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# INTEREST FREE AND INTEREST-BEARING MONEY DEMAND: POLICY INVARIANCE AND STABILITY

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## Abstract

This paper, using quarterly Iranian data for the period 1966-1998, extends the literature by investigating the stability of the interest-free money demand function. The study also examines the stability of economic agents' behavior in demanding interest-bearing and interest-free money. It was found that, contrary to interest-bearing demand for money, both short- and long-run demand for interest-free money functions are stable and their coefficients are invariant with respect to policy and other exogenous shocks, as well as changes in regime.

## 1. Introduction

To monetary authorities, the knowledge of the demand-for-money function, among many other functions in the economy, is necessary, though not sufficient, for understanding the way in which the economy responds to changes in exogenous factors, or at least to the supply of money. Most importantly, for a monetary policy to be effective, demand for money must be stable. For example, if the relationship between the demand for money and its determinants shifts around unpredictably, the central bank loses the ability to derive results from the implementation of its policies. In such a case, variations in the demand-for-money function themselves are an independent source of disturbance to the economy.

As far as the issue of monetary policy is concerned, the stability of interest elasticity of demand for money becomes relatively more important than the other factors affecting the demand-for-money function. Recently, many studies found a stable money demand in different countries. For example, Stock and Watson (1993) as well as Ball (2001) find a stable long-run demand for U.S. M1. Peytrignet and Stahel (1998) find a stable M2 and M3 demand for Switzerland, Muscatelli and Spinelli (2000) find a stable long-run M2 demand for Italy, and Buch (2001) finds a stable demand for M1 and M2 for Hungary and a stable M1 demand for Poland. However, many other studies found unstable demand for money. For instance, Lieberman (1980) finds unstable interest elasticity for demand for money for the U.S and U.K. Ripatti (1998) finds unstable demand for M1 and M3 for Finland. Bahmani-Oskooee and Bohl (2000) find unstable M1, M2 and M3 demand functions for Germany, and Hamori and Tokihisa (2001) find unstable demand for M2 in Japanese data.

The interest rate may be one of the major factors subject to speculation, if not the only one, of the demand-for-money function. Furthermore, since money may be demanded as an inventory to smooth differences between income and expenditure streams, and as one among several assets in a portfolio, both actual and expected interest rates may have a strong impact on the economic agents' behavior related to the demand for money. One may, therefore, argue that money demand would be more stable if this major source of instability would be eliminated.

Recently, taking Tunisia as a case study, Darrat (1988) shows that demand for money, in the Islamic interest-free system, is relatively more stable, the monetary authority can control more effectively interest-free monetary assets, and only these assets have a reliable link with the ultimate policy objective. Yousefi *et al.* (1997), following Darrat's (1988) approach, but using Iranian data, confirm Darrat's conclusion on the stability of demand for money, but contradict Darrat's finding on monetary aggregate/price link. Darrat (2000), using the data of Yousefi *et al*. and correcting the misspecification error made by these authors, concludes again that only the interest-free banking system provides a reliable link between money growth and inflation both in the short- and long-run. Hassan and Al-Dayel (1998/9), employing data of 15 Islamic countries, supported Darrat's findings for most of those countries. However, these studies analyze the stability of only short-run interest-bearing and interest-free demand-for-money functions. Hassan and Mazumder (2000) extend the analysis by investigating short-run stability of the velocity of interest-free money vis-à-vis interest-bearing money as well as long-run policy controllability of these two kinds of money supply for six African countries (Algeria, Egypt, Morocco, Nigeria, Sudan and Tunisia), three Asian countries (Indonesia, Malaysia and Pakistan), four Gulf countries (Bahrain, Kuwait, Qatar and Saudi Arabia) as well as Iran, Jordan, Syria and Turkey. Their findings support the relative effectiveness of interest-free banking in these countries in terms of stable and smooth velocity of money, controllability of monetary aggregates, and stronger linkage between monetary policy instruments and ultimate policy goals of these countries.

Darrat (2002) further extends the analysis by testing the relative efficiency and policy usefulness of interest-bearing and interest-free monetary systems over the long run for both Iran and Pakistan. The motivation of these papers (Darrat (1988), Yousefi *et al.* (1997), Darrat (2000), Hassan and Al-Dayed (1998/9), Hassan and Mazumber (2000) and Darrat (2002)) comes from the fact that the Islamic economic system prohibits receipts and payments of pre-determined (fixed) interest on any financial transactions. Namely, the basic premise of Islamic banking lies on the sharing of profit or loss among depositors, investors and banks.

However, none of the existing studies directly investigated the stability of long-run demand for interest-free money in contrast to demand for interest-bearing money. Furthermore, since the coefficients of money demand may be constant, but may not be invariant to policy shocks, as mentioned by Lucas (1976), a possible extension of this literature is to investigate whether demand for interest-free money is invariant to policy or other exogenous shocks. Note that constancy and invariance are two different concepts. The coefficients of the demand-for-money equation can be constant during the historical period, but may vary in response to regime changes or to changes in the distribution of variables (e.g., tax rates, interest rates, etc.) which are under the control of a governmental agency such as a central bank or executive. In other words, if agents in demanding money are forward looking then policy or regime changes will result in the change of the agents' behavior and may lead to policy ineffectiveness. The above-mentioned studies investigated the stability of the backward-looking model for interest-free in contrast to interest-bearing demand for money. If agents are forward looking in demanding money, then the demand for money is not policy invariant. In such a case, testing the stability of a backward-looking demand for money may fail to identify the true constant-parameter expectation-generating equation.

This issue was recently given special attention. For example, Favero and Hendry (1992) find M1 money demand in the U.S. is invariant to policy and other shocks. Hurn and Muscatelli (1992) as well as Engle and Hendry (1993) find M1 and M4 demand-for-money functions in U.K. are invariant to policy shocks. To the best of my knowledge, no study has so far investigated whether coefficients of interest-free demand for money in contrast to interest-bearing demand for money are invariant to policy changes. The goal of this paper is to extend this literature by investigating the long-run stability and invariance of demand for interest-free vis-à-vis interest-bearing money. Furthermore, it is also interesting to verify whether the agents' reaction to equilibrium error is the same for any size of deviation from the equilibrium path for interest-free vis-à-vis interest-bearing demand for money. Specifically, if the agents' reaction to a small deviation from equilibrium can be ignored while their reaction to a large deviation is drastically large, then the error-correcting term will be nonlinear in the error correction model (ECM). This paper will also investigate this important issue.

This study uses the extended data of Yousefi *et al.* (1997). Iranian data was chosen because of the following reasons: (i) Iranian data may be more appropriate in testing the stability of interest-free demand for money as Iran officially announced an interest-free financial system, and (ii) Iran has experienced a wider range of real and monetary shocks over the sample period than any other country, because of periods of political upheaval and several changes in monetary policy regimes. In particular, during the post-revolutionary period, there have been dramatic changes in monetary authorities towards the interest rate. Note that parameters of demand for money may vary because of (a) changes in the environment, (b) changes in economic policy control rules and (c) changes in the environment, which alter expectations (Favero and Hendry, 1992). In fact, all of these conditions are relevant to the economy of Iran. Therefore, Iran provides us with an interesting testing ground for demand-for-money functions. Consequently, this study should be of interest to monetary economists in general. The data used in Yousefi *et al.* (1997) ends late 1992 while the data in this study ends in 1998Q4.[1] As a matter of fact, it turns out the data since late 1992 are very informative about money demand.

As Figure 1 shows, the country's inflation rate (measured as the annual growth rate of GDP deflator) went up to a record level of about 60 percent before falling to less than 10 percent. This was mostly due to the end of the Iraq-Iran war. The superexogeneity test developed by Engle and Hendry (1993) was used to investigate the behavior of economic agents for interest-bearing and interest-free demand-for-money functions. The Maximum Likelihood test developed by Johansen and Juselius (1991) and the Dynamic OLS test developed by Stock and Watson (1993) were used to estimate these long-run demand-for-money functions. The latter test was also used to verify the stability of the long-run demand functions.

It was found that both short- and long-run demand for interest-free money functions are stable and their coefficients are invariant with respect to policy and other exogenous shocks as well as changes in regime. By contrast, short- and long-run interest-bearing demand-for-money functions were found to be unstable. It was also found that agents in demanding interest-bearing assets are forward-looking and their expectations are formed rationally in Iranian financial markets. This result implies that the coefficients of interest-bearing demand for money are not policy invariant. Namely, a monetary policy shock results in a change of the coefficients of interest-bearing demand for money so that the effectiveness of the shock will be uncertain. Finally, it was found while the agents' reaction to equilibrium errors for interest-free demand for money is always the same for any error size, they may react differently to different magnitudes of deviation from the desired level of interest-bearing demand for interest-bearing money. Namely, it was found the short-run dynamic demand for interest-bearing money is nonlinear. The nonlinear part of the error in the demand for interest-bearing money may be ignored for a small error equilibrium while agents react drastically to any large equilibrium error size.

The findings of this paper are important to both policy makers and academicians since, according to Darrat (2002), there are about two hundred interest-free financial institutions in over sixty countries, including non-Muslim countries like Australia, the Bahamas, Canada, Germany, France, Luxembourg, South Africa, Switzerland, the U.K. and the U.S. Furthermore, since the stability of demand for money is essential for the effectiveness of monetary policy and, as it was mentioned earlier, the fact that many studies found an unstable demand for money in many countries, the issue of the interest-free, but risk-sharing banking system becomes more important even in the Western world. Finally, according to

[1] The source of data is *International Financial Statistics,* International Monetary Fund (CD-ROM March 2001). Many thanks to Professor Sohrab Abizadeh who provided me with the series used in Yousefi *et al.* (1997). The series used in Yousefi *et al.* (1997) were also obtained from *International Financial Statistics* CD-ROM. There are some missing observations in earlier years, which are filled from the series provided by Professor Abizadeh. The missing observations are: for M1 series, from the second quarter of 1984 to the first quarter (inclusive) of 1986; for Consumer Price Index, from the third quarter of 1986 to the second quarter (inclusive) of 1988, and finally, for quasi-money (interest-bearing time and saving deposits), the last quarter of 1978 and 1984 as well as from the second quarter of 1985 to the first quarter (inclusive) of 1986. As in Yousefi *et al.* (1997), quarterly data was used and, following Yousefi *et al.* (1997), quarterly data on GDP and GDP deflator was generated according to Diz's (1970) specifications. For a simple and very clear explanation on generating quarterly GDP data from annual observations, see the appendix in Yousefi *et al.* (1997).

the findings of this paper, if the central bank uses money-growth targets to reduce inflation, it should focus on the interest-free monetary aggregate.

The following section deals with the velocity of money, the theoretical model as well as the long-run estimation results. Section 3 is devoted to conditional and marginal models, as well as the superexogeneity test and the long-run stability test results. The final section is devoted to concluding remarks. The appendix is devoted to the full derivation of the superexogeneity test.

## 2. Demand for Money

### *2.1 The Velocity of Money*

On March 21, 1984, the Iranian government banned the payment of interest on all lending and borrowing activities with the exception of ordinary transactions of the Central Bank with the government, government institutions, public enterprises and banks as long as these institutions use their own resources. However, banks were allowed, based on their profitability, to pay a return on saving and time deposits. This led to minimum rates of return, depending on the term to maturity, on time deposits. As of the second quarter of 2001, these minimum rates have remained constant since the introduction of Islamic banking in Iran in 1984. These rates are as follows: short-term 8 percent; special short-term 10 percent; one-year 14 percent; two-year 15 percent; three-year 16 percent and five-year 18.5 percent (Central Bank of the Islamic Republic of Iran, 2000-2001, p. 23). For a detailed explanation on these issues, see Yousefi *et al.* (1997) and references therein.

Following Yousefi *et al.* (1997) and Darrat (2002) and based on banking institutions mentioned above we can consider M1 (i.e., demand deposits - which do not pay interest in Iran - plus currency with the public) as interest-free money supply in Iran. Furthermore, M2 (i.e., M1 plus quasi-money, defined as saving and term deposits, which pay interest) as interest-bearing money supply. It is interesting to verify if the addition of six-year data will change the velocity of money analysis reported in tables 1(a) and 1(b) of Yousefi *et al.* (1997). Let us define the velocity of money as: V1 = gdp/ M1 and V2 = gdp/ M2 where gdp is the nominal GDP. Table 1 reports the behavior of velocity of M1 and M2.

A comparison of the results reported in Table 1 and of those reported in tables 1(a) and 1(b) of Yousefi *et al.* (1997) indicates minor differences between similar sub-periods. This is due to a revision in IMF financial statistics. Similar to the finding of Yousefi *et al.* (1997), the volatility of both velocities has fallen drastically after the introduction of the Islamic banking system. The addition of six years of data stresses this fact further. Namely, while velocity of M1 varies within a range of 2.98-11.15 (with a standard deviation of 1.95) over the pre-Islamic banking period, it varies within a range of 2.98-5.89 (with a standard deviation of 1.01) after that period. A similar result is observed for the velocity of M2, that is, while the velocity varies within a range of 1.66-4.98 (with a standard deviation of 0.91) over the pre-Islamic banking period it varies within a range of 1.65-2.77 (with a standard deviation of 0.37) after that period. Following Yousefi *et al.* (1997), we can conclude that velocity is less volatile over the Islamic banking period.

### *2.2 Long-run demand for money*

In order to stay within the framework of this literature, the following typical demand function, which was used by Darrat (1988) and Yousefi *et al.* (1997) will be estimated.

$$lrm_t = \beta_1 \, lrgdp_t + \beta_2 \, R_t + u_t, \qquad (1)$$

where lrm is the logarithm of real money, lrgdp is the logarithm of real GDP, R refers to yields expected on real assets or the interest rate and u is the disturbance term which is assumed to be white noise with zero mean. β's are parameters to be estimated. Following Darrat (1988) and Yousefi *et al.* (1997), we assume expectations are static and the actual inflation rate (growth rate of GDP deflator) as a good proxy for return to real assets. Equation (1), consequently, will be

$$lrm_t = \beta_1 \, lrgdp_t + \beta_2 \, infgdp_t + u_t, \qquad (2)$$

where infgdp is the growth rate of GDP deflator. Equation (2) is a long-run semi-log linear demand for money. Note that demand for money for M1 and M2 in both Darrat (1988) and Yousefi *et al.* (1997) is a short-run relationship, but here we have a long-run version of their demand-for-money equation. Namely, there is no lag dependent variable in the equation. It should also be mentioned that in a recent study Bahmani-Oskooee (1996) includes the logarithm of exchange rate (once the official rate and once the black market rate) in the above equation. He finds only long-run demand for M2 is a function of the black market exchange rate. Neither demand for M1 nor M2 is a function of the official exchange rate. However, as it was mentioned earlier, for the sake of comparison and consistency with the stability study on interest-free demand for money, the demand function used by both Darrat (1988) and Yousefi *et al.* (1997) will be estimated. Furthermore, Equation (2) is similar to the model used by Stock and Watson (1993) and Muscatelli and Spinelli (2000) if we assume interest rates were zero in their model, and is similar to the one used by Chen (1997).

As stationarity test results (reported in Table 2) indicate, all variables, except 'infgdp' are integrated of degree one (non-stationary). They are, however, first-difference stationary.[2] Consequently, we will first verify if long-run relationships exist between the level of M1 and M2 and their determinants, as specified in

[2] This result is similar to what was found, for example, for Italy by Muscatelli and Spinelli (2000).

Equation (2).[3] If a cointegrating relation exists then short-term departures from equilibrium relationship between these variables are eliminated over the long run by market forces and monetary or fiscal policies. Namely, a long-run stable demand for money for M1 and M2 may exist.

In determining the lag length one should verify if the lag length is sufficient to get white noise residuals. LM(1) and LM(4) will be employed to confirm the choice of lag length. The order of cointegration (r) will be determined by using Trace and $\lambda_{max}$ tests developed in Johansen and Juselius (1991). Following Cheung and Lai (1993), both tests were adjusted in order to correct a potential bias possibly generated by small sample error; see footnote to tables 3 and 4 for the formulas. Tables 3 and 4 report the result of $\lambda_{max}$ and Trace tests for lag length of five and four quarters (k=4 and 5) for M1 and M2, respectively. According to diagnostic tests reported in the tables, there is no autocorrelation. The only non-congruency is non-normality for M1. However, as was mentioned by Johansen (1995), a departure from normality is not very serious in cointegration tests, see also, for example, Hendry and Mizon (1998). The significant non-normality statistic is, however, due to outliers in 1972, 1973 and 1979. According to the result of Table 3, the $\lambda_{max}$ test rejects r=0 at the 10 percent level while we cannot reject r≤1, implying that r=1. According to Trace test, we reject the null hypothesis of r=0 at the 10 percent level while we cannot reject the null hypothesis of r≤1, implying that r=1. Consequently, at least one cointegrating relationship exists between M1 and its determinants at the 10 percent level.

According to the result of Table 4, the $\lambda_{max}$ test rejects r=0 at the 5 percent level while we cannot reject r≤1, implying that r=1. According to Trace test, we reject the null hypothesis of r=0 at the 5 percent level while we cannot reject the null hypothesis of r≤1, implying that r=1. Consequently, at least one cointegrating relationship exists between M2 and its determinants at the 5 percent level. Here demand for money M2 has a stronger long-term relationship than demand for M1. These long-run relationships are reported in Table 5 under the heading Maximum Likelihood.

The Stock and Watson's (1993) dynamic OLS (DOLS) estimator was also used to estimate Equation (2) for M1 and M2.[4] The estimation result is reported under the heading DOLS. Since $infgdp_t$ is stationary the short-run part of the DOLS estimator was accordingly modified and Equation (2) for M1 and M2, using the modified version, was estimated (see Footnote *** of Table 5). The estimation result is reported under the heading DOLSM.

The coefficient of income and inflation rate has the correct sign for both demand functions according to Johansen and Juselius's (1991) Maximum Likelihood estimator result. According to the exclusion test ($\chi^2$) the estimated coefficient of income is not statistically significant. However, since all variables, except $infgdp_t$, are non-stationary, unless there is weak exogeneity, the asymptotic distribution of the estimator of $\beta$ does not permit the use of the usual $\chi^2$ distribution, even though the estimator of $\beta$ is consistent (Johansen, 1992).[5] Table 6 reports the weak exogeneity tests. According to the result, both variables $lrgdp_t$ and $infgdp_t$ are individually and jointly weakly exogenous for $\beta$ in the long-run demand for M1 and M2 relationships and therefore, the use of $\chi^2$ may be permitted. Furthermore, this result also means that the marginal models of $lrgdp_t$ and $infgdp_t$ do not react to equilibrium errors.

The DOLS Wald test result also confirms a long-run relationship for both interest-free and interest-bearing demand for money. However, the coefficient of the inflation rate has the wrong sign for both interest-free and interest-bearing demand-for-money functions, though it is not statistically significant. The existence of a cointegrating relationship between the levels of variables in Equation (2), for both M1 and M2, indicates that a valid error correction model (ECM) for both M1 and M2 exists. However, the existence of an ECM for the demand for money does not only indicate that economic variables determining the demand for money adjust to past equilibrium errors, but it may also be due to changes of economic agents' forecasts of future income, monetary policy and the inflation rate. Then the ECM parameters may no longer be invariant to the process of forcing variables (policy and other exogenous shocks) as mentioned by Lucas (1976). Furthermore, it is also possible for the contemporaneous variables in the ECM to be endogenous due to the violation of weak exogeneity of the variable.

In either case, at least one of the parameters may vary with changes in the expectation process. That is, at least one of the variables in ECM fails to be superexogenous in the sense of Engle *et al.* (1983) and Engle and Hendry (1993). In such a case monetary policy may not be effective, as the parameters of demand for money will vary with a change of regime or any policy shock. Under the Islamic banking system, the interest rate exposure is completely eliminated

[3] Note that in a multivariate cointegrating relationship we need at least two variables to be non-stationary.

[4] Stock and Watson's (1993) test (DOLS) is based on the following regression: lrmt = β0 + β1 lrgdpt + β2 infgdpt + δ1(L) Δ infgdpt + δ2(L) Δlrgdpt + ut,, where $\delta_1(L)$ and $\delta_2(L)$ have two leads and lags as suggested by Stock and Watson for the number of observations of 100 or close to 100.

[5] This is due to the fact that as the number of observations becomes infinity large, the mean of the variables approaches to its true value, and the distribution of, say, $((E(x_t) - x_t) / \sqrt{n})$, for x=lrm1, lrm2 and lrgdp approaches quickly to the normal, but the variance of the estimator may explode quite fast as $n \to \infty$. Thus, no matter how large the sample is the standard central limit theorem may not apply.

due to the sharing of profit or loss among depositors, banks and investors. Consequently, banks' balance sheets are more stable. This leads to a stable demand for money as argued in this literature,for example, Darrat (1988). Furthermore, because of the stability of the banking system's balance sheet and risk-sharing strategy, one would also expect the demand for money to be invariant to policy shocks. The next section concentrates on this issue.

## 3. Conditional and Marginal Models: Superexogeneity Results and Long-run Stability

Having established in the previous section that a long-run relationship to describe demand for M1 and M2 exists, we need to verify if variables in each of these demand-for-money relationships are also superexogenous, i.e. these demand-for-money relationships are invariant to policy or other exogenous shocks. This requires a superexogeneity test of the variables. It is therefore necessary to specify the ECM implied by our cointegrating vectors. The ECM term generated from the long-run relationship estimated with the Maximum Likelihood Estimation technique and reported in Table 5 will be used for the following reasons: (a) to be consistent with the superexogeneity test literature (e.g., Favero & Hendry (1992) & Engle & Hendry(1993)), (b) contrary to the Dynamic OLS estimator result, inflation variable has the right sign.

### *3.1 Error-Correction Results: Conditional Models*

To be consistent with the literature, we assume, in determining the lag length, that agents incorporate current available information as well as past information up to three years. Consequently, the lag length of 12 was chosen.[6] Given this lag length, the parsimonious ECM was obtained by engaging in a general-to-specific modeling procedure. Following Granger (1986), we should note that: (a) the inclusion of a constant in ECM makes the mean of error zero, and (b) if small equilibrium errors can be ignored by agents, while they react substantially to large ones, the error correcting equation is nonlinear.[7] It should be noted that error term EC is a generated regressor and its t-statistic should be interpreted with caution (Pagan (1984) and (1986)). To cope with this problem, following Pagan (1984) and (1986), the instrumental variable estimation technique was implemented, where the instruments are fourth and twelfth lagged values of the error terms for both M1 and M2.

[6] It should be noted that in ECM we allow agents to be backward looking (reacting to previous deviations from equilibrium) while they may also be forward looking.

[7] Escribano (1985) originally developed a nonlinear error-correction model, in a restricted form, for the demand for money. His model was used, among many others, by Hendry & Ericsson (1991) and recently Teräsvirta & Eliasson (2001) developed two unrestricted versions of the model. This paper, however, uses a data-determined unrestricted nonlinear error-correction model.

Tables 7 and 8 report the estimation results of the ECM model for M1 and M2, respectively. Except for the normality test for M1 none of the diagnostic checks is significant. The significant non-normality statistic for M1 is due to three large outliers in 1974 (significant increase in the country's oil income), 1979 (revolution) and 1988. According to Hansen's stability L test result (5 percent critical value=0.47, Hansen (1992), Table 1), all of the coefficients, except the coefficient of $\Delta lm1_{t-4}$ for M1 and $\Delta infgdp_{t-3}$ for M2, are stable.

Furthermore, the joint Hansen's (1992) stability $L_c$ test result is 2.44 (<2.75 for 11 degrees of freedom) for M1 and 2.44 (<3.15 for 13 degrees of freedom) for M2, which indicates that we cannot reject the null of joint stability of the coefficients together with the estimated associated variance. However, as indicated by the Chow test result, reported in the last row of tables 7 and 8 for a break point (1984 first quarter), there is a structural break in the demand for interest-bearing money, but not in the interest-free demand for money. Note that in March 1984 the Islamic banking system was implemented in Iran, and by choosing the first quarter of 1984 as a break point we can almost split the sample into two equal sub-samples. The above result is consistent with the findings of Darrat (1988), Yousefi *et al.* (1997), Hassan and Mazumder (2000) and Darrat (2002).[8]

The only contemporaneous variable in both short-term demand equations is the change in quarterly inflation rate with the correct sign. All possible kinds of non-linear specifications, that is, squared, cubed and fourth powered of the equilibrium errors (with statistically significant coefficients) as well as the products of those significant equilibrium errors were included. According to our estimation results, the error-correction term is significant for both M1 and M2, but the impact of equilibrium error on the growth of money demand for M2 is nonlinear. Namely, the agents' reaction to equilibrium errors (departure from the desired level for M2) varies for different error sizes. For a small equilibrium error the nonlinear part may not be as important, but for a very large error the agents' reaction will be drastic, even though the coefficient of the nonlinear part is smaller than the coefficient of the linear part. To the best of my knowledge, there is no study so far on the error correction model for interest-free demand for money in the literature. The result on demand for M2 is consistent with, for example, Hendry and Ericsson (1991) and Ericsson *et al.* (1998) for U.K. It is also consistent with Bahmani-Oskooee and Bohl (2000) for Germany even though they used a linear EC model.

In sum, it was found in this section that while a linear and stable error-correction model for interest-free demand for money exists, the error correction model for

[8] Note that none of these studies estimated an error correction model for the demand for money, but investigated short-run stability of interest-free vis-à-vis interest-bearing demand for money.

interest-bearing demand for money is nonlinear and may not be stable. Having established that an ECM for both interest-free and interest-bearing demand for money exists, we need to verify whether the coefficients of these money demand equations are invariant to the process of forcing variables. This requires the contemporaneous variable in the ECM, that is, $\Delta infgdp_t$, to be superexogenous. Consequently, we need to establish a marginal model for variable $\Delta infgdp_t$.

### *3.2 Marginal Model*

There have been several potential regime changes over the sample period: (i) the Iranian Revolution in April 1979; (ii) the introduction of the Islamic banking system on March 21, 1984[9], (iii) the introduction of the first privately owned financial institution after the revolution; "In September 1997 the first non-bank credit institution, 'Credit Institution for the Development of Construction Industry', was established by the private sector [...] According to Constitution, private sector cannot own and operate banks, thus non-bank credit institutions have been created to promote competition and provision of services. The main advantages of these institutions lie in the lower transaction costs of their operations, quicker decision-making ability, customer orientation and prompt provision of services." (Central Bank of the Islamic Republic of Iran (1997-98), p. 10); (iv) the introduction of inflation rate target by Central Bank of the Islamic Republic of Iran. Starting March 21,1995, the Central Bank determined ceilings for banking facilities to curb inflation rate (Central Bank of the Islamic Republic of Iran, 1995-96). The imposition of credit ceiling facilities was removed in March 1998 (Central Bank of the Islamic Republic of Iran (1999-2000a), Appendix III).[10]

Dummy variables were created for step changes, that is, (i) Rev = 1 for 1979 (second quarter) and after, zero otherwise, (ii) Zero=1 for 1984 (first quarter) and after, zero, otherwise, (iii) War = 1 for period 1980 (fourth quarter)-1988 (third quarter) and zero otherwise,[11] (iv) Private=1 for period 1997 (third quarter)-1998 (fourth quarter) and zero otherwise, and (v) Inflation = 1 for period 1995 (second quarter)-1998 (first quarter) and zero otherwise.

Level and interactive combinations of these dummy variables were tried for the impact of these potential shift events in the marginal models for $\Delta infgdp_t$ and any first round significant effects were retained. The resulting marginal model is reported in Table 9. According to this result the marginal model passes the diagnostic checks for residual autocorrelation, heteroskedasticiy and the RESET tests. However, it fails for the normality of the residual. Such failure is common in the estimation of marginal equations (Hurn and Muscatelli, 1992; and Metin, 1998).

Overall, the estimated equation seems a reasonable marginal model for the analogues of conditional mean of $\Delta infgdp$. Clearly, there is evidence of the structural break in this equation, that is, possible break points are due to the introduction of the Islamic banking system, the introduction of inflation target and the Iraq-Iran war. Note that non-constancy of the marginal model is related to the concept of superexogeneity, which implies that the parameters of the conditional model remain constant if agents are not forward looking.

In the marginal model the 'Zero', 'Inflation' and 'War' dummy variables are significant. All dummy variables affect the slope. According to the estimation results, the relationship between the change in inflation rate and the growth of income is stronger after the introduction of Islamic banking and during the implementation of the inflation rate target by the Central Bank. Furthermore, during such implementation, the overall impact of the growth of M1 on the inflation rate, as would be expected, has been negative. During the war period, however, the reverse happened.

### *3.3 Superexogeneity Test and Results*

In this section we need to verify if the contemporaneous variable $\Delta infgdp_t$ in our ECM models fails to be superexogenous. Letting $Z_t=\Delta infgdp_t$ and following Engle *et al.* (1983), Engle and Hendry (1993) and Psaradakis and Sola (1996), we can write the relationship between $\Delta\log(Mi_t)$ and $Z_t$ as:

$$\Delta\log(Mi_t) = \alpha_0 + \psi_0 Z_t + (\delta_0 - \psi_0)(Z_t - \eta^Z{}_t) + \delta_1 \sigma_t^{ZZ}(Z_t - \eta^Z{}_t) \qquad (3)$$
$$+\psi_1 (\eta^Z{}_t)^2+\psi_2 (\eta^Z)^3+\psi_3 \sigma_t^{ZZ} \eta^Z{}_t+\psi_4 \sigma_t^{ZZ} (\eta^Z{}_t)^2+ \psi_5(\sigma_t^{ZZ})^2\eta^Z+z'_t\gamma+ ui_t$$

where $\alpha_0$, $\psi_0$, $\psi_1$, $\psi_2$, $\psi_3$, $\psi_4$, $\psi_5$, $\delta_0$ and $\delta_1$ are regression coefficients of $\Delta\log(Mi_t)$, for i=1 or 2, on $Z_t$ conditional on $z'_t\gamma$, and term $ui_t$ is assumed to be, as before, white noise, normally, identically and independently distributed. Vector z includes past values of $\Delta\log(Mi_t)$, $Z_t$, and other explanatory variables in the ECM as well as current and past values of other valid conditioning variables in the ECM. Furthermore, $\eta^Z{}_t=E(Z_t \mid I_t)$ and $\sigma_t^{ZZ}=E[(Z_t - \eta^Z{}_t)^2 \mid I_t]$ are the conditional moments of $Z_t$, given information set $I_t$ which includes the past values of $\Delta\log(Mi_t)$, for i=1 or 2, and $Z_t$ as well as the current and past values of other valid conditioning variables included in $z_t$. See the appendix for a detailed

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[9] For a detailed explanation see Yousefi *et al.* (1997).

[10] Note that "... in accordance with article 19 of the Interest-Free Banking Act of 1983 which stipulates that short-term credit policies need to be approved by the government and long-term credit policies have to be incorporated within the Five Year Development Plan documents and approved by the parliament." (Central Bank of the Islamic Republic of Iran (1999-2000b), Appendix II,- p. 25).

[11] This dummy variable was created to reflect the impact of the Iraq-Iran war. It is true that the impact of a war may not be similar to the implication of a specific policy, but governments usually behave differently in reaction to the cost of a war than the usual government expenditure. Economic agents also react differently, and if they are forward looking they should also behave differently in expectations of possible post-war expansionary policy, which would reflect the rebuilding of the country.

explanation of the formulation of superexogeneity and invariance hypothesis associated with the conditional model (ECM) for M1 and M2.

Note that $Z_t$ can be a control/target variable that is subject to policy interventions. Under the null of weak exogeneity, $\delta_0$-$\psi_0$=0. Under the null of invariance, $\psi_1=\psi_2=\psi_3=\psi_4=\psi_5=0$ in order to have $\psi_0=\psi$. Finally, if we assume that $\sigma_t^{ZZ}$ has distinct values over different, but clearly defined regimes, then under the null of constancy of $\delta$, we need $\delta_1$=0. If all these hypotheses are accepted the contemporaneous variable in the ECM is superexogenous and coefficients of the money demand equation (ECM) for M1 or M2 are invariant to policy shocks.

From the marginal model, reported in Table 9, estimates of $\eta^Z$ and $\sigma_t^{ZZ}$, for Z=$\Delta$infgdp$_t$ were calculated. As for $\sigma_t^{ZZ}$, since the error for $\Delta$infgdp$_t$ variable, according to ARCH test, is not heteroskedastic, a five-period moving average of the variance of the error was tried. All of these constructed variables were then included in the ECM reported in tables 7 and 8. The models were re-estimated and the estimation results on these constructed variables are given in Table 10. Except for the normality test for M1 none of the diagnostic checks reported in the table is significant. The significant non-normality statistic for M1, as before, is due to three large outliers in 1974, 1979 and 1988.

The individual F-test is on the null hypothesis that the coefficient of each variable is zero. The F-test on the null hypothesis that all constructed variables are jointly zero is given in the last row of the table. As the estimation result in Table 10 shows, the joint F-test on the null hypothesis that coefficients of these constructed variables are jointly zero is not significant for M1, indicating that these variables together should not be included. This result immediately implies that the contemporaneous variable ($\Delta$infgdp$_t$) in the conditional model, reported in Table 7, is superexogenous, and the interest-free demand for money is invariant to policy shocks.

Moreover, the F-test on the null hypothesis that coefficients of all constructed variables are jointly zero is significant at the conventional level for M2, indicating that these variables together should be included, and, therefore, variable $\Delta$infgdp$_t$ is not superexogenous in the conditional model for the interest-bearing demand for money (M2). However, the coefficient of (Z-$\eta^Z$) of the contemporaneous variable $\Delta$infgdp$_t$ is statistically insignificant, implying this variable is weakly exogenous and the inferences on the parameters of the agents' model (conditional model for M2) reported in Table 8 are efficient.

The coefficient of $\sigma^{ZZ}$(Z-$\eta^Z$) is not significant implying that the null of constancy cannot be rejected for this variable.[12] Since the coefficient of $(\eta^Z)^3$ is significant, the null of invariance with respect to policy changes is rejected for $\Delta$infgdp$_t$ variable in the conditional model for M2. As it was noted before, constancy and invariance are different concepts. Parameters could be constant over time, but not invariant with respect to policy changes. Furthermore, as mentioned by Engle and Hendry (1993), all three conditions must be satisfied in order to ensure superexogeneity. The failure of the invariance condition, therefore, justifies the result of the joint F-test, rejecting the null hypothesis that all coefficients of the constructed variables are jointly zero. In general, we reject the null hypothesis that the conditional model for interest-bearing demand for money is invariant to policy changes. Hence, the above result means that, for a given growth of real income, the monetary policy that alters the inflation rate will affect coefficients of interest-bearing money demand.

It should be noted that even when superexogeneity holds (as for the demand for M1), policy can and, in fact, does impact agents' behavior by affecting the variables entering the conditional model, *albeit* not through the parameters of that model. In our model the policy might well affect the inflation rate and income and so the demand for M1. More explicitly, as mentioned by Ericsson *et al.* (1998, p. 320), "...under super exogeneity, the precise mechanism that the government adopts for such a policy does not affect agents' behavior, except insofar as the mechanism affects actual outcomes."

Following Psaradakis and Sola (1996), I also simplified both conditional models by sequentially deleting variables with insignificant coefficients. The final specification for M1 included $(\eta^Z)^3$ with the coefficient of -18.36 and a t-ratio of -1.018, so the superexogeneity of $\Delta$infgdp$_t$ variable in M1conditional model is further verified. As for the conditional model of M2 the final specification included $(\eta^Z)^3$ with the coefficient of -33.40 and a t-ratio of –3.16, so the superexogeneity of $\Delta$infgdp$_t$ variable in M2 conditional model is again rejected.

Furthermore, as it was mentioned by Psaradakis and Sola (1996), since structural invariance implies that the determinant of parameter non-constancy in the marginal process should not affect the conditional model, I also tested the significance of the dummy variables in both conditional models. None of the dummy variables was significant in the conditional model for M1. The final estimate of the conditional model for M2 included dummy variable War with a coefficient of –0.02 and a t-ratio of –3.01 as well as ($\Delta$infgdp$_t$)*( Zero$_t$) with the coefficient of –0.30 with a t-ratio of –3.46 and ($\Delta$infgdp$_{t-4}$)*(Zero$_{t-4}$) with the

[12] Note that the coefficient of the contemporaneous variable being constant within a specified regime does not contradict the Chow test result reported in Table 8, which indicates the change of the regime resulted in a structural break in the dynamic demand for M2.

coefficient of 0.30 with a t-ratio of 2.61. This result further supports the rejection of structural invariance of the interest-bearing demand for money.

Although the t-test result on a generated variable is not reliable, I also estimated, for the sake of curiosity, both error correction models using the actual values of the error terms. The results were not materially different from the reported results. For the sake of brevity these results are not reported, but are available upon request. However, a brief elaboration on these unreported results might be fruitful. The White test result on demand for M1 (White=83.38, p-value=0.06) was weaker than the one reported. All other results including the superexogeneity test result were the same as or very close to the reported results.

The demand for M2, however, had a different specification. Namely, all the coefficients of the lag values of the change in inflation rate were statistically insignificant and had to be dropped from the regression. The standard error of the regression was higher than what is reported, that is, it was 0.05 rather than 0.03. Consequently, the model estimated by the instrumental variable technique (reported) variance-dominates the unreported model. Furthermore, the error was heteroskedastic in the ARCH sense (ARCH=12.77, p-value 0.025). I, therefore, used robust-error estimation to correct the standard errors. It should also be mentioned that the superexogeneity test result, in rejecting the null hypothesis that the contemporaneous variable in the demand for M2 is superexogenous, was stronger ($\chi^2$=24.36, p-value=0.0009)[13] than what is reported. Namely, if the actual values of errors rather than the instruments were used, the rejection of the null of policy invariance for demand for M2 would be stronger. All other results were the same as or close to the reported results.

In sum, it was shown in this section that while the coefficients of interest-free demand for money are constant and invariant to policy shocks, the interest-bearing demand for money is unstable and has coefficients that are not invariant to policy interventions. To the best of my knowledge, there is no study so far in the literature that investigated the policy invariance of interest-free demand for money and therefore no comparison is possible.

### *3.4 Long-Run Money Demand Stability*

Having established that the coefficients of interest-free demand for money are constant and invariant to policy shocks, while the interest-bearing demand for money is unstable with or without policy interventions, we need to investigate the stability of long-run demand for both interest-free and interest-bearing demand for money. Following Ball (2001), I will use Stock and Watson's (1993) DOLS to test for the long-run stability. Since, according to the results in the previous section, the short-run dynamic of demand for interest-bearing money is not stable, I will allow the intercept and the slope of the short-run variables $\Delta infgdp_t$ and $\Delta lrgdp_t$ to be different in pre- and post-Islamic banking systems. However, I will also assume the short-run dynamic of this demand for money to be constant and will report both results. The estimation result is given in Table 11. The coefficient on income is close to one for the entire period for both long-run demands for money. This result is consistent with what was documented by Ball (2001) and Stock and Watson (1993) for the U.S. interest-bearing demand for money over the 1903-1994 and 1903-1989 periods, respectively.

The coefficient of inflation rate except for the recent sub-sample has the wrong sign. According to the stability test result, there is little evidence that the long-run coefficients of interest-free demand for money are unstable. However, the stability test result for the coefficients of the interest-bearing demand for money produces strong evidence against stability. Since again, to the best of my knowledge, there is no study so far in the literature that investigated the stability of long-run interest-free demand-for-money function, no comparison is possible for the former result. However, the latter result is consistent with the findings of Ball (2001) and Stock and Watson (1993) for the U.S. interest-bearing demand for money and Hamori and Tokihisa (2001) for Japan. In sum, it was found in this section that the long-run coefficients of interest-bearing demand for money are not stable while the coefficients of long-run interest-free demand for money are stable.

## 4. Concluding Remarks

This paper, using quarterly Iranian data for the period 1966-1998, extended the literature by investigating the long-run stability of interest-free demand-for-money function. It also examines whether the coefficients of both interest-free and interest-bearing demand-for-money functions are invariant with respect to policy shocks.

It was found that both short- and long-run interest-free demand-for-money functions are stable and their coefficients are invariant with respect to policy and other exogenous shocks as well as changes in regime. By contrast, interest-bearing demand for money was found to be unstable. Furthermore, agents in demanding interest-bearing assets are forward looking and their expectations are formed rationally in the financial markets in Iran. Finally, it was found that while the agents' reaction to equilibrium errors for interest-free demand for money is always the same for any error size, they may react differently to different magnitudes of deviation from the desired level of interest-bearing demand for money. Namely, the error correction model for interest-bearing demand for money is nonlinear. This paper has the following

[13] Note that when robust-error estimation technique is used $\chi^2$ rather than F-test is a valid test for the exclusion of variables in the regression.

policy implication: if money-growth targets are used to reduce inflation, then the interest-free monetary aggregate is more reliable than other aggregates.

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**Figure 1:**

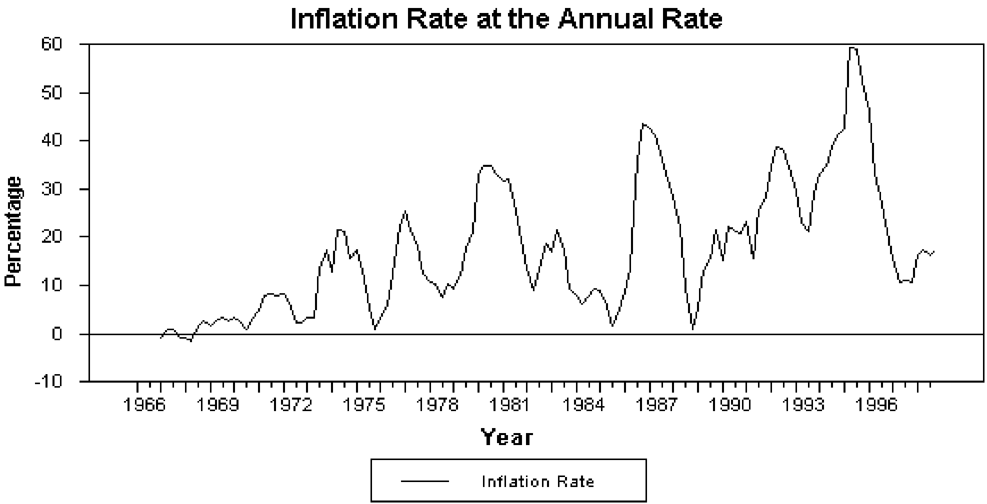
Inflation Rate at the Annual Rate
Percentage
60
50
40
30
20
10
0
-10
1966
1969
1972
1975
1978
1981
1984
1987
1990
1993
1996
Year
Inflation Rate

**Table 1: Velocity of M1 and M2***

| Period | Minimum | Maximum | Mean | Standard Deviation |
|---|---|---|---|---|
| **V1=gdp/M1** | | | | |
| 1966:1-98:4 | 2.98 | 11.15 | 5.26 | 1.80 |
| 1966:1-83:4 | 2.98 | 11.15 | 6.04 | 1.95 |
| 1984:1-98:4 | 2.98 | 5.89 | 4.31 | 1.01 |
| **V2=gdp/M2** | | | | |
| 1966:1-98:4 | 1.65 | 4.98 | 2.67 | 0.86 |
| 1966:1-83:4 | 1.66 | 4.98 | 3.11 | 0.91 |
| 1984:1-98:4 | 1.65 | 2.77 | 2.15 | 0.37 |

Notes: * gdp is the nominal GDP. M1 (demand deposits which do not pay interest in Iran plus currency with the public) is interest-free money supply while M2 is M1 plus quasi-money, i.e., saving and term deposits which pay interest.

**Table 2: Stationary Tests, 1966, Q1 - 1998Q4* (Absolute Values)**

| Variables | Augmented Dickey-Fuller $\tau$-Stat. | Phillips-Perron Z-Stat. |
|---|---|---|
| **Levels:**** | | |
| lrm1 | 1.49 | 1.19 |
| lrm2 | 1.91 | 1.49 |
| lrgdp | 2.36 | 1.87 |
| infgdp | 3.99[b] | 9.27[a] |
| **Changes of:** | | |
| lrm1 | 4.81[a] | 15.09[a] |
| lrm2 | 5.05[a] | 12.74[a] |
| lrgdp | 4.31[a] | 7.14[a] |

Notes: * All tests include constant and trend. The critical value for Augmented Dickey-Fuller $\tau$ test (lag-length = 4) and for Phillips-Perron non-parametric Z test (window size = 4) is 3.43 at 5 percent and 3.99 at 1 percent. The number of observations is 131.

a=Significant at 1 percent.

b=Significant at 5 percent.

** lrm1 is the log of the real M1, lrm2 is the log of real M2, and lrgdp is the log of real GDP, where all deflated at Consumer Price Index (CPI). infgdp is the annualized growth rate of GDP deflator.

**Table 3*: Tests of the Cointegration Rank - M1: Period 1966Q1-1998Q4**

| $H_0$=r | Eigenv.=$\hat{\lambda}$ | $\lambda_{max}^{(1)}$ | $\lambda_{max}$90 | $\lambda_{max}95^{(2)}$ | Trace$^{(3)}$ | Trace90 | Trace95$^{(4)}$ |
|---|---|---|---|---|---|---|---|
| 0 | 0.1293 | 15.35 | 13.39 | 25.54 | 27.53 | 26.70 | 29.38 |
| 1 | 0.0730 | 9.41 | 10.60 | 18.96 | 12.17 | 13.31 | 15.34 |
| 2 | 0.0333 | 3.77 | 2.71 | 12.25 | 3.76 | 2.71 | 3.84 |
| Diagnostic tests**: | | | | | | | |
| LM(1) | p-value = 0.27 | | | | | | |
| LM(4) | p-value = 0.33 | | | | | | |
| Normality | p-value = 0.00 | | | | | | |

Notes: (1) $\lambda_{max}$ is adjusted to correct the small sample bias error. Namely, N is replaced by (N–kp). $\lambda_{max}$ = -(N-kp) ln(1- $\hat{\lambda}_r$), where N is the number of observations, k is the number of lag length and p is the number of endogenous variables, see Cheung and Lai (1993).

(2) The source is Osterwald-Lenum (1992), Table 2, p. 469.

(3) Trace test is adjusted to correct the small sample bias error. Namely, N is replaced by (N–kp). Trace test = -(N-kp) $\sum_{i=r+1}^{P} \ln(1-\hat{\lambda}_i)$ , see Cheung and Lai (1993). Both Trace and $\lambda_{max}$ tests were developed in Johansen and Juselius (1991).

(4) The source is Johansen (1995), Table 15.3, p. 215.

* The regression does not include time trend. Lag length is 5.

** LM(1) and LM(4) are one and four-order Lagrangian Multiplier test for autocorrelation, respectively. The significant non-normality statistic is due to outliers in 1972, 1973 and 1979.

**Table 4*: Tests of the Cointegration Rank - M2: Period 1966Q1-1998Q4**

| $H_0$=r | Eigenv. =$\hat{\lambda}$ | $\lambda_{\mu\alpha\xi}^{(1)}$ | $\lambda_{\mu\alpha\xi}$90 | $\lambda_{\mu\alpha\xi}95^{(2)}$ | Trace$^{(3)}$ | Trace90 | Trace95$^{(4)}$ |
|---|---|---|---|---|---|---|---|
| 0 | 0.2190 | 28.42 | 13.39 | 25.54 | 41.60 | 26.70 | 29.38 |
| 1 | 0.0727 | 8.68 | 10.60 | 18.96 | 13.19 | 13.31 | 15.34 |
| 2 | 0.0384 | 4.51 | 2.71 | 12.25 | 4.50 | 2.71 | 3.84 |
| Diagnostic tests**: | | | | | | | |
| LM(1) | p-value = 0.08 | | | | | | |
| LM(4) | p-value = 0.30 | | | | | | |
| Normality | p-value = 0.13 | | | | | | |

Notes: (1) $\lambda_{max}$ is adjusted to correct the small sample bias error. Namely, N is replaced by (N–kp). $\lambda_{max}$ = -(N-kp) ln(1- $\hat{\lambda}_r$), where N is the number of observations, k is the number of lag length and p is the number of endogenous variables, see Cheung and Lai (1993).

(2) The source is Osterwald-Lenum (1992), Table 2, p. 469.

(3) Trace test is adjusted to correct the small sample bias error. Namely, N is replaced by (N–kp). Trace test = -(N-kp) $\sum_{i=r+1}^{P} \ln(1-\hat{\lambda}_i)$. Both Trace and $\lambda_{max}$ tests were developed in Johansen and Juselius (1991).

(4) The source is Johansen (1995), Table 15.3, p. 215.

* The regression does not include time trend. Lag length is 4.

** LM(1) and LM(4) are one and four-order Lagrangian Multiplier test for autocorrelation, respectively.

**Table 5*: Long-Run Demand for Money 1966Q1-1998Q4**

| Estimator | Maximum Likelihood | | DOLS** | | DOLSM*** | |
|---|---|---|---|---|---|---|
| Dependent | Coeff | $\chi^2$ (P-value) | Coeff | LSE | Coeff | LSE |
| Variable=lrm1 | 4.37 | 2.24 (0.13) | 1.03 | 0.49 | 1.05 | 0.46 |
| lrgdp | -0.27 | 7.80 (0.01) | 0.01 | 0.02 | 0.002 | 0.01 |
| infgdp | | | | | | |
| | | | Wald stat. (P-value) | | Wald stat. (P-value) | |
| | | | 8.0 (0.02) | | 5.6 (0.05) | |
| Dependent | | | | | | |
| Variable=lrm2 | Coeff | $\chi^2$ (P-value) | Coeff | LSE | Coeff | LSE |
| lrgdp | 1.37 | 7.70 (0.01) | 1.17 | 0.37 | 1.19 | 0.34 |
| infgdp | -0.05 | 15.05 (0.00) | 0.01 | 0.02 | 0.002 | 0.01 |
| | | | Wald stat. (P-value) | | Wald stat. (P-value) | |
| | | | 17.4 (0.00) | | 12.6 (0.00) | |

Notes: * lrm1 is the log of the real interest-free money (M1), lrm2 is the log of the real interest-bearing money (M2), lrgdp is the log of real GDP and infgdp is the change (percentage and annualized) of log GDP deflator. Coeff is the abbreviation for coefficient and LSE is the long-run standard error of the coefficient.

** DOLS is Stock and Watson's (1993) dynamic OLS estimator. The statistics are based on the regression, $lrm_t = \beta_0 + \beta_1\, lrgdp_t + \beta_2\, infgdp_t + \delta_1(L)\, \Delta infgdp_t + \delta_2(L)\, \Delta lrgdp_t + u_t$, where $\delta_1(L)$ and $\delta_2(L)$ have two leads and lags as suggested by Stock and Watson for the number of observations of 100 or close to 100. The Wald statistic tests the hypothesis that $\beta_1 = \beta_2 = 0$ and has a $\chi^2$ distribution. The covariance matrix was computed using an AR(2) spectral estimator.

*** Since infgdp is level stationary the short-run part of the DOLS was modified and the statistics are based on the regression, $lrm_t = \beta_0 + \beta_1\, lrgdp_t + \beta_2\, infgdp_t + \delta_1(L)\, infgdp_t + \delta_2(L)\, \Delta lrgdp_t + u_t$, where as before $\delta_1(L)$ and $\delta_2(L)$ have two leads and lags. The Wald statistic tests the hypothesis that $\beta_1 = \beta_2 = 0$ and has a $\chi^2$ distribution. The covariance matrix was computed using an AR(2) spectral estimator.

**Table 6*: Test For Weak Exogeneity of the Variables of the, Long-Run Parameters**

| Variables | p-value |
|---|---|
| Null: The variable is weakly exogenous for the long-run coefficients: | |
| Equation for M1 | |
| lrgdp | 0.45** |
| infgdp | 0.05** |
| Both variables lrgdp and infgdp are jointly weakly exogenous for the long-run coefficients: | |
| Equation for M2 | 0.09** |
| lrgdp | 0.87** |
| infgdp | 0.43** |
| Both variables lrgdp and infgdp are jointly weakly exogenous for the long-run coefficients | 0.67** |

Notes: * lrgdp is the log of real GDP and infgdp is the change of log GDP deflator (percentage and annualized).

** Cannot reject the null.

**Table 7*: Error Correction Model: Instrumental-Variable, Dependent Variable = $\Delta lm1_t$**

| Variable | Coefficient | Standard Error | Hansen's (1992) stability $L_i$ test (5 percent critical value = 0.47) |
|---|---|---|---|
| $\Delta lm1_{t-4}$ | 0.44 | 0.09 | 0.68 |
| $\Delta lrgdp_{t-3}$ | 0.33 | 0.11 | 0.15 |
| $\Delta lrgdp_{t-6}$ | -0.25 | 0.10 | 0.03 |
| $\Delta infgdp_t$ | -0.29 | 0.13 | 0.07 |
| $\Delta infgdp_{t-1}$ | -0.45 | 0.16 | 0.07 |
| $\Delta infgdp_{t-2}$ | -0.47 | 0.15 | 0.25 |
| $\Delta infgdp_{t-3}$ | -0.28 | 0.14 | 0.39 |
| $EC_{t-1}$ | -0.002 | 0.0004 | 0.26 |
| Q2 | -0.05 | 0.02 | 0.61 |
| Q4 | -0.05 | 0.02 | 0.11 |
| Hansen's (1992) stability $L_i$ test on variance of the ECM | | | 0.50 |
| Joint (coefficients and the error variance) Hansen's (1992) stability $L_c$ test (5 percent critical value (df=11)=2.75 | | | 2.44 |
| Chow test (break point=84Q1) | | | F-test (9, 99)=1.43 (p value=0.19) |

Notes: * Period=1966Q1-1998Q4. Mean of dependent variable=0.014. $\Delta$ means the first difference, $\Delta lm1$ is the change of log M1 and $\Delta lrgdp$ is the change of the log of real GDP. To be consistent with the rest of the variables in ECM, $\Delta infgdp$ is the first difference of the quarterly change of log GDP deflator and EC is the error correction term, but it is calculated from the cointegration equation which contains quarterly inflation rate rather than percentage-annualized rate. Q2 and Q4 are dummy variables for the second and fourth quarters of the year, respectively.

The estimation method is Instrumental-Variable estimation technique. The instruments are fourth and eighth lags of the error term. $\overline{R}^2$=0.47, $\sigma$=0.05, DW=1.92, Godfrey(5)=0.84 (significance level=0.54), White=79 (significance level=0.11), ARCH(5)=6.69 (significance level=0.24), RESET=0.20 (significance level=0.90) and Normality($\chi^2$=2)=24.00 (significance level=0.00). The significant non-normality statistic is due to outliers in 1974, 1979 and 1988.

Note that $\overline{R}^2$, $\sigma$ and DW, respectively, denote the adjusted squared multiple correlation coefficient, the residual standard deviation and the Durbin-Watson statistic. White is White's (1980) general test for heteroskedasticity, ARCH is five-order Engle's (1982) test, Godfrey is five-order Godfrey's (1978) test, REST is Ramsey's (1969) misspecification test, Normality is Jarque and Bera's (1987) normality statistic, $L_i$ is Hansen's (1992) stability test for the null hypothesis that the estimated coefficient or variance of the error term is constant and $L_c$ is Hansen's (1992) stability test for the null hypothesis that the estimated coefficients as well as the error variance are jointly constant.

**Table 8*: Error Correction Model: Instrumental-Variable Dependent Variable = $\Delta lm2_t$**

| Variable | Coefficient | Standard Error | Hansen's (1992) stability $L_i$ test (5 % critical value = 0.47) |
|---|---|---|---|
| Constant | -0.08 | 0.02 | 0.10 |
| $\Delta lm2_{t-3}$ | -0.18 | 0.09 | 0.09 |
| $\Delta lm2_{t-7}$ | -0.25 | 0.08 | 0.08 |
| $\Delta lrgdp_{t-3}$ | 0.40 | 0.08 | 0.09 |
| $\Delta infgdp_t$ | -0.29 | 0.08 | 0.29 |
| $\Delta infgdp_{t-1}$ | -0.38 | 0.09 | 0.07 |
| $\Delta infgdp_{t-2}$ | -0.44 | 0.11 | 0.17 |
| $\Delta infgdp_{t-3}$ | -0.35 | 0.10 | 0.98 |
| $\Delta infgdp_{t-4}$ | -0.40 | 0.11 | 0.18 |
| $EC_{t-1}$ | -0.03 | 0.01 | 0.11 |
| $(EC)^3_{t-3}$ | -0.001 | 0.0003 | 0.04 |
| Q2 | -0.07 | 0.01 | 0.18 |
| Hansen's (1992) stability $L_i$ test on variance of the ECM | | | 0.18 |
| Joint (coefficients & error variance) Hansen's (1992) stability $L_c$ test (5 % critical value (df=13)=3.15) | | | 2.44 |
| Chow test (break point=84Q1) | | | F-test (10,94)=2.01 (p value=0.04) |

Notes:* Period=1966Q1-1998Q4. Mean of dependent variable=0.05. Δ means the first difference, Δlm2 is the change of log M2, Δlrgdp is the change of the log of real GDP, Δinfgdp is the first difference of the quarterly change of log GDP deflator and EC is the error correction term from the cointegration equation which contains quarterly inflation rate (see the footnote of Table 7 for more explanation). Q2 and Q4 are dummy variables for the second and fourth quarters of the year, respectively.

The estimation method is Instrumental-Variable estimation technique. The instruments are fourth and eighth lags of the error term. $\overline{R}^2$=0.59, σ=0.03, DW=1.75, Godfrey(5)=0.85 (significance level=0.53), White=76 (significance level=0.85), ARCH(5)=2.75 (significance level=0.74), RESET=0.48 (significance level=0.70) and Normality($\chi^2$=2)=0.78 (significance level=0.68).

Note that $\overline{R}^2$, σ and DW, respectively, denote the adjusted squared multiple correlation coefficient, the residual standard deviation and the Durbin-Watson statistic. White is White's (1980) general test for heteroskedasticity, ARCH is five-order Engle's (1982) test, Godfrey is five-order Godfrey's (1978) test, REST is Ramsey's (1969) misspecification test, Normality is Jarque and Bera's (1987) normality statistic, $L_i$ is Hansen's (1992) stability test for the null hypothesis that the estimated coefficient or variance of the error term is constant and $L_c$ is Hansen's (1992) stability test for the null hypothesis that the estimated coefficients as well as the error variance are jointly constant.

**Table 9*: Marginal Model:Dependent Variable = $\Delta infgdp_t$**

| Variable** | Coefficient | Standard Error |
|---|---|---|
| Constant | 0.02 | 0.003 |
| $\Delta infgdp_{t-1}$ | -0.43 | 0.07 |
| $\Delta infgdp_{t-2}$ | -0.48 | 0.08 |
| $\Delta infgdp_{t-3}$ | -0.27 | 0.08 |
| $\Delta infgdp_{t-6}$ | -0.15 | 0.07 |
| $(\Delta lrgdp)(Zero)_{t-8}$ | 0.29 | 0.12 |
| $(\Delta lrgdp)(Inflation)_{t-1}$ | 1.32 | 0.55 |
| $(\Delta lrgdp)(Inflation)_{t-2}$ | 0.76 | 0.38 |
| $(\Delta lrgdp)(Inflation)_{t-3}$ | 2.18 | 0.87 |
| $(\Delta lrgdp)(Inflation)_{t-5}$ | 1.25 | 0.60 |
| $(\Delta lrm1)(Inflation)_{t-1}$ | -0.50 | 0.24 |
| $(\Delta lrm1)(Inflation)_{t-3}$ | -0.64 | 0.33 |
| $(\Delta lrm1)(Inflation)_{t-7}$ | 0.64 | 0.32 |
| $(\Delta lrm1)(War)_{t-4}$ | 0.35 | 0.15 |
| $Q3_t$ | 0.06 | 0.01 |

Notes: * Period=1966Q1-1998Q4, Δ means the first difference, Mean of dependent variable=-0.0001. Δinfgdp is the first difference of the change of log GDP deflator and Δlrgdp is the change of log of real GDP. Zero is dummy variable, which is equal to one for the first quarter of 1984 and is zero, otherwise. Inflation is a dummy variable, which is equal to one for period 1995Q2-1998Q1 and is zero otherwise. War is a dummy variable, which is equal to one for period 1980Q4-1988Q3 and is zero otherwise. Q3 is a dummy variable for the third quarter of the year.

The estimation method is Instrumental-Variable estimation technique. The instruments are fourth and eighth lags of the error term. $\overline{R}^2$=0.71, σ=0.3, DW=2.09, Godfrey(5)=0.79 (significance level=0.58), White=52 (significance level=1.00), ARCH(5)=1.65 (significance level=0.90), RESET=0.95 (significance level=0.42), Normality ($\chi^2$=2)=8.82 (significance level=0.01).

Note that $\overline{R}^2$, σ and DW, respectively, denote the adjusted squared multiple correlation coefficient, the residual standard deviation and the Durbin-Watson statistic. White is White's (1980) general test for heteroskedasticity, ARCH is five-order Engle's (1982) test, Godfrey is five-order Godfrey's (1978) test, REST is Ramsey's (1969) misspecification test, Normality is Jarque and Bera's (1987) normality statistic, $L_i$ is Hansen's (1992) stability test for the null hypothesis that the estimated coefficient or variance of the error term is constant and $L_c$ is Hansen's (1992) stability test for the null hypothesis that the estimated coefficients as well as the error variance are jointly constant.

**Table 10: Superexogeneity Tests**

| | F(1,101) (P-Value) | F(1, 97) (P-Value) |
|---|---|---|
| Variable ($Z = \Delta$infgdp)* | M1** | M2*** |
| $Z - \eta^Z$ | 0.99 | 0.49 |
| | (0.32) | (0.48) |
| $\sigma^{ZZ} (Z - \eta^Z)$ | 0.49 | 1.57 |
| | (0.49) | (0.21) |
| $(\eta^Z)^2$ | 0.61 | 1.03 |
| | (0.44) | (0.31) |
| $(\eta^Z)^3$ | 4.07 | 9.83 |
| | (0.05) | (0.002) |
| $\sigma^{ZZ} \eta^Z$ | 0.002 | 0.39 |
| | (0.96) | (0.53) |
| $\sigma^{ZZ} (\eta^Z)^2$ | 1.69 | 0.65 |
| | (0.20) | (0.42) |
| $(\sigma^{ZZ})^2\eta^Z$ | 0.14 | 0.02 |
| | (0.70) | (0.88) |
| F-Statistics (7, 101 for M1), (7, 97 for M2) on coefficients of all variables | 1.09 | 2.29 |
| | (0.38) | (0.03) |

Notes: * $\Delta$infgdp is the first difference of the change of log GDP deflator. $\eta^Z$ is the conditional mean of Z and $\sigma^{ZZ}$ is the conditional variance of Z.

$$\Delta lmi_t = \alpha_0 + \psi_0 Z_t + (\delta_0 - \psi_0)(Z_t - \eta^Z_t) + \delta_1 \sigma_t^{ZZ} (Z_t - \eta^Z_t) + \psi_1 (\eta^Z)^2 + \psi_2 (\eta^Z)^3 + \psi_3 \sigma_t^{ZZ} \eta^Z + \psi_4 \sigma_t^{ZZ} (\eta^Z)^2 + \psi_5 (\sigma_t^{ZZ})^2\eta^Z + z'_t\gamma + u_t, \ (i=1, 2)$$

** The estimation method is OLS estimation technique: $\overline{R}^2$=0.47, $\sigma$=0.05, DW=1.91, Godfrey(5)=0.77 (significance level=0.58), White=115 (significance level=0.99), ARCH(5)=8.79 (significance level=0.12), RESET=0.29 (significance level=0.83) and Normality($\chi^2$=2)=33 (significance level=0.00). The significant non-normality statistic is due to outliers in 1974, 1979 and 1988.

*** The estimation method is OLS estimation technique: $\overline{R}^2$=0.62, $\sigma$=0.03, DW=1.75, Godfrey(5)=1.35 (significance level=0.24), White=116 (significance level=1.00), ARCH(5)=5.02 (significance level=0.41), RESET=0.56 (significance level=0.64) and Normality($\chi^2$=2)=1.72 (significance level=0.42).

Note that $\overline{R}^2$, $\sigma$ and DW, respectively, denote the adjusted squared multiple correlation coefficient, the residual standard deviation and the Durbin-Watson statistic. White is White's (1980) general test for heteroskedasticity, ARCH is five-order Engle's (1982) test, Godfrey is five-order Godfrey's (1978) test, REST is Ramsey's (1969) misspecification test and Normality is Jarque and Bera's (1987) normality statistic.

**Table 11*: Pre-interest Free vs Post-interest Free Environment Estimates (DOLS)**

| | 1966Q1-1983Q4 | | 1984Q1-1998Q4 | | 1966Q1-1998Q4 | |
|---|---|---|---|---|---|---|
| Dependent Variable = lrm1 | Coeff | LSE | Coeff | LSE | Coeff | LSE |
| Lrgdp | 0.78 | 0.24 | 0.09 | 0.53 | 1.03 | 0.49 |
| infgdp | 0.03 | 0.01 | -0.01 | 0.01 | 0.01 | 0.02 |
| $\chi^2$ for sub-sample stability ** | 3.89 (p-value=0.14) | | | | | |
| Dependent Variable= lrm2 | | | | | | |
| lrgdp | 0.97 | 0.36 | 0.27 | 0.36 | 1.17 | 0.37 |
| infgdp | 0.02 | 0.02 | -0.002 | 0.006 | 0.01 | 0.02 |
| $\chi^2$ for sub-sample stability: | | | | | | |
| Assuming constan shrt-run dynamic** | 6.19 (p-value=0.045) | | | | | |
| Allowing for short-run dynamic instability*** | 9.91 (p-value=0.003) | | | | | |

Notes: * lrgdp is the log of real GDP and infgdp is the change of log GDP deflator. Coeff is the abbreviation for coefficient and SE is the long-run standard error of the coefficient.

** The statistics are based on the regression, $lrm_t = \beta_0 + \beta_1 \, lrgdp_t + \beta_2 \, infgdp_t + \alpha_1 (lrgdp_t - lrgdp_{t-\tau})1(t>\tau) + \alpha_2 (infgdp_t - infgdp_{t-\tau})1(t>\tau) + \delta_1(L) \Delta infgdp_t + \delta_2(L) \Delta lrgdp_t + u_t$, where 1(.) is the indicator function, $\delta_1(L)$ and $\delta_2(L)$ have two leads and lags as suggested by Stock and Watson (1993) for a number of observations of 100 or close to 100, and $\tau$ = 1984Q1. The Wald statistic tests the hypothesis that $\alpha_1 = \alpha_2 = 0$ and has a $\chi^2$ distribution. The covariance matrix was computed using an AR(2) spectral estimator.

*** Based of the result of the Chow test (reported in Table 8) the statistics now are based on the regression, $lrmt = \beta0 + Zerot + \beta1 \, lrgdpt + \beta2 \, infgdp \, t + \alpha1 (lrgdpt - lrgdpt\text{-}\tau)1(t>\tau) + \alpha2 (infgdp \ t - infgdp \ t\text{-}\tau)1(t>\tau) + \delta1(L) \Delta infgdpt + \delta2(L) \Delta lrgdpt + \delta3(L)(\Delta infgdpt)(Zerot) + \delta4(L)(\Delta lrgdpt)(Zerot) + ut$, where Zerot(=1 for 1984Q1-1998Q4 period and zero, otherwise) is a dummy variable to reflect the break point in the first quarter of 1984. 1(.) is the indicator function, $\delta1(L)$ to $\delta4(L)$, as before have two leads and lags, and $\tau$ = 1984Q1. The Wald statistic tests the hypothesis that $\alpha1 = \alpha 2 = 0$ and has a $\chi2$ distribution. The covariance matrix was computed using an AR(2) spectral estimator.

**Appendix**: **Formulating superexogeneity and invariance hypothesis for the conditional model – Derivation of Equation (3)**

Let $Z_t$=$\Delta$infgdp$_t$ and vector z includes past values of $\Delta\log(Mi_t)$, $Z_t$, and current and past values of other valid conditioning variables. Define, respectively, the conditional moments of $\Delta\log(Mi_t)$ and $Z_t$ as $\eta^{Mi}_t=E(\Delta\log(Mi_t)\mid I_t)$, $\eta^Z_t=E(Z_t\mid I_t)$, $\sigma_t^{MMi}=E[(\Delta\log(Mi_t)-\eta^{Mi}_t)^2\mid I_t]$ and $\sigma_t^{ZZ}=E[(Z_t-\eta^Z_t)^2\mid I_t]$, and let $\sigma_t^{MZi}=E[(\Delta\log(Mi_t)-\eta^{Mi}_t)(Z_t-\eta^Z_t)\mid I_t]$, for i=1, 2. The information set $I_t$ consists of the past values of $\Delta\log(Mi_t)$ (for i=1, 2) and Z as well as the current and past values of z. Consider the joint distribution of $\Delta\log(Mi_t)$ and $Z_t$ conditional on information set $I_t$ to be normally distributed with mean $\eta i_t=[\eta^{Mi}_t, \eta^Z_t]$ and a non-constant error covariance matrix $\Sigma=\begin{bmatrix}\sigma^{MMi} & \sigma^{MZi}\\ \sigma^{ZMi} & \sigma^{ZZ}\end{bmatrix}$. Then following Engle *et al.* (1983), Engle and Hendry (1993) and Psaradakis and Sola (1996) we can write the relationship between $\Delta\log(Mi_t)$ and $Z_t$ as:

$$\Delta\log(Mi_t)\mid Z_t, I_t \sim N[\delta_t(Z_t-\eta^Z_t)+\eta^{Mi}_t, \Omega i_t],\ (i=1, 2), \qquad (A1)$$

where the set of coefficients $\delta_t$ includes the regression coefficients of $\Delta\log(Mi_t)$ on $Z_t$ conditional on $z'_t\gamma$, and $\Omega i_t=\sigma_t^{MMi}-(\sigma_t^{MZi})^2/\sigma_t^{ZZ}$ denotes the conditional variance.

Note that $Z_t$ (i.e., inflation rate) is a control/target variable that is subject to policy interventions. Although the parameter of $Z_t$ is assumed to be constant over the sample period it is possible that this parameter changes under interventions affecting DGP (data generating process) of this control/target variable. In this case agents have a forward-looking behavior and the demand-for-money equation is unstable. Hence, the parameters of interest in the analysis will be $\psi$ and $\gamma$ in the behavioral relationship (A2), which relates the conditional means of $\Delta\log(Mi_t)$ and $Z_t$ to the set of variables $z_t\in I_t$.

$$\eta^{Mi}_t=\psi_t(\varphi_t)\,\eta^Z_t+z'_t\gamma,\ (i=1, 2), \qquad (A2)$$

where $\varphi_t$ is the set of coefficients of marginal density of $Z_t$. However, we have allowed for the possibility that the parameters included in $\psi_t$ might vary with changes in the parameters included in $\varphi_t$. Consequently, the form of Equation (A2) may itself be time varying.

It should also be mentioned that for Equation (A2) to be forward looking (coefficients not to be invariant to policy shocks) we do not need $\gamma$ to be variable, since otherwise vector $z_t$ would merely be classified as part of an extended vector $Z_t$. Substituting (A2) in (A1) yields

$$\Delta\log(Mi_t)\mid Z_t, I_t\sim N[\psi_t(\varphi_t)Z_t+z'_t\gamma+\{\delta_t-\psi_t(\varphi_t)\}(Z_t-\eta^Z_t), \Omega i_t],\ (i=1, 2). \qquad (A3)$$

If $Z_t$ is superexogeneous, Equation (A3) will be reduced to a conventional dynamic equation and will be a stable and backward-looking relationship between $\Delta\log(Mi_t)$ and its determinants. The superexogeneity of $Z_t$ requires the following three conditions to be met (Engle and Hendry, 1993):

(a) Weak exogeneity of $Z_t$ for $\psi_t$ and $\gamma$ requires that $\eta^Z_t$ and $\sigma_t^{ZZ}$ do not enter conditional model (A3), implying that $\delta_t=\psi_t(\varphi_t)$.

(b) Constancy of the regression coefficients of conditional model (A3) entails $\delta_t=\delta$ for all t.

(c) Invariance of the coefficient of variable $Z_t$ in the conditional model to the potential changes in $\varphi_t$ in the marginal equation requires $\psi_t(\varphi_t)=\psi_t$ for all t, where the set of parameters $\psi_t$ may vary over time without depending on variations in the coefficients of marginal equation, i.e., $\varphi_t$.

If (a) is satisfied the parameters in conditional model (A3) are uniquely determined and are variation free. Namely, these parameters remain constant within a regime. Furthermore, if the variables in conditional model (A3) are strongly exogenous (i.e., in addition to weak exogeneity condition, they are not Granger caused $\Delta\log(Mi_t)$), Equation (A3) can be used for future prediction of the demand for money conditional on the future values of Z and z's. Restrictions (a), (b) and (c) together entail that $\delta=\psi=$ constant. If (a) fails then the appearance of the mean and variance of variable $Z_t$ in the conditional model also results in the failure of (b), since the changes of these moments will change the parameters of the conditional model.

It should be noted that (a) and (b) alone do not entail (c). Namely, if variable $Z_t$ is weakly exogenous for $\psi$ and $\gamma$ and is constant over the historical period, the economic agents may still be forward looking and will change their behavior as an intervention alters $\psi$ post sample. Consequently, demand for money will not be invariant to policy shocks. Note also that each of the above restrictions alone is a necessary condition to validate a constant parameter and invariant conditional model. All three conditions together constitute the superexogeneity of $Z_t$ for the coefficients of interest of money equations. It should also be mentioned that weak exogeneity is neither necessary nor sufficient for structural invariance of the conditional model, see Engle *et al.* (1983).

In arriving at an expression that can be used to test for superexogeneity, I follow Engle and Hendry (1993) and allow $\psi_t(\varphi_t)$ in (A2) to be a function of the first and second moments of $Z_t$ and approximate

$$\psi_t(\varphi_t)\,\eta^Z_t=\psi_0\,\eta^Z_t+\psi_1(\eta^Z_t)^2+\psi_2\,\sigma_t^{ZZ}+\psi_3\,\sigma_t^{ZZ}\,\eta^Z_t, \qquad (A4)$$

assuming $\eta^Z_t$ is non-zero. Substituting (A4) in (A2) and the resulting equation into (A3) yields

$\Delta\log(Mi_t) \mid Z_t, I_t \sim N[\psi_0 Z_t + z'_t\gamma + (\delta_t - \psi_0)(Z_t - \eta^Z_t) + \psi_1 (\eta^Z_t)^2 + \psi_2 \sigma_t^{ZZ} + \psi_3 \sigma_t^{ZZ} \eta^Z_t, \Omega]$, (i=1, 2). (A5)

Furthermore, to develop a formal testing procedure, following Engle and Hendry (1993), I expand $\delta_t = \sigma_t^{PZ} / \sigma_t^{ZZ} = \delta_0 + \delta_1 \sigma_t^{ZZ}$, and substitute it in (A5) to get

$$\Delta\log(Mi_t) = \alpha_0 + \psi_0 Z_t + z'_t\gamma + (\delta_0 - \psi_0)(Z_t - \eta^Z_t) + \delta_1 \sigma_t^{ZZ} (Z_t - \eta^Z_t) \qquad \text{(A6)}$$

$$+ \psi_1 (\eta^Z_t)^2 + \psi_2 \sigma_t^{ZZ} + \psi_3 \sigma_t^{ZZ} \eta^Z_t + ui_t.$$

Equation (A6) is a modified version of ECM that allows verifying whether or not agents are forward looking. The error term $ui_t$ is assumed to be white noise, normally, identically and independently distributed. Under the null of weak exogeneity, $\delta_0$-$\psi_0$=0. Under the null of invariance, $\psi_1=\psi_2=\psi_3=0$ and so $\psi_0 = \psi$. Finally, if we assume that $\sigma_t^{ZZ}$ has distinct values over different, but clearly defined regimes, then under the null of constancy of $\delta$, we need $\delta_1=0$. If all these hypotheses are accepted the equation will be reduced to results reported in Table 7 or Table 8 and the demand for interest-free or interest-bearing money, respectively, is invariant to policy intervention. Furthermore, to ensure the strength of these tests I will use the extended test shown by Psaradakis and Sola (1996, Equation 13). This entails substituting $\sigma_t^{ZZ}$ by $(\eta^Z_t)^3$ and adding extra terms $\sigma_t^{ZZ}(\eta^Z_t)^2$ and $(\sigma_t^{ZZ})^2\eta^Z_t$ to conditional model (A6) to get Equation (3). Clearly under the null of invariance, for the extended model, we need $\psi_1=\psi_2=\psi_3=\psi_4=\psi_5=0$ in order to have $\psi_0 = \psi$.

$$\Delta\log(Mi_t) = \alpha_0 + \psi_0 Z_t + (\delta_0 - \psi_0)(Z_t - \eta^Z_t) + \delta_1 \sigma_t^{ZZ} (Z_t - \eta^Z_t) \qquad (3)$$

$$+\psi_1 (\eta^Z_t)^2 + \psi_2 (\eta^Z_t)^3 + \psi_3 \sigma_t^{ZZ} \eta^Z_t + \psi_4 \sigma_t^{ZZ} (\eta^Z_t)^2 + \psi_5 (\sigma_t^{ZZ})^2 \eta^Z_t + z'_t\gamma + ui_t, \text{ (i=1, 2).}$$