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

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Abstract

This paper extends the literature on interest-free banking systems by modeling money demand equations under profit-sharing scheme. Since Iran has followed profit-sharing banking system since early 1980 the models are estimated on quarterly Iranian data for the period 1966-2001. Unlike prior research, this paper focuses on whether the estimated equations are policy invariant in addition of being stable in the short- and long run. Our empirical results persistently suggest that both money demand models, M1 and profit-sharing monetary aggregates, and especially the demand for profit-sharing deposits, are stable and policy invariant despite the numerous shocks that have characterized Iran in recent years. These results provide another piece of evidence supportive of the merit of the interest-free-profit-sharing banking system, and suggest that profit-sharing monetary aggregates represent a credible instrument for monetary policy-making in Iran.

Keywords: Profit-sharing deposits, Interest-free banking system, Policy-invariance, Super-exogeneity, Long- run stability, Central Bank of Iran

JEL Codes: E41, E52

1. Introduction

The concept of profit-sharing in the banking industry, as opposed to the alternative 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 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.[2] With only a few exceptions [Darrat (1988, 2002) and Yousefi *et al.* (1997)], 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 demand function is required by almost all theories of macroeconomic activities and particularly for the smooth operation of an effective monetary policy. An unstable

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[1] The principle underlying the interest-free banking system is not without some support from prominent economists. For example, Friedman (1969) argues that zero nominal interest rates are necessary for optimal resource allocation, while Cole and Kocherlakota (1998) suggest that this "Friedman rule" of zero interest rates is both necessary and sufficient condition for optimality. Commenting on possible causes of the East Asian financial crisis, Wilson (1998) argues that at the heart of the problem is that the funds flowing to the region were not participatory.

[2] See, for example, Bashir (1983), Khan (1986), Darrat (1988, 2002), Khan and Mirakhor (1990), Chapra (1992), Yousefi *et al.* (1997), and Aggarwal and Yousef (2000).

function undermines 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 New Classical models [Sargent and Wallace (1975)], in some Neo-Keynesian models [Mankiw (1991)], and in models of real business cycles [King *et al.* (1991).

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 we show in this paper, the introduction of an expected rate of profit (instead of the fixed rate of interest) into money demand models provides money demand relationships with an internal source of stability.

This paper focuses on the Iranian experience with the profit-sharing banking scheme. Compared to most other countries that have experimented with interest-free banking systems, Iran provides an interesting case since the elimination of all interest-based financial dealings is perhaps most closely and consistently practiced in Iran, and for a relatively long time (since the mid-1980's). Moreover, Iran has also witnessed several changes in policy regimes and undergone numerous exogenous shocks during the past two decades which makes Iran an almost ideal case to test whether its money demand equations have endured all such shocks and regime changes.

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 short-run *and* long-run money demand equations for Iran. 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 be policy invariant as well. As Lucas (1976) points out, temporal stability and policy invariance are two 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. Interest-Free and Profit-Sharing Money Demand Models 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, CBI imposed a minimum "profit" rate for bank depositors to ensure the attractiveness of such deposits. Various reports of 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: short-term 7%, one-year 13% and five-year 17%. With an annual inflation rate running at about 35%, one apparent purpose of these minimum profit rates is to mitigate 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:

$$E\,\{\sum_{t=0}^{\infty}\beta^{t}\,U(c_t, c^*_t, S_t)\}, \tag{1}$$

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 U.S. dollar represents 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. Specifically,

$$S_t = S\,(m_t, m^*_t) \tag{2}$$

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 $S_{m^*} = \partial S(m, m^*)/\partial m^*>0$.

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

$$\tau_t + y_t + (1 + \pi_t)^{-1} m_{t-1} + q_t (1 + \pi^*_t)^{-1} m^*_{t-1} + (1 + \pi_t)^{-1} (1 + r_t) d_{t-1} + q_t (1 + \pi^*_t)^{-1} (1 + r^*_{t-1}) d^*_{t-1} = c_t + q_t c_t^* + m_t + q_t m_t^* + d_t + q_t d_t^* \tag{3}$$

where τ_t is the real value of any lump-sum taxes/transfers received/paid by consumers, 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, y_t is the current real endowment (income) received by the individual, m^*_{t-1} is foreign real money holdings at the start of the 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 $E(r_{t+1} \mid I_t) = r_t^e$, and d_t* is the real foreign one-period time (non-checking) deposits which pays a predetermined risk-free interest rate r_t*. Assume further that d_t and d_t* are the only two storable assets.

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 λ_t = the marginal utility of wealth at time t. Substituting S_t from (2) into (1), and then maximizing the resulting preferences with respect to m, c, m*, c*, d and d*, subject to budget constraint (3) will yield the first-order conditions:

$$U_{ct} + \lambda_t = 0 \tag{4}$$

$$U_{c^*t} + \lambda_t q_t = 0 \tag{5}$$

$$U_{st} S_{mt} + \lambda_t - \beta\lambda^e_{t+1} (1 + \pi^e_{t+1})^{-1} = 0 \tag{6}$$

$$U_{st} S_{m^*t} + \lambda_t q_t - \beta\lambda^e_{t+1} q^e_{t+1} (1 + \pi^{*e}_{t+1})^{-1} = 0 \tag{7}$$

$$\lambda_t - \beta\lambda^e_{t+1} (1 + r^e_{t+1}) (1 + \pi^e_{t+1})^{-1} = 0 \tag{8}$$

$$\lambda_t q_t - \beta\lambda^e_{t+1} q^e_{t+1} (1 + r^*_t) (1 + \pi^{*e}_{t+1})^{-1} = 0. \tag{9}$$

Note that $x^e_{t+1} = E(x_{t+1} \mid I_t)$ is the conditional expectations of x_{t+1}, given current information I_t. From (4) and (5) we can write:

$$U_{ct}/U_{c^*t} = 1/q_t. \tag{10}$$

Equation (10) indicates that the marginal rate of substitution between domestic and foreign goods is equal to their relative price. Solving (5), (7) and (9) yields:

$$U_{c^*t} (1 + r^*_t)^{-1} + U_{st} S_{m^*t} = U_{c^*t}. \tag{11}$$

Equation (11) implies that the expected marginal benefit of adding to foreign currency holdings at time t must equal the marginal utility from consuming foreign goods at time t.

Note that the holdings of foreign currency directly yield utility through its services (U_{st} S_{m*t}). Furthermore, from (9) we have $U_{c*t} = \beta\lambda^{e}_{t+1}\, q^{e}_{t+1}\, (1 + r^{*}_{t})\, (1 + \pi^{*e}_{t+1})^{-1}$ which implies that real foreign currency invested in foreign deposits is expected to have a value of $\beta\lambda^{e}_{t+1}q^{e}_{t+1}(1+r^{*}_{t})(1+\pi^{*e}_{t+1})^{-1}$. Consequently, total marginal benefit of money at time t is $U_{c*t}\,(1 + r^{*}_{t})^{-1} + U_{st}\, S_{m*t}$.

Similarly, from (4), (6) and (8), we have:

$$U_{ct}\,(1 + r^{e}_{t})^{-1} + U_{st}\, S_{mt} = U_{ct} \qquad (12)$$

Equation (12) implies that the expected marginal benefit from adding to domestic currency holdings at time t must equal the marginal utility of consuming domestic goods at time t. To construct a parametric example of equation (12), substitute equation (2) into (1) and assume the resulting indirect utility has an instantaneous function as:

$$U(c_t, c^{*}_{t}, m_t, m^{*}_{t}) = (1 - \sigma)^{-1}[c^{\alpha 1}{}_{t}\, c^{*\alpha 2}{}_{t}\, m^{\eta 1}{}_{t}\, m^{*\eta 2}{}_{t}]^{1-\sigma}, \qquad (13)$$

where σ , α_1, α_2, η_1 and η_2 are positive parameters. The demand for domestic real balances, using equations (12) and (13) will be:

$$m_t = (\eta_1 c_t) / \alpha_1\, r^{e}_{t+1}\, (1 + r^{e}_{t+1})^{-1} \qquad (14)$$

From (14), we have $m_{ct} = \partial m_t/\partial c_t > 0$ and $m_{ret+1} = \partial m_t/\partial r^{e}_{t+1} < 0$. Equation (14) can be rewritten as:

$$\log(m_t) = \log(\eta_1) + \log(c_t) - \log(\alpha_1) - \log[r^{e}_{t+1}\, (1 + r^{e}_{t+1})^{-1}] \qquad (15)$$

Let domestic real consumption (c_t) be some constant proportion (ω) of domestic real income (y_t). Furthermore, assume the current relevant information for estimating r^{e}_{t+1} includes current inflation rate (π_t), foreign interest rate (r^{*}_{t}), and real exchange rate (q_t). We may hypothesize:

$$\log[r^{e}_{t+1}\, (1 + r^{e}_{t+1})^{-1}] = \theta_2\, \pi_t + \theta_3\, r^{*}_{t} + \theta_4 \log(q_t) + u_t, \qquad (16)$$

where θ'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 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 θ_3 may be indeterminate. For θ_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 θ_4 negative over the short run. Substituting $c_t=\omega\, y_t$, and (16) into (15), yield the following final m1 demand equation:

$$\log m1_t = \beta_0 + \beta_1 \log y_t + \beta_2\, \pi_t + \beta_3\, r^*_t + \beta_4 \log q_t + u_{t,} \qquad (17)$$

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. Furthermore, 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; π is the CPI inflation rate, r* is the London interbank offer rate, 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 βs are the parameters to be estimated.

We now turn our attention to deriving an estimable equation for the demand for real profit-sharing monetary aggregate (d_t). From equations (10) and (13), we have:

$$c^*_t = \alpha_2 c_t / q_t \alpha_1. \tag{18}$$

Using equations (10), (11), (13) and (18), we can write:

$$m^*_t = (\eta_2 c_t) / \alpha_1 q_t r^*_t (1 + r^*_t)^{-1}. \tag{19}$$

Assume τ_t=0 and $d_t^* = \nu_0\ r^*_t{}^{\nu 1}\ y_t^{\nu 2}$, where ν's are constant parameters. Substituting $c_t(=\omega y_t)$, $\tau_t(=0)$, $d_t^*(= \nu_0\ r^*_t{}^{\nu 1}\ y_t^{\nu 2})$, along with equations (14), (18) and (19), into budget line (3), we can derive:

$$d_{t-1} = \frac{X_t}{R_t}, \tag{20}$$

where $R_t = \dfrac{(1 + r_t)}{(1 + \pi_t)}$ is the real profit rate and $X_t = \omega y_t + q_t\ \eta_1 \alpha_2 \omega y_t / \alpha_1 + (\eta_1 \omega y_t)/\alpha_1 r^e_{t+1} (1 + r^e_{t+1})^{-1} + q_t (\eta_2\ \omega y_t) / \alpha_1 r^*_t (1 + r^*_t)^{-1} + d_t + q_t \nu_0\ r^*_t{}^{\nu 1}\ y_t^{\nu 2} - y_t - (1 + \pi_t)^{-1} (\eta_1\ \omega y_{t-1}) / \alpha_1\ r^e_t (1 + r^e_t)^{-1} - q_t (1 + \pi^*_t)^{-1} (\eta_2\ \omega y_{t-1}) / \alpha_1 r^*_{t-1} (1 + r^*_{t-1})^{-1} - q_t (1 + \pi^*_t)^{-1} (1 + r^*_{t-1}) \nu_0\ r^*_{t-1}{}^{\nu 1}\ y_{t-1}^{\nu 2}$.

Note that, from equation (20), we can obtain $d_t = \dfrac{X_{t+1}}{R_{t+1}}$, $d_{t+1} = \dfrac{X_{t+2}}{R_{t+2}}$, $d_{t+2} = \dfrac{X_{t+3}}{R_{t+3}}$ and so on. Substitute $d_{t+1}(= \dfrac{X_{t+2}}{R_{t+2}})$ into $d_t(= \dfrac{X_{t+1}}{R_{t+1}})$ to eliminate d_{t+1}. By successive elimination of this type, we can arrive at an equation for d_t as a function of current and expected future values of y, q, R, π* and r*, provided the transversally condition is satisfied. Note that the present value of d_t approaches zero as $t \to \infty$. From $d_t = \dfrac{X_{t+1}}{R_{t+1}}$, it can be easily shown that $\partial d_t/\partial y_t > 0$, $\partial d_t/\partial r_t < 0$, $\partial d_t/\partial r^*_t, < 0$ and $\partial d_t/\partial \pi^*_t, > 0$, but the sign of $\partial d_t/\partial q_t$ is indeterminate. As before, we assume investors use currently available information on these variables to forecast their future values.

The final demand equation for profit-sharing deposits can be approximated in the following form:

$$\log qm_t = \gamma_0 + \gamma_1 \log y_t + \gamma_2 \pi_t + \gamma_3 \pi^*_t + \gamma_4 r^*_t + \gamma_5 \log q_t + u_{t,} \qquad (21)$$

where γ's are the parameters, and qm denotes d. As it was shown above, $\gamma_1>0$, $\gamma_2<0$, $\gamma_3>0$, $\gamma_4<0$, and γ_5= indeterminate. Note that π* is the U. S. inflation rate as a proxy for foreign inflation, and r* is the London interbank (LIBOR) rate to represent foreign interest rates.

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.

We should emphasize that our both money demand equations ((17) and (21)) suggest the expected rate of profit in the banking system as a key opportunity cost for holding money. Only Bashir (2002) recently outlines a model with some similar features, though in the context of a closed-economy model. Our proposed model is also different from recent Caganian-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) and Yousefi *et al.* (1997), but similar to Stock and Watson (1993) and Muscatelli and Spinelli (2000), our money demand equations are long-run variant.

2.2. Data and Cointegration Test Results

Our data on Iran are quarterly spanning the period 1966Q1-2001Q4, and come from the international *Financial Statistics CD-ROM,* of the International Monetary Fund[3]. Table (1) provides some data description and summary statistics. According to Augmented Dickey-Fuller and the Phillips-Perron test results all variables, except for the inflation and real exchange rates, are nonstationary in levels, but they achieve stationarity when converted to first-differences. Stationarity test results are reported in Table (2) of Kia and Darrat (2003).

Tables 1 about here

Since the model contains at least two 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)].

Tables 2 about Here

[3] A few years have missing observations which we obtain from the series used by Yousefi *et al.* (1997). 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.

Table (2) reports the results from the λ_{max} and trace tests for equations (17) and (21) for the lag length of 4 and 6, respectively.[4]. The λ_{max} test rejects r = 0 at the 5% level, while r ≤1 is not rejected, implying that r=1. According to the trace test, we reject the null hypothesis of r ≤1 at the 5% level, while we cannot reject the null hypothesis of r ≤2, implying that r=2. To further investigate the number of cointegrating ranks, we estimate eigenvalues of the companion matrix. We find that all roots are either equal to or less than one. The two largest roots are 0. 0.9729 ≈ 1, followed by a complex pair of roots with modulus 0.8623 ≠ 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, as the trace test did, that r=2.

However, since r=2 the system is not identified. Following Johansen (1995b) and Kia (2003), we investigate possible economic hypotheses underlying the multiple cointegrating vectors. We first focus on the cointegrating relation between inflation and real exchange rates. That is, we test if the following long-run inflation/exchange rate relationship exists:

$$\pi_t = \chi_0 + \chi_1 \log q_t + u_{t,} \qquad (22)$$

where χ's are coefficient and 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. Higher nominal exchange rate will also depress foreign prices of exports, resulting in higher demands for exports. Higher demands for exports will exert pressures on domestic resources, putting further upward pressures on domestic prices. This suggests χ_1>0. With

[4] According to diagnostic test results reported in the table, these lag lengths are sufficient to eliminate autocorrelation. The only non-congruency is non-normality. However, Johansen (1995a) shows that departures from normality are not alarming in cointegration tests.

this sign restriction, the system becomes overidentified and the rank condition is not satisfied. To ensure the rank condition, we further impose a zero restriction on the constant of the m1 demand equation (17). These restrictions ensure generic empirical and economic identifications [Johansen and Juselius (1991)].

The resulting estimation of equation (22) as well as the restricted long-run demand for m1 (the figures in parentheses beneath the estimated parameters are standard errors) are:

$$\underset{}{\pi_t} = \underset{(1135.81)}{-\,9198.56} + \underset{(134.48)}{707.43}\log q_t \tag{23}$$

$$\log m1_t = \underset{(0.18)}{1.61}\log y_t - \underset{(0.01)}{0.04}\,\pi_t - \underset{(0.03)}{0.04}\,r^*_t - \underset{(0.15)}{0.57}\log q_t. \tag{24}$$

All 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 long-run inflation equation and 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.

As for equation (21) a lag length of 6 was required to ensure the error term is not autocorrelated, see LM test results in Table (2). According to λ_{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 $r \leq 1$ at the 5% level, but cannot reject the null of $r \leq 2$, implying that $r=2$. However, all roots of the estimated eigenvalues of the companion matrix are either equal to unity or inside the unit disc, where the largest root is

$0.9817 \approx 1$, followed by a complex root with modulus $0.9444 \neq 1$, each of which implies only one root. Thus, we may conclude that r=1 as the λ_{max} test suggests.

We report below the estimated long-run real demand for profit-sharing monetary aggregate (qm), where the figures in brackets beneath the estimates are the corresponding p-values for chi-squared exclusion tests:

$$\log qm_t = - 0.97 + 0.64 \log y_t - 0.11\ \pi_t + 0.57\ \pi^*_t + 0.43\ r^*_t + 0.63 \log q_t. \quad (25)$$

$$[0.90]\ \ [0.60] \qquad [0.00] \quad [0.00] \quad [0.00] \quad [0.24]$$

The estimated coefficient on real income has the correct positive sign, but is statistically insignificant. The remaining coefficients also have the correct theoretical signs and, except for the constant term and the coefficient of the real exchange rate, are statistically significant including the coefficient on foreign interest rates.

Two features of the results are puzzling. The estimated coefficient of real income fails to achieve statistical significance, and 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 (Q*ard Hasan*) that should not respond to "material" yields.[5] One possible explanation of these results is that, with non-stationary variables, the calculated chi-squared statistics are not very reliable for measuring statistical significaunce.[6]

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[5] According to data from CBI, the portion of goodwill loans in the profit-sharing monetary deposits increased from 11% in March 1995 to almost 17% in March 2001. Note that Iranian banks do not pay any yield on goodwill deposits since they are obliged to use such funds in the form of interest-free loans to individuals. However, private conversations suggest that public banks in Iran usually offer up to 3% yield on goodwill deposits.

[6] As the sample size increases, the mean of a non-stationary variable approaches its true value and the distribution of, say, $((E(x_t) - x_t) / \sqrt{n})$, for x=log y, r* and log q quickly approaches normality. However,

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, 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 first, fourth and fifth lagged values of the error terms for M1 first, third and fourth lag values of the error term for profit-sharing deposits.

Tables 3 and 4 about Here

As the specification test result reported in tables (3) and (4) indicates none of the diagnostic checks is significant. According to Hansen's stability L test result (5% critical value=0.47, Hansen (1992), Table 1), all of the coefficients are stable. Furthermore, the joint Hansen's (1992) stability L_c test result is 2.29 (<3.58 for 16 degrees of freedom) for M1 and 0.73 (<1.36 for 5 degrees of freedom) for profit-sharing deposits, which in both cases indicates that we cannot reject the null of joint stability of the coefficients together with the estimated associated variance. It is interesting to note that both individual and joint stability tests indicate a stronger result for profit-sharing aggregate.

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the variance of the estimator may also quickly explode as $n \rightarrow \infty$. Thus, the standard central limit theorem

The only contemporaneous variables in the short-term demand for M1 are the change in quarterly inflation rate and the growth rate of real exchange rate with the correct sign. For profit-sharing aggregate, however, the contemporaneous variable is only the change in inflation rate and has again a correct sign. All possible kinds of 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 were included. According to our estimation results, the error-correction term is significant for both aggregates and the impact is linear, but it takes longer (two quarters) for the M1 market to adjust to disequilibrium error. Furthermore, the error-correction term resulted from deviation from the long-run equilibrium relationship (23) was not statistically significant and so it was dropped from the EC model.

The results reported in tables (3) and (4) accord well with our theoretical model and the underlying theory. However, as one would expect in an Islamic system at least over the short run, the foreign interest rate was not statistically significant in either aggregates and was dropped. Interestingly, according to the results reported in Table (4) in demanding profit-sharing deposits agents use current change in the inflation rate as well as the previous change in the level of their deposits to forecast future profits of the banking system. Furthermore, the coefficient on the EC term in both of the two money demand equations is negative (error correcting) and proved to be 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'

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may not apply even in large samples.

forecasts of future real income, the inflation rate, the profit-sharing rate and/or monetary policy moves (forward-looking behavior). Under the latter adjustment 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 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 profit-sharing deposits, our focus next is on whether these estimated money demand equations can be reliable tools for policy analysis. That is, we test if the estimated money demand equations are invariant to policy changes and other exogenous shocks which require that the variables in the money demand equations be 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 super-exogeneity of these variables in turn requires the estimation of marginal models for these variables against the backdrop of several possible regime changes.

A perusal of Iranian modern history indicates that there have been six major regime changes over the past three decades. They are: (i) the revolution of April 1979; (ii) the ban on fixed interest rate in the banking system that began in March 1984, (iii) the Iran/Iraq 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 1979: II- 2001: IV, and = 0, otherwise, Zero = 1 from 1984: I- 2001: IV, and = 0, otherwise, War = 1 from 1980: IV-1988: III, and = 0, otherwise, Ue = 1 from 1993: I, and = 0, otherwise, Inflation = 1 from 1995: II-1998: I, and = 0, otherwise, and Private = 1 from 1997: III-2001: IV, and = 0, otherwise.

Tables 5 and 6 about Here

Tables (5) and (6) report 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 evince no major violations of key assumptions[8]. 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. 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 test 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 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 (=Δlog $m1_t$ or Δlog q_t) and Z_t as:

$$X_t = \alpha_0 + \psi_0 Z_t + (\delta_0 - \psi_0)(Z_t - \eta^Z{}_t) + \delta_1 \sigma_t^{ZZ} (Z_t - \eta^Z{}_t) + \psi_1 (\eta^Z{}_t)^2 + \psi_2 (\eta^Z)^3$$
$$+ \psi_3 \sigma_t^{ZZ} \eta^Z{}_t + \psi_4 \sigma_t^{ZZ} (\eta^Z{}_t)^2 + \psi_5 (\sigma_t^{ZZ})^2 \eta^Z + z'_t\gamma + u_t \qquad (26)$$

where α_0, ψ_0, ψ_1, ψ_2, ψ_3, ψ_4, ψ_5, δ_0 and δ_1 are regression coefficients on Z_t conditional on $z'_t\gamma$, and u_t is a white-noise disturbance term. The 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 $\eta^Z{}_t = E(Z_t \mid I_t)$ and $\sigma_t^{ZZ} = E[(Z_t - \eta^Z{}_t)^2 \mid I_t]$ are the conditional moments of Z_t, given 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, δ_0-ψ_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 δ, and assuming that σ_t^{ZZ} has distinct values over different (but definable) regimes, δ_1 must equal zero. If all these hypotheses are not rejected, the contemporaneous variables in the ECMs become super-exogenous and the estimated ECMs can be considered invariant to policy shocks[9].

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[8] The only exception is perhaps non-normality of the real exchange rate, which is a common problem in most marginal models [see Hurn and Muscatelli (1992), and Metin (1998)].
[9] The appendix in Kia and Darrat (2003) provides further details.

We estimate η^Z and σ_t^{ZZ} for Z_t from the marginal models reported in Tables (4) and (5). Since the errors for the Z_t variable appear homoskedastic according to an ARCH test, we experimented with 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.

Table 7 about Here

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 both 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.

Given the importance of the above conclusion, we pursue additional tests to check the robustness of the results 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 specification for M1 includes $(\sigma^{ZZ})^2\eta^Z$ for the growth of the real exchange rate with a coefficient of -9.53 (t=-2.15), further corroborating super-exogeneity of $\Delta\pi_t$ variable in the conditional demand model of M1. However, the statistically significant coefficient for $(\sigma^{ZZ})^2\eta^Z$ for the growth of the real exchange rate may weaken the super-exogeneity of this contemporaneous variable in the conditional model of M1. As for the conditional model of profit-sharing aggregate, the final specification includes

η^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 provide additional support to the conclusion that the profit-sharing aggregate possesses a stronger policy invariance property compared to the traditional M1 aggregate.

Secondly, we note that structural invariance implies that 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 the dummy variables is significant in any conditional model, again attesting to the robustness of our finding that both estimated money demand models are policy invariant and can be reliably used for policy analysis in Iran. We finally note that since the M1 and the profit-sharing aggregates are policy invariant, one would expect their sum (i.e., M2) to have similar desirable properties. Results from estimating and testing the demand for M2 (available upon request) confirm this presumption.

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 plots 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 rejection of the null hypothesis and suggest unstable cointegrating vectors. The blue-colored curve (BETA_Z) plots actual disequilibrium as a function of all short-run dynamics including seasonal dummy variables, while the black-colored curve (BETA_R) plots "clean" disequilibrium that corrects for short-run effects. We hold up the first fifteen years for the initial estimation. As both figures show, the demand for

Figures 1 and 2 about here

the two aggregates appear 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 the ban on 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 whether or not adjustments are made for short-run dynamics.

4. Concluding Remarks

We investigate 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, this paper examines the demand for two alternative aggregates; namely, the interest-free M1 and profit-sharing deposits. Unlike previous studies, our focus is on whether the estimated money demand models are policy invariant especially in the face of numerous exogenous shocks and policy regimes changes that have plagued Iran in recent years. Besides being temporally stable (backward-looking behavior), we show that estimated money demand equations must also be policy invariant (forward-looking behavior) 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 recently 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), Chapra (1992)] indicating that the profit-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. To the contrary, the introduction of profit-sharing banking system has apparently strengthened Iran's financial stability and provided the Central Bank with credible and reliable monetary policy instruments in their important and ongoing fight against inflationary pressures.

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

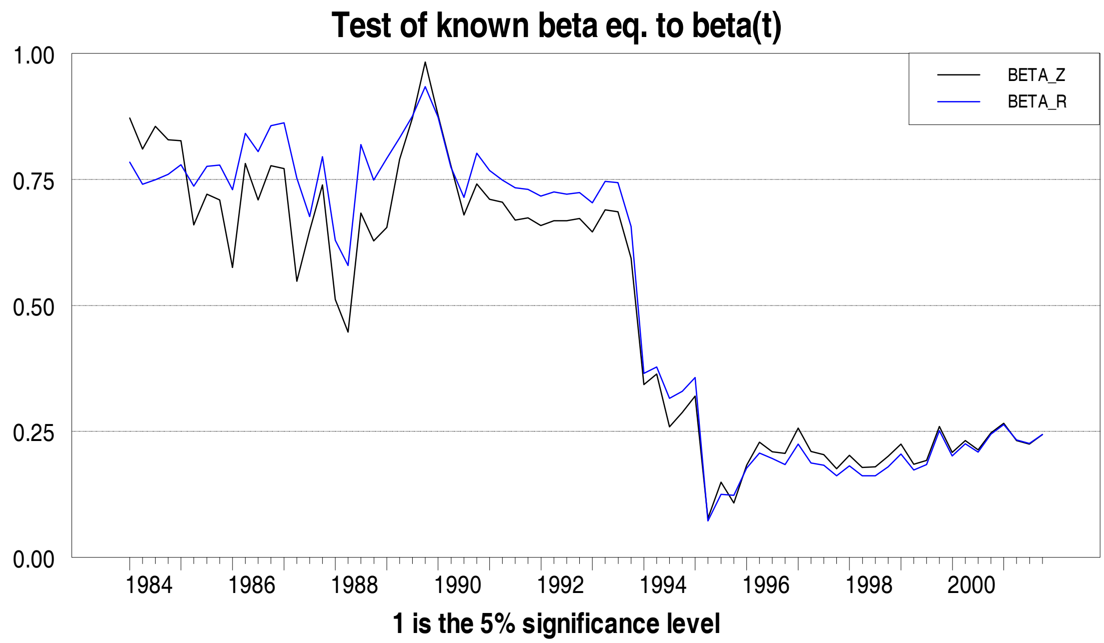

Figure 2: Recursive Likelihood Ratio Tests for Profit-Sharing Monetary Aggregate

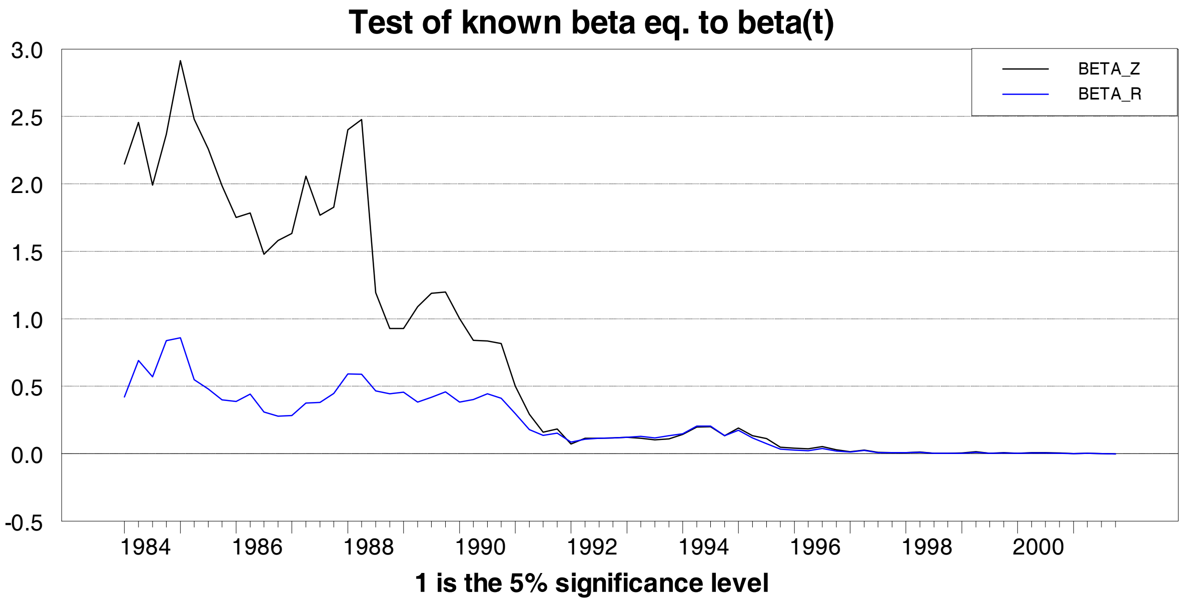

Table 1: Data Description and Summary Statistics
Sample Period: 1966:Q1 -- 2001:Q4

Variables	Mean	Standard Deviation	Minimum	Maximum
Log m1	5.57	0.70	3.94	6.18
Log qm	5.52	0.79	3.37	6.30
Log y	7.27	0.47	6.03	7.85
π	15.04	16.58	-25.00	78.38
Log q	8.28	0.59	7.22	9.25
r^*	7.71	3.15	2.14	18.50
π^*	4.86	2.70	0.79	12.89

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 log of real GDP, π 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/U.S exchange rate (domestic price per U.S. dollar) multiplied by the CPI in the U.S. divided by the Iranian CPI, r^* is the London interbank LIBOR interest rate, and π^* is the United States inflation rate representing foreign inflation for Iran. Nominal magnitudes are deflated by the CPI to obtain real figures.

Table 2: Test Results of the Cointegration Rank				
H_0=r	**λ_{max}**	**C. V. 95%**	**Trace**	**C. V. 95%**
m1 System				
0	36.4	28.14	75.07	53.12
1	17.73	22.00	65.20	34.91
2	14.42	15.67	20.93	19.96
3	6.50	9.24	6.50	9.24
Profit-Sharing Deposits System (qm)				
0	40.03	28.14	82.24	53.12
1	21.55	22.00	42.20	34.91
2	13.07	15.67	20.65	19.96
3	7.38	9.24	7.38	9.24

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 = $\Delta \log m1_t$)

Variable	Coefficient	Standard Error	Hansen's stability L_i test (5% critical value = 0.47)
Constant	0.03	0.01	0.05
$\Delta \log m1_{t-1}$	-0.17	0.06	0.25
$\Delta \log m1_{t-3}$	-0.19	0.06	0.11
$\Delta \log m1_{t-4}$	0.24	0.06	0.48
$\Delta \log m1_{t-5}$	-0.11	0.06	0.25
$\Delta \log y_{t-3}$	0.37	0.06	0.28
$\Delta \pi_t$	-0.001	0.0002	0.19
$\Delta \pi_{t-1}$	-0.001	0.0003	0.03
$\Delta \pi_{t-2}$	-0.001	0.0003	0.11
$\Delta \pi_{t-3}$	-0.001	0.0003	0.18
EC_{t-2}	-0.03	0.01	0.08
$\Delta \log q_t$	0.13	0.04	0.15
Oil	-0.14	0.03	Before the stability test $\Delta \log m1$ was adjusted for theses dummy variables to avoid non-invertible matrix.
Q2	-0.05	0.01	
Q3	-0.04	0.01	
Q4	-0.03	0.01	
Hansen's stability L_i test on variance of the ECM	0.19		
Joint (coefficients and the error variance) Hansen's stability L_c test (5% critical value) =3.58	2.30		

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. The instruments are first, fourth and fifth lag of the EC term from the m1 equation.

Summary Statistics: $\overline{R}^2$=0.72, σ=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 (χ^2=2) =2.50 (significance level=0.29). To ensure normality of the disturbance term, we include dummy variables accounting for outliners observed in the data for 1969Q4, 1976Q2, 1972Q2, and 1979Q1. The estimated coefficients of the dummy variables are not reported but are available upon request.

Table 4: Error-Correction Model: Instrumental-Variable Estimations (Dependent Variable = $\Delta \log qm_t$)

Variable	Coefficient	Standard Error	Hansen's stability L_i test (5% critical value = 0.47)
Constant	0.07	0.004	0.35
$\Delta \log qm_{t-1}$	0.35	0.06	0.14
$\Delta \pi_t$	-0.002	0.0002	0.09
EC_{t-4}	-0.01	0.002	0.30
Hansen's stability L_i test on variance of the ECM	0.08		
Joint (coefficients and the error variance) Hansen's stability L_c test (5% critical value) =3.58	0.73		

Notes: See notes to Table 5. The instruments used are first, third and fourth lag of the EC term for the profit-sharing deposits equation.

Summary Statistics: $\overline{R}^2$=0.61, σ=0.03, DW=2.09, Godfrey (5) =0.80 (significance level=0.57), White=18.51 (significance level=1.00), ARCH (5) =2.20 (significance level=0.82), RESET=0.11 (significance level=0.96) and Normality (χ^2=2) =0.41 (significance level=0.82). Note that to ensure normality of the disturbance term we also included dummy variables accounting for outliners observed in 1975Q1, 1978Q4, 1980Q1, Q4, 1984Q2 and 1985Q2.

Table 5: Marginal Model
(Dependent Variable = $\Delta\pi_t$)

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

Notes: Zero is a dummy variable representing the introduction of the interest-free banking system in Iran and is equal to one for 1984: I – 2001: IV and is zero otherwise. Rev is a dummy variable to account for the revolution in Iran. It is equal to one for 1979: II – 2001: IV 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.

Summary Statistics: $\overline{R}^2$=0.66, σ=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 (χ^2=2) =2.62 (significance level=0.27).

Table 6: Marginal Model
(Dependent Variable = Δlog q_t)

Variable	Coefficient	Standard Error
Constant	0.04	0.01
Δlog q_{t-1}	0.68	0.22
Δlog q_{t-11}	-0.64	0.12
(Δlog q)(Rev)$_{t-1}$	-0.78	0.23
(Δlog q)(Rev)$_{t-3}$	-0.41	0.08
(Δlog q)(Zero)$_{t-11}$	0.56	0.14
(Δlog q)(War)$_{t-6}$	-0.28	0.13
Rev	0.13	0.02
Zero	-0.13	0.02
Q1	-0.07	0.02
Q4	-0.08	0.02

Notes: See notes to Table 7. War is a dummy variable to capture the Iraq-Iran war. It is equal to one for 1980: IV – 1988: III, and zero otherwise. Q1 and Q4 are dummy variables for the first and fourth quarters of the year. The estimation method is OLS.

Summary Statistics: $\overline{R}^2$=0.45, σ=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 (χ^2=2) =477 (significance level=0.00).). To mitigate 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

Variable Z	Δlog $m1_t$		Δlog qm_t
Contemporaneous Variable	Δlog q_t	$\Delta\pi_t$	$\Delta\pi_t$
$Z - \eta^Z$	0.32 (0.02)	0.00 (0.14)	-0.00 (0.76)
$\sigma^{ZZ}(Z - \eta^Z)$	2.03 (0.58)	0.00 (0.28)	0.00 (0.61)
$(\eta^Z)^2$	-0.23 (0.67)	-0.00 (0.25)	0.00 (0.36)
$(\eta^Z)^3$	-1.83 (0.32)	0.00 (0.96)	0.00 (0.99)
$\sigma^{ZZ}\eta^Z$	-43.79 (0.22)	-0.00 (0.87)	0.00 (0.99)
$\sigma^{ZZ}(\eta^Z)^2$	14.17 (0.90)	0.00 (0.74)	0.00 (0.93)
$(\sigma^{ZZ})^2\eta^Z$	619.39 (0.50)	-0.00 (0.78)	0.00 (0.68)
F-Statistics (14, 98 for m1), (7, 118 for qm)	1.48 (0.13)		1.10 (0.37)

Notes:

Summary Statistics for Δlog ml: $\overline{R}^2$=0.74, σ=0.03, DW=2.01, 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 (χ^2=2) =0.23 (significance level=0.89).

Summary Statistics for Δlog qm: $\overline{R}^2$=0.61, σ=0.03, DW=2.15, Godfrey (5) =1.15 (significance level=0.33), White=51 (significance level=1.00), ARCH (5) =1.35 (significance level=0.93), RESET=0.01 (significance level=0.99) and Normality (χ^2=2) =0.13 (significance level=0.93).