Modeling Money Demand under the Profit-Sharing Banking Scheme:
Some Evidence on Policy Invariance and Long-Run Stability

Amir Kia

And

Ali F. Darrat
Modeling Money Demand under the Profit-Sharing Banking Scheme: Some Evidence on Policy Invariance and Long-Run Stability

Amir Kia*
Finance and Economic Department
Utah Valley State College
Orem, UT 84058-5999 USA

Ali F. Darrat
Department of Economics and Finance
Louisiana Tech University
Ruston, LA 71272 USA

Abstract

This paper extends the literature on profit-sharing banking systems by modeling money demand behavior in Iran. We estimate demand for M1 and profit-sharing deposits over the period 1966-2001. We focus on whether the estimated money demand equations are policy invariant in addition to being temporally stable in the short and long run. Our empirical results persistently suggest that the demand equation for profit-sharing deposits is particularly stable and policy invariant in Iran despite numerous policy and non-policy shocks. These results lend support to the profit-sharing banking system and suggest that profit-sharing monetary aggregates are a credible instrument for monetary policy-making.

Keywords: profit-sharing deposits, interest-free banking system, policy-invariance, super-exogeneity, long-run stability

JEL Codes: E41, E52

* Corresponding author. Telephone (801) 863-6898, Fax (801) 864-8060, Email: kiaam@uvsc.edu. An earlier draft of this paper was presented to the 21st Symposium on Banking and Monetary Economics (Nice, France) in June 2004, and to the 2004 Canadian Economics Association Annual Meetings (Toronto, Canada). The authors wish to thank participants in both conferences for helpful comments.
Modeling Money Demand under the Profit-Sharing Banking Scheme: Some Evidence on Policy Invariance and Long-Run Stability

1. Introduction

The concept of interest-free (profit-sharing) in the banking industry, as opposed to the alternative and more common concept of predetermined (fixed) interest rates, has gained some popularity since the early 1980s. Recent data reveal that there are at least 300 banks and non-bank financial institutions operating under some form of the profit-sharing principle in different parts of the world. These financial institutions have been growing at an annual rate of about 10% with total assets exceeding $200 billion (Hassoune, 2002). Recently, major financial institutions like the Citibank have begun offering similar (interest-free) financial services to an increasing customer base\(^1\).

Parallel to the growth and the popularity of profit-sharing banks, there has been an equally impressive volume of research on the nature and structure of these banks and on their efficiency relative to the more traditional interest-based banks (see, for example, Bashir, 1983, Khan, 1986, Khan and Mirakhor, 1990, and Chapra, 1992). With only a few exceptions (e.g., Darrat, 1988, 2002), most prior research on the subject is essentially theoretical, void of any empirical evidence.

One way to empirically examine the merit of profit-sharing banking scheme is to investigate the nature of the aggregate money demand function in a country that has had a long actual experience with this banking system. A well-behaving and stable money

---

\(^1\) Friedman (1969) argues that optimal resource allocation requires zero nominal interest rates, and Cole and Kocherlakota (1998) suggest that a zero interest rate is both a necessary and sufficient condition for optimality. Commenting on possible causes behind the East Asian financial crisis of the 1990s, Wilson (1998) contends that one main factor was that funds flowing into the region were not participatory.
demand function is required by almost all theories of macroeconomic activities and particularly for the smooth operation of an effective monetary policy. An unstable function undermines the monetary policy which becomes itself a source of economic disturbance. As Hoffman et al. (1995) argue, the importance of a well-behaving money demand function is a basic tenant not only for the monetarist theory (Friedman, 1956), but also in Neo-Classical models (Sargent and Wallace, 1975), in some Neo-Keynesian models (Mankiw, 1991), and also in models of real business cycles (King et al., 1991).

It might be argued that the stability of money demand is an important prerequisite for achieving price and output stability only in countries that seriously use monetary aggregates (not interest rates) as intermediate targets. However, it is primarily the instability of money demand that convinced many countries to move to an interest rate targeting. Consequently, if the issue of money demand instability can be resolved, then the case for an interest rate instrument would lose much of its appeal (Kia, 2005). Note also that the stability of money demand is often assumed in applied macroeconomics even under policy regime changes and mounting evidence for unstable money demand relationships in many countries.

For instance, some recent studies report unstable interest-rate elasticity of money demand for several countries, including Finland (Ripatti, 1998), Germany (Bahmani-Oskooee and Bohl, 2000), and Japan (Hamori and Tokihisa, 2001). Interest rates, perhaps more so than any other determinants of money demand, are subject to speculative behavior and could have been the culprit behind the observed instability of money demand. In addition, since money balances may be held to smooth out differences between streams of income and expenditure, both actual and expected interest rates influence agents’ portfolio behavior. It is possible, therefore, that money demand
relationships could become less unstable in the absence of interest rates. As discussed below, the use of a flexible expected rate of profit (instead of the fixed rate of interest) in modeling money demand may provide the estimated relationships with an internal source of stability. Such a flexible profit rate acts as a shock absorber which, unlike the fixed rate of interest, can mitigate structural breaks in the underlying demand relationships resulting from exogenous shocks.

In the absence of predetermined interest rates, for a given risk, the profit rate alone is the determining factor in the performance of a portfolio. Therefore, one source of rebalancing, i.e., the interest rate, is eliminated. In this case, there is no need for agents to transfer their invested funds from one kind of deposit (or fixed-income asset) to another to rebalance their portfolios as predetermined interest rates change. Specifically, the speculative part of the demand for money will be eliminated. Another source of instability of the demand for money (mostly broad aggregate) is the balance sheet repricing. For example, when interest rates rise, a bank’s profit, under the conventional banking system, will rise if the bank is able to reprice its assets at higher rates before it reprices its liabilities. Alternatively, when rates fall, the bank’s profit will rise if the bank is able to reprice its liabilities at lower rates before it reprices its assets.

Consequently, balance sheet repricing, as an important element in profit maximization, is also a source of disturbance in the money market as fixed rates fluctuate. This supply side instability causes the demand for money to be unstable, since major parts of the broad monetary aggregate, i.e., time deposits, cannot be supplied if there is no demand for them. Since there is no interest exposure risk under the profit-risk-sharing banking system, there is no need for the balance sheet repricing. Therefore, one would expect, under the conventional banking system, where the interest exposure risk
exists, the money demand to be less stable than under the profit-risk-sharing banking system.

However, under the profit-risk-sharing banking system, the central bank will lose one of its monetary policy tools, i.e., interest rate, but can rely on a more powerful tool, i.e., to control money supply. The central bank can keep the money supply at its optimal level. At this level, the stable money demand allows the central bank to always operate at the optimum money supply where the consumer surplus is maximized. Assuming demand for money is stable, Friedman (1969) shows the optimum level of money supply can be achieved when the interest rate is zero. Actually, the monetary policy in Iran, where the banking system is under the profit-risk-sharing scheme, is implemented by controlling the money supply. It should be mentioned that the stability of demand for money, while it is a necessary condition for the financial stability, is not a sufficient condition. However, in this study we investigate only the stability of the demand for money.

This paper focuses on the Iranian experience with the profit-risk-sharing banking scheme. Compared to other countries that have experimented with interest-free banking systems, Iran provides an interesting case since the prohibition of interest-based financial transactions is most closely and consistently enforced in Iran, and for a relatively long time (since the mid-1980s). In addition, Iran has also witnessed several changes in policy regimes and undergone many other exogenous shocks during the past two decades which makes this country an ideal case to test whether its underlying money demand equations have endured all such shocks and regime changes. It should be noted at the outset that, like most other developing countries, Iran has a heavily regulated financial system in which financial innovations emerge rather slowly.
Our empirical analysis on Iranian money demand departs from previous empirical research in this area in at least two main respects. First, we formally estimate both short-run and long-run money demand equations. Second, prior research in this area focuses on whether the estimated money demand equations are temporally stable, but overlooks the additional important requirement that the estimated equations should also be policy invariant. As Lucas (1976) points out, temporal stability and policy invariance are distinctly different concepts. Estimated parameters of a given money demand equation may remain constant over time, but the parameters could still vary in response to a policy regime change or other exogenous shocks in the economy. If asset holders are forward looking, then any regime change would alter the agents’ behavior which will then undermine policy effectiveness. Therefore, estimated money demand models should be tested for policy-invariance prior to their use for policy analysis. In contrast to the forward-looking behavior underlying policy-invariance, the more common concept of parameter stability is predicated on backward-looking behavior. While a few studies (e.g., Favero and Hendry, 1992 and Engle and Hendry, 1993) examine this issue for developed countries, research on policy invariance of money demand in developing countries is scant, and in the case of profit-sharing money demand, this research is virtually non-existent.

The rest of the paper is organized as follows: Section 2 formulates the short- and long-run money demand models and reports the empirical results; Section 3 focuses on results from the policy-invariance and stability tests; Section 4 provides concluding remarks and outlines key policy implications.
2. Modeling Profit-Sharing Money Demand in an Open Economy

2.1. Model Specification

On March 21, 1984, the Iranian government started implementing tight restrictions on the payment of fixed interest rate on most financial transactions in the country. In the case of private banks and non-bank credit institutions, the Central Bank of Iran (CBI) banned all fixed rates of interest on the asset and liability sides of these institutions, requiring them to bear market-based profit rates. However, for government-owned banks, the CBI imposed a minimum “profit” rate for bank depositors to ensure the attractiveness of such deposits. Various reports of the CBI suggest that the minimum rates from 1984 until 2001 were as follows: short-term 8%; special short-term 10%; one-year 14%; two-year 15%; three-year 16% and five-year 18.5%. However, since May 2001, these minimum rates have been reduced to the following: short-term 7%, one-year 13% and five-year 17%. With an annual inflation rate running at about 35%, the apparent reason for these minimum profit rates is to compensate deposit holders for the erosion in the value of financial obligations resulting from such high inflation rates.

Consider an economy with a single consumer, representing a large number of identical consumers. The consumer maximizes the following utility function:

\[
E \left\{ \sum_{t=0}^{\infty} \beta^t U(c_t, c^*_t, S_t) \right\},
\]

where \(c_t\) and \(c^*_t\) are single, non-storable, real domestic and foreign consumption goods, respectively. \(S_t\) is the flow of services per unit of time derived from the holdings of domestic and foreign real cash balances, \(E\) is the expectation operator, and \(0<\beta<1\). The utility function is assumed to be increasing in all its arguments, strictly concave and continuously differentiable. The demand for monetary services will always be positive if
we assume \( \lim_{s \to 0} U_s(c, c^*, S) = \infty \), for all \( c \) and \( c^* \), where \( U_s = \partial U(c, c^*, S) / \partial s \). Assume the flow of services derived from the holding of real cash balances is a function of both domestic and foreign stocks of real cash balances. Assume also that the $US represents the foreign currency and that, following Stockman (1980), Lucas (1982), Guidotti (1993) and Hueng (1999), purchases of domestic and foreign goods are made with domestic and foreign currencies, respectively.\(^2\) Specifically,

\[
S_t = S(m_t, m_t^*),
\]

(2)

where \( m \) is domestic real money \((M/p)\), and \( m^* \) is foreign real money \((M^*/p^*)\). Furthermore, assume \( S_m = \partial S(m, m^*) / \partial m > 0 \), and \( S_{m^*} = \partial S(m, m^*) / \partial m^* > 0 \), implying that as the holding of domestic and foreign currencies respectively increases, the services of these currencies go up.

The consumer maximizes his utility function (1) subject to the following budget constraint:

\[
\tau_t + y_t + (1 + \pi_t)^{-1} m_{t-1} + q_t (1 + \pi_t^*)^{-1} m^*_{t-1} + (1 + \pi_t)^{-1} (1 + r_t) d_{t-1} + \\
q_t (1 + \pi_t^*)^{-1} (1 + r^*_{t-1}) d^*_{t-1} = c_t + q_t c_t^* + m_t + q_t m_t^* + d_t + q_t d_t^*,
\]

(3)

where \( \tau_t \) is the real value of any lump-sum taxes/transfers received/paid by consumers, \( y_t \) is the current real endowment (income) received by the individual, \( \pi_t \) and \( \pi_t^* \) are, respectively, domestic and foreign rates of inflation, \( q_t \) is the real exchange rate, defined as \( e_t p_t^* / p_t \), \( e_t \) is the nominal market (non-official) exchange rate (domestic price of foreign currency), \( p_t^* \) and \( p_t \) are the foreign and domestic price levels of foreign and domestic goods, respectively, \( m^*_{t-1} \) is the foreign real money holdings at the start of the

\(^2\) By this assumption neither these authors nor we mean that individuals need foreign currencies in order to buy imported goods at home. This assumption simply means that the importers have to pay foreign currencies to purchase foreign-produced goods in order to import.
period, $d_t$ is the one-period real domestic term deposit which is expected, conditional on current information $I_t$, to pay the rate of profit of $\mathbb{E}(r_{t+1} \mid I_t) = r_t^e$, and $d_t^*$ is the real foreign one-period time (non-checking) deposits which pay a predetermined risk-free interest rate $r_t^*$. Assume further that $d_t$ and $d_t^*$ are the only two storable assets. Note that in Iran any investor, domestic or foreign, can have US$ deposits in any bank and can transfer the funds outside the country without any restriction. The rate on these deposits is LIBOR plus one percentage point (CBI, 2001/2002).

The above model is standard with the exception that the rate of return on the one-period asset is not predetermined as commonly assumed. Define $U_c = \partial U(c, c^*, m, m^*)/\partial c$, $U_{c^*} = \partial U(c, c^*, m, m^*)/\partial c^*$, $U_s = \partial U(c, c^*, m, m^*)/\partial S$, and $\lambda_t = $ the marginal utility of wealth at time $t$. Substitute $S_t$ from (2) into (1), and assume the resulting indirect utility has an instantaneous function of the form:

$$U(c_t, c_t^*, m_t, m_t^*) = (1 - \sigma)^{-1} [c_t^{\alpha_1} c_t^{\alpha_2} m_t^{\eta_1} m_t^{\eta_2}]^{1-\sigma},$$  \hspace{1cm} (4)

where $\sigma$, $\alpha_1$, $\alpha_2$, $\eta_1$ and $\eta_2$ are positive parameters. Relegating full details of the model derivations to an appendix available upon request, the demand for domestic real balances, using first-order conditions from maximizing (4) subject to (3), can be written as:

$$m_t = (\eta_1 c_t) / \alpha_1 r_{t+1}^e (1 + r_{t+1}^e)^{-1}. \hspace{1cm} (5)$$

From (5), we have $m_{ct} = \partial m_t / \partial c_t > 0$ and $m_{ret+1} = \partial m_t / \partial r_{t+1}^e < 0$. Thus, (4) becomes:

$$\log (m_t) = \log(\eta_1) + \log(c_t) - \log(\alpha_1) - \log[r_{t+1}^e (1 + r_{t+1}^e)^{-1}]. \hspace{1cm} (6)$$

Let the domestic real consumption ($c_t$) be some proportion ($\omega$) of the domestic real income ($y_t$). Assume further the current relevant information for estimating $r_{t+1}^e$ are the current inflation rate ($\pi_t$), the foreign interest rate ($r_t^*$), and the real exchange rate ($q_t$). We may hypothesize:
\[
\log \left[ r_{t+1}^c (1 + r_{t+1}^c)^{-1} \right] = \theta_2 \pi_t + \theta_3 r^* + \theta_4 \log (q_t) + u_t, \tag{7}
\]

where \(\theta\)'s are constant coefficients and \(\theta_2 > 0, \theta_3 \geq 0, \theta_4 > 0\), and \(u_t\) is a white noise disturbance term with zero mean. In the case of Iran, \(\theta_2 > 0\) since the CBI guarantees a minimum profit rate for non-checking accounts as an inducement for bank customers in a highly inflationary environment.

In a profit-sharing system, the majority of economic agents do not formulate their expectations on the basis of a predetermined rate of interest, \(r^*\). Consequently, we assume \(\theta_3 \geq 0\). However, \(r^*\) may still be a driving force in forming expectations of the future rate of profit through arbitrage activities of those agents that are not strictly adhering to the ban on fixed interest rates. Accordingly, the sign of \(\theta_3\) may be indeterminate. For \(\theta_4\), a higher real exchange rate should reduce the demand for imports but increase the demand for exports, leading to a higher profit at least over the long-run, i.e., \(\theta_4 > 0\). However, the short-run demand for imports is inelastic, possibly making \(\theta_4\) negative over the short run. Substituting \(c_t = \omega y_t\), and (7) into (6) yields the following final m1 demand equation:

\[
\log m_{1,t} = \beta_0 + \beta_1 \log y_t + \beta_2 \pi_t + \beta_3 r^* + \beta_4 \log q_t + u_t, \tag{8}
\]

where \(\beta_0 = \log(\eta_1) - \log(\alpha_1), \beta_1 = \log(\omega) > 0, \beta_2 = -\theta_2 < 0, \beta_3 = \theta_3 \leq 0, \) and \(\beta_4 = \pm \theta_4, \) or \(\beta_4 < 0\) over the long term; and \(\beta_4 > 0\) over the short term. Note that \(\log m\) is the log of real narrow money stock (defined as currency plus interest-free demand deposits); \(\log y\) is the log of real GDP; \(\pi\) is the CPI inflation rate, \(r^*\) is the London inter-bank offered rate (LIBOR), \(\log q\) is the log of real exchange rate using the CPI of the U.S. as the foreign price and the CPI in Iran as the domestic price; \(u\) is a disturbance term assumed to be white noise with zero mean; and the \(\beta\)s are the parameters to be estimated.
Some rationale for using the above variables is in order. The LIBOR is used to approximate the foreign rate since maximum interest rates paid on foreign currency deposits in Iran are closely linked to LIBOR; namely, LIBOR plus one percentage point (CBI, 2001/2002). Furthermore, the CPI in the U.S. is used to represent the foreign price for Iran due to the very high association between changes in the CPIs of both countries (the correlation coefficient over the sample period is 0.9983 ≈ 1). Indeed, consumer prices in Iran closely follow adjustments in the rial/$US exchange rate.

We now turn our attention to deriving an estimable equation for the demand for real profit-sharing monetary aggregate \( (d_t) \). From the first-order conditions we can derive demand for foreign currency \( m^* \) as:

\[
m^*_t = \left( \eta_2 c_t / \alpha_1 q_t \right) \left( 1 + r^*_t \right)^{-1}.
\]

Assume \( \tau_t = 0 \) and \( d_t^* = \nu_0 r^*_t^{\nu_1} y_t^{\nu_2} \), where \( \nu \)'s are constant.\(^3\) Substituting \( c_t (= \omega y_t) \), \( \tau_t (= 0) \), \( d_t^* (= \nu_0 r^*_t^{\nu_1} y_t^{\nu_2}) \), along with equations (4) and (9), into budget line (3), we can write:

\[
d_t-1 = \frac{X_t}{R_t},
\]

where \( R_t = \frac{(1 + r_t)}{(1 + \pi_t)} \) is the real profit rate and \( X_t = \omega y_t + q_t \eta_1 \omega y_t / \alpha_1 r_t^* \left( 1 + r_t^* \right)^{-1} + q_t \nu_0 r_t^{\nu_1} y_t^{\nu_2} - y_t - (1 + \pi_t)^{-1} \left( \eta_1 \omega y_{t-1} / \alpha_1 r_t^* \right) \left( 1 + r_t^* \right)^{-1} + q_t \left( 1 + \pi_t^* \right)^{-1} \left( 1 + r_{t-1}^* \right) \nu_0 r_{t-1}^{\nu_1} y_{t-1}^{\nu_2}.
\]

Note that, from Equation (10), we can obtain \( d_t = \frac{X_{t+1}}{R_{t+1}} \), \( d_{t+1} = \frac{X_{t+2}}{R_{t+2}} \), \( d_{t+2} = \frac{X_{t+3}}{R_{t+3}} \) and so on. Substitute \( d_{t+1}(= \frac{X_{t+2}}{R_{t+2}}) \) into \( d_t(= \frac{X_{t+1}}{R_{t+1}}) \) to eliminate \( d_{t+1} \). By similar successive

\(^3\) By the relationship \( d_t^* = \nu_0 r^*_t^{\nu_1} y_t^{\nu_2} \), we simply assume that the interest elasticity of foreign deposits (\( \nu_1 \)) and the income elasticity of these deposits (\( \nu_2 \)) are constant and may be different.
eliminations, we derive an equation for \( d_t \) as a function of current and expected future values of \( y, q, R, \pi^* \) and \( r^* \), provided the transversally condition is satisfied. Note that the present value of \( d_t \) approaches zero as \( t \to \infty \). From \( d_t = \frac{X_{t+1}}{R_{t+1}} \), we can easily show that 
\[
\begin{align*}
\frac{\partial d_t}{\partial y_t} &> 0, \\
\frac{\partial d_t}{\partial r_t} &< 0, \\
\frac{\partial d_t}{\partial r^*_t} &< 0 \text{ and } \frac{\partial d_t}{\partial \pi^*_t} > 0, \end{align*}
\]
but the sign of \( \frac{\partial d_t}{\partial q_t} \) is indeterminate.

The final demand equation for profit-sharing deposits can be written as:
\[
\log q_{mt} = \gamma_0 + \gamma_1 \log y_t + \gamma_2 \pi_t + \gamma_3 \pi^*_t + \gamma_4 r^*_t + \gamma_5 \log q_t + u_t, \tag{11}
\]
where \( \gamma \)'s are the parameters, and \( q_m \) denotes \( d \). As it was shown above, \( \gamma_1 > 0, \gamma_2 < 0, \gamma_3 > 0, \gamma_4 < 0, \text{ and } \gamma_5 = \text{indeterminate} \). Note that \( \pi^* \) is the U.S. inflation rate as a proxy for foreign inflation, and \( r^* \) is the London interbank rate to represent foreign interest rates.

Furthermore, the only difference between equations (8) and (11) is the absence of \( \pi^* \) in Equation (8), which traces back by Equation (7). As the latter equation indicates, the expected rate of profit is a function of the domestic rate of inflation, the foreign rate of interest (because investors can invest in $US denominated deposits in Iran) and the real exchange rate. There is no theoretical reason or empirical evidence, which indicates the foreign inflation rate can influence directly the rate of profit of the banking system in Iran. If, however, there is any, it could be indirectly through the predetermined foreign rate (\( r^* \)) and the real exchange rate (\( q \)).

According to the underlying theory, under a strict ban of fixed interest rates, the profit-sharing rate and the expected inflation rate are the relevant opportunity costs of holding money. The situation is not much different even when fixed interest rates are allowed in other developing countries since authorities (rather than markets) determine such rates. Therefore, researchers are typically compelled to drop interest rates from empirical money demand models.
Note that both money demand equations (8) and (11) suggest the expected rate of profit in the banking system as a key opportunity cost for holding money. Only Bashir (2002) outlines a model with some similar features, albeit in the context of a closed-economy model. Our model is also different from Cagan-type models, including those of Tallman et al. (2003) and Nagayasu (2003), since we allow for both domestic and foreign inflation rates in determining money holdings. Unlike the short-run money demand equations estimated in Darrat (1988, 2002), but similar to Stock and Watson (1993) and Muscatelli and Spinelli (2000), our money demand equations are long-run models.

2.2. Data and Co-integration Test Results

Our data on Iran are quarterly observations spanning the period 1966Q1-2001Q4. Except for a few missing observations for some variables, all the data come from the international Financial Statistics CD-ROM, of the International Monetary Fund. Table 1 provides data description and summary statistics. According to the Augmented Dickey-Fuller and Phillips-Perron testing procedures, all variables, except for the inflation and real exchange rates, are nonstationary in levels, but they achieve stationarity when converted to first-differences (details are available upon request).

Since the model contains some variables that are integrated of degree one, our next step is to investigate if cointegration exists among the variables. We use the Johansen and Juselius (1991) test to check if at least one cointegrating vector exists between each of the monetary aggregates and their determinants in Iran (conditional on the exogenous foreign interest rate and foreign inflation rate). We use the Lagrange Multiplier (LM) testing procedure to ensure that the lag profiles used in the tests are sufficiently long to yield residuals, which are not autocorrelated. We also adjust the
resulting test statistics to correct for potential finite-sample biases (Cheung and Lai, 1993).

Table 2 reports the results from the $\lambda_{\text{max}}$ and trace tests for equations (8) and (11) for the lag length of 4 and 6, respectively. The $\lambda_{\text{max}}$ test rejects $r=0$ at the 5% level, while $r \leq 1$ is not rejected, implying that $r=1$. According to the trace test, we reject the null hypothesis of $r \leq 2$ at the 5% level, while we cannot reject the null hypothesis of $r \leq 3$, implying that $r=3$. Since these two test results are different, we further investigate the number of cointegrating ranks by estimating eigenvalues of the companion matrix. We find that all roots are either equal to or less than one. The two largest roots are $0.9729 \approx 1$, followed by a complex pair of roots with modulus $0.8623 \neq 1$, implying two unit roots. Since the number of common stochastic trends in the model should correspond to the number of unit roots equal or close to unity in the companion matrix, we may conclude that $r=2$.

With $r=2$, the system is unidentifiable. As in Johansen (1995b) and Kia (2003), we investigate the possible economic interpretation underlying multiple cointegrating vectors. Since the inflation and real exchange rates proved stationary series, the relation linking them must also be stationary. Thus, we focus on the relation between these two variables. That is, we test if the following inflation/exchange rate relation exists:

$$\pi_t = \chi_0 + \chi_1 \log q_t + u_t, \quad (12)$$

---

4 According to the results from diagnostic tests reported in the table, these lag lengths are sufficient to eliminate autocorrelation. The only non-congruency is non-normality. However, Johansen (1995a) argues that departure from normality is not particularly damaging in co-integration tests.
where $u_t$ is the disturbance term. For a given foreign price, a higher nominal exchange rate makes imports more expensive, which will raise domestic prices. A higher nominal exchange rate would also depress foreign prices of exports, leading to a higher demand for exports. The resulting pressures on domestic resources would further increase domestic prices. Hence, we expect that $\chi_1 > 0$. With this restriction, the system is over identified and the rank condition is not satisfied. To resolve this problem, we also impose a zero restriction on the constant in the $m_1$ demand Equation (7). These restrictions ensure generic, empirical and economic identifications (Johansen and Juselius, 1991).

Note that a generic identification is related to the estimability of a statistical model, while an empirical identification relates to the estimated parameter values and an economic identification is related to the economic interpretability of the estimated coefficients of an empirically identified structure.

We report below estimates from Equation (12) as well as from the restricted long-run demand for $m_1$ (figures in parentheses are standard errors):

\[
\pi_t = -9198.56 + 707.43 \log q_t, \quad (13)
\]

\[
\log m_1 = 1.61 \log y_t - 0.04 \pi_t - 0.04 r^*_t - 0.57 \log q_t. \quad (14)
\]

All estimated coefficients have the correct signs and, except for the coefficient of foreign interest rate, are highly statistically significant. Based on a chi-squared test, we cannot reject the hypothesized inflation equation and the fact that the rank condition is satisfied (the associated chi-squared statistic $= 3.78$, p-value $= 0.15$). As one would expect in an economy dominated by profit-sharing rates, it is not surprising to find the coefficient of the predetermined foreign interest rate to be statistically insignificant. Note that in Equation (13) the inflation rate is in percentage, while the real exchange rate is in...
log. Therefore, in order to interpret this estimation result, we will multiply the log of the real exchange rate by 100. In this way, the estimated coefficient of $\log q_t$ will be 7.07 implying a one percent increase in the real exchange rate results in an increase in the inflation rate by 7.07% over the long run in the country.

As for Equation (11), a lag length of 6 was required to ensure white-noise errors, see LM test results in Table 2. According to $\lambda_{\text{max}}$ test, reported in Table 2 we reject $r=0$ at the 5% level, while we cannot reject $r\leq 1$, implying that $r=1$. The trace test rejects the null hypothesis of $r\leq 2$ at the 5% level, but cannot reject the null of $r\leq 3$, implying that $r=3$. As for the case of $m_1$ system, the result of these two statistics is different. However, all roots of the estimated eigenvalues of the companion matrix are either equal to unity or inside the unit disc, where the two largest roots are $0.9817 \approx 1$ and $0.9444 \approx 1$, followed by a complex root with modulus $0.8706 \neq 1$, implying two unit roots. Thus, we may conclude that $r=2$.

With $r=2$, the system becomes unidentified. To find identified relationships, we assume that the absolute purchasing power parity (PPP) exists between Iran and the United States. We also impose two other restrictions on the demand equation; namely, the constant term and the coefficient of the real exchange rate are zero. Below are the estimated PPP relationship and the identified long-run real demand for profit-sharing money ($q_m$), where standard errors are in brackets:

$$
\log q_t = 8.89 \quad (15)
$$

$$
\log q_m = 1.16 \log y_t - 0.09 \pi_t + 0.38 \pi^*_t - 0.29 r^*_t \quad (16)
$$

All estimated coefficients have the correct signs and are highly statistically significant. Based on a chi-squared test, we cannot reject the absolute purchasing power
parity relationship between Iran and the United States as well as the rank condition (the associated chi-squared statistic with 5 df is 10.57, p-value = 0.06). Note that, following Kia (1996), we assume that the goods as well as foreign exchange markets are imperfect, i.e., buying (offer) and selling (bid) prices are different and transaction costs (t) exist. Assume \( p^o \) and \( p^b \) are offer and bid prices in Iran, respectively, \( p^*o \) and \( p^*b \) are offer and bid prices in the United States, respectively, and \( e^o \) and \( e^b \) are offer and bid prices of the exchange rate, respectively. Similar to what is proved by Kia (1996) for the interbank markets, we can show arbitrage activities result in two PPP relationships for Iran: (i) \( p^o(1+t) = p^*b e^b \) for exports to the U.S. and (ii) \( p^b = p^*o e^o(1+t) \) for imports from the U.S. Since \( t \) is a constant fraction of the offer price, it is very small. Therefore, it is plausible to assume \( p^*e^n \) or \( p^*e^n t^n \) for \( n \geq 2 \) to be zero, where \( p \) and \( p^* \) are mid-points of domestic and foreign prices, respectively. Thus, we can easily show that \( p = p^*e - p^*e t \), or \( p(1+t) = p^*e \), noting that \( 1/(1+t) = 1 - t + t^2 - \ldots \) and \( t^n \) for \( n \geq 2 = 0 \). We can, therefore, have \( \log(1+t) = \log(p^*e/p) = \log(q) = \text{constant if PPP holds between Iran and United States over the long run.} \)

As we can see, the coefficient of the scale variable is close to one. We, therefore, impose the additional restriction of \( \gamma_1 = 1 \). The estimation results are:

\[
\log q_t = 8.88 \\
[0.12]
\]

\[
\log q_{mt} = \log y_t - 0.05 \pi_t + 0.26 \pi^*t - 0.20 r^*_t. \\
[0.01] [0.04] [0.02]
\]

Again, all estimated coefficients have the correct signs and are highly statistically significant. The rank condition as well as the new restriction are both satisfied (the associated chi-squared statistic with 6 df is 12.05, p-value = 0.06). We will use the error terms resulting from equations (17) and (18) to estimate the corresponding error
correction models. Note that one feature of the result appears puzzling. The estimated coefficient of foreign (LIBOR) interest rate proves significant. The latter finding is particularly puzzling since a large portion of these profit-sharing deposits is goodwill loans (Qard Hasan) that are insensitive to financial returns\(^5\).

2.3. Estimates of Short-Run Money Demand Equations

Tables 3 and 4 assemble the results from estimating ECMs for M1 and the profit-sharing deposits, respectively. In estimating ECMs, several concerns are important.

First, to avoid biased results, we allow for a lag profile of three years (12 quarters) in the estimated ECMs for the two alternative monetary aggregates. Second, having too many coefficients can also lead to inefficient estimates. To guard against this problem and ensure parsimonious estimations, we select the final ECMs on the basis of Hendry’s General-to-Specific approach. Third, observe that the error term EC is a generated regressor whose t-statistic should be interpreted with caution (Pagan, 1984 and 1986). To address this problem, we follow Pagan and apply the instrumental variable estimation technique. The instruments include first, fourth and fifth lagged error terms for both error terms for M1, and first, third and fourth lagged for error terms generated by the long-run demand for the profit-sharing deposits, and first, fourth and fifth lagged error terms for the error term generated from the PPP equation.

\(^5\) Based on data from the CBI, goodwill loan portions in profit-sharing monetary deposits increased from 11% in March 1995 to almost 17% in March 2001. Note that banks in Iran do not pay any yield on goodwill deposits since they are restricted to using such funds in the form of interest-free loans to individuals. However, private conversations suggest that some banks still offer up to 3% yield on goodwill deposits.
The specification test results reported in the tables suggest that the estimated equations are statistically adequate. According to Hansen’s stability L test, all of the coefficients are stable [the 5% critical value=0.47, see Table 1 in Hansen (1992)]. Furthermore, the joint Hansen stability Lc test result is 2.30 (<3.58 with DF=16) for M1 and 2.76 (<2.96 with DF=12) for profit-sharing deposits, which in both cases fails to reject the null of joint stability of the coefficients and the estimated associated variance.

The only contemporaneous variables in the short-term M1 demand are changes in the inflation rate and the growth rate of real exchange rate with the correct sign. For the profit-sharing aggregate, the contemporaneous variable is only changes in the inflation rate, which also bears the correct sign. We checked all possible non-linear specifications, i.e., squared, cubed and fourth powered of the equilibrium errors (with statistically significant coefficients) as well as the products of those significant equilibrium errors. We find only the EC term generated from Equation (14) for M1, and the EC term generated from Equation (18) for profit-sharing demand to be significant and the impact of both EC terms to be linear.

The results reported in tables 3 and 4 accord well with our theoretical model. As is expected in a banking system that views fixed interest rate as usury, the foreign interest rate proved statistically insignificant for M1 equation and was thus dropped. However, the coefficient of foreign rate is statistically significant and has a correct sign, according to our model, but with a very small magnitude for the profit-sharing aggregates. Note that in Iran any individual investor, Iranian citizen or otherwise, can hold Rial and/or $US denominated deposits. The return to the investment, in the real sector or the $US denominated deposits, is tax exempt. This fact clearly explains the statistically significant, but very small coefficient of the foreign rate in the profit-sharing aggregate.
According to the results in Table 4, agents appear to use current and previous changes in the inflation rate and previous changes in the foreign inflation as well as the interest rate to forecast future profits of the banking system. Furthermore, the coefficients on the EC terms in both demand equations bear the correct negative sign (error correcting) and are statistically significant.

Clearly, the existence of an ECM for money holdings does not necessarily ensure that model adjustments occur only for past equilibrium errors (backward-looking behavior). Such adjustments can also occur due to changes in the economic agents’ forecasts of future real income, the inflation rate, the profit-sharing rate, monetary policy moves and other domestic or foreign exogenous shocks (forward-looking behavior). Under such a scenario, the estimated ECM becomes susceptible to exogenous shocks from the forward-looking behavior of money holders. This lack of invariance will characterize the estimated model if one or more of the contemporaneous variables fail to be super-exogenous in the sense of Engle et al. (1983) and Engle and Hendry (1993). Under these circumstances, the estimated ECM parameters will vary with any change in the policy regime and/or other exogenous shocks.

3. Test Results for Super-Exogeneity and Long-Run Stability

Having identified statistically adequate long-run demand equations for real M1 and the profit-sharing deposits, we focus next on whether the estimated money demand equations are a reliable guide for policy analysis. Thus, we check if the estimated money demand equations are invariant to policy changes and other exogenous shocks which imply that the contemporaneous variables in the estimated equations are super-exogenous.
3.1. Estimating Marginal Models

As the previous section shows, the only contemporaneous variables remaining in the final ECM money demand models are the inflation and real exchange rates in the real M1 equation and the inflation rate in the real profit-sharing deposits. For the estimated ECMs to be policy invariant, these contemporaneous variables must be super-exogenous. Testing the super-exogeneity of these variables in turn requires the estimation of marginal models for these variables against the backdrop of several possible regime changes as well as other exogenous shocks.

A perusal of the Iranian modern history indicates that there have been five major regime changes and one exogenous shock (war) over the past three decades. They are: (i) the revolution of April 1979; (ii) the ban on fixed interest rates in the banking system that began in March 1984; (iii) the Iraq/Iran war over the period 1980-1988; (iv) the unification of official and market-determined foreign exchange rates since late March 1993; (v) the introduction of inflation targeting by the CBI during the period March 1995-March 1998; and (vi) the introduction of the privately owned financial institutions in September 1997. We use the following dummy variables to represent these potential regime shifts: Rev = 1 from 1979Q2-2001Q4, and = 0, otherwise, Zero = 1 from 1984Q1-2001Q4, and = 0, otherwise, War = 1 from 1980Q4-1988Q3, and = 0, otherwise, Ue = 1 from 1993Q1, and = 0, otherwise, Inflation = 1 from 1995Q2-1998Q1, and = 0, otherwise, and Private = 1 from 1997Q3-2001Q4, and = 0, otherwise.

Tables 5 and 6 about here

Tables 5 and 6 display the final empirical results from the marginal model for the inflation and real exchange rates, respectively. Diagnostic tests reported in the tables suggest that the estimated models are adequate and generally evince no major violations.
of key assumptions. Based on the significance of the dummy coefficients, there is strong
evidence for a structural break due to the interest-rate ban in the estimated marginal
model of inflation, and there is also a structural break due to the Iraq/Iran war in the
estimated marginal model of the exchange rate. The instability of marginal models
implies that the parameters of the associated conditional models remain stable, but only if
economic agents are not forward-looking. We provide some evidence on this issue next.

3.2. Super-Exogeneity Test Results

We examine if the contemporaneous variables in the two estimated money
demand equations are super-exogenous as required by the policy invariance hypothesis.
Let \( Z_t \) represent the contemporaneous stationary (first-difference) inflation rate or the
growth of the real exchange rate. Following Engle et al. (1983), Engle and Hendry (1993)
and Psaradakis and Sola (1996), we write the relationship between the demand for
various monetary aggregates \( X_t (=\Delta \log m_1 t \) or \( \Delta \log q_t \)) and \( Z_t \) as:

\[
X_t = \alpha_0 + \psi_0 Z_t + (\delta_0 - \psi_0) (Z_t - \eta^Z_t) + \delta_1 \sigma_t^{ZZ} (Z_t - \eta^Z_t) + \psi_1 (\eta^Z_t)^2 + \psi_2 (\eta^Z_t)^3
\]

\[
+ \psi_3 \sigma_t^{ZZ} \eta^Z_t + \psi_4 \sigma_t^{ZZ} (\eta^Z_t)^2 + \psi_5 (\sigma_t^{ZZ})^2 \eta^Z + z'_{it} \gamma + u_t, \tag{19}
\]

where \( \alpha_0, \psi_0, \psi_1, \psi_2, \psi_3, \psi_4, \psi_5, \delta_0 \) and \( \delta_1 \) are regression coefficients on \( Z_t \) conditional on
\( z'_{it} \gamma \), and \( u_t \) is a white-noise disturbance term. Vector \( z \) includes all past values of \( X_t, Z_t, \)
and other possible explanatory variables in the ECM, plus current and past values of
other relevant conditioning variables. The terms \( \eta^Z_t = \text{E}(Z_t \mid I_t) \) and \( \sigma_t^{ZZ} = \text{E}[(Z_t - \eta^Z_t)^2 \mid I_t] \)
are the conditional moments of \( Z_t \), given the information set \( I_t \) which includes past values
of \( X_t, Z_t, \) as well as current and past values of other relevant conditioning variables. \( Z_t \) can
be a control/target variable that is subject to policy interventions. With the null of weak
exogeneity, \( \delta_0 - \psi_0 = 0 \), and with the null of invariance, \( \psi_1 = \psi_2 = \psi_3 = \psi_4 = \psi_5 = 0 \) in order for
\( \psi_0 = \psi \). Under the null of constant \( \delta \), and assuming that \( \sigma_t^{ZZ} \) has distinct values over
different (but definable) regimes, $\delta_1$ must equal zero. If none of these hypotheses is rejected, the contemporaneous variables in the ECMs become super-exogenous and the estimated ECMs can be considered invariant to policy shocks.

We estimate $\eta^Z$ and $\sigma_{zz}^2$ for $Z_t$ from the marginal models reported in tables 5 and 6. Since the errors for the $Z_t$ variable appear homoskedastic according to an ARCH test, we tried a five-period moving average of the error variance, and incorporated the constructed variables in the ECMs reported in tables 3 and 4. Again, most diagnostic tests suggest the adequacy of the estimated models.

Results in Table 7 for the real M1 and profit-sharing demand equations show that all contemporaneous variables are super-exogenous. Specifically, the joint F-test on the null hypothesis that the coefficients on the constructed variables are jointly zero is not significant in either demand equations, implying that the two monetary aggregates are policy invariant. The results further suggest that the demand for profit-sharing deposits is especially policy invariant.

Table 7 about here

Given the importance of the above results, we pursue additional tests to check their robustness to reasonable model adjustments. First, following Psaradakis and Sola (1996), we adjust the conditional money demand models by sequentially deleting variables with insignificant coefficients. Results from the modified models persist in suggesting that the contemporaneous variables in both ECMs are super-exogenous. The final M1 specification includes $(\sigma_{zz}^2)^2 \eta^Z$ for the growth of the real exchange rate with a coefficient of -9.53 ($t=-2.15$), further supporting super-exogeneity for the $\Delta\pi_i$ variable in the conditional M1 demand model. However, the statistically significant coefficient for $(\sigma_{zz}^2)^2 \eta^Z$ for the growth in the real exchange rate could weaken the super-exogeneity of
this contemporaneous variable in the conditional M1 model. As for the conditional model of the profit-sharing aggregate, the final specification includes $\eta^Z$ with a coefficient of -$0.00084$ ($t=-0.637$), which confirms the super-exogeneity of the $\Delta \pi_t$ variable in the profit-conditional model. These findings further support the conclusion that the profit-sharing aggregate possesses a stronger policy invariance property compared to the traditional M1 aggregate.

The second sensitivity test regards the fact that, under structural invariance, the determinant of parameter non-constancy in the marginal process should not affect the conditional model (Psaradakis and Sola, 1996). Hence, we examine the significance of the dummy variables in the two conditional models. The results indicate that none of these variables is significant in any conditional model, again confirming the robustness of our finding that both estimated money demand models are policy invariant and are useful guides for policy analysis in Iran. Further, under the policy invariance of the M1 and the profit-sharing aggregates, their sum (M2) should also have similar desirable properties. Estimates of M2 demand equation (available upon request) support this presumption.

Finally, although the inflation and real exchange rates proved stationary, the estimated ECMs dictate the use of their first differences. Thus, our results should be checked for possible over-differencing. Consequently, we reestimated the ECMs and their corresponding superexogeneity equations with the levels of these variables. The results (available upon request) altered none of our results reported in the text.

3.3. Long-Run Stability of Money Demand Models

Our final task in this paper is to examine the stability of the long-run demand models of the two alternative monetary aggregates in Iran. Hansen and Johansen (1993) outline a procedure that tests for the constancy of cointegrating vectors in the context of
FIML estimations. Holding the short-run dynamics of the tested model constant at the full sample estimates, the procedure treats these estimates as the null hypothesis in consecutive recursive tests. In this way, any rejection of the null of a stable cointegrating vector should emanate from a breakdown in the long-run relation, rather than from any possible shift in the underlying short-run dynamics (Hoffman et al., 1995).

Figures 1 and 2 plot the calculated values of the recursive test statistics for the real M1 and profit-sharing deposits models, respectively. Note that these statistics are recursive likelihood-ratios normalized by the 5% critical value. Thus, calculated statistics that exceed unity imply the rejection of the null hypothesis and suggest unstable cointegrating vectors. The broken curve (BETA_Z) plots the actual disequilibrium as a function of all short-run dynamics including seasonal dummy variables, while the solid curve (BETA_R) plots the “clean” disequilibrium that corrects for short-run effects. It is the corrected series that is actually tested for stationarity and thus determines the number of cointegrating ranks in the maximum likelihood procedure. Thus, if the solid line remains within the range of 0.0 and 1.0 all the betas are statistically constant over the sample period, see Hansen and Johansen (1993) as well as Hansen and Juselius (1995).

Figures 1 and 2 about here

We hold up the first fifteen years for the initial estimation. As both figures suggest, the demand for the two aggregates appears stable over the long run when the models are corrected for short-run effects. Note that the long-run demand for real M1 is stable even without adjustments for short-run dynamics. This is because, unlike the profit-sharing aggregate, the nature of the M1 aggregate is similar in the initial hold-up period as well as in the rest of the period. That is, M1 was interest free before and after banning interest rate transactions in Iran. In contrast, Figure 2 shows that without
adjustments for short-run effects, the cointegrating parameters for the profit-sharing aggregate are unstable until about 1990, after which they become highly stable. Since this aggregate was interest-bearing for almost 13 years of the initial period (up to 1979), a longer hold-up period is required for the initial estimation. As Figure 2 also suggests, with an initial period of 1966-1990, the profit-sharing aggregate proves stable over the long run irrespective of any adjustment for the short-run dynamics. It should also be mentioned that although the solid line remains within the range of 0.0 and 1.0 and betas are statistically constant over the sample period we can see a sudden change in the test result for m1 in 1993 and 1995, see Figure 1. These changes in the LR test statistics could be due to the policy regime changes of the unification of official and market-determined foreign exchange rates in late March 1993 and the introduction of inflation targeting by the CBI during the period March 1995-March 1998.

4. Concluding Remarks

We examine the behavior of money demand in the Iranian economy using quarterly data spanning the period 1966-2001. Since the mid-1980s, interest-based financial transactions have been banned in Iran. Consequently, we test the demand for two alternative monetary aggregates; namely, the interest-free M1 and profit-sharing deposits. Unlike previous studies, the focus of this paper is on whether the estimated money demand models are policy invariant despite the numerous exogenous shocks and policy regime changes that have plagued Iran in recent years. We show that estimated money demand equations, besides being temporally stable, must also be policy invariant in order for these equations to be useful for monetary policy analysis.

The evidence which persistently emerges from a whole range of empirical models and tests suggests that the estimated demand for M1 and profit-sharing deposits in Iran
behave remarkably well and proved to be temporally stable both in the short and in the long run. Perhaps more importantly, the estimated money demand equations are also invariant to changes in policy regimes and other exogenous shocks that have characterized Iran over the past three decades. These findings prove robust and they stand up to various adjustments in model specifications. The results for Iran are broadly consistent with those reported by Darrat (2000, 2002).

The results further suggest that, of the two alternative monetary aggregates, the demand for profit-sharing deposits possesses the most stable and policy invariant function. This empirical finding accords well with theoretical evidence (e.g., Khan, 1986 and Chapra, 1992) indicating that the profit-risk-sharing banking scheme insulates the monetary system from interest-rate exposure risk and minimizes financial instability.

It is thus reasonable to argue that the elimination of fixed interest rates from the Iranian banking system and its replacement with the profit-sharing scheme since 1984 has not hampered the financial stability of the country. Indeed, the introduction of a profit-risk-sharing banking system has apparently strengthened Iran’s financial stability and provided the Central Bank with credible and reliable monetary policy instruments to fight against inflationary pressures.
References


Figure 1: Recursive Likelihood Ratio Tests for Interest-Free Monetary Aggregate

![Figure 1](image1.png)

Figure 2: Recursive Likelihood Ratio Tests for Profit-Sharing Monetary Aggregate

![Figure 2](image2.png)
### Table 1: Data Description and Summary Statistics
**Sample Period: 1966Q1-2001Q4**

<table>
<thead>
<tr>
<th>Variables</th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>log m1</td>
<td>5.57</td>
<td>0.70</td>
<td>3.94</td>
<td>6.18</td>
</tr>
<tr>
<td>log qm</td>
<td>5.52</td>
<td>0.79</td>
<td>3.37</td>
<td>6.30</td>
</tr>
<tr>
<td>log y</td>
<td>7.27</td>
<td>0.47</td>
<td>6.03</td>
<td>7.85</td>
</tr>
<tr>
<td>(\pi)</td>
<td>15.04</td>
<td>16.58</td>
<td>-25.00</td>
<td>78.38</td>
</tr>
<tr>
<td>log q</td>
<td>8.28</td>
<td>0.59</td>
<td>7.22</td>
<td>9.25</td>
</tr>
<tr>
<td>(r^*)</td>
<td>7.71</td>
<td>3.15</td>
<td>2.14</td>
<td>18.50</td>
</tr>
<tr>
<td>(\pi^*)</td>
<td>4.86</td>
<td>2.70</td>
<td>0.79</td>
<td>12.89</td>
</tr>
</tbody>
</table>

Notes: log m1 is the logarithm of real M1 (non-interest demand deposits plus currency with the public), log qm is the logarithm of real profit-sharing monetary aggregate (saving and term deposits that are based on profit-sharing), log y is the logarithm of real GDP, \(\pi\) is the inflation rate measured by the annualized percentage of the CPI (quarterly inflation rate multiplied by 400), log q is the logarithm of real exchange rate defined as the nominal market rial/$US exchange rate (domestic price per $US) multiplied by the CPI in the U.S. divided by the Iranian CPI, \(r^*\) is the London interbank LIBOR interest rate, and \(\pi^*\) is the United States inflation rate representing foreign inflation for Iran. Nominal magnitudes are deflated by the CPI to obtain real figures.
<table>
<thead>
<tr>
<th></th>
<th>$H_0=r$</th>
<th>$\lambda_{\text{max}}$</th>
<th>C. V. 95%</th>
<th>Trace</th>
<th>C. V. 95%</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>m1 System</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>0</td>
<td></td>
<td>36.4</td>
<td>28.14</td>
<td>75.07</td>
<td>53.12</td>
</tr>
<tr>
<td>1</td>
<td></td>
<td>17.73</td>
<td>22.00</td>
<td>65.20</td>
<td>34.91</td>
</tr>
<tr>
<td>2</td>
<td></td>
<td>14.42</td>
<td>15.67</td>
<td>20.93</td>
<td>19.96</td>
</tr>
<tr>
<td>3</td>
<td></td>
<td>6.50</td>
<td>9.24</td>
<td>6.50</td>
<td>9.24</td>
</tr>
<tr>
<td><strong>Profit-Sharing Deposits System (qm)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>0</td>
<td></td>
<td>40.03</td>
<td>28.14</td>
<td>82.24</td>
<td>53.12</td>
</tr>
<tr>
<td>1</td>
<td></td>
<td>21.55</td>
<td>22.00</td>
<td>42.20</td>
<td>34.91</td>
</tr>
<tr>
<td>2</td>
<td></td>
<td>13.07</td>
<td>15.67</td>
<td>20.65</td>
<td>19.96</td>
</tr>
<tr>
<td>3</td>
<td></td>
<td>7.38</td>
<td>9.24</td>
<td>7.38</td>
<td>9.24</td>
</tr>
</tbody>
</table>

**Specification Tests: m1 System**
- LM(1)  p-value = 0.05
- LM(4)  p-value = 0.11
- Normality p-value = 0.00

**Specification Tests: Profit-Sharing Deposits System**
- LM(1)  p-value = 0.12
- LM(4)  p-value = 0.54
- Normality p-value = 0.00

Notes: The maximal eigenvalue test statistics are corrected for small sample bias using the procedure outlined in Cheung and Lai (1993), while the trace statistics are corrected using the Johansen and Juselius (1991) procedure. The 95% critical values come from Osterwald-Lenum (1992). The lag length is 4 for m1 and 6 for qm. LM (1) and LM (4) are the Lagrangian Multipliers to test for autocorrelation of the first- and fourth-order, respectively. Normality is the Jarque and Bera test.
Table 3: Error-Correction Model: Instrumental-Variable Estimations
(Dependent Variable = Δlog m1t)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>Hansen’s stability L₄ test (5% critical value = 0.47)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.03</td>
<td>0.01</td>
<td>0.05</td>
</tr>
<tr>
<td>Δlog m₁₋₁</td>
<td>-0.17</td>
<td>0.06</td>
<td>0.25</td>
</tr>
<tr>
<td>Δlog m₁₋₃</td>
<td>-0.19</td>
<td>0.06</td>
<td>0.11</td>
</tr>
<tr>
<td>Δlog m₁₋₄</td>
<td>0.24</td>
<td>0.06</td>
<td>0.48</td>
</tr>
<tr>
<td>Δlog m₁₋₅</td>
<td>-0.11</td>
<td>0.06</td>
<td>0.25</td>
</tr>
<tr>
<td>Δlog y₋₃</td>
<td>0.37</td>
<td>0.06</td>
<td>0.28</td>
</tr>
<tr>
<td>Δπ₋₁</td>
<td>-0.001</td>
<td>0.0002</td>
<td>0.19</td>
</tr>
<tr>
<td>Δπ₋₂</td>
<td>-0.001</td>
<td>0.0003</td>
<td>0.03</td>
</tr>
<tr>
<td>Δπ₋₃</td>
<td>-0.001</td>
<td>0.0003</td>
<td>0.11</td>
</tr>
<tr>
<td>Δπ₋₄</td>
<td>-0.001</td>
<td>0.0003</td>
<td>0.18</td>
</tr>
<tr>
<td>EC₋₂</td>
<td>-0.03</td>
<td>0.01</td>
<td>0.08</td>
</tr>
<tr>
<td>Δlog q₋₁</td>
<td>0.13</td>
<td>0.04</td>
<td>0.15</td>
</tr>
<tr>
<td>Oil</td>
<td>-0.14</td>
<td>0.03</td>
<td></td>
</tr>
<tr>
<td>Q2</td>
<td>-0.05</td>
<td>0.01</td>
<td></td>
</tr>
<tr>
<td>Q3</td>
<td>-0.04</td>
<td>0.01</td>
<td></td>
</tr>
<tr>
<td>Q4</td>
<td>-0.03</td>
<td>0.01</td>
<td></td>
</tr>
</tbody>
</table>

Before the stability test Δlog m₁ was adjusted for these dummy variables to avoid non-invertible matrix.

Hansen’s stability L₄ test on variance of the ECM = 0.19

Joint (coefficients and the error variance)
Hansen’s stability L₄ test (5% critical value, for the degrees of freedom 16)=3.58

Notes: Oil is a dummy variable to account for the oil shock of the fourth quarter of 1973 and the first quarter of 1974. Q2, Q3 and Q4 are seasonal dummy variables for the second, third and fourth quarters of the year, respectively. EC is the error correction term which is defined as EC₁ = log m₁ - 1.61 log y₁ + 0.04 π₁ + 0.04 r*₁ + 0.57 log q. The instruments are first, fourth and fifth lags of the EC term from the m₁ equation. The error term generated from the inflation equation, which is ECπ₁ = logq₁ – 8.88, was not statistically significant and so it was dropped.

Specification Tests:
- $R^2=0.72$, $\sigma=0.03$, $DW=2.01$, Godfrey (5)=0.73 (significance level=0.62), White=127 (significance level=1.00), ARCH (5)=9.69 (significance level=0.08), RESET=0.84 (significance level=0.47) and Normality (Jarque-Bera $\chi^2$)=2.50, (significance level=0.29). To ensure the normality of the disturbance term, we include dummy variables accounting for outliers observed in the data for 1969Q4, 1972Q2, 1976Q2 and 1979Q1. The estimated coefficients of the dummy variables are not reported but are available upon request.
<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>Hansen’s stability $L_4$ test (5% critical value = 0.47)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta\log y_{t-2}$</td>
<td>0.19</td>
<td>0.04</td>
<td>0.80</td>
</tr>
<tr>
<td>$\Delta\pi_t$</td>
<td>-0.001</td>
<td>0.0002</td>
<td>0.42</td>
</tr>
<tr>
<td>$\Delta\pi_{t-1}$</td>
<td>-0.002</td>
<td>0.0002</td>
<td>0.25</td>
</tr>
<tr>
<td>$\Delta\pi_{t-2}$</td>
<td>-0.001</td>
<td>0.0002</td>
<td>0.24</td>
</tr>
<tr>
<td>$\Delta\pi_{t-3}$</td>
<td>-0.001</td>
<td>0.0002</td>
<td>0.10</td>
</tr>
<tr>
<td>$\Delta\pi_{t-4}$</td>
<td>-0.001</td>
<td>0.0002</td>
<td>0.23</td>
</tr>
<tr>
<td>$\Delta\pi_{t-5}$</td>
<td>-0.0007</td>
<td>0.0001</td>
<td>0.11</td>
</tr>
<tr>
<td>$\Delta\pi_{t-6}$</td>
<td>-0.0004</td>
<td>0.0001</td>
<td>0.15</td>
</tr>
<tr>
<td>$\Delta r^*_{t-3}$</td>
<td>-0.006</td>
<td>0.002</td>
<td>0.31</td>
</tr>
<tr>
<td>$\Delta\pi^*_{t-3}$</td>
<td>-0.005</td>
<td>0.001</td>
<td>0.15</td>
</tr>
<tr>
<td>$EC_{t-3}$</td>
<td>-0.03</td>
<td>0.002</td>
<td>0.03</td>
</tr>
</tbody>
</table>

Hansen’s stability $L_4$ test on variance of the ECM 0.29

Joint (coefficients and the error variance) Hansen’s stability $L_4$ test (5% critical value, for the degrees of freedom 12)=2.96 2.76

Notes: The instruments used are first, third and fourth lags of the EC term for the profit-sharing deposit equation, where $EC_t = \log q_{mt} - \log y_t + 0.05 \pi_t - 0.26 \pi^*_t + 0.20 r^*_t$. The error term generated from PPP, which is $ECPP_{Tt} = \log q_t - 8.88$, was not statistically significant and so it was dropped.

Specification Tests: $R^2=0.81$, $\sigma=0.02$, $DW=2.03$, Godfrey (5)=1.69 (significance level=0.13), White=89.59 (significance level=1.00), ARCH (5)=5.89 (significance level=0.32), RESET=0.12 (significance level=0.95) and Normality (Jarque-Bera $\chi^2$)=3.57, significance level=0.17). Note that to ensure the normality of the disturbance term we also included dummy variables accounting for outliners observed in 1975Q1, 1978Q4, 1979Q1, 1980Q1, 1980Q4, 1984Q2 and 1985Q2. Before conducting the stability test $\Delta\log q_m$ was adjusted for these dummy variables to avoid non-invertible matrix.
**Table 5: Marginal Model**  
*(Dependent Variable = $\Delta \pi_t$)*

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>Standard Error</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-15.03</td>
<td>2.17</td>
</tr>
<tr>
<td>$\Delta \pi_{t-1}$</td>
<td>-0.62</td>
<td>0.08</td>
</tr>
<tr>
<td>$\Delta \pi_{t-2}$</td>
<td>-0.36</td>
<td>0.07</td>
</tr>
<tr>
<td>$(\Delta \pi)(Zero)_{t-1}$</td>
<td>1.32</td>
<td>0.55</td>
</tr>
<tr>
<td>$\Delta \log q_{t-1}$</td>
<td>31.66</td>
<td>11.67</td>
</tr>
<tr>
<td>$(\Delta \log q)(Rev)_{t-3}$</td>
<td>56.90</td>
<td>12.06</td>
</tr>
<tr>
<td>$\Delta \log rgd_{t-1}$</td>
<td>49.67</td>
<td>20.38</td>
</tr>
<tr>
<td>Q1</td>
<td>14.58</td>
<td>3.42</td>
</tr>
<tr>
<td>Q3</td>
<td>17.79</td>
<td>3.22</td>
</tr>
<tr>
<td>Q4</td>
<td>23.41</td>
<td>3.79</td>
</tr>
</tbody>
</table>

Notes: Zero is a dummy variable representing the introduction of the interest-free banking system in Iran and is equal to one for 1984Q1-2001Q4 and is zero otherwise. Rev is a dummy variable to account for the revolution in Iran. It is equal to one for 1979Q2-2001Q4 and zero otherwise. Q1, Q2 and Q3 are seasonal dummy variables for the first, third and fourth quarters of the year. The estimation method is OLS.

**Specification Tests:** $R^2=0.66$, $\sigma=11.61$, $DW=2.12$, Godfrey (5)=0.43 (significance level=0.86), White=45.40 (significance level=0.97), ARCH (5)=5.58 (significance level=0.35), RESET=1.02 (significance level=0.36), Normality (Jarque-Bera ($\chi^2$)=2.62, significance level=0.27).
Table 6: Marginal Model  
(Dependent Variable = $\Delta \log q_t$)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>Standard Error</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.04</td>
<td>0.01</td>
</tr>
<tr>
<td>$\Delta \log q_{t-1}$</td>
<td>0.68</td>
<td>0.22</td>
</tr>
<tr>
<td>$\Delta \log q_{t-11}$</td>
<td>-0.64</td>
<td>0.12</td>
</tr>
<tr>
<td>$(\Delta \log q)(Rev)_{t-1}$</td>
<td>-0.78</td>
<td>0.23</td>
</tr>
<tr>
<td>$(\Delta \log q)(Rev)_{t-3}$</td>
<td>-0.41</td>
<td>0.08</td>
</tr>
<tr>
<td>$(\Delta \log q)(Zero)_{t-11}$</td>
<td>0.56</td>
<td>0.14</td>
</tr>
<tr>
<td>$(\Delta \log q)(War)_{t-6}$</td>
<td>-0.28</td>
<td>0.13</td>
</tr>
<tr>
<td>Rev</td>
<td>0.13</td>
<td>0.02</td>
</tr>
<tr>
<td>Zero</td>
<td>-0.13</td>
<td>0.02</td>
</tr>
<tr>
<td>Q1</td>
<td>-0.07</td>
<td>0.02</td>
</tr>
<tr>
<td>Q4</td>
<td>-0.08</td>
<td>0.02</td>
</tr>
</tbody>
</table>

Notes: War is a dummy variable to capture the Iraq-Iran war. It is equal to one for 1980Q4-1988Q3, and zero otherwise. Q1 and Q4 are dummy variables for the first and fourth quarters of the year. The estimation method is OLS.

Specification Tests: $R^2=0.45$, $\sigma=0.07$, DW=2.01, Godfrey (5)=0.03 (significance level=0.99), White=34.16 (significance level=1.00), ARCH (5)=9.50 (significance level=0.09), RESET=0.93 (significance level=0.42), Normality (Jarque-Bera ($\chi^2$)=477, significance level=0.00). To mitigate the non-normality of the disturbance term, we include dummy variables accounting for outliners observed in 1988Q3, 1996Q2 and 1999Q2. The estimated coefficients of these dummy variables are not reported, but are available upon request.
Table 7: Super-exogeneity Test Results

<table>
<thead>
<tr>
<th>Variable Z</th>
<th>Δlog m1_t</th>
<th>Δlog qm</th>
</tr>
</thead>
<tbody>
<tr>
<td>Z – η^Z</td>
<td>0.32</td>
<td>-0.95</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.35)</td>
</tr>
<tr>
<td>σ^2^Z (Z – η^Z)</td>
<td>2.03</td>
<td>-0.48</td>
</tr>
<tr>
<td></td>
<td>(0.58)</td>
<td>(0.63)</td>
</tr>
<tr>
<td>(η^Z)^2</td>
<td>-0.23</td>
<td>0.49</td>
</tr>
<tr>
<td></td>
<td>(0.67)</td>
<td>(0.63)</td>
</tr>
<tr>
<td>(η^Z)^3</td>
<td>-1.83</td>
<td>-0.01</td>
</tr>
<tr>
<td></td>
<td>(0.32)</td>
<td>(0.99)</td>
</tr>
<tr>
<td>σ^2^Z η^Z</td>
<td>-43.79</td>
<td>-0.29</td>
</tr>
<tr>
<td></td>
<td>(0.22)</td>
<td>(0.77)</td>
</tr>
<tr>
<td>σ^2^Z (η^Z)^2</td>
<td>14.17</td>
<td>-0.81</td>
</tr>
<tr>
<td></td>
<td>(0.90)</td>
<td>(0.42)</td>
</tr>
<tr>
<td>(σ^2^Z)^2 η^Z</td>
<td>619.39</td>
<td>0.34</td>
</tr>
<tr>
<td></td>
<td>(0.50)</td>
<td>(0.11)</td>
</tr>
<tr>
<td>F-Statistics (14, 98 for m1), (7, 111 for qm)</td>
<td>1.48</td>
<td>0.61</td>
</tr>
<tr>
<td></td>
<td>(0.13)</td>
<td>(0.74)</td>
</tr>
</tbody>
</table>

Notes:
Specification Tests for Δlog m1: \( R^2 = 0.74, \sigma = 0.03, DW = 2.01, \text{Godfrey (5)} = 0.50 \) (significance level=0.80), White=59.63 (significance level=1.00), ARCH (5)=8.02 (significance level=0.16), RESET=0.53 (significance level=0.66) and Normality (Jarque-Bera (\chi^2)=0.23, significance level=0.89).

Specification Tests for Δlog qm: \( R^2 = 0.80, \sigma = 0.02, DW = 2.08, \text{Godfrey (5)} = 1.61 \) (significance level=0.15), White= 134.96 (significance level=1.00), ARCH (5)= 6.86 (significance level=0.23), RESET=0.19 (significance level=0.90) and Normality (Jarque-Bera (\chi^2)=3.00, significance level=0.22).