School of Economics, University of Adelaide, Australia, E-mail: reza.siregar@adelaide.edu.au. The authors wish to thank Ross McLeod, an anonymous referee of this journal and the participants of Australian Conference of Economists, Adelaide (September 30 October 4, 2002) for insightful discussions on various parts of the paper. Data supports from the Econometrics Study Unit at the National University of Singapore are greatly appreciated. Usual disclaimers apply.

I. Introduction

One of the most striking and consistent features of financial crisis has been the considerable inflationary pressure that plagued the effected nations (Darrat, 1985). In this respect, the financial crisis of 1997 in East Asia is no exception. The five most crisis-effected economies of East Asia (Indonesia, Korea, Malaysia, Philippines and Thailand) experienced a rise in their domestic general price levels (Figure 1). During the peak of the crisis in 1998, these economies have seen their annual inflation rates to rise by at least four to six percentage points from their average rates in 1996, with the exception of the Philippines and Indonesia.

The magnitude of the rise in Inflation in Indonesia however was among the sharpest that the East Asian economies has ever witnessed in the recent decades. During the height of the crisis in 1998, the average annual inflation rate in Indonesia was around 58 percent, with its highest rate of close to 80 percent. After falling to a relatively low rate in early 2000, inflation rate has increased

Figure 1. Inflation Rate of Selected East Asian Crisis Effected Economies, January 1991 - September 2001 (Year on Year, in %).

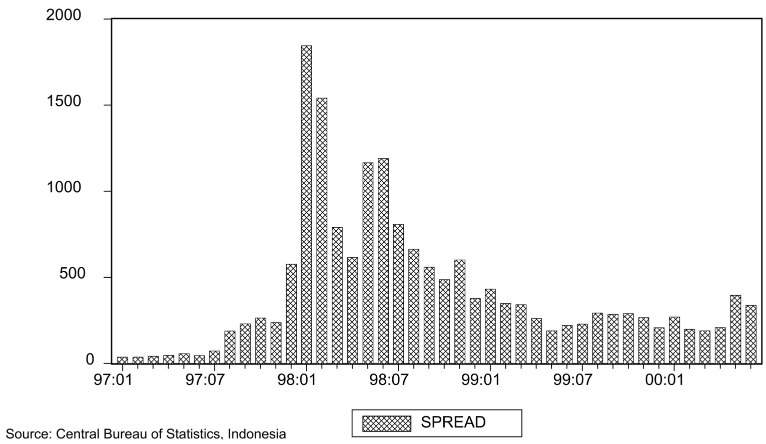


significantly in late 2000 and hovered between 12 to 15 percent in late 2001. In comparison to the chronic or acute inflation rate of many Latin American countries, the inflation rate in Indonesia during the post-1997 financial crisis has been relatively moderate. However for a country that has been committed to and successfully kept its annual inflation rate at a single digit, the 1998 price increase was in fact the worst inflation that Indonesia had experienced in nearly 30 years.[1]

In a recent report, the government of Indonesia and the International Monetary Fund (IMF) have underlined the instability of rupiah and the overall rise in uncertainties in the foreign exchange market among the root causes of the strong inflationary pressures during the last few years (IMF (2002)). Reflecting the significant level of uncertainties in the foreign exchange market, the spread between the buying and the selling rate of the rupiah nominal exchange rate against the US dollar widened from less than Rp.100 during the first few months of 1997 to more than Rp1500 on February 1998 (Figure 2). Accordingly, among the cornerstones of monetary policy agreed between IMF and the government of Indonesia, particularly since the second IMF Letter of Intents (LOIs) signed on January 15, 1998, were to stabilize the exchange rate and to contain the inflationary impact of the large depreciation of rupiah against the US dollar (Soesastro and Basri, 1998, pp.40).

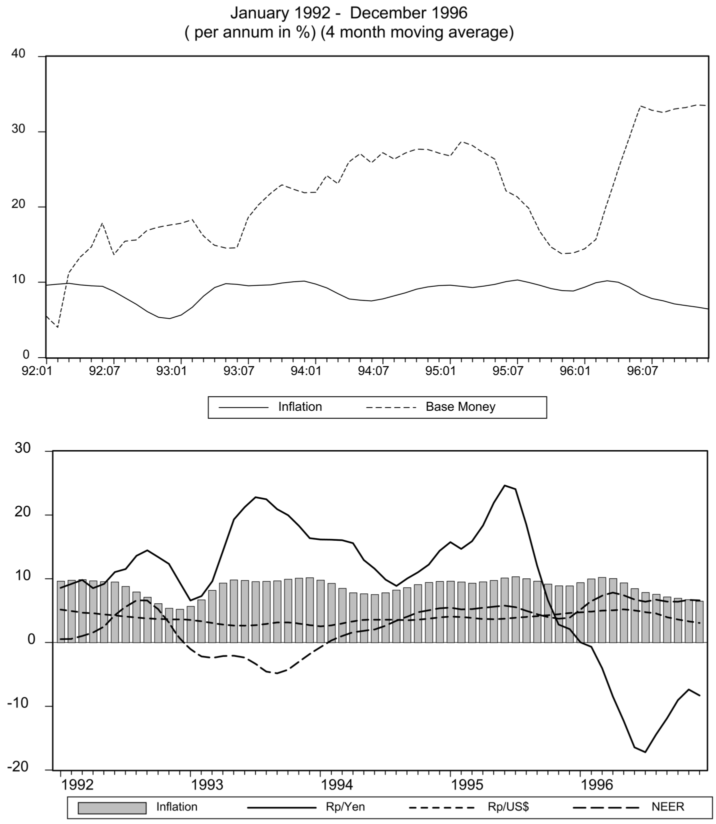
The other cornerstone of the Letter of Intents between IMF and the government

Figure 2. Monthly Spread Between Selling and Buying Exchange Rate of Rupiah Against the US dollar (in Unit of Rupiah).



[1]In fact, from 1990 to 1996, the monetary authority of Indonesia had successfully managed to keep inflation at an annual average of less than 9 percent.

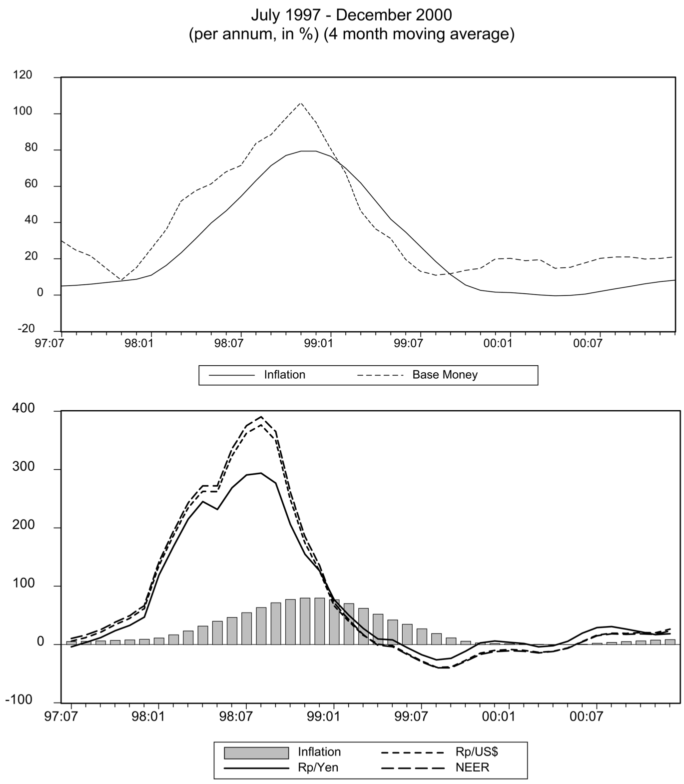
Figure 3. Inflation Rate, Growth Rate of Base Money and Change in Nominal Rupiah, January 1992 - December 1996 (per annum in %) (4 month moving average).



of Indonesia is monetary control. The agreement stipulates the limits on broad money (M2) growth, to be achieved through controlling base money (M0) quarterly growths. In late 1997 and early 1998, Indonesia had experienced a rapid growth in the base money due to the liquidity supports provided to troubled banks and the impact of depositor runs on banks (Figure 4). In its third agreement with the IMF signed on March 1998, the government of Indonesia acknowledged the necessary tight monetary policy to reduce inflationary pressures in the economy (Johnson, 1998).

Amid sustained overshooting of the central banks targets for base money growth and volatile rupiah until late 2001, the effectiveness of the monetary policy has, however, continued to be on the center of debates among the academics, policy makers and members of parliament in the country (Siregar (2001)). IMF(2002) shows that the level of base money and the inflation rate in December 2001 was

Figure 4. Inflation Rate, Growth Rate of Base Money and Change in Nominal Rupiah, July 1997 - December 2000 (per annum, in %) (4 month moving average).



about 5.1 percent and 4.7 percent higher than its target set in April 2001, respectively. Looking at Figure 4, the evidences clearly indicate that inflation has indeed reflected closely the growth rates of base money since early 1998.[2] The acceleration in the inflation rate emerged only within few months after the expansion of base money. We can also trace close co-movements between inflation and the fluctuations of rupiah, albeit not as strongly as between inflation and money supply. It is interesting to note however that there were hardly any traceable co-movements between these key monetary variables during the pre-crisis period (Figure 3).

Given the evidences presented in Figure 4, the monetarist approach to investigate the sources of inflation in Indonesia will be adopted in this paper. Among the alternative hypotheses of the sources of inflation, the monetary model

[2]McLeod (2001) shows a similar finding.

has perhaps received the most attention, particularly since early 1980s. Lim (1987), McNellis (1987) and Morrison (1987) have found bilateral exchange rates, which have been devalued occasionally, as one of the determinants of inflation in the Philippines, Latin American countries and Portugal, respectively.[3] Similarly, Calvo *et al.* (1995) find that the policy of real exchange rate targeting has led to some combinations of persistently high inflation and domestic interest rates in Brazil, Chile and Columbia. Another recent studies that have confirmed the similar conclusions are Bahmani-Oskooee and Malixi (1992) and Alba and Papell (1998). In addition to the exchange rate factor, most of those early studies also find (domestic and foreign) interest rate and money supply as the other two key monetary factors that have contributed significantly to the inflationary pressure.

With the exception of Rana and Dowling (1985), a limited number of empirical investigations have however been done to investigate the roles of monetary variables in explaining the inflation rate in Indonesia. Adopting the model developed by Calvo *et al.* (1995), Siregar (1999) confirmed the inflationary consequence of the exchange rate regime adopted in Indonesia, but only for the period from January 1990 to July 1995. McLeod (1997), on the other hand, elaborated the adverse consequences of base money on the domestic price level in Indonesia during the pre-crisis period (up to 1995).

As for the analysis on the post-1997 financial crisis, we come to know only few studies. McLeod (2001) concludes that the inflation rate clearly reflects the pattern of base money growth from May 1997 to late 1999. A preliminary study prepared by Basri et.al. (2002) finds the evidences that both money growth and exchange rate movements have contributed to a rapid inflation rate during the post-1997 financial crisis in Indonesia. Confirming the findings of Basri et.al (2002), Ramakrishnan and Vamvakidis (2002) identify the exchange rate, foreign inflation and monetary growth as the main variables with a significant predictive power for inflation in Indonesia during the period of 1980-2000.

With the objective to further understand the sources of the recent strong inflationary pressures in Indonesia, the aim of our paper is to extend the works of the previous studies in a number of ways. Firstly, given our monthly and quarterly observation sets from 1987 to 2001, we aim to compare and contrast the sources of inflation rate during the pre-and post-crisis period. Ramakrishnan and Vamvakidis (2002) acknowledge that the post-crisis is the most relevant for understanding

[3]Latin American countries included are Argentina, Brazil, Chile, Equador, Peru, and Uruguay.

inflation process in Indonesia under the floating exchange rate regime. Yet, given their quarterly observations, they do not have enough degrees of freedom to break the observations into the pre-and post-crisis sets. We deal with this problem by constructing both quarterly and monthly sample sets. Our study comes to two contrasting conclusions on the sources of inflation during the pre- and post-crisis. These outcomes are critical for the policy making process in the country.

Secondly, most of the recent studies in Indonesia are largely empirical, with limited theoretical frameworks. In this study, we introduce a basic monetary model that encompasses a number of monetary variables (including money supply and exchange rate). The basic framework, in turn, provides a testable empirical model.

Lastly, those recent studies on inflation rate in Indonesia only consider the fluctuation of nominal exchange rate of rupiah against the US dollar. We construct three different series for the exchange rate variable, namely the bilateral nominal rate of rupiah against the US dollar, the bilateral nominal rate of rupiah against the yen, and the nominal effective exchange rate of rupiah against the currencies of the seven major world economies (the United States, Japan, Germany, the United Kingdom, Canada, France and Italy). Having these three different indicators, we are not only able to confirm the robustness of the test results. But more importantly, we also have the opportunity to analyze the implications of the exchange rate policy regimes adopted in the country on the domestic inflation.

As will be shown in this study, the soft-US dollar pegged policy, adopted during the pre-1997 financial crisis period in Indonesia, had successfully managed a stable local currency against the US dollar. However, it is not without any cost. The volatility rates of nominal exchange rate of rupiah against yen and against a basket of major currencies of seven key economies (the United States, Japan, Germany, United Kingdom, Canada, France and Italy) are found to be significantly higher than the volatility of rupiah against the US dollar.[4]

In contrast, with the adoption of a more relaxed exchange rate regime (especially against the US dollar) during the post-1997 crisis, particularly until late 1998, our empirical results find that all three series of rupiah exchange rate have become significantly more volatiles, confirming the findings of McKinnon (2000, 2001), Hernandez and Montiel (2001) and Siregar and Rajan (2003)). Given the vast

[4]Siregar and Rajan (2003) show that the rise in the volatility of rupiah (against the US dollar, the Japanese yen and the nominal effective exchange rate) has adversely affected the performance of its export sectors. A similar conclusion can be drawn for the case of Thailand baht (Rahmatsyah, Rajaguru and Siregar (2002)).

changes in the behaviors of rupiah during the pre-and post-1997 crisis, it is therefore important to investigate any inflationary consequences of the three different series of rupiah exchange rate to ensure the robustness of our test results and to analyze possible consequences of different exchange rate regimes on the domestic price.

The outline of the paper is as follows. Next section will present basic theoretical frameworks of the empirical model. Section 3 discusses the relevant empirical tests and findings. To understand further the two key monetary roots of the domestic inflation, section four analyzes more closely the fluctuations of rupiah and base money during the post-1997 crisis. Brief concluding remarks section ends the paper.

II. Working Model

Monetarists advocate that the rate of inflation (Δp_t) should equal the growth rate of the nominal money supply (Δm_t^s) minus the growth rate of real money demand ($\Delta(m^d/p)$) (Abel and Bernanke (2001), Deme and Fayissa (1995) and Darrat and Arize (1990)).

$$\Delta p = (\Delta m_t^s) - \left(\Delta \frac{m^d}{p}\right)_t \tag{1}$$

All variables are in the logarithmic forms. Δ denotes the first difference operation, and t captures time.

The basic real money demand function can be expressed as the following:

$$\left(\frac{m^d}{p}\right)_t = f(y_t, r_t) \tag{2}$$

That is real money demand is a function of income (y_t) and prevailing domestic interest rate (r_t).[5] However recent studies have shown that in an open and financially liberalized economy, the impacts of external factors in the demand for money are found to be significant (Arango and Nadiri (1981), Girton and Roper (1981), Miles (1981), Bordo and Choudhri (1982), Cuddington (1983), Khalid (1999) and Sriram (2001)). To incorporate the external factors, we follow early studies and specify the following simple real money demand function.

[5]For a good review of money demand, please refer to Chapter 3 of McCallum (1989).

$$\left(\frac{m^d}{p}\right)_t = f(y_t, r_t, rf_t, ed_t) \tag{3}$$

where: (ed_t) is the expected depreciation rate of the local currency. It is proxied as the actual depreciation of the local currency during the last period.[6] (rf_t) is the foreign interest rate variable.

Substituting equation (3) into equation (1) will yield the following general expression for domestic inflation:

$$\Delta p_t = f(\Delta y_t, \Delta r_t, \Delta rf_t, \Delta ed_t, \Delta m_t^s) \tag{4}$$

Equation (4) suggests that the level of domestic inflation is going to be influenced by the changes in the level of domestic income, domestic and foreign interest rates, expected depreciation of the local currency (the exchange rate factor) and domestic money supply.

The following first order conditions should hold.

$$\frac{\partial \Delta p_t}{\partial \Delta ed_t} > 0 \tag{5}$$

Given no other changes, a rise in (ed_t) lowers money demand. Therefore, there will be a relatively higher supply of money than demand for money in the domestic economy. Inflation is therefore expected to rise (Equation 1).

$$\frac{\partial \Delta p_t}{\partial \Delta rf_t} > 0 \tag{6}$$

Similarly, a rise in foreign interest rate (rf_t) will lower demand for money in domestic economy, as the opportunity cost of holding money increases. Given everything else in the economy remains unchanged, price level is expected to rise.

$$\frac{\partial \Delta p_t}{\partial \Delta y_t} < 0 \tag{7}$$

The rise in output / income should increase demand for money (Equation 2). Given money supply remains unchanged, the rise in the level of money demand relative to money supply will lead to a decline in inflation rate (Equation 1). Hence, a rise in output will eventually cause inflation rate to decline.

$$\frac{\partial \Delta p_t}{\partial \Delta r_t} > 0 \tag{8}$$

[6](ed_t) is positive (negative) if there was a depreciation (appreciation) of the local currency last period.

A rise in the domestic interest rate will increase the opportunity cost of holding money, hence demand for money should fall (Equation 2). With the supply of money unchanged, the fall in money demand should increase domestic inflation (Equation 1).

$$\frac{\partial \Delta p_t}{\partial \Delta m_t^s} > 0 \tag{9}$$

Lastly, as clearly indicated by Equation 1, an increase in money supply, given everything else remains unchanged, should lead to a higher domestic inflation.

III. Data and Empirical Testing

A. Data

Variable *ed* represents the expected depreciation of rupiah against the US dollar and the Japanese yen, and the nominal effective exchange rate. *ed* at time (t) is represented as the actual change of the bilateral and nominal effective exchange rate at time (t-1). A positive *ed* implies a depreciation of rupiah against the major global currencies (and vice versa). The bilateral nominal exchange rates are adopted from the International Financial Statistics, the International Monetary Fund for various years.

The nominal effective exchange rate (*neer*) is a GDP-weighted of seven major world economies currencies against rupiah. As stated, those countries are the United States, Japan, Germany, United Kingdom, Canada, France and Italy.

Each weight is the ratio of each countrys annual GDP over the total sum of the seven countries GDPs. The nominal effective exchange rate is the total sum of the bilateral nominal exchange rate of each major currency multiplied by its own GDP-weight. The GDP series and the bilateral nominal exchange rate series are adopted from the International Financial Statistics, the International Monetary Fund for various years.

The base money series (m^s) and the nominal domestic interest rate are gathered from the database of Bank Indonesia. For the domestic interest rate (*r*), we adopted the 3 months rate of the Certificate of Bank Indonesia. The nominal foreign interest rate is the US three months deposit rate, taken from the International Financial Statistics CD-ROM.

The income variable for Indonesia is the real GDP of the country. This series is from the database of the Econometrics Study Unit of the National University of Singapore.

Inflation rate is calculated as the change in the consumer price index (CPI). The CPI series is sourced from the International Financial Statistics CD-ROM.

B. Unit Root Testing

It is well known that the data generating process for most macroeconomic time series are characterised by unit roots, which puts the use of standard econometric methods under question. Therefore, it is important to analyse the time series properties of the data in order to avoid the spurious results. To ensure the robustness of the test results, three most commonly used unit-root tests are applied here, namely the Augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and KPSS unit root tests on the relevant variables. The test for integration based on the ADF test involves formulating the ADF regression (Dickey and Fuller, 1979).

$$\Delta z_t = \mu + \gamma t + \rho z_{t-1} + \sum_{i=1}^{p} \delta_i \Delta z_{t+i} + \varepsilon \tag{10}$$

where t is the time trend. The lag length p is selected to ensure that the residuals are white noise. In essence, the test of whether the variable z_t is non-stationary is equivalent to the test of the significance of ρ, i.e., H_0: ρ=0, in equation (10). Alternatively, Phillips and Perron (1988) proposed a nonparametric method of controlling for higher-order serial correlation in the series. The test regression for the PP (Phillips and Perron) test is the AR(1) process:

$$\Delta z_t = \mu + \rho z_{t-1} + \varepsilon_t \tag{11}$$

While the ADF test corrects for higher order serial correlation by adding lagged differenced terms on the right hand side of equation (11), the PP test makes the correction to the t-statistic of the r coefficient from the AR(1) regression to account for the serial correlation. As opposed to both ADF and PP tests for which the test statistic is constructed under the null hypothesis that the series is non-stationary, Kwiatkowski *et al*., 1992 (KPSS) propose the test procedure in which the null hypothesis is stationary. The KPSS test is based on the following model:

$$z_t = \xi t + r_t + \varepsilon_t \tag{12}$$

where: ε_t is a stationary random error and r_t is a random walk: $r_t = r_{t-1} + u_t$. The initial value r_0 is treated as fixed and it serves the role of an intercept. The stationary hypothesis is that the variance of the residuals in the random walk component (u_t) is zero. These three tests make a good combination, as the null hypothesis in one test is the alternative hypothesis in the other one. Note, in all of these tests, the number of lags is determined by the Akaike Information Criteria (AIC) and the Schwarz Criteria (SC).

The unit root test results on the log-forms of the relevant variables are reported in Table 1. Note here, given the availability of the data series, we test output (y) variable based on quarterly observations. As for the rest of the variables, we apply the monthly series. We break the monthly observation set into pre-and post-crisis

Table 1. Unit Root Test Results: January 1987 June 1997 (Post-crisis period (Monthly Data Base)).

		ADF statistics		PP test		KPSS			
Series		Test statistic	Lags	Test statistic	Lags	Test statistic	Lags	Test type	Order of integration
p	Level	-3.02	1	-2.82	4	0.24***	4	trend & drift	I(1)
	First difference	-6.67***	3	-9.58***	4	0.06	4	with drift	
ed1 (against US$)	Level	-5.38***	3	-10.69***	4	0.20	4	with drift	I(0)
ed2 (against the yen)	Level	-7.76***	3	-7.68***	4	0.11	4	With drift	I(0)
ed3 (neer)	Level	-9.20***	0	-9.16***	4	0.15	4	With drift	I(0)
r	Level	-1.63	1	-2.01	4	0.58**	4	with drift	I(1)
	First difference	-8.78***	1	-14.57***	4	0.09	4	no drift	
r_f	Level	-2.12	8	-1.30	4	1.07***	4	with drift	I(1)
	First difference	-7.91***	0	-7.99***	4	0.21	4	no drift	
m0	Level	-1.16	2	-2.26	4	0.57***	4	trend & drift	I(1)
	First difference	-11.78***	1	-19.58***	4	0.34	4	with drift	

Note:
1) ***, ** and * represents the rejection of null at 1%, 5% and 10% levels of significance respectively
2) Lag lengths for the ADF test regression is choosen such that Akaike Information Criteria (AIC) or the Schwarz Criteria (SC) is minimized.
3) Truncation lag to evaluate the serial correlation for the Newey-West correction for both PP and KPSS test is computed by $q = floor(4/(T/100)^{2/9})$.

Table 1 (contd). Unit Root Test Results: July 1997 December 2001 (Post-crisis period (Monthly Data Base)).

Series		ADF statistics Test statistic	Lags	PP test Test statistic	Lags	KPSS Test statistic	Lags	Test type	Order of integration
p	Level	-3.02	1	-2.82	4	2.23***	4	trend & drift	I(1)
	First difference	-6.67***	3	-9.58***	4	0.34	4	with drift	
ed1 (against US$)	Level	-5.38***	3	-10.69***	4	0.25	4	with drift	I(0)
ed2 (against the yen)	Level	-5.36***	1	-5.98***	4	0.32	4	With drift	I(0)
ed3 (neer)	Level	-9.20***	0	-9.16***	4	0.25	4	With drift	I(0)
r	Level	-1.63	1	-2.01	4	0.62**	4	with drift	I(1)
	First difference	-8.78***	1	-14.57***	4	0.20	4	no drift	
r_f	Level	-2.12	8	-1.30	4	0.39*	3	with drift	I(1)
	First difference	-7.91***	0	-7.99***	4	0.24	4	no drift	
m0	Level	-1.16	2	-2.26	4	0.19**	4	trend & drift	I(1)
	First difference	-11.78***	1	-19.58***	4	0.13	4	with drift	

Note:

4) ***, ** and * represents the rejection of null at 1%, 5% and 10% levels of significance respectively
5) Lag lengths for the ADF test regression is choosen such that Akaike Information Criteria (AIC) or the Schwarz Criteria (SC) is minimized.
6) Truncation lag to evaluate the serial correlation for the Newey-West correction for both PP and KPSS test is computed by $q = floor(4/(T/100)^{2/9})$.

Table 1 (contd). Unit Root Test Results: Quarter 1, 1987 Quarter 1, 1997 (Pre-crisis period)

Series		ADF statistics Test statistic	Lags	PP test Test statistic	Lags	KPSS Test statistic	Lags	Test type	Order of integration
y	Level	-2.31	1	-2.79	3	0.15	2	With drift	I(1)
	First difference	-2.86***	0	-2.88***	2	0.19	2	no drift	

Note:

7) ***, ** and * represents the rejection of null at 1%, 5% and 10% levels of significance respectively
8) Lag lengths for the ADF test regression is choosen such that Akaike Information Criteria (AIC) or the Schwarz Criteria (SC) is minimized.
9) Truncation lag to evaluate the serial correlation for the Newey-West correction for both PP and KPSS test is computed by $q = floor(4/(T/100)^{2/9})$.

periods and avoid structural breaks on the series associated with the transition from the pre- to post-crisis period. Accordingly, we have applied the unit-root tests for these periods separately as opposed to the unit-root tests in the presence of structural breaks. For the quarterly data, we focus only on the pre-crisis. The use of monthly or quarterly series should not affect the final outcomes of the tests (Pierse and Snell (1995), and Marcellino (1999)).

The results reveal that the variables such as rate of inflation (Δp), and the expected depreciation of the local currency do not contain unit roots and thus stationary processes. But output (y), domestic interest rate (r), foreign interest rate (rf) and the money supply ($m0$) are found to be an I(1) series (non-stationary processes at the level and stationary processes at the first difference).

C. Short Run Dynamics

The main focus of this study is to analyze the role of various macroeconomic policies on inflation rate. Given the unit-root test results, it is irrelevant to examine the long-run equilibrium relationship between the variables, as the dependent variable is stationary. Furthermore, due to the asymptotic properties of the Johansen cointegration test statistics, the application of the Johansen cointegration test on a limited sample size has been frequently argued to be unstable. Early studies, such as Sephton and Larsen (1991), Barkoulas and Baum (1997), Cheung and Lai (1993), Choi (1992), Choi and Chung (1995), Lahiri and Mamingi (1995) showed that the Johansen test statistics are biased toward finding cointegration too often.[7]

Given the stability problem of the Johansen test under the limited sample periods as we have for the pre- and post-1997 crisis (of around 9 years for the pre-crisis and 5 years for the post-crisis), we therefore proceed with the employment of the Autoregressive Distributed Lag model (ARDL), without incorporating any error correction component in the regression.[8] The Autoregressive Distributed lag model (ARDL) has been constructed by treating inflation as an endogenous variable. The non-stationary explanatory variables are differenced appropriately to remove the unit roots. Hence our working model, based on the unit-root test

[7]Sephton and Larsen (1991), SL henceforth, showed that inference based on Johansen cointegration tests of foreign exchange market efficiency suffers from structural instability. Burkoulas and Baum (1997) re-examine the evidence found in SL to longer data sets. Instead of using a data set of less than 10 years, Burkoulas and Baum (1997) expand the sample up to 20 years and conclude that the stability of the test results increases as a longer observation period included in the test.

results, will be as follows.

$$\Delta p_t = a + \sum_{i=0} \alpha_i \Delta m^s_{t-i} + \sum_{i=0} \beta_i \Delta y_{t-i} + \sum_{i=0} \delta_i ed_{t-i} + \sum_{i=0} \theta_i \Delta r_{r-i} + \sum_{i=0} \gamma_i \Delta rf_{t-i} + \varepsilon_t \quad (13)$$

$$\sum_{i=0} \alpha_i > 0, \ \sum_{i=0} \delta_i > 0, \ \sum_{i=0} \theta_i > 0, \ \sum_{i=0} \gamma_i > 0 \text{, and } \sum_{i=0} \beta_i > 0$$

The expected signs of the coefficient estimates are consistent with Equations 5-9. *a* and ε are a constant and an error term variable, respectively. We assume that the error term to be a white noise process. Δ denotes the first difference operation, and all the variables are in the log-forms.

Up to eight lags for the monthly observations and four lags for the quarterly observations of the dependent variables are included in the initial estimation, and then sequentially we exclude the statistically insignificant lags of the variables.[9] Two different regressions are estimated to establish the role of macroeconomic polices on rate of inflation: (1) pre-crisis period and (2) post-crisis period. The pre-crisis regressions are done in both quarterly and monthly data.[10] The model for the post-crisis period is based on monthly data as the use of quarterly data for the post-crisis period highly suffers from the lack of degrees of freedom. Furthermore, since the output variable is not available in the monthly frequencies we omit them from the post-crisis analysis.

Table 2-4 report the overall results. We find the signs of the estimated coefficients are in general consistent with the theory discussed in section 2, except for the output variable for the pre-crisis.[11] The diagnostic statistics, including the R^2 statistics adjusted for degrees of freedom, the Durbin-Watson (DW), the F-statistics (and its probability), and the Engles ARCH test for heteroscedasticity, are

[8]Just for the sake of completeness, we test for the cointegration relationship of the pre-crisis model with variables all at levels. We find two cointegration relationships. Note, given the exchange rate factor (ed) is I(0), at least 2 cointegration relationships should be found to confirm the long-run relationship (Johansen and Juselius (1992), and Rahmatsyah, Rajaguru and Siregar (2002)). Some of the the normalized cointegration coefficients of the Johansen test are however theoretically inconsistent and statistically insignificant. Due to the stability problem discussed above and for the sake of brevity, we do not post the cointegration test results and the full ARDL with the error correction component. However the test results can be made available upon request to the authors.

[9]The numbers of lags are chosen to ensure that we have enough degrees of freedom. Our test results have shown that no significant results are found beyond the lags that we have imposed.

[10]The pre-crisis covers the period of quarter 1, 1987 to quarter 2, 1997, or January 1987 June 1997. Availability of data dictates our sample observations. The post-crisis set includes observations from July 1997 to December 2001.

Table 2. ARDL Results (With the expected depreciation of bilateral nominal exchange rate of rupiah against the US dollar).

1). ARDL Results : Quarter 1, 1987 Quarter 1, 1997 (Pre-crisis)

$$\Delta p_t = -0.31\Delta p_{t-2} + 0.05ed_{t-3} + 0.19\Delta y_{t-4} + 0.009\Delta M^s_{t-4} + 0.02$$

(0.012)*** (0.01)*** (0.01)** (0.003)*** (0.004)***

R-squared= 0.23, DW=2.17; F-stat=2.81; Prob(F-stat) = 0.000; ARCH (Prob) = 0.83

2). ARDL Results : January 1987 - June 1997 (Pre--crisis)

$$\Delta p_t = 0.228\Delta p_{t-1} - 0.207\Delta p_{t-2} + 0.198\Delta p_{t-3} - 0.269\Delta p_{t-4} + 0.018\Delta M^8_{t-2} - 0.007$$

(0.089)** (0.089)** (0.089)** (0.084)*** (0.009)* (0.001)***

R-squared= 0.17, DW=1.99; F-stat=4.40; Prob(F-stat) = 0.001; ARCH (Prob) = 0.84

3). ARDL Results : July 1997 - December 2001 (Post-crisis)

$$\Delta p_t = 0.255\Delta p_{t-4} + 0.054\Delta r_{t-2} + 0.058\Delta M_{t-8} + 0.158\Delta M^s_t + 0.162\Delta M^s_{t-2} + 0.052\Delta M^8_{t-3}$$

(0.089)*** (0.025)** (0.023)** (0.023)*** (0.028)*** (0.027)*

$$0.037\Delta ed_{t-1} + 0.024\Delta ed_{t-2}$$

(0.014)** (0.013)**

R-squared= 0.74, DW=1.96; F-stat= 17.79; Prob(F-stat) = 0.000; ARCH (Prob) = 0.903

Note: ( ) is standard error; *Significant at 10%; **Significant at 5%; ***Significant at 1%; DW= Durbin-Watson

Table 3. ARDL Results (With the expected depreciation of bilateral nominal exchange rate of rupiah against the Japanese Yen).

1). **ARDL Results : Quarter 1, 1987 Quarter 1, 1997 (Pre-crisis)**

$$\Delta p_t = -0.30\Delta p_{t-2} + 0.07ed_{t-3} + 0.21\Delta y_{t-4} + 0.018\Delta M^s_{t-4} + 0.02$$

(0.013)*** (0.02)*** (0.09)** (0.007)*** (0.004)***

R-squared= 0.33, DW=2.04; F-stat=4.58; Prob(F-stat) = 0.000; ARCH (Prob) = 0.36

2). **ARDL Results : January 1987 - June 1997 (Pre-crisis)**

$$\Delta p_t = 0.228\Delta p_{t-1} - 0.207\Delta p_{t-2} + 0.198\Delta p_{t-3} - 0.269\Delta p_{t-4} + 0.018\Delta M^8_{t-2} - 0.007$$

(0.089)** (0.089)** (0.089)** (0.084)*** (0.009)* (0.001)***

R-squared= 0.17, DW=1.99; F-stat=4.40; Prob(F-stat) = 0.00106; ARCH (Prob) = 0.84

3). **ARDL Results : July 1997 - December 2001 (Post-crisis)**

$$\Delta p_t = 0.277\Delta p_{t-1} - 0.261\Delta p_{t-2} + 0.248\Delta p_{t-4} + 0.070\Delta ed_t + 0.024\Delta ed_{t-1} + 0.038\Delta ed_{t-2} + 0.025\Delta ed_{t-3}$$

(0.129)** (0.0116)** (0.075)*** (0.009)*** (0.013)* (0.012)*** (0.009)***

$$+0.055\Delta r_{t-2} + 0.054\Delta M^s_{t-s} + 0.088\Delta M^s_{t-1} + 0.096\Delta M^s_{t-2} + 0.052\Delta M^s_{t-3} + 0.049\Delta M^s_{t-4}$$

(0.01)*** (0.015)*** (0.016)*** (0.022)*** (0.022)** (0.018)***

R-squared= 0.91, DW=2.16; F-stat=34.36; Prob(F-stat) = 0.000; ARCH (Prob) = 0.349

Note: ( ) is the standard error; * Significant at 10%; **Significant at 5%; ***Significant at 1%; DW= Durbin-Watson.

presented for each regression. The F-statistics indicate that the probability is at least 95 percent that one or more of the independent variables are non-zero. The Durbin-Watson statistics indicate that the serial correlations are not a problem in

[11]We still include the output variable in the final regression as the coefficient estimate is found to be significant.

Table 4. ARDL Results (With the expected depreciation of nominal effective exchange rate of rupiah).

1). ARDL Results : Quarter 1, 1987 Quarter 1, 1997 (Pre-crisis)

$$\Delta p_t = -0.35\Delta p_{t-2} + 0.22\Delta y_{t-4} + 0.02\Delta M^s_{t-1} + 0.07\Delta ed_t + 0.02$$

(0.13)** (0.09)** (0.007)*** (0.04)* (0.003)***

R-squared= 0.34, DW=2.09; F-stat=4.74; Prob(F-stat) = 0.003; ARCH (Prob) = 0.63

2). ARDL Results : January 1987 - June 1997 (Pre--crisis)

$$\Delta p_t = 0.228\Delta p_{t-1} - 0.207\Delta p_{t-2} + 0.198\Delta p_{t-3} - 0.269\Delta p_{t-4} + 0.018\Delta M^s_{t-2} - 0.007$$

(0.089)** (0.089)** (0.089)** (0.084)** (0.009)* (0.001)***

R-squared= 0.17, DW=1.99; F-stat=4.40; Prob(F-stat) = 0.00106; ARCH (Prob) = 0.84

3). ARDL Results : July 1997 - December 2001 (Post-crisis)

$$\Delta p_t = 0.531\Delta p_{t-1} + 0.051\Delta r_{r-1} + 0.057\Delta M^s_t + 0.101\Delta M^s_{t-1} + 0.072\Delta M^s_{t-2} + 0.051\Delta ed_{t-1}$$

(0.064)*** (0.017)** (0.017)*** (0.019)*** (0.021)*** (0.009)***

R-squared= 0.85, DW=2.14; F-stat=51.69; Prob(F-stat) = 0.000; ARCH (Prob) = 0.118

Note: ( ) is the standard error; *Significant at 10%; **Significant at 5%; ***Significant at 1%; DW= Durbin-Watson

any of the regression results. In addition, the ARCH results conclude the absence of heteroscedasticity in general.

Several key findings warrant further analysis. For the pre-crisis period, we do not find any of the interest rate variables contributes significantly to the changes in the domestic price level. The R-square is only around 30 percent, reflecting the low explanatory power of the independent variables. Furthermore, while only the quarterly tests show that expected depreciation of rupiah significantly determines the inflation rate, both the quarterly and the monthly regressions confirm the important contribution of the money supply in explaining fluctuations of the inflation rate in Indonesia at 5% and 1% significant level. Indicating the robustness of the test results, each set of regressions (with bilateral nominal exchange rates of rupiah against the US dollar and yen, and the nominal effective exchange rate) arrives at the same overall conclusion.

During the post-1997 financial crisis, each of our monthly regression results robustly confirms the significant roles of expected depreciation of rupiah, money supply and domestic interest rate in explaining changes in the inflation rate in the country. In general, we find more significant lag variables of the key explanatory variables, namely the money supply and the expected depreciation during the post-crisis. Reflecting the rise in the number of significant lag variables, the R-squares for the post-crisis period are much larger than the pre-crisis period, ranging from 70 percent to 90 percent. The big gaps between the R-squares confirm the much more significant explanatory powers of the monetary model in explaining the inflationary pressures during the post-1997 crisis than during the pre-crisis period. We will analyze further the empirical results posted in Tables 2-4 in section four of the paper. But first, the next two sub-sections of the paper will quickly review two important diagnosis tests.

3.3.1 Testing the Implicit Assumption of Exogeneity

The validity of the econometrics test results posted in the previous tables crucially depends on the implicit assumption that the right-hand side variables in Equation (13) are statistically exogenous to inflation. To test for the statistical exogeneity, we employ the one-sided procedure to test for causality in the sense of Granger (1969). This one-sided Granger causality test is chosen here from a number of alternative causality techniques in the light of the Monte Carlo evidence reported by Geweke, Meese, and Dent (1983).[12]

To be consistent with the ARDL tests, we also break the periods into pre-and post-crisis periods, and consider only the significant variables as posted in Table 2-4. Furthermore, since the Granger test is narrowly interpreted here as a test for statistical exogeneity of particular variables within a given model, it seemed more prudent to maintain the same lag specifications as in the early results shown in Table 2-4 when applying the Granger test.[13]

3.3.2 Stability Test

In addition to exogeneity test, we also conduct the commonly used Chow-stability test (Chow, 1960) for each of the regressions. Following Farley, Huinich, and McGuire (1975), we split the observation sets at its midpoint to maximize the empirical power of the test. In general, our test results confirm that our estimated equations are structurally stable. For the sake of brevity, we do not report the test results. But the results can be made available upon request.

D. Variance Decomposition

In addition to the ARDL models, we formulate the vector autoregressive models (VAR) to evaluate the variability in inflation rates by the means of key policy variables such as expected depreciations and the money supply. In the traditional

[12]The same procedure was also employed by Darrat and Arize (1990).

[13]We experimented with different lag structures and consistent overall results were obtained. From the test results, we can conclude that the implicit assumption of exogeneity for the explanatory variables is generally found to be applicable in our cases, except for the post-crisis domestic interest rate.

Table 5. Granger Causality Test
A. Quarterly Data
Pre-Crisis: Quarter 1, 1985 Quarter 2, 1997

Null Hypothesis	F-Statistics	Probability
ΔP Does not Granger Cause ΔY	1.7741	0.1681
ΔP Does not Granger Cause ΔM	0.3287	0.5693
ΔP Does not Granger Cause *ED (against the US dollar)*	0.9846	0.4282
ΔP Does not Granger Cause *ED (against the Japanese yen)*	1.2901	0.2859
ΔP Does not Granger Cause *ED (Nominal Effective Exchange Rate of Rupiah)*	0.3398	0.5628

B. Monthly Data
Pre-Crisis Sample: January 1987 June 1997

Null Hypothesis	F-Statistics	Probability
ΔP Does not Granger Cause ΔM	1.7938	0.1831

Post-Crisis: July 1997 December 2001

Null Hypothesis	F-Statistics	Probability
ΔP Does not Granger Cause ΔM	0.6211	0.5416
ΔP Does not Granger Cause Δr	2.7646	0.0522
ΔP Does not Granger Cause *ED (against the US dollar)*	0.0261	0.8722

multivariate time series framework of Sims (1980), dynamic analysis of VAR models often incorporates forecast error variance decompositions. Variance decomposition separates the variation in an endogenous variable into the component shocks to the VAR. In each of the variance decomposition test, we only include the significant explanatory variables reported in each ARDL test. The objective here is to roughly estimate further the explanatory powers of the significant independent variables listed in Table 2-4. The optimal lag length for the VAR models are determined by both AIC and SC criteria.[14] Furthermore, in order to generate the robust variance decompositions we need to ensure that the VAR residuals assumptions are satisfied. For this purpose, we have conducted the multivariate diagnostic tests to examine the VAR assumptions such as residual autocorrelations (Vector Portmanteau (8), Vector AR 1-2 test), normality and heteroskedasticity.[15] The results are reported in Table 6a-6c. The results suggest that the residuals are not autocorrelated. In addition, the vector heteroskedasticity

[14]For the sake of brevity, the estimated coefficients and the relevant test statistics of VAR models are not reported here and it can be made available from authors upon request.

[15]We have reported the test statistics for the lag length of 8. However, the results are consistent for all lags.

Table 6a. Multivariate Diagnostics Tests on VAR residuals (With the expected depreciation of bilateral nominal exchange rate of rupiah against the US dollar).

	Pre-crisis Quarterly	Precrisis Monthly	Post crisis Monthly
Vector Portmanteau (8)	84.65 (0.79)	13.86 (0.98)	82.38 (0.14)
Vector AR 1-2 test:	1.04 (0.43)	0.72 (0.68)	2.57 (0.06)
Vector Normality	13.37 (0.10)	4.83 (0.31)	9.952 (0.27)
Vector Heteroskedasticity	0.32 (1.00)	0.53 (0.96)	82.02 (0.41)

Note: Values in the parentheses are p-values

Table 6b. Multivariate Diagnostics Tests on VAR residuals (With the expected depreciation of bilateral nominal exchange rate of rupiah against the Japanese Yen).

	Pre-crisis Quarterly	Precrisis Monthly	Post crisis Monthly
Vector Portmanteau (8)	87.12 (0.73)	13.86 (0.98)	74.01 (0.97)
Vector AR 1-2 test:	1.20 (0.26)	0.72 (0.68)	0.61 (0.80)
Vector Normality	8.16 (0.42)	4.83 (0.31)	4.81 (0.77)
Vector Heteroscadsticity	0.48 (0.99)	0.53 (0.96)	81.24 (0.44)

Note: Values in the parentheses are p-values

Table 6c. Multivariate Diagnostics Tests on VAR residuals (With the expected depreciation of nominal effective exchange rate of rupiah).

	Pre-crisis Quarterly	Precrisis Monthly	Post crisis Monthly
Vector Portmanteau (8)	102.21 (0.31)	13.86 (0.98)	79.18 (0.94)
Vector AR 1-2 test:	0.98 (0.52)	0.72 (0.68)	1.05 (0.51)
Vector Normality	2.87 (0.94)	4.83 (0.31)	8.41 (0.39)
Vector Heteroscadsticity	0.54 (0.99)	0.53 (0.96)	77.24 (0.57)

Note: Values in the parentheses are p-values

test results conclude the absence of heteroskedasticity in general. Moreover, vector normality test statistics show that the residuals from VAR models follow multivariate normal distribution.

It has been criticized in the literature that the traditional orthogonal variance decomposition techniques obtained by multiplying Cholesky decomposition matrix is sensitive to ordering of the variables. To overcome this problem, we have also employed the *generalized* forecast error variance decomposition technique developed by Pesaran and Shin (1998) and Koop *et al.* (1996) which is invariant to the ordering of the variables.[16] We observed in this exercise that both techniques lead to similar conclusion. However, it is not surprising because of the weak correlations of the VAR residuals observed from the both Vector Portmanteau and

[16]For the descriptions of generalized variance decomposition techniques, see Pesaran and Shin (1998) and Ewing (2002).

Vector AR 1-2 tests reported in Table 6a-6c. The generalized forecast error variance decomposition results are posted in Table 7a-7c.[17]

Several points are worth to be highlighted. The shares of the growth rates of the money supply and the expected depreciation of rupiah in explaining the variances in the domestic inflation are very modest during the pre-1997 crisis. The statistics show that at least 82 percent (97 percent) of the quarterly (monthly) variances of the inflation can be explained by its own shocks.[18] This result reveals the importance of the inflationary inertia on inflationary dynamics in Indonesia during

Table 7a. Variance Decomposition of Δp (With the expected depreciation of bilateral nominal exchange rate of rupiah against the US dollar).

(Pre-Crisis Period on Quarterly Data)

Period	Δp	*Ed*	Δm	Δy
1	100.00	0.08	3.48	8.37
4	96.02	0.08	5.87	8.81
8	95.12	0.08	6.22	9.71
12	95.11	0.08	6.23	9.71
16	95.11	0.08	6.23	9.71
20	95.11	0.08	6.23	9.71

(Pre-Crisis Period on Monthly Data)

Period	Δp	Δm
1	100.00	1.70
4	97.41	20.53
8	97.35	20.95
12	97.35	20.95
16	97.35	20.95
20	97.35	20.95

(Post-Crisis Period on Monthly Data

Period	Δp	Δm	*ed*	Δr
1	100.00	11.94	1.63	0.86
4	48.77	38.72	11.86	0.40
8	48.52	38.37	11.81	0.59
12	48.50	38.39	11.81	0.59
16	48.50	38.39	11.81	0.59
20	48.50	38.39	11.81	0.59

[17] For the sake of brevity, the traditional variance decompositions results are not reported here.

[18] To get more insights into the proper interpretations of the test results, refer to Pesaran and Shin (1998), Koop et.al. (1996) and Ewing (2002).

Table 7b. Variance Decomposition of Δp (With the expected depreciation of bilateral nominal exchange rate of rupiah against the Japanese Yen)
(Pre-Crisis Period on Quarterly Data)

Period	Δp	*Ed*	Δm	Δy
1	100.00	6.27	0.95	8.40
4	95.84	5.98	1.43	9.31
8	94.88	6.00	1.45	10.36
12	94.83	6.00	1.44	10.38
16	94.83	6.00	1.44	10.38
20	94.83	6.00	1.44	10.38

(Pre-Crisis Period on Monthly Data)

Period	Δp	Δm
1	100.00	1.70
4	97.41	20.53
8	97.35	20.95
12	97.35	20.95
16	97.35	20.95
20	97.35	20.95

(Post-Crisis Period on Monthly Data)

Period	Δp	Δm	ed	Δr
1	100.00	13.55	1.82	0.49
4	55.72	40.07	10.45	0.25
8	55.91	39.65	10.27	0.44
12	55.94	39.66	10.26	0.44
16	55.94	39.67	10.26	0.44
20	55.94	39.67	10.26	0.44

the pre-crisis period. These findings are indeed consistent with the low-R squares reported in Table 2-4 on the pre-crisis regressions. Looking at Figure 3, hardly any consistent patterns between inflation and the growth rates of base money and expected depreciations of rupiah can also be traced.

Despite the significant t-statistics for the pre-crisis quarterly regressions (Table 2-4), the shares of the various estimates of the expected depreciation of rupiah (i.e. against the US dollar, the yen and the nominal effective exchange rate) are relatively small. Among the three measures of rupiah rate, we find its expected depreciation against the yen contributed the most to the inflation rate in the country. Furthermore, the VAR test results suggest that shares of output variable are more significant than the monetary variables in explaining the inflation rate. However, given the inconsistent sign of the coefficient estimate for income (Table 2-4), these results are arguably inconclusive. As for the pre-crisis monthly result,

Table 7c. Variance Decomposition of Δp (With the expected depreciation of nominal effective exchange rate of rupiah)

(Pre-Crisis Period on Quarterly Data)

Period	Δp	*Ed*	Δm	Δy
1	100.00	0.11	3.85	14.48
4	83.93	0.53	5.20	12.11
8	82.16	0.52	5.28	14.53
12	82.13	0.52	5.28	14.59
16	82.13	0.52	5.28	14.60
20	82.13	0.52	5.28	14.60

(Pre-Crisis Period on Monthly Data)

Period	Δp	Δm
1	100.00	1.70
4	97.41	20.53
8	97.35	20.95
12	97.35	20.95
16	97.35	20.95
20	97.35	20.95

(Post-Crisis Period on Monthly Data)

Period	Δp	Δm	ed	Δr
1	100.00	12.24	1.58	0.82
4	49.12	38.82	11.74	0.40
8	48.87	38.48	11.71	0.58
12	48.85	38.50	11.71	0.58
16	48.85	38.50	11.71	0.58
20	48.85	38.50	11.71	0.58

we find the shock to the base money contributed about 20 percent of the variations in the price level.

Consistent with the sharp rise in the R-squares of the ARDL tests, the combined variances of the monetary variables (growth of money, expected depreciation and the changes in the domestic interest rate) have contributed more than 50 percent of the monthly variations in the domestic inflation rate during the post-1997 crisis. In each set, we can also conclude that the role of the exogenous shock to the money supply is indeed by far the most significant one, contributing as much as 40 percent of the forecast error in the inflation rate. As for the shock to the exchange rate, we find its contribution to be significantly less, at around 10 to 11 percent. These findings are fully supported by evidences shown in Figure 4 and test results of Tables 2-4. Unfortunately, given the lack of monthly data on the output variable, we cannot access the impact of the real GDP growth rates of the economy on the inflation during the post-1997 crisis.

IV. Rupiah and Base Money

Preceding discussions have confirmed the important roles of base money and exchange rate movements in explaining inflation in Indonesia during the pre- and post-1997 financial crisis. But clearly, the test results have shown that the roles of these two factors are much more significant during the post-crisis period. To understand further the fluctuations of rupiah and base money during the volatile years of 1997-2001, we will examine briefly the recent trends of the two key monetary variables in Indonesia.

A. Rupiah

Drawing on the work of Frankel and Wei (1994), McKinnon (2000, 2001) concluded that after a temporary adoption of a more flexible regime during the height of the crisis (July 1997 to December 1998), the Southeast Asian-5 economies (Indonesia, Malaysia, the Philippines, Thailand and Singapore) had reverted to their pre-crisis US dollar soft pegged exchange rate policies since 1999. Lim (2002) extends the study to cover observations until November 2001 and confirms the McKinnon results.[19]

To further examine the recent trends of rupiah during the pre-and post-crisis period, we employ different types of ARCH models to estimate the volatility rates of the currency. The GARCH specification that we consider takes the form:

$$\ln NER_t = a_0 + a_1 \ln NER_{t-1} + a_2 dummy_t + e_t \text{, where } e_t \sim N(0,h_t) \qquad (14)$$

$$h_t = \alpha + \beta e_{t-1}^2 + \gamma h_{t-1} + \delta dummy_t + u_t. \qquad (14b)$$

Where u_t is a white noise process with $E(u_t)$=0 and $E(u_t u_\tau) = \begin{cases} \sigma_u^2 \; for \;\; t=\tau \\ 0 \;\; otherwise \end{cases}$.

(Ln*NER*) represents the nominal effective exchange rate and the bilateral nominal

[19]Hernandez and Montiel (2001), who analyse the evidence regarding post-crisis exchange rate policies pursued in the Asia-5 economies, conclude as follows. contrary to the views of some observers…there has indeed been a change in *de facto* exchange rate regimes in all five of these countries between the pre- and post-crisis periods. While none of them have adopted "soft pegs" with unfettered capital movements, neither have they moved to the extreme corner solutions of "hard" pegs or clean floats. In other words, all of them have continued to manage their exchange rates in an active manner..and have thus occupied the supposed "hollow middle" of exchange rate policy (p.16).

exchanges rate of rupiah against the US dollar and the Japanese yen in the log forms. The conditional variance equation (Eq. 14b) described above is a function of three terms: (1) the mean α; (2) news about volatility from the previous period, measured as the lag of the squared residual from the mean equation: e^2_{t-1} (the ARCH term); and (3) the last periods forecast error variance, h_{t-1} (the GARCH term). In addition, we add the dummy variable to capture the crisis period and the shift in the exchange rate policy. It is equal to zero up to July 1997, and equals to one from August 1997 to December 2001. As mentioned before, Indonesia abandoned its rigid policy in August 1997 and freed the rupiah to fluctuate.

Different types of ARCH models such as ARCH, GARCH and EGARCH models were estimated on the data. However, the GARCH(1,1) model is found to be superior in generating the volatility for the nominal rupiah against US dollar and the nominal effective exchange rate. On the other hand, ARCH(1) model is found to be superior in generating the volatility for the nominal exchange rate against the Japanese yen. The GARCH(1,1) and ARCH(1) estimates are reported in Table 8 and Figure 5.

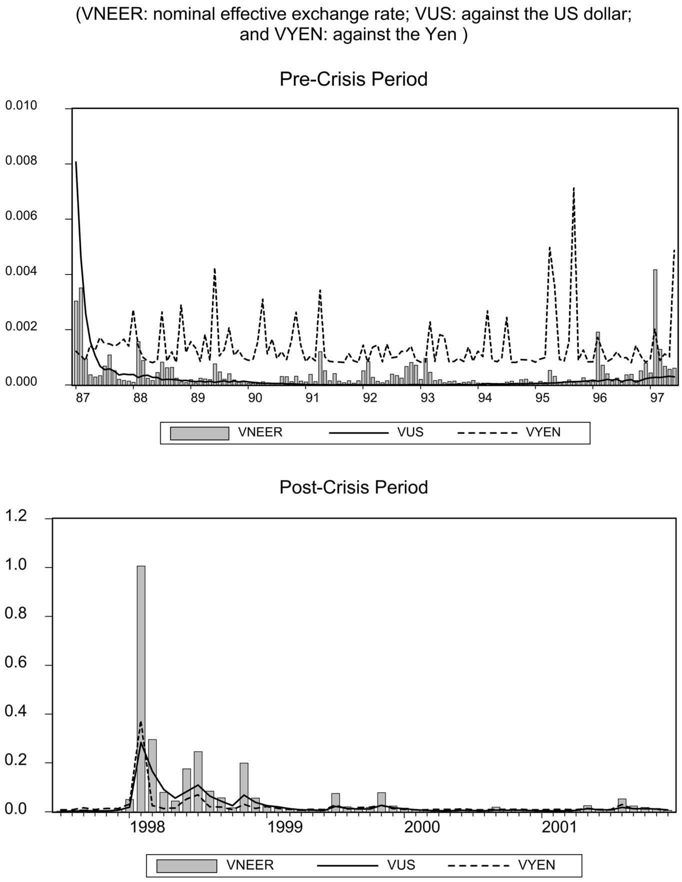
Few interesting points should be highlighted from the GARCH(1,1) and ARCH(1) results. Confirming the early findings of Hernandez and Montiel (2001), McKinnon (2000) and Lim (2002), we find the volatility of nominal rupiah against the US dollar to be very moderate during the pre-crisis. However, as in the case of Thailand baht (Rahmatsyah, Rajaguru, and Siregar (2002)), the soft-US dollar pegged policy adopted during the pre-1997 crisis has allowed substantially more

Table 8. GARCH (1,1) and ARCH(1) Volatility of Rupiah
Monthly Nominal Effective Exchange Rate and Nominal Exchange Rates against US dollars and Yen (January 1985-December 2001)

	α	β	γ	δ
Against the US dollar	0.000000339 (1.207)	0.4834 (9.610)***	0.5369 (10.663)***	0.0013 (1.759)*
Against the Japanese yen	0.00081 (6.294)***	0.5573 (5.329)***		0.0073 (3.114)***
Nominal Effective Exchange Rate	0.00019 (1.699)*	1.6402 (6.646)***	0.2452 (4.485)***	0.0014 (1.533)

*significant at 10%; **significant at 5%; and ***significant at 1%
Note
a). Numbers inside ( ) are the t-statistics
b). For the nominal exchange rate of rupiah against the US dollar, we find GARCH(1,1) model.
c). For the nominal exchange rate of rupiah against the Japanese yen, we find ARCH(1).

Figure 5. Volatility of Bilateral Nominal Exchange rate (VNEER: nominal effective exchange rate; VUS: against the US dollar; and VYEN: against the Yen).



severe volatilities of the nominal effective exchange rate and the nominal rupiah exchange rate against the yen. In fact, the average conditional variances of rupiah against the yen and the nominal effective exchange rate were at least 100 percent and 20 percent, respectively, higher than the prevailing rates against the US dollar during the period from January 1990 to December 1996, respectively (Figure 5).

During the post-1997 crisis period, the GARCH(1,1) conditional variances of rupiah against the US dollar, the yen and the NEER have risen between 1400 percent to 2100 percent from the pre-crisis rates. More importantly, the average conditional variance for the nominal rupiah against the US dollar is found to be moderately higher than the average conditional variance for the nominal rupiah against the yen. Figure 5 also shows the fall in overall volatilities since 1999. The average conditional variances of the three different measures of rupiah nominal exchange rate during the post-1999 period are still however generally higher than the pre-1997 crisis period.

Furthermore, it is also relevant to note here that the coefficients for dummy in all three rupiah series are positive and significant, except for the NEER variable (Table 8). The positive coefficients confirm that the adoption of less rigid exchange rate policy allowed the rupiah to be more volatile against the world currencies, particularly at the height of political and financial crisis in 1998.

Based on those findings on the volatilities of rupiah during the pre-and post-1997 financial crisis and the empirical results posted in Table 2-4, we can safely conclude that the adoption of a more flexible exchange rate during the crisis period has adverse implication on the inflation rate in Indonesia. Our findings on the rupiah exchange rate provide a supporting evidence for the “fear of floating” phenomena posted by Calvo and Reinhart (2000a and 2000b). The two studies indicated that adverse consequences of exchange-rate volatilities on trade and inflation are found to be more damaging to the emerging market economies than developed economies. As a result, the developing economies (such as the East Asian countries) are more reluctant to tolerate large exchange rate movements --- by adopting a more flexible exchange rate policy and abandoning the soft-US dollar pegged policy.

B. Base Money

On November 1, 1997, the day after the first IMF agreement was signed, the government of Indonesia announced the liquidation of 16 banks. Although the decision had already been foreshadowed, it created shock waves that resulted in a total loss of confidence in the banking system (Soesastro and Basri (1998)). One of the aftermaths of the closure of the banks was the rise in the levels of monetary aggregates during the last few months of 1997 and first seven months of 1998. The expansion reflected the liquidity support provided to troubled banks and the impact of depositor runs on banks. The consequence of the banking sector bailouts prompted an increasing use of seigniorage, and would eventually require infusions of liquidity to prevent systemic runs.

Within a month after the announcement of the closures of the 16 banks, the level of base money has grown by more than 36%. Figure 4 shows that by the end of July 1998, the base money had experienced an unprecedented increase of more than 115% from its level in November 1997. For the sake of comparison, between 1991 and 1996, the annual growth rate of base money in Indonesia had been averaging only around 25%, with the highest growth in 1996 at 38% and the lowest in 1991 at around 15%. Consequently, as our test results suggest, the rapid expansion of base money played the most significant role in generating strong inflationary pressures during the post-crisis period.

V. Brief Concluding Remarks

In this study, we construct a simple and testable monetary model to uncover the source of inflation in Indonesia, particularly during the pre- and post-1997 financial crisis. Based on the working monetary model, our empirical results have shown that a significant rise in the expected depreciation of rupiah and a loose management of base money, particularly during the early stage of the 1997 financial crisis, have indeed been among the fundamental roots of the strong inflationary pressures in Indonesia during the recent years. In addition, we also find that the adoption of a more flexible exchange regime in August 1997 has allowed the rupiah to be more volatile and inflationary. However, we find limited evidences on the roles of monetary variables in explaining the inflation rate during the pre-1997 period.

In its 2002 budget plan, the government of Indonesia has announced a target of an annual inflation rate of 8%. Our empirical results supports McLeod (2001) which argues implicitly that there is no reason why the target inflation rate cannot be met, provided the central bank sticks to the targeted growth rate of base money. As our empirics have shown, the success of the country to manage its inflation during the pre-crisis is largely due to its ability to keep the money supply growing at a respectable rate of around 25 percents. However, maintaining a conservative monetary policy stand when the financial institutions are effectively collapsed is a complex task for the government of Indonesia (Alamsyah *et.al* (2001) and Siregar (2001)). It requires not only the commitment by the monetary authorities, but also the political will of both central and local provincial governments. Recent efforts by the country to push for a greater autonomy for the local governments, including in the managements of the budgets, have created concerns over the management of price stability at the provincial levels (SMERU 2001). It is clear from the recent crisis however that failure to achieve price stability has been proven to be very costly for the economy in general.

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