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Stability of the Demand for Real Narrow Money in Indonesia: Evidence from the Pre and Post Asian Crisis Era

Reza Anglingkusumo[1]
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Abstract

The stability of the demand for real M1 in Indonesia is empirically examined using quarterly data between 1981 and 2002. A cointegrated VAR methodology that isolates the period of structural breaks in the data generating process of the variables, caused by the Asian crisis, is used. The results show that the nominal M1 demand function is long run homogenous in the price level and the price level itself is endogenous in the equation for nominal M1. Therefore, a reparameterization towards the real M1 demand function is necessary. In the pre and post Asian crisis era, the demand function for real M1 in Indonesia is empirically stable and consists of a small number of variables. In the long run, the real private household consumption spending forms the permanent part of the demand for real M1 balances. Meanwhile, in the short run, the opportunity cost of holding real M1 balances, measured by the 1-month nominal interest rate of time deposits in commercial banks, and agents' seasonal preference for real money balances, are key determinants of the demand for real M1 balances. In addition, there is evidence of a co-breaking relationship between the real M1 balances and the real private household consumption spending in Indonesia during the Asian crisis.

JEL classifications: *E41, C12*
Key Words: *money demand, cointegrated VAR, structural breaks, co-breaking, Asian crisis, Indonesia*

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1. Introduction

In this paper the stability of the demand function for real narrow money (M1) in Indonesia is examined. We concentrate on the demand for real M1 instead of other measures of real money balances because of the following reasons[2]. First of all, it is difficult to find measures of interest rates that can properly gauge the opportunity costs of holding real broad money (i.e. real M2) balances[3]. This is because most interest rates in Indonesia reflect only the own rate of broad money[4]. Therefore, the examination of the demand for real broad money in Indonesia is naturally constrained by the data availability. Second, the non-cash component makes up only a small portion of the monetary base; meanwhile, the cash component is already included in the definition of M1[5]. As a consequence, analysing the demand for real base money can be seen as a "special case" of the more general demand for real narrow money balances (real M1). Therefore, we will not assess the demand for the narrower monetary aggregates such as the real base money or the real coins and notes in circulation, and focus instead on the demand for real M1 balances.

Previous studies on money demand in Indonesia have been conducted by *inter alia* Price and Insukindro (1994), Iljas (1998), Dekle and Pradhan (1999) and more recently in Simorangkir (2002) and the IMF country report for Indonesia (IMF, Country Report 02/154, 2002). Our study adds to the literature in the following sense. First, we take into account the possibility of structural breaks in the data caused by the effects of the 1997 Asian crisis. In doing so, the empirical analysis uses a cointegrated VAR framework that takes into account the possibility of structural breaks such as proposed by Johansen-Mosconi-Nielsen (2000). Second, we empirically analyse the long run homogeneity property of the price level in the nominal money demand system to justify the money demand analysis in real terms. Third, we examine the exogeneity property of the conditioning variables in the long run demand for real money function, i.e. the property of weak exogeneity, so that the final empirical specification can be used for inferential purposes.

The rest of this paper is organized as follows. Section 2 presents the basic theoretical model of the demand for real money balances. Section 3 describes the data and their time series properties. Section 4 briefly describes the econometric framework. Section 5 reports the findings. Section 6 concludes.

[2] The measures of monetary aggregates in Indonesia can be categorised as follows: (1) coins and notes in circulation (cash), (2) base money (cash plus the reserves of the commercial banks), (3) M1 (cash plus the non-interest bearing demand deposits at the commercial banks), and (4) M2 (M1 plus the interest bearing time deposits at the commercial banks).

[3] As will be shown in Section 2, the opportunity cost of holding real money balances is a key determinant of the demand function for real money.

[4] In this regard, it must be noted that the relative shares of the long term private and government bonds in the private agents' portfolios are relatively very small in Indonesia and it is only recently that the long-term bond starts to play a role in agents' portfolio decisions. Therefore, the analysis on the demand for M2 in Indonesia will need to wait until more bond yield data, as gauges of the opportunity cost of holding real M2 balances, become available.

[5] The shares of the non-cash component of base money are approximately 30% on average in 1980s, 20% in 1990s, and 40% since late 1990s.

2. The Demand for Real Money Balances: A Conceptual Framework

In their conceptualisation of the demand for real money balances, empirical economists tend to put a small number of arguments or variables on the right hand side of the real money balances equation and specify the real money demand as:

$$M^d/P = f(y, \boldsymbol{R}, \pi^e; u) \tag{1}$$

where y stands for real income, $\boldsymbol{R}$ is a vector of returns on financial assets, π^e represents expected inflation, and u is a vector of other variables. The variable y is often referred to as the scale variable. Variables in $\boldsymbol{R}$ and π^e are the opportunity costs of holding real money balances. The former is the vector of opportunity costs in terms of the lost yields on the alternative financial assets other than money. The latter is the opportunity cost in terms of the expected forgone benefits from consumable goods and services. The early developer of the theory, i.e. Friedman (1960, 1969, 1971), prefered a simple representation of the demand for real money as expressed in (1).

A more contemporary approach to the money demand conceptualisation argues that people hold real money balances because it is an argument (determinant) in their utility function, and therefore money will enter into their inter-temporal decision making process. For the purpose of our study, this is the theoretical workhorse that we follow, for it has a well-founded microeconomic foundation. One of the strengths of this workhorse is that it provides a clear interpretation of the scale variable to be used in specifying the real money demand function next to a clear mathematical expression for empirical analysis. This so-called money in the utility function approach assumes that money is a store of value and a medium of exchange that provides a service to its users in the form of efficiency in conducting economic transactions. Following the exposition in Obstfeld and Rogoff (1996, p. 530-535), we formally present the concept by starting with the following inter-temporal utility function, U, of a representative private agent in period t:

$$U_t = \sum_{s=t}^{\infty} \beta^{s-t} u\left(C_s, M_s / P_s\right) \tag{2}$$

where M_s=nominal money stock, P_s=price level, and C_s=real consumption in period s, $0 < \beta < 1$ is the discount factor indicating a time preference, and u is the function operator of U. $\partial U / \partial C > 0$, $\partial U / \partial (M / P) > 0$, and $u(C,M/P)$ is strictly concave. The individual's budget constraint in period t is given by:

$$B_{t+1} + M_t/P_t = (1 + r)B_t + M_{t-1}/P_t + Y_t - C_t - T_t \tag{3}$$

where B_t is the net private holdings of bonds issued by domestic firms or government in real terms, T_t is the lump-sum tax in real terms, r is the constant real interest rate, and Y_t is real income. Thus, the agent's total sources of financing (his/her wealth) in period $t+1$, or the term on the left hand side of (3), is the sum of real money balances, M_t/P_t, and future value of the bonds, B_{t+1}. This wealth is accumulated by the interest earnings on bonds $(1 + r)B_t$ and carried over into nominal money balances, M_{t-1}/P_t, and discretionary savings, $Y_t - C_t - T_t$. Substituting (3) into (2) yields the following expression for the inter-temporal decision problem of the agent:

$$Max\ U_t = \sum_{s=t}^{\infty} \beta^{s-t}\ u[-B_{s+1} - M_s/P_s + (1+r)B_s + M_{s-1}/P_s + Y_s - T_s, M_s/P_s] \tag{4}$$

Taking partial derivatives with respect to B_{t+1} and M_t results in two *Euler* equations:

$$u_C\,(C_t,\,(M_t/P_t)) = (1 + r)\,\beta\, u_C(C_{t+1},\,(M_{t+1}/P_{t+1})) \quad (5)$$

$$(1/P_t)\,u_C\,(C_t,\,(M_t/P_t)) = (1/P_t)\,u_{M/P}(C_t,\,(M_t/P_t)) + (1/P_{t+1})\,\beta\, u_C\,(C_{t+1},\,(M_{t+1}/P_{t+1})) \quad (6)$$

Combining (6) and (5) and collecting terms gives the following expression for the marginal rate of substitution (MRS) of real consumption and real money:

$$MRS_{M/P,\,C} = u_{M/P}(C_t,\,M_t/P_t)/\,u_C\,(C_t,\,M_t/P_t) = 1-((P_t/P_{t+1})/(1+r)) \quad (7)$$

Before continuing, it is necessary to explicitly write a model for the interest rate. Irving Fisher has reasoned that the nominal interest rate is simply the real rate of interest plus an allowance for inflation expectations that would prevail in the next period. Formally this can be written as:

$$1 + i_{\,t+1} = (1 + r)(P_{t+1}/P_t) \quad (8)$$

where i is the nominal interest rate on a financial instrument (for instance a saving bond). Equation (8) states that the real rate of return on the real and nominal bonds must be the same, and hence the nominal interest rate will move one-on-one with the expected inflation. At a higher expected inflation lenders will be willing to lend money only if the nominal interest rate is adjusted to incorporate an inflation risk premium. If (8) holds and r is constant then (7) is equivalent to:

$$u_{M/P}(C_t,\,M_t/P_t)/u_C\,(C_t,\,M_t/P_t) = 1-((P_t/P_{t+1})/(1+r) = i_{\,t+1}\,/\,(1 + i_{\,t+1}) \quad (9)$$

This equation says that the marginal rate of substitution between real money balances and real consumption positively depends upon the nominal interest rate. Assuming we have the following one-period utility function, such that the inter-temporal elasticity of substitution between real consumption and real money equals one:

$$u\,(C_t,\,M_t/P_t) = [\,C_t^{\gamma}\,(M_t/P_t)^{1-\gamma}\,]^{1-(1/\sigma)}\,/\,[1-(1/\sigma)] \quad (10)$$

we can then rewrite (9) as:

$$(M_t/P_t) = [(1-\gamma)\,/\,\gamma\,]\,[1 + (1/\,i_{\,t+1})]\,C_t \quad (11)$$

where $0 < \gamma < 1$ is the elasticity of $u(.)$ with respect to real consumption and $\sigma>0$ defines the degree of risk aversion of the representative agent. This is the mathematical expression for the demand for real money balances in the form of (1) where the real private consumption spending enters as the measure of real transaction (scale variable), instead of the total real income (output). Hence the demand for real money balance is unitary elastic to the changes in the real private consumption spending and negatively affected by the changes in the opportunity cost of holding transaction money (i.e. the changes in the nominal interest rate of the saving's bond). Adding a stochastic term and rewriting (11) in its log linear form yields the following specification of the long run demand for real money balances that can be tested empirically:

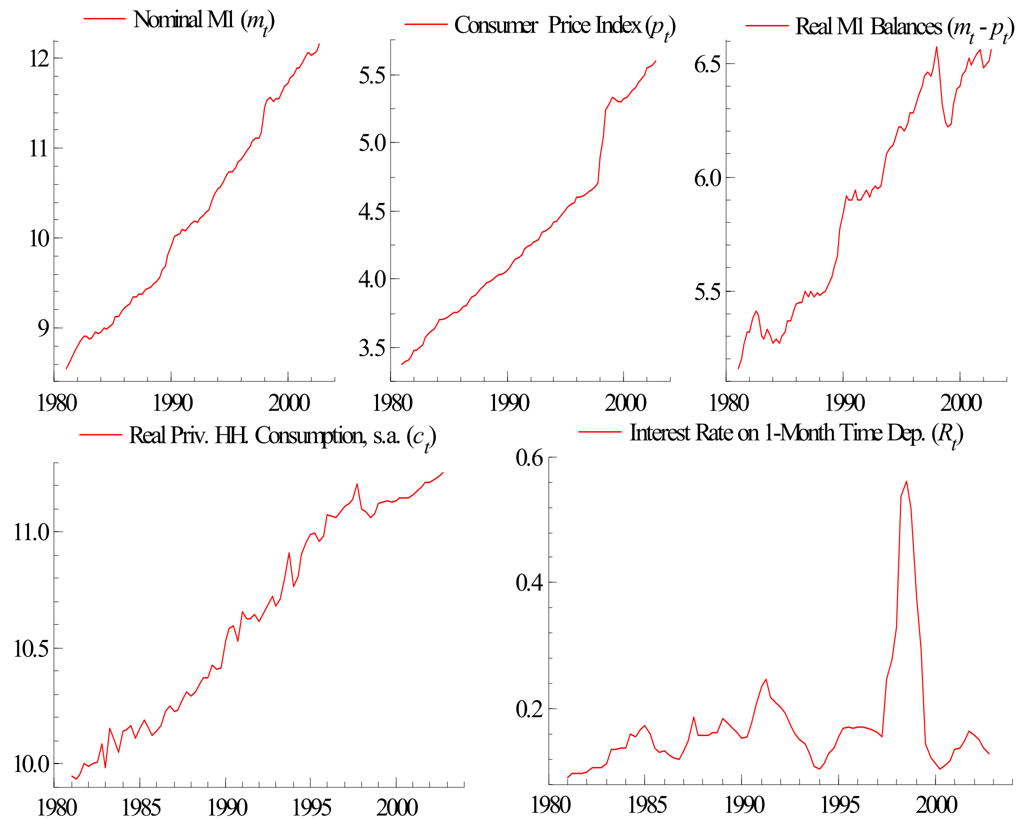
$$\log M_t - \log P_t = \delta_0 + \log C_t - \delta_1\, i_{t+1} + \varepsilon_t \quad (12)$$

where ε_t is stationary error term and δ_0 is a constant parameter.

3. The Data and Their Time Series Properties

We use quarterly data from 1981 until 2002. The notation, with small letters denoting a variable in logarithms is as follows: p_t is the log price level measured by the consumer price index (CPI), c_t is the log real private household consumption spending, m_t is the log nominal narrow money measured by M1, and R_t is the 1 month nominal interest rate of time deposits in commercial banks.

Figure 1. Time Series Plots of The Variables
(Q1/1981- Q4/2002, in logarithms except for R_t)



The original frequency of m_t and p_t is monthly. Therefore, the series are transformed into quarters by taking 3-month moving averages. This type of data transformation seems to sufficiently eliminate the seasonal patterns in the time series. Meanwhile, the original frequency of c_t is quarterly and it appears to have strong seasonal components prior to 1990. Accordingly, c_t is deseasonalized using the standard Census X-12 method without trading day adjustments. Bank Indonesia (BI) 's Directorate of Monetary Statistics provided the data on m_t, p_t, and c_t. Data on R_t is obtained from various issues of BI Annual Reports. Prior to 1989 we use the 1-month time deposits in state-owned banks as the proxy for R_t[6].

One of the effects of the Asian crisis on Indonesian data was that some if not all of the above series suffered from major structural breaks. A brief inspection of Figure 1 could easily point to this, especially if one observes the period after the outbreak of the Asian crisis in late 1997 and the period following the stabilization attempts implemented as a part of the IMF stabilization program in 1998-1999. These observations serve as a motivation to conduct a money demand analysis that takes into account the presence of structural breaks.

[6] It must be noted that before October 1988, state owned banks were the main players in the banking industry in Indonesia. In October 1998 a deregulation package was launched allowing privately owned banks to enter the industry.

4. The Econometric Framework

As mentioned above, the presence of structural breaks in the data generating process of our variables requires a money demand analysis that takes into account the presence of structural breaks. Fortunately for us, Johansen et al (2000) have developed a VECM (Vector Error Correction Modelling) methodology for testing the presence of cointegration relationships in a vector of variables, allowing for possible breaks in the deterministic trends of the long run relationships. The applicability of this methodology to our data seems quite promising. Accordingly, below we briefly describe the methodology of Johansen et al (2000).

We analyze the long run demand for money in a VECM framework that allows for breaks around time T_j, where $j = 1, \ldots, q$, as given by equation (2.6) in Johansen et al. (2000)[7], with slight modifications, namely:

$$\Delta X_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} X_{t-1} \\ tE_t \end{pmatrix} + \mu E_t + \sum_{i=k_{0,j}}^{k_{i,j}} \Gamma_i \Delta X_{t-i} + \sum_{i=1}^{k} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + \varepsilon_t \qquad (13)$$

X_t is the $(p \text{ x } 1)$ vector of variables, and D_j is an indicator function for the i-th observation in the j-th period; such that, $D_{j,t-i} = 1$ if $t = T_{j-1} + i$. Meanwile, $E_{j,t} = \sum_{i=k+1}^{Tj-Tj-i} D_{j,t-i} = 1$ for $T_{j-i}+k+1 \leq t \leq T_j$ and zero otherwise, capturing the effective sample of the j-th period. k is the number of lags in the VAR. Accordingly, $k_{0,j}$ and $k_{i,j}$ indicate the positions of the dummy variables. The standard choice of Johansen et al (2000) is $k_{0,j} = 1$ and $k_{i,j} = k$. In our case we choose $k_{0,j} < 1$ and $k_{i,j} > k$ for reasons described in Section 5.1. below. This, however, does not affect the applicability of Johansen et al (2000). The effect of the dummy variables $D_{j,t-1}, \ldots, D_{j,t-k}$ is to cause the estimated $\varepsilon_{Tj-1+1}, \ldots, \varepsilon_{Tj-1+k}$ to equal zero in the least squares estimation of (13). The break dummies and corresponding drift parameters for each of the sub sample periods are gathered in $E_t = (E_{1,t}, \ldots, E_{q,t})'$, $\mu = (\mu_1, \ldots, \mu_q)$, $\gamma = (\gamma_1', \ldots, \gamma_q')$, with dimensions $(qx1)$, (pxq), and (qxr) respectively, where q is the number of sample periods, p is the number of variables in X_t and r is the rank of $\Pi = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}'$. μ is the matrix of drift parameters for the different sub-periods and γ is the matrix of slope parameters for the cointegrating vector in the different sub-periods. Π is the usual $(p \text{ x } r)$ long run matrix $\alpha\beta'$ as in the conventional Johansen cointegrated VAR analysis (Johansen, 1995) but it is now augmented by $\alpha\gamma'$ to capture the possible shifts in the time trends in the cointegration relationships for different sub sample periods. α is a $p \text{ x } (r + q)$ matrix of long run adjustment coefficients and β is the $(r \text{ x } p)$ matrix of cointegrating vectors. The matrices Γ_i are the $(p \text{ x } p)$ matrices of short-run adjustment coefficients. ε_t is a vector of NID errors.

Assuming that the variables in X_t are not $I(2)$ processes, Johansen et al (2000, p. 220) show that the linear combinations of X_t produce cointegration relationships throughout all sub samples: $(\beta' X_t + \gamma' tE_t)$ is $I(0)$ and shows no trend. The rank of

[7] This set up generalizes Perron (1997) analysis to a multivariate setting with more than one break.

$\Pi = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}'$ is determined using the Johansen method of reduced rank regression as described in Johansen (1995) with extensions detailed in Johansen et al (2000). The presence of structural breaks in the deterministic trend and the inclusion of the indicator and shift dummies demand new asymptotic tables of the trace-test statistics for testing the rank of Π, different from those in Johansen and Juselius (1990), Johansen (1995) and Osterwald-Lenum (1992). Extending on the earlier manuscript, i.e. Johansen and Nielsen (1993), Johansen et al (2000, Table 4) provide formulae for the estimated response surface to calculate the critical values for cointegration analysis in the presence of at most 2 structural breaks with known dates at fixed proportions of the total sample. The asymptotic distributions needed for the critical values of the trace statistics can be simulated as in Johansen and Nielsen (1993) or approximated by a Γ-distribution and depend on the number of non-stationary relations, the location of break points, and vary across deterministic trend specifications. Equation (3.11) in Johansen et al (2000) gives the formula to approximate the first two log moments of the approximating Γ-distribution. Using such a formula and the explanations laid out in Johansen et al (2000, p. 226-230), we approximate the mean and variance as well as the percentiles of the Γ-distribution for one structural break at approximately 80% towards the end of the sample, as this is relevant in our example. However, in reality there are 2 breaks in our sample, one at the beginning and one at the end of the crisis, but the crisis period is too short to estimate separate trends in all 3 sub-periods, given the order of the VAR. Example of these critical values are presented in Table 1 below.

Our cointegrated VAR and its corresponding VECM analyses reported in the next section use $q=2$, i.e. 2 sub-samples corresponding to the pre and post crisis era, and assume the timing of the break is known. A procedure for selecting the timing of the break in a multivariate setting has been put forward by Inoue (1999), however, the framework by Johansen et al (2000) and their calculations of critical values assume a prior knowledge on the timing of the break, hence our analysis uses similar assumptions.

Finally, we conduct a test for weak exogeneity for the estimation of cointegrating vector β. Implications of such a test for the purpose of inference in a cointegrated VAR analysis have been studied by *inter alia* Boswijk (1992), Johansen (1992), Ericsson, Hendry, and Mizon (1998), and Hendry and Juselius (2000). Concise reviews are available in Favero (2001) and Harris and Sollis (2003).

5. Empirical Results[8]

5.1. The Cointegration Test of Johansen et al (2000)

To start, we examine the cointegration properties of the vector of variables $X_t = [m_t\text{-}p_t,\ c_t,\ R_t]$, using the VECM representation of equation (13) with $p=3$ and $q=2$. Lag length $k = 5$ is chosen as it is the minimum lag which removes the serial correlation in the residuals. As reported in Panel A of Table 1 joint tests for both serial correlation and heteroskedasticity of the VAR(5) residuals point to the non-

[8] To derive the empirical results we used the PcGive 9.3 and PcFiml softwares, Doornik and Hendry (2000, vol I and II).

rejection of the null. Despite the rejection of the null of non-Gaussian residuals we keep the VAR(5) as our multivariate representation of the data.

The choice for the timing of the indicator dummies, D_j, is motivated by the following information. First of all, we show in Anglingkusumo (2005b) that in a two regime Markov switching error correction model for quarterly changes in price level, i.e. quarterly CPI inflation, the period of monetary crisis in Indonesia runs from Q1/1998 until Q4/1999. Hence, p_t had its first trend break in Q1/1998 when it jumped significantly and its second trend break in Q4/1999 when it finally stabilized. However, the variable R_t needs an earlier break date than Q1/1998, i.e. in Q4/1997, and a later date for the second break, i.e. Q1/2000. Meanwhile, m_t reveals that this variable breaks 1 quarter earlier than p_t. Therefore, in order to safely capture the crisis breaks we use Q4/1997 as the starting date of the crisis break and Q1/2000 as the ending date[9], i.e. T_j indicates Q4/1997. Unfortunately, this makes the period between the breaks too short to identify in the Johansen et al (2000) framework. Accordingly, we reduce the number of periods in the analysis to two *(q=2)* and we neutralize the effect of the period between the breaks by introducing extra dummy variables. In this regard, the indicator dummies, $D_{j,t-i}$, are applied from Q4/1997 until Q1/2000, i.e. $i=0,....,9$ where $i=0$ is in Q4/1997 and $i=9$ is in Q1/2000. The inclusion of these [0,1] dummies effectively isolates the corresponding observations where the dummies are applied. As such, the empirical analysis is primarily focused on two sub-sample periods, namely the pre and post crisis periods. We set E_1 equal to 0 for the period from Q4/1997 onwards, otherwise it equals 1. Meanwhile E_2 is 0 prior to Q2/2000 and 1 otherwise. This means E_1 and E_2 in (13) are the pre and post crisis intercepts, respectively; and, tE_1 and tE_2 are the pre and post crisis trends, respectively.

Using the critical values from Johansen et al (2000), the cointegration analysis indicates $r=1$ at 5% as reported in Panel A of Table 1. The trace statistic for $r \leq 0$ is 59.86 and above the 5% critical value of 55.31. Meanwhile, the trace statistic for $r \leq 1$ is clearly below the critical value. Therefore, we choose the alternative that $r = 1$. The cointegration relationship is normalized with the long run coefficient of variable m_t-p_t, i.e. β_{m-p}, restricted to equal 1. In this way the long run relationship can be regarded as the long run demand for real money balances as implied by equation (12). Panel B of Table 1 shows that the long run coefficient of c_t, i.e. β_c, is approximately equal to 1, entailing that the theoretical assumption of the unit elasticity of m_t-p_t with respect to c_t is possibly not violated. However, one can readily notice that the long run coefficient of R_t, i.e. β_R, is very small and could possibly be zero in the long run, indicating the possible absence of R_t in forming the long run monetary equilibrium. Meanwhile, the coefficients of the pre and post crisis trend components, i.e. γ_1 and γ_2 respectively, are approximately 0. This implies the absence of a broken trend in the long run money demand equation across the two periods. Moreover, the significant negative equilibrium correction coefficient of m_t-p_t, i.e. α_{m-p}, points to a self-adjusting equilibrium real M1, and the insignificant equilibrium correction coefficients for c_t and R_t, i.e. α_c and α_R, point to weak exogeneity of both variables for the estimation of β. Next, we conduct several tests for restrictions on the long run vector $[\beta\gamma]'$ and the loading vector α to see if the above conjectures can be empirically validated . The results are presented in Panel C of Table 1 and discussed in the next sub-section.

[9] Moreover, in Anglingkusumo (2005b) we address the issue of long run price homogeneity in a system involving a vector of three variables consisting m_t, p_t, and c_t, hence the use of these two dates to enclose the crisis breaks will serve our purpose.

Table 1. Cointegration Test Results for Vector $X_t = [m_t\text{-}p_t, c_t, R_t]$
Sample Q2/1982 - Q4/2002, t=1,…,83

[A]. The Johansen et al (2000) Cointegration Test VAR (*k=5*), *p=3*, *q=2*, and a break point at ~ 80% of end sample			
H_0: rank ≤	Trace (LR) Statistics	Critical Values [5%/10%]	Diagnostics on estimated VAR (5) Residuals :
			Vector AR (5) test: $F(45,110)$ / *p-value* = 1.0028 / 0.48
$r \leq 0$	59.86	55.31 / 51.70	Vector Normality test:
$r \leq 1$	23.84	34.48/31.53	$\chi^2(6)$ / *p-value* = 74.317 / 0.00
$r \leq 2$	7.16	17.46/15.20	Vector Heteroskedasticity test: $F(204,91)$ / *p-value* = 0.297 / 1.00
[B] The Normalized *CI (1,1)*$_{r=1}$ Relationship			
Vector *[βγ]'* $\beta_{m\text{-}p}$ = 1 (identifying), β_c = -1.23 (0.17), β_R = - 0.11 (0.34), γ_1 = 0.0041 (0.0035), γ_2 = 0.0065 (0.0088)			Vector *α* $\alpha_{m\text{-}p}$ = -0.33 (0.08), α_c = 0.24 (0.11), α_R = 0.034 (0.046)
[C]. VECM Restrictions on the Long Run Vector *[βγ]'* and the Loading Vector *α* $\beta_{m\text{-}p}$ = 1 is the identifying restriction			
Restrictions $H_1 \rightarrow$ $\gamma_1 = 0, \gamma_2=0$ *[LR~$\chi^2(2)$/p-value = 1.55 (0.45)]*			
Restrictions $H_2 \rightarrow \gamma_1 = 0, \gamma_2=0, \alpha_c = 0$ *[LR~$\chi^2(3)$/p-value = 3.66 (0.30)]*			
Restrictions $H_3 \rightarrow \gamma_1 = 0, \gamma_2=0, \alpha_R = 0$ *[LR~$\chi^2(3)$/p-value = 2.29 (0.51)]*			
Restrictions $H_4 \rightarrow \gamma_1 = 0, \gamma_2=0, \alpha_c = \alpha_R = 0$ *[LR~$\chi^2(4)$/p-value = 4.18 (0.38)]*			
Restrictions $H_5 \rightarrow$ $\beta_c = -1, \gamma_1 = 0, \gamma_2=0, \alpha_c = \alpha_R = 0$ *[LR~$\chi^2(5)$/p-value = 7.8 (0.16)]*			
Restrictions $H_6 \rightarrow$ $\beta_R = 0, \gamma_1 = 0, \gamma_2=0, \alpha_c = \alpha_R = 0$ *[LR~$\chi^2(5)$/p-value = 5.10 (0.40)]*			
Restrictions $H_7 \rightarrow \beta_c=-1, \beta_R=0, \gamma_1= 0, \gamma_2=0, \alpha_c=\alpha_R=0$ *[LR~$\chi^2(6)$/p-value = 7.81 (0.25)]*			

Notes: [A] The cointegration test follows Johansen et al (2000) using equation (13). LR-test rank *(Π)* = *r* against rank *(Π)* = *3*. Numbers in [.] denote critical values at 95th / 90th percentiles of the approximating *Γ*-distribution calculated using the response surface parameters of Table 4 of Johansen et al (2000). [B] ML estimate of $\beta = (\beta_{m\text{-}p}, \beta_c$, and $\beta_R)$ when $r=1$, $\gamma = (\gamma_1, \gamma_2)$, and $\alpha=(\alpha_{m\text{-}p}, \alpha_c, \alpha_R)$; 1 and 2 denote the pre- and post crisis sub-periods respectively. [C] LR-test for (joint) restrictions on *α, β,* and *γ*. Numbers in (.) denote asymptotic *p-values*.

5.2. Tests for Weak Exogeneity and Long Run Restrictions

In Panel C of Table 1 we consider 7 types of joint restrictions implying 7 hypotheses. Restrictions H_1 imply that the trends in the long run money demand of the pre and post crisis periods are equal zero given $\beta_{m\text{-}p} = 1$. The result suggests that the restrictions hold (*p-value* = 0.45). This entails a trendless *CI(1,1)* relationship in X_t before and after the crisis. Restrictions H_2 and H_3 test whether c_t and R_t are individually weakly exogeneous variables given restrictions H_1. The results show both restrictions hold (*p-values* equal 0.30 and 0.51, respectively). Next, restrictions H_4 test the joint restrictions that $\alpha_c = \alpha_R = 0$ given restrictions H_1. The result shows a *p-value* = 0.38. We conclude that both variables c_t and R_t are weakly exogenous conditioning variables in the long run cointegration relationship where $m_t\text{-}p_t$ serves as the normalizing variable. Hence, the presence of a long run real M1 demand function is empirically verified.

We continue to examine a more complete form of the long run real M1 demand function where we also check the stability of the estimated coefficients. Restrictions H_5 imply the long run coefficient of c_t in $[\beta\gamma]'$ is = -1. The test result shows that given restrictions H_4, the restriction $\beta_c = -1$ is valid (*p-value* = 0.16). However, this result is not so easily accepted in the post crisis sample as shown by the recursive plot in Figure 2a. Restrictions H_6 imply the long run coefficient of R_t in $[\beta\gamma]'$ is zero given restrictions H_4. The test points to the validity of having $\beta_R = 0$. This result holds recursively at 10% (see Figure 2b). Restrictions H_7 imply both $\beta_c = -1$ and $\beta_R = 0$ in $[\beta\gamma]'$ given restrictions H_4. The test shows a *p-value* of 0.25, indicating valid restrictions. Combining these results we conclude that the long run coefficient of c_t is unity, hence the theoretical assumption of the unitary elasticity of c_t in equation (12) is empirically verified. However, the result that $\beta_R = 0$ implies that variable R_t does not matter in agents' long run real M1 demand function. In other words, the long run changes in Indonesian R_t have not been large enough to identify this effect. It is however still interesting to examine whether such an opportunity cost of holding real M1 balances also does not matter in the short run. We address this in Sub-section 5.4.

$$\Delta X_t = \mu E_t + \sum_{i=1}^{4} \Gamma_i \Delta X_{t-i} + \alpha \quad [\beta\gamma]' \quad X_{t-1}$$

$$\begin{pmatrix} \Delta(m-p)_t \\ \Delta c_t \\ \Delta R_t \end{pmatrix} = \mu \begin{pmatrix} E_{1,t} \\ E_{2,t} \end{pmatrix} + \sum_{i=1}^{4} \Gamma_i \begin{pmatrix} \Delta(m-p)_{t-i} \\ \Delta c_{t-i} \\ \Delta R_{t-i} \end{pmatrix} + \begin{pmatrix} \alpha_{m-p} \\ \alpha_c \\ \alpha_R \end{pmatrix} [\beta_{m-p} \;\; \beta_c \;\; \beta_R \;\; \gamma_1 \;\; \gamma_2] \begin{pmatrix} (m-p)_{t-1} \\ c_{t-1} \\ R_{t-1} \\ tE_{1,t} \\ tE_{2,t} \end{pmatrix}$$

$$+ \quad \Sigma_{i=0}^{9} \kappa_{2,i} D_{2,t-i} \quad + \quad \varepsilon_t \sim NID(0,\Sigma)$$

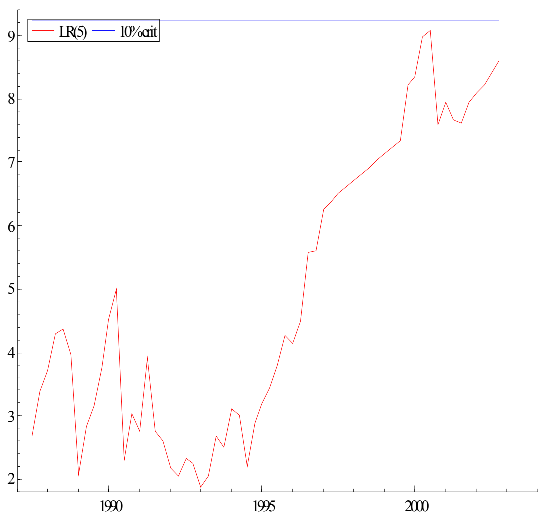
Where $\mu = (\mu_1, \mu_2)$; $\alpha_c = \alpha_R = \beta_R = \gamma_1 = \gamma_2 = 0$; $\beta_{m-p} = -\beta_c = 1$
Sample: Q2/1982 - Q4/2002, $t=1,\ldots,T = 83$, *and* $T_1 = 63$

(14)

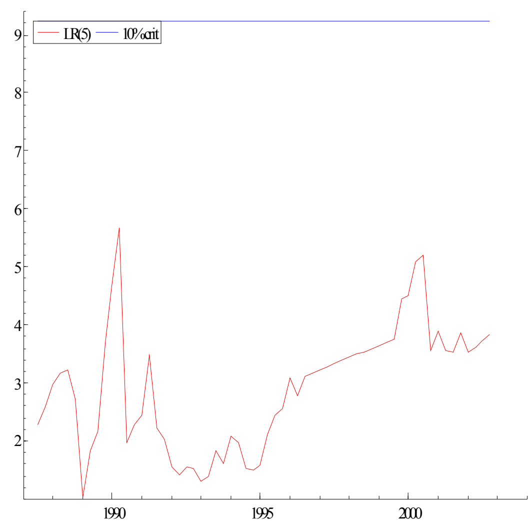
At this stage we have identified a VECM in the form of system of equations (14), which is an example of equation (13). This system of equations states that since we can neglect the effect of R_t on m_t-p_t, as $\beta_R = 0$, the long run stochastic trend of m_t-p_t is simply composed of the stochastic trend of c_t. In other words, m_t-p_t and c_t share a common trend.

Figure 2. Recursive Results of the LR- χ^2 Tests for *[βγ]'* and *α*
(Initialization 20 Quarters)

(a) H_5: $\beta_{m-p} = 1, \beta_c = -1, \gamma_1 = \gamma_2 = 0, \alpha_c = \alpha_R = 0$



(b) H_6 : $\beta_{m-p} = 1, \beta_R = 0, \gamma 1 = \gamma 2 = 0, \alpha_c = \alpha_R = 0$



Note: Recursive plot of LR-test statistic for H_5 and H_6 in Table 1, for increasing samples

5.3. Testing for Long Run Price Homogeneity

Figure 3 illustrates the restricted cointegration relationship $(m\text{-}p)_t - c_t$ (Figure 3a) and the estimate of the restricted long run demand for real M1 (Figure 3b). It is quite clear from Figure 3a that the monetary disequilibrium is a stationary *I(0)* variable in the long run. Its sample mean is around – 4.74. This sample mean can be taken as the estimate of δ_0 in equation (12). From Figure 3b one notices that the fitted line of the long run demand for real M1 balances is simply c_t since we have shown above that, in the long run, restrictions H_7 holds empirically.[10] Hence, Figure 3b reaffirms our conclusion that the long run stochastic trend of real M1 balances is fully composed of c_t.

The above conclusion, however, rests on the assumption that money is long run homogenous in prices, and therefore a simple money demand system in real terms can be identified. In other words, based upon restrictions H_7 we can write the long run money demand relationship as: $m_t - \eta\, p_t - c_t - 0.t(E_{1,t} + E_{2,t}) - 4.74$, where η is restricted to equal 1. This type of restriction is also implicitly assumed in the theoretical expositions leading to equation (12) in Section 2. It is therefore necessary to empirically verify this assumption. To do so we analyze the following vector of three variables, $V_t = [m_t, p_t, c_t]$[11]. The Johansen et al (2000) cointegration test is performed using a similar set up of dummy variables as in Sub-section 5.1. and lag length 5 is used. The diagnostics on the residuals are reported in Panel A of Table 2.

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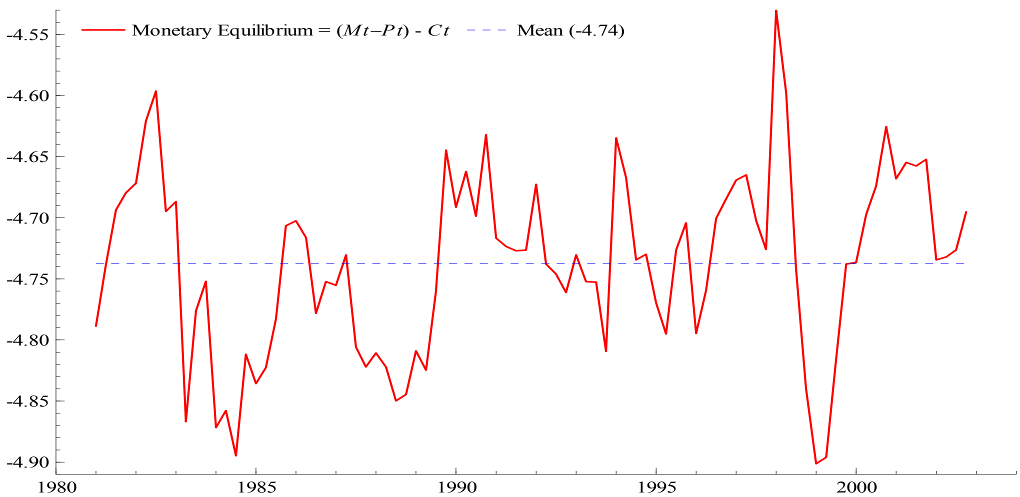
[10] With a caveat that restrictions H_5: $\beta_{m-p} = 1, \beta_c = -1, \gamma_1 = \gamma_2 = 0, \alpha_c = \alpha_R = 0$ are not as easily accepted for the post crisis period as the pre-crisis period.
[11] Since we have shown previously that R_t has no role in determining the real M1 demand in the long run, we omit this variable in V_t.

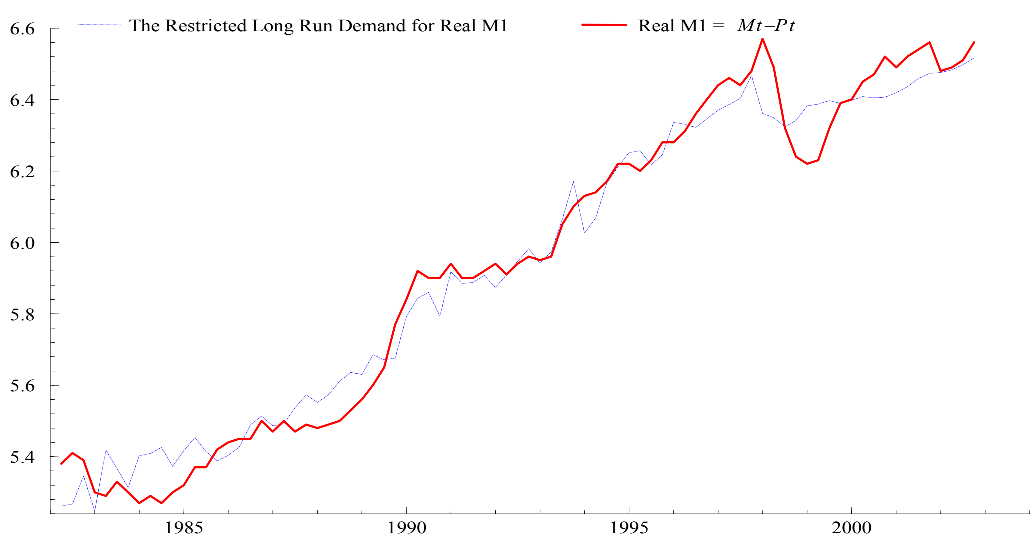
They suggest the absence of serial correlation and the presence of homoskedastic residuals. It appears also that residuals' distribution is close to being normal. The results of the cointegration test using the VAR representation of equation (13) are presented in Panel A of Table 2. They show that the variables in V_t are cointegrated with $r=1$.

Figure 3. The Restricted Cointegration Relationship and Long Run Demand for Real M1

(a) The Restricted Cointegration Relationship [βγ]'



(b) The Restricted Long Run Demand for Real M1



Note: The restricted long run demand for real M1 in Figure 3b is obtained using restrictions H_7.

Table 2. The Tests Results for Vector $V_t = [m_t, p_t, c_t]$
Sample Q2/1982 - Q4/2002, t=1,….,83

[A]. The Johansen et al (2000) Cointegration Test **VAR (*k=5), p=3, q=2*, break point at ~ 80% of end sample**			
H_0: rank ≤	Trace (LR) Statistics	Critical Values [5%/10%]	Diagnostics on VAR (5) Residuals:
			Vector AR 5 test F(45,110) / *p-value* = 0.956 / 0.55
$r \leq 0$	70.94	55.31 / 51.70	Vector Normality: $\chi2(6)$ / *p-value* = 10.697 / 0.098
$r \leq 1$	29.36	34.48/31.53	Vector Hetero: F(204,91) / *p-value* = 0.344 / 1.00
$r \leq 2$	8.81	17.46/15.20	
[B]. VECM(4) Restrictions on the Long Run Vector $[\beta\gamma]'$ $\beta_m = 1$ (Identifying Restriction)			
Restrictions H_{11}→ $\gamma_1 = 0, \gamma_2=0, \beta_p = -1$ [LR~$\chi^2(3)$/*p-value* = 6.01 (0.11)]			
Restrictions H_{21}→ $\gamma_1 = 0, \gamma_2=0, \beta_p = -1, \alpha_p=0$ [LR~$\chi^2(4)$/*p-value* = 30.87 (0.00)]			
Restrictions H_{31}→ $\gamma_1=0, \gamma_2=0, \beta_p = -1, \alpha_p= \alpha_m=0$ [LR~$\chi^2(4)$/*p-value* = 37.87 (0.00)]			

Notes: [A] See Table 1. [B] See note [C] in Table 1.

Given the cointegration test result we examine the long run homogeneity of the price level by testing several restrictions on the vector $[\beta\gamma]'$. The results are reported in Panel B of Table 2. The first set of restrictions, H_{11}, imply a trendless $CI(1,1)_{r=1}$ relationship and the long run restriction that the coefficient of the price level, β_p, equals –1. The test result points to the plausibility of such a long run restriction (*p-value* = 0.11). Restrictions H_{21} entail the weak exogeneity of the price level in the equation for m_t. The result shows that the price level is not a weakly exogeneous variable in the equation for m_t. In fact, from restrictions H_{31} we find that both m_t and p_t are endogenous in the long run equation for m_t and the two should be modeled jointly. Hence we conclude that price level is long run homogenous in V_t and a reparameterization into a real money demand system is both plausible and necessary. In the short run analysis of this model we maintain the parameterization in $\Delta(m\text{-}p)_t$.

5.4. The Short Run Money Demand Equation

For inferential purposes, the weak exogeneity of both c_t and R_t as verified above is sufficient to map the $CI(1,1)_{r=1}$ relationship in X_t into a single $I(0)$ equation, or simply put, to examine only the equation for $\Delta(m\text{-}p)_t$ in the system of equations (14). Using the general to specific reduction technique as illustrated in Doornik and Hendry (2000, vol. II), we arrive at the short run money demand equation for $\Delta(m\text{-}p)_t$ as reported in (15) below.

$$\Delta (m\text{-}p)_t = -\underset{[0.30\,/\,-2.71]^{***}}{0.812}\, E_t^{pre\text{-}crisis} - \underset{[0.29\,/\,-2.73]^{***}}{0.811}\, E_t^{post\text{-}crisis} + \underset{[0.12\,/\,3.85]^{***}}{0.474}\, \Delta (m\text{-}p)_{t-4} - \underset{[0.30\,/\,-2.78]^{***}}{0.845}\, \Delta R_{t-1}$$

[s.e./t-stat]

$$- \underset{[0.06\,/\,-2.75]^{***}}{0.173}\, [(m\text{-}p)_{t-1} - c_{t-1} - 0\, (\, tE_t^{\,pre\text{-}crisis} + tE_t^{\,post\text{-}crisis})]$$

$$- \underset{[0.29\,/\,-2.40]^{*}}{0.715}\, D_t(1997/4) - \underset{[0.30\,/\,-2.40]^{*}}{0.722}\, D_t(1998/1) - \underset{[0.28\,/\,-2.89]^{***}}{0.831}\, D_t(1998/2)$$

$$- \underset{[0.29\,/\,-2.68]^{***}}{0.782}\, D_t(1998/3) - \underset{[0.30\,/\,-2.98]^{***}}{0.900}\, D_t(1998/4) - \underset{[0.03\,/\,-3.20]^{***}}{0.127}\, D_t(1999/1)$$

$$- \underset{[0.05\,/\,-2.00]^{**}}{0.107}\, D_t(1999/2) + \underset{[0.04\,/\,1.42]}{0.063}\, D_t(1999/3) - \underset{[0.05\,/\,-0.75]}{0.041}\, D_t(1999/4)$$

$$- \underset{[0.03\,/\,0.228]}{0.007}\, D_t(2000/1) + \quad \varepsilon_t \sim NID\,(0,\ \sigma^2) \qquad (15)$$

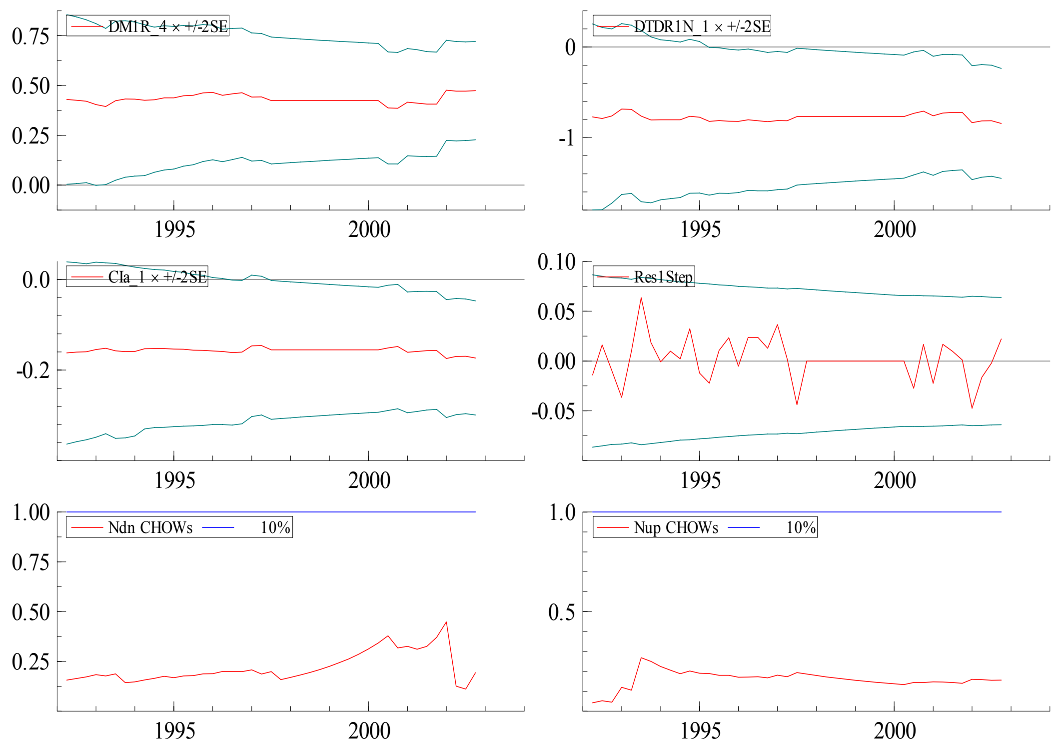
Effective Sample: Q2/1982 – Q4/2002

$R^2 = 0.56185$	*AR 1-5 test: F(5,63)/p-val. = 0.30 / 0.90*
RSS = 0.0694 ;	*ARCH 1-4 test : F(4,60)/p-val. = 0.82 / 0.51*
S.E. = 0.0319	*Normality test:* $\chi^2(2)$*/ p-val. = 4.84 / 0.08*
F(14, 68) = 6.228 (0.000)	*Heteroskedasticity test: F(17, 50)/p-val.=0.33 / 0.99*
N = 83	*RESET test: F(1, 67)/p-val. = 2.44 / 0.12*

Notes: E_t *and* tE_t *are the sub-sample intercepts and trends, respectively. The* D_t*s are [0,1] indicator dummies, wherein the estimated* ε_t *for the dummied period equals zero. ***, **, and * indicate significance at 1 %, 5%, and 10%.*

As shown by the diagnostics, equation (15) does not suffer from residual serial correlation and heteroskedasticity, or non-linearity in terms of its explanatory variables. Yet there is a trace of non-Gaussian errors. In terms of the coefficients, this specification yields sensible results. In the short run, agents' demand for real M1 balances adjusts as a result of the seasonal preference for real M1 balances (modeled by last year's money growth), and the last quarter changes in the 1-month time deposit rate in commercial banks with a semi elasticity coefficient around –0.85. In the long run, real M1 adjust to its equilibrium path with an adjustment coefficient around –0.17.

Figure 4. The Stability of the *I(0)* Real M1 Equation
(Recursive Estimates with 20 Quarters Initialization)



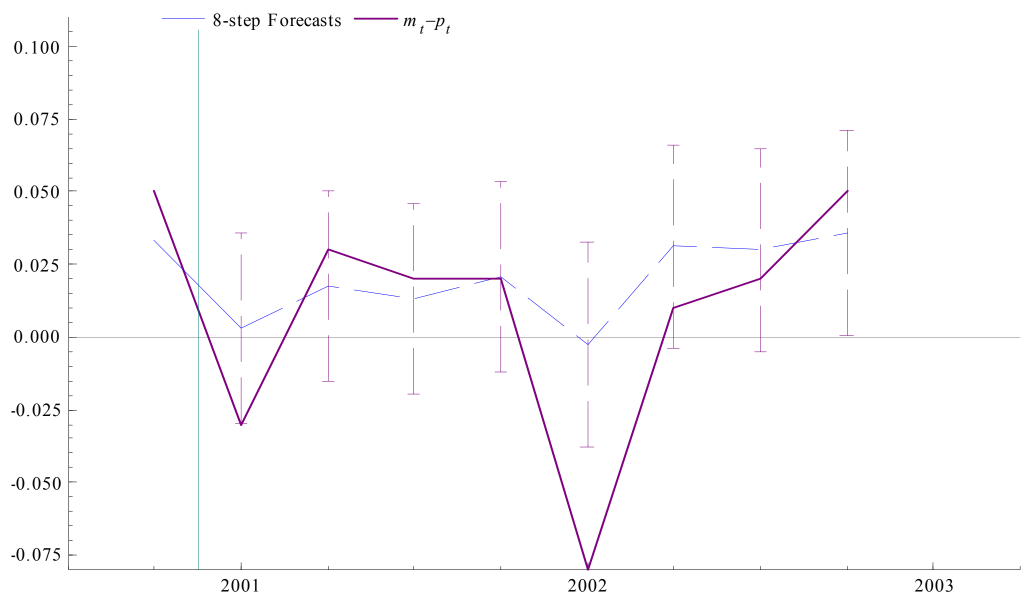
***Notes:** DM1R_4 = $\Delta(m\text{-}p)_{t-4}$, DTDR1N_1 = ΔR_{t-1} , and CIa_1 = $EqC_{t-1} = [(m\text{-}p)_{t-1} - c_{t-1} - 0\ (tE_t^{pre\text{-}crisis} + tE_t^{post\text{-}crisis})]$, are the recursive parameter's plot.* Res1Step = Recursive 1-step residuals, Ndn Chows =Recursive Chow Break Point Test, Nup Chows = Recursive Chow forecast test. The critical value for all the recursive LR-$\chi 2$ test is 10%. 2SE is 2 standard errors.

The graphical assessment of the stability of equation (15) is reported in Figure 4. The stability of the equation throughout the two sub sample periods can be studied from the recursive 1-step residuals plot (Res1Step). There is no observed period of instability during the pre and post crisis period. The recursive Chow break point (Ndn CHOWS) and forecast tests (1up CHOWS and Nup CHOWS) also point to a stable equation throughout the two sub sample periods at a 10% critical value. With respect to the parameters, equation (15) yields parameter values that are constant and significant for the coefficients of $\Delta(m\text{-}p)_{t-4}$ and the long run adjustment. The significant result for ΔR_{t-1} is also confirmed in the plot of recursive parameter estimates. However, the significance of the coefficient of ΔR_{t-1} only appears in the aftermath of the crisis. This may suggest agents' increasing consiousness on the opportunity costs of holding real money balances, after experiencing a period of high nominal interest rate during the Asian crisis era.

Another way to examine the stability of equation (15) is to test for a general restriction that the coefficients of $E_t^{pre\text{-}crisis}$ and $E_t^{post\text{-}crisis}$ are the same. If they are in fact the same, then the intercept of equation (15) before and after the crisis does not change, and hence the demand for real M1 balances in the two periods does not suffer from an intercept break. An *LR* test is conducted to see if the difference of the two coefficients is zero. This results in an $LR\text{-}\chi^2$ test value of 0.0206 with a *p-value* of 0.8858, suggesting that the difference between the two sub-periods is in fact equal zero. This result also indicates the presence of a co-breaking relationship among the elements of the long run vector in equation (15) which was already depicted in Figure 3b above[12]. In sum we find that (a) the demand function for real M1 does not suffer from an intercept break, and (b) real money and real consumption in the long run relationship co-break during the crisis.

Figure 5 illustrates the post crisis 1 to 8-step ahead forecasts and standard errors of of equation (15) estimated up to Q4/2000. In other words, we use the pre-crisis observations heavily to predict the post crisis period. As can be seen from the figure, the forecast tracks the behavior of the post crisis real M1 balances quite well. This last point strengthens the previous conclusions that equation (15) has been stable in terms of parameters and intercepts before and after the Asian crisis.

Figure 5. The Post Crisis 1 to 8 Step Ahead Forecasts with a 1 s.e. Band for Q1/2001 – Q4/2002 using Equation (15) Estimated up to Q4/2000



6. Conclusions

In this paper we have specified a demand function for real narrow money (real M1) in Indonesia for the period between Q1/1981 and Q4/2002. The estimated demand function for real M1 is plausible due to the empirical long run homogeneity of the price level variable. Moreover, the endogeneity of the price level variable in the nominal M1 demand equation makes it necessary to analyze the money demand function in real terms. The empirical real M1 demand function estimated in this paper is conditioned on a small set of weakly exogenous variables, namely the real private household consumption spending and the 1-month nominal interest rate of the time deposits in commercial banks. The former can be regarded as the permanent part of

[12] For the notion of co-breaking see Hendry (1997) and Hendry and Mizon (1998).

the demand for real M1 balances. Its dynamic path is the sole determinant of the stochastic trend of the real M1 balances in the long run. The latter can be viewed as agents' opportunity cost of holding real M1 balances. It affects the demand for real M1 balances only in the short run with a negative semi elasticity. The short run demand for real M1 balances is also affected by agents' seasonal preference for the demand for real money balances. The empirical results further show that the residuals of the estimated long run demand for real M1 balances have no long run trend before and after the crisis. This cointegrating relationship has a constant that can be interpreted as a non-declining (constant) velocity. After mapping the cointegrated long run real M1 demand system into an *I(0)* equation we find that the demand function for real M1 balances in Indonesia does not suffer from an intercept break, and real M1 and real consumption in the long run relationship co-break during the Asian crisis.

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