

A Long Run Structural Macroeconometric Model for Iran

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We employ the modelling strategy of Garratt, Lee, Pesaran and Shin (2003a) to estimate a structural cointegrating VARX* model for Iran in which core macroeconomic variables of the Iranian economy are related to current and lagged values of a number of key foreign variables. The long run macroeconomic relations for real money balances, interest rates, output, prices and exchange rates are identified and tested within this framework over the period 1979Q1-2007Q4. We make use of generalised impulse response functions to analyze the dynamic properties of the model following a shock to exogenous variables (oil prices and foreign interest rates). We also examine via the persistence profiles, the speed of adjustments to the long run relations following a system-wide shock. The results show that money demand relation and UIP-PPP (international parity conditions jointly) are not rejected within the model. Furthermore, these two long run relations have well-behaved persistence profiles in which the effects of system wide-shocks on the long run relations are transitory and die out eventually. However, both UIP-PPP and the money demand relations exhibit sluggish rates of adjustments to shocks. We also provide evidence for the excessive importance of oil price shocks for Iranian economy in our impulse response analysis.

Keywords: cointegrated vector autoregression (VARX*), long run relations, Iranian economy, oil price, shock, foreign interest rate shock.

JEL classification: C10, C22, E20, E30

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1. Introduction

International economic interdependencies among countries mean that national economic issues have to be considered from a global perspective. However, most national econometric models do not have a coherent global dimension and interdependencies are only allowed through an informal off-model way, which casts doubt on the plausibility of the results. Garratt, Lee, Pesaran and Shin (2003b and 2006) contribute towards this issue by developing a transparent approach with a theoretically coherent foundation. The authors improve on modelling frameworks of King and Watson (1991), Gali (1992), Mellander, Vredin, and Warne (1991) and Crowder, Hoffman, and Rasche (1999), and develop a structural cointegrating VAR approach suitable for modelling an economy in a global context. This new strategy, which has been used extensively in the literature¹, offers a practical device to testing the long-run relationships suggested by economic theory while at the same time it allows for a flexible modelling of short-run dynamics. Production, arbitrage, solvency and portfolio balance conditions, or alternatively DSGE models are used to identify the long-run relations.

Many macroeconomic models have been estimated for the Iranian economy in the past, including large simultaneous equation systems and structural or non-structural vector autoregressives (VARs), though most of these models only consider domestic variables in their specifications for technical reasons. See for example Noferesti and Arabmazar (1993), Valadkhani (1997, 2007), Mehrara and Oskoui (2007), Elyasiani and Zhao (2008), among others. Iran, being engaged in economic activities with other countries, is strongly influenced by international oil price movements and economic growth in the rest of the world. Monetary policy positions taken by other countries are also likely to affect Iran's macroeconomy, though little is known or have been done about the significance of such factors in shaping Iran's economic activities.

This paper employs the modelling device of Garratt, Lee, Pesaran and Shin (2003b and 2006) to specify a structural macroeconomic model for Iran in a global context. We augment a typical cointegrated vector autoregressive model (VAR) with weakly exogenous (or long-

1. Assenmacher-Wesche and Pesaran (2008), Pesaran and Smith (2006), and Chen , Schneider, and Frohn (2007).

run forcing¹) variables and refer to it as Iran's VARX* model. We estimate this model subject to exact and over-identifying restrictions using quarterly data over the period 1979Q1 to 2007Q4. Theory driven long-run relationships are identified and tested within the model and imposed when acceptable. We also provide a detailed analysis of the dynamic properties of this model with a particular emphasis on the monetary policy following the approach of Dees *et al.* (2007). In particular, we carry out an analysis of the dynamic impacts of oil price shocks on core macroeconomic variables of Iran using the generalized impulse response functions. In addition, we employ a sieve bootstrap procedure for simulation of the model as a whole, which is then used in testing the structural stability of the parameters, and to obtain confidence intervals for the responses of the economy to various shocks such as changes in international oil prices and foreign monetary policy stances taken by other countries.

This paper is organized as follows. Section 2 introduces the data, and presents the theory based long run restrictions that can be tested in the context of Iran's VARX* model. Section 3 sets out the econometric methodology. Section 4 investigates the error-correcting properties of the model, and reports the empirical results, and Section 5 concludes. The data sources are provided in the appendix.

2. Modelling Choices

We specify a VARX* model that relates the core macroeconomic variables of the Iranian economy to current and lagged values of a number of key foreign variables, reflecting the interlinkages of Iran with the rest of the world. As shown in Pesaran and Smith (2006), the VARX* model can be derived as the solution to a small open economy Dynamic Stochastic General Equilibrium (DSGE) model. Therefore, it is possible in principle to impose short- and long-run DSGE-type restrictions on the VARX* model, although in this paper we shall focus on the long-run relations and leave the short-run parameters unrestricted. We incorporate those key relations from economic theory that can be expected to have an important effect on the Iranian economy. One of these long run restrictions is the money demand, which postulates a relationship between the real money stock, real output and the interest rate. Another is the Fisher equation,

1. In a global context, the foreign variables will be long run forcing for the core domestic variables.

which establishes a long-run relation between the interest rate and the inflation rate. We expect Iran to be influenced by exchange rate movements. Therefore, purchasing power parity, which links the domestic price level to the nominal exchange rate and the foreign price level, is also included. In addition, we consider the price of oil as the most important commodity price with direct and indirect impacts on the world economy in our analysis. Finally, we consider long-run relations between domestic and foreign output and interest rates to account for international business/interest rate cycles. It must be mentioned that the present model has its own restrictions including the fact that it mainly focuses on the monetary sector in Iran and does not consider all remaining sectors of the Iranian economy. The model does not account for credit market imperfections, labour market frictions, and real wage rigidities as well. Developing such a complicated model is beyond the scope of this research.

2.1. Data on the Core Variables and Long Run Equilibrium Conditions

The data span from the first quarter of 1979 to the last quarter of 2007. The endogenous (domestic) variables include real GDP, y_t , real money balances and quasi money, m_{2t} , short term interest rate which is constructed as the average of expected rates of return on housing services, and short term (special) investment deposit rates, r_t , Consumer Price Index (CPI), p_t , the quarterly rate of inflation, $\pi_t = p_t - p_{t-1}$, and finally the effective nominal exchange rate, e_t , (defined as domestic price of foreign currency basket, so that an increase in e_t represents a depreciation of home currency). The rest of the worlds variables consist of CPI, p_t^* , real GDP, y_t^* , and the nominal short term (three-month) interest rate, r_t^{s*} . The latter three variables, together with the nominal oil price, p_t^{oil} , are regarded as exogenous (long-run forcing). Except for the interest rates, all series are in logarithms. Interest rates are expressed as $0.25 * Ln(1 + R_t/100)$ where R_t is measured in percent per annum.

The foreign (star) variables are constructed as weighted averages of the corresponding variables in twenty-eight major trading partners of Iran. For instance, the foreign output is computed as

$$y_t^* = \sum_{j=0}^N w_{jt} y_{jt}$$

where y_{jt} is the logarithm of real output of country j , and w_{jt} is the time-varying trade weight between country j and Iran (the share of country j in total trade of Iran in U.S. dollars, such that $w_{jj} = 0$ and

$$\sum_{j=0}^N w_{jt} = 1).$$

The quarterly trade weights are computed as averages of Iran’s imports from and exports to the country in question divided by the total trade of all 28 countries, which cover about 85 percent of Iran’s total foreign trade. The countries included in our model are listed in Table (A.2.1), appendix A. Figure (1) shows the trade weights for Iran’s 8 largest trade partners. Japan is the most important trading partner of Iran accounting for a trade share of about 17 percent. More than 40 percent of the Iran’s trade is with the Euro area countries with the most important ones being Germany, Italy and France, who individually account for between 7 to 13 percent of Iran’s trade. Other countries in our data set with which Iran’s total trade is more than 5 percent are China, UK, Korea, and Turkey. A detailed description of the variables, their sources, and the construction of the foreign variables is also given in appendix A.

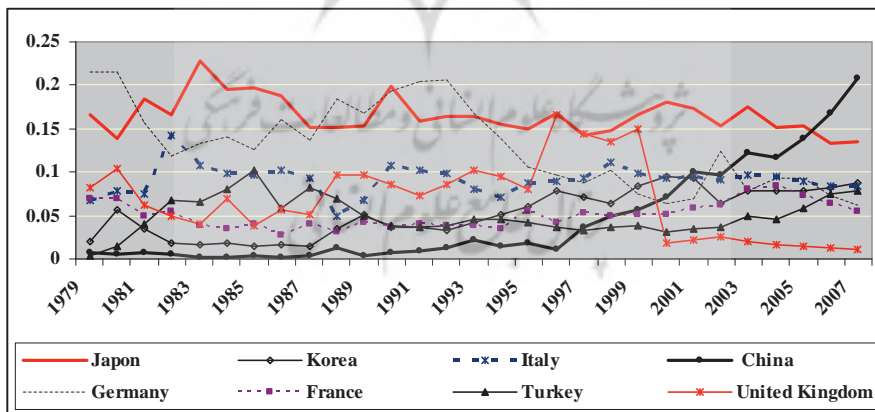


Figure 1. The major trading partners of Iran

The first step in our exercise is to test the long run relations that might exist amongst Iran's variables m_t , y_t , r_t , p_t , e_t , and their foreign counterparts, y_t^* , r_t^* , p_t^* . Theory-based long run equations can be derived either from equilibrium conditions of a canonical open economy DSGE model, see Pesaran and Smith (2006), or from arbitrage and solvency relationships, see Garratt, *et al.* (2003a, 2006). The long run relations mostly considered in the literature involve the purchasing power parity (PPP) condition, based on international goods market arbitrage, the output gap (OG), which is derived from the neoclassical growth model, the Fisher equation, based on arbitrage process between holdings of bonds and investing in physical assets, the uncovered interest parity (UIP) condition, based on arbitrage between domestic and foreign bonds, and the real money demand (MD) equation which is derived from the long-run solvency conditions and assumptions about the determinants of the demand for domestic and foreign assets.

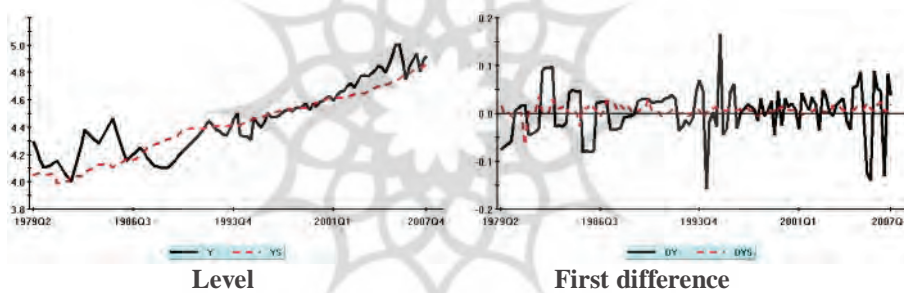


Figure 2. Domestic (Y) and foreign output (YS)

The case of Iran's economy deserves surely a more thorough investigation rather than simply adopting the six long run relations derived for a small open market economy in Garratt, *et al.* (2003, 2006). During the period from 1979Q1 to 2007Q4, Iran's economy went through a turbulent epoch. In the first two years of this period, the real GDP decreased almost by 25% of its level at 1979Q1 partly due to the redistributive and political conflicts that undermined the production and investment incentives after the Revolution of February 1979. During the war with Iraq from 1980 to 1988, Iran's economy also suffered a decline in real GDP (see Fig. 2). In mid 1990s, the decline in the real GDP in Iran coincided with the decline in the revenue of oil exports. The growth of the real GDP, especially since

2002, also coincided with the rapid growth in the revenue of oil exports. These particularities resulted in the persistent deviation of Iran's real GDP from its foreign counterpart (see Fig. 2). Using solely an "output gap" relation to characterize the difference between the real GDP in Iran and the real GDP in the rest of the world is thus not appropriate from both the theoretical and the empirical points of view. Therefore, this relationship will not be tested in the present work. In what follows, we focus on PPP, UIP and MD relationships.

We combine the two arbitrage conditions (the PPP and UIP) into a single equation, as the empirical literature is more supportive of such a combined relationship than of either PPP or UIP separately. The theoretical basis for this approach comes from the exchange rate overshooting model of Dornbush (1976) and Frankel (1979). Johansen and Juselius (1992) first tested the PPP-UIP relation jointly for the United Kingdom over the period 1972Q1 to 1987Q2. They rejected the hypothesis that the PPP relationship is stationary by itself. However, they found support for the combined version. Pesaran *et al* (2000) establish that combining the PPP and UIP equations into a single cointegrating vector performs better than attempting to attribute each relationship to different cointegrating vectors in the system. They argue that the joint treatment allows for gradual convergence to PPP, because deviations from PPP are more persistent than deviations from UIP. Other papers that support the joint test of PPP and UIP include Juselius (1995), Macdonald and Marsh (1997, 1999) and Juselius and Macdonald (2004, 2007). They argue that the combined PPP-UIP equation could also be used to test for possible exchange rate misalignments as it proposes a scheme for the exchange rate equilibrium using a single cointegrating vector involving the exchange rate, prices, and interest rates. In view of this literature, we consider the following two long-run relationships as possible candidates for the Iranian economy:

$$p_t - p_t^* - e_t = b_{10} + \beta_{16}r_t^{s*} - \beta_{15}r_t^s + \zeta_{1,t} \sim I(0),$$

$$m_t - y_t = b_{20} + \beta_{25}r_t^s + \zeta_{2,t} \sim I(0),$$

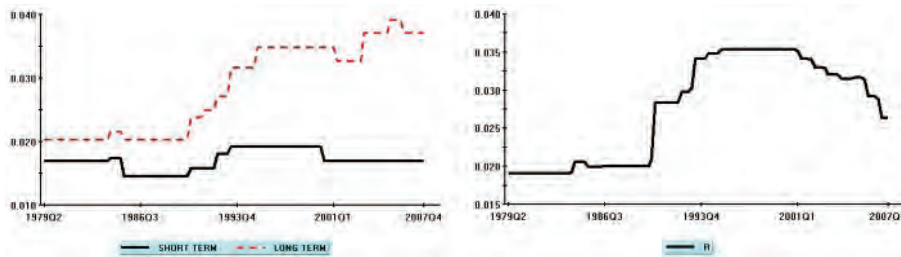


Figure 3. Short- and long-term deposit rates and the average of interest rates

Figure (3) depicts the short-run and long-term interest rates in Iran. These rates are clearly not market determined, as they are imposed administratively, and are relatively stable over time. To create more variation in our interest rate series, we construct a new measure by taking the average of expected rates of return (on exports, trade, services and miscellaneous, constructions and housing, manufacturing and mining, and agriculture), short term investment deposit rate, and one-year special deposit rate.

3. System Approach: Econometric Methodology

Pesaran and Shin (2002) and Pesaran, Shin and Smith (2000, hereafter PSS) modify and generalize Johansen's (1991, 1995) approach to estimation and hypothesis testing in vector error correction models (VECM) by changing its identification procedure. Their approach does not involve the empirical identification implicit in Johansen's reduced rank regression. Instead, they emphasize on the use of economic theory in the long run to solve the identification problem. PSS also develop the econometric techniques for analysing VAR models with weakly exogenous I(1) variables (VARX*). They allow for the possibility of drawing distinction between endogenous and exogenous variables and, in turn, specifying a VARX* error correcting model. It is plausible that some of the variables considered in our model are I(1) exogenous, such as foreign price, foreign interest rate and the oil price which justifies the use of VARX* in this research.

In error-correction form, the VARX*(p, p*) model can be written as

$$(1) \quad \Delta z_t = -\Pi z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + a_0 + a_1 t + u_t$$

where $z_t = (x_t', x_t^{*'})'$ consists of a $m_x \times 1$ vector of endogenous variables (x_t) and a $m_{x^*} \times 1$ vector of weakly exogenous variables (x_t^*), with $m_x + m_{x^*} = m$. a_0 denotes a $m \times 1$ vector of constants and a_1 is a vector of trend coefficients. Π is a $m \times m$ matrix of long-run multipliers, while $\{\Gamma_i\}_{i=1}^{p-1}$ are the matrices that summarise the short-run dynamics.

This model can be partitioned into a conditional model for the endogenous variables, Δx_t and a marginal model for the evolution of exogenous variables, Δx_t^* in which the parameter matrices and vectors Π, Γ_i, a_0, a_1 as well as the error term u_t are partitioned as:

$\Pi = (\Pi_x', \Pi_{x^*}')'$, $\Gamma_i = (\Gamma_{xi}', \Gamma_{x^*i}')'$, for $i = 1, \dots, p-1$, $a_0 = (a_{x0}', a_{x^*0}')'$, $a_1 = (a_{x1}', a_{x^*1}')'$, and $u_t = (u_{xt}', u_{x^*t}')'$. Consequently, the variance-covariance matrix of u_t is written as

$$\Sigma = \begin{pmatrix} \Sigma_{xx} & \Sigma_{xx^*} \\ \Sigma_{x^*x} & \Sigma_{x^*x^*} \end{pmatrix}$$

so that

$$u_{xt} = \Sigma_{xx^*} \Sigma_{x^*x^*}^{-1} u_{x^*t} + v_t$$

where $v_t \sim iid(0, \Sigma_{xx} - \Sigma_{xx^*} \Sigma_{x^*x^*}^{-1} \Sigma_{x^*x})$ is uncorrelated with u_{x^*t} by construction. The conditional model for the endogenous variables is written in terms of $z_{t-1}, \Delta x_t^*, \Delta z_{t-1}, \Delta z_{t-2}, \dots$; as follows:

$$(2) \quad \Delta x_t = -\Pi_x z_{t-1} + \Lambda \Delta x_t^* + \sum_{i=1}^{p-1} \Psi_i \Delta z_{t-i} + c_0 + c_1 t + v_t,$$

whereas the marginal model for the exogenous variables is

$$(3) \quad \Delta x_t^* = \sum_{i=1}^{p-1} \Gamma_{x^*i} \Delta z_{t-i} + a_{x^*0} + u_{x^*t},$$

in which

$$\Lambda \equiv \Sigma_{xx^*} \Sigma_{x^*x^*}^{-1}, \quad \Psi_i \equiv \Gamma_{xi} - \Sigma_{xx^*} \Sigma_{x^*x^*}^{-1} \Sigma_{x^*i}, \quad \text{and } c_1 \equiv a_{x1} - \Sigma_{xx^*} \Sigma_{x^*x^*}^{-1} a_{x^*1}.$$

for $i = 1, \dots, p-1$, $c_0 \equiv a_{x0} - \Sigma_{xx^*} \Sigma_{x^*x^*}^{-1} a_{x^*0}$

For the Iranian economy, it is reasonable to assume that x_t^* variables are weakly exogenous or long run forcing which renders $\Pi_{x^*} = 0$. This means that the information available from the marginal model is redundant for efficient estimation of the parameters of the conditional model. This restriction also implies that the variables x_t^* are I(1) and not cointegrated.

In all subsequent empirical analysis, we work in the context of a VARX* model with unrestricted intercepts and restricted trend coefficients such that $c_1 = \Pi_x \gamma$, where γ is a $m \times 1$ vector of free coefficients. These restrictions ensure that the solution of the model in levels of z_t will not contain quadratic trends (Garratt, *et al.* (2006)).

Finally, the above VECM model can be written as

$$(4) \quad \Delta x_t = -\Pi_x [z_{t-1} - \gamma(t-1)] + \Lambda \Delta x_t^* + \sum_{i=1}^{p-1} \Psi_i \Delta z_{t-i} + \bar{c}_0 + v_t,$$

Where $\bar{c}_0 = c_0 + \Pi_x \gamma$.

In what follows, having selected the order of the underlying VAR model (using model selection criteria such as the Akaike Information Criterion (AIC) or the Schwarz Bayesian Criterion (SBC)), we test for the number of cointegrating relations among the variables in z_t . We then compute maximum likelihood (ML) estimates of the model's parameters subject to exact and over-identifying restrictions on the long run coefficients.

4. Empirical Results

This section applies the econometric techniques described in Pesaran, *et al.* (2000) and Garratt, *et al.* (2006) to the VARX* model of Iran. The preliminary stage in our analysis is to establish the order of integration of the variables. Then we select the order of VAR using AIC and SBC and perform the cointegration tests. Having decided the

number of cointegrating vectors, we estimate the exact-identified model and test the validity of the over-identified long-run restrictions. Finally, we present the results from the VECM estimations and analyze the dynamics using persistence profiles and generalized impulse response functions.

4.1. Unit Root Properties of the Variables in the Model

Before estimating the model, it is important to establish the unit root properties of the variables under investigation to obtain a sensible interpretation of the long-run relations. The standard tests for unit roots (such as the Augmented Dickey - Fuller (1979) or the Phillips - Perron (1988) tests) provide important information on the nature of the persistence of the time series under consideration. The results of the Augmented Dickey-Fuller (ADF) and Philips-Perron (PP) tests, applied to the levels and first differences of variables over the sample period, are reported in Table (B.1) in appendix B.

Both sets of tests strongly support the view that $y_t, e_t, m_t, r_t, r_t^*$, and the price of oil, $Poil$, are I(1). The unit root hypothesis is rejected when applied to the first differences of these variables. However, the order of integration of domestic and foreign price variables is ambiguous. Application of ADF test to Δp_t and Δp_t^* yield mixed results. Domestic and foreign price levels appear in the PPP-UIP relationship which raises the potential difficulty of mixing I(1) and I(2) variables. Haldrup (1998) argues that it is often useful to transform the time series a priori to obtain variables that are unambiguously I(1) rather than working with a mixture of I(1) and I(2) variables. In the present context, we construct and work with a new variable, the relative price $rp = p_t - p_t^*$, rather than the two price levels p_t and p_t^* separately. As shown in Table (B.1), this new variable is clearly I(1) according to both tests.

In summary, it would be possible to consider all variables of vector z_t as I(1) on the basis of the unit root statistics reported. Moreover, it is easily established that the two exogenous variables are not cointegrated.

4.2. Estimation and Testing of the Long-Run Relationships

The first stage of our modelling sequence is to select the order of the underlying VAR. Here we find that a VAR of order one appears to be appropriate over the period 1979Q1 to 2007Q4 when using the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC) as the model selection measures. The maximum lag length is set to four based on the number of observations available. Considering a higher number of lags did not seem appropriate as increasing the lag order raises the number of coefficients to be estimated in a VAR quickly.

Table 1. Akaike and Schwarz Information Criteria for lag order selection

Lag length	Log likelihood	AIC	SBC
0	863.23	838.23	804.36
1	1210.9	1160.9	1093.1
2	1226.3	1151.3	1049.7
3	1238.6	1138.6	1003.1
4	1257.7	1132.7	963.37

We proceed with the cointegration analysis using a VARX*(1,1) model with an unrestricted intercept and a restricted time trend. The lag lengths of the domestic and foreign variables are set to 1; while r_t^* and the price of oil are treated as weakly exogenous I(1) or long-run forcing. Since the long-run relations may or may not contain linear trends, we test for co-trending, that is whether the trend coefficients are zero in the two cointegrating vectors. A war and revolution (WR) dummy is also included in our model to capture the effects of structural breaks (1979 Revolution and the 1980-1988 Iran-Iraq war). This variable takes the value of 1 between 1979Q1 and 1988Q2 and zero otherwise. We determine the number of cointegrating vectors given by Johansen's 'trace' and 'maximal eigenvalue' statistics. Table (2) shows the results with the associated 90% and 95% simulated critical values. The maximal eigenvalue statistic indicates the presence of one cointegrating relationship at the 5% significance level, which does not support our a priori expectations of two cointegrating vectors. However, as shown by Cheung and Lai (1993), the maximum eigenvalue test is generally less robust to the presence of skewness and excess kurtosis in the errors than the trace test. Given that we have evidence of non-normality in the residuals of the underlying VAR

model used to compute the test statistics, it is more appropriate to base our cointegration tests on the trace statistic, which supports the presence of two cointegrating relations at the 5% significance level.

Table 2. Cointegrating rank test statistics for the core model

H_0	H_1	Test statistic	95% Critical Value	90% Critical Value
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix				
$r = 0$	$r = 1$	55.28	51.07	46.91
$r \leq 1$	$r = 2$	40.42	42.82	39.64
$r \leq 2$	$r = 3$	28.80	36.54	33.51
$r \leq 3$	$r = 4$	22.88	29.16	26.69
$r \leq 4$	$r = 5$	12.31	20.82	18.45
Cointegration LR Test Based on Trace of the Stochastic Matrix				
$r = 0$	$r = 1$	159.69	127.21	122.48
$r \leq 1$	$r = 2$	104.41	94.89	90.33
$r \leq 2$	$r = 3$	63.99	67.57	63.15
$r \leq 3$	$r = 4$	35.19	42.34	39.70
$r \leq 4$	$r = 5$	12.31	20.82	18.45

The choice of two cointegrating vectors can be double-checked based on the satisfactory performance of the VARX* model in terms of stability, persistence profiles, VECM estimation results and the impulse responses.

Two restrictions must be imposed on each of the two cointegrating vectors (four restrictions in total) to exactly identify the long-run relationships. Having identified the long-run relations, we test for co-trending. The log-likelihood ratio statistic for jointly testing the two co-trending restrictions takes the value of 6.83, lower than the bootstrapped critical values at the 10% and 5% significance levels, 6.84 and 8.49, respectively. Therefore, we cannot reject the hypothesis that there are no linear trends in the cointegrating relations. The bootstrapped critical values are computed based on 3000 replications of the LR statistic¹ to overcome the small sample properties of the data.

1. For testing the long run restrictions, we use Microfit 5.0. For further technical details, see Pesaran and Pesaran (2009), Section 22.10. For generating the bootstrap critical values, we employ Gauss 6.0.

The next step in our exercise is to impose economically meaningful over-identifying restrictions that are in accordance with theoretical priors. We impose two restrictions on β , namely the PPP-UIP relation and the money demand (MD) equation:

$$\begin{aligned}(p_t - p_t^*) - e_t &= b_{10} + \beta_{16}r_t^{s*} - \beta_{15}r_t^s + \zeta_{1,t}, \\ m_t - y_t &= b_{20} + \beta_{25}r_t^s + \zeta_{2,t}\end{aligned}$$

These two long run relationship can be written as:

$$\xi_t = \beta'z_t - b_0 - b_1t - b_2WR$$

where

$$\begin{aligned}z_t &= (y_t, e_t, rp, m_t, r_t^s, r_t^*, p_t^0)' \\ b_0 &= (b_{10}, b_{20})', \quad b_1 = (0, 0)', \quad b_2 = (0, 0)', \quad \xi_t = (\xi_{2t}, \xi_{6t})' \\ \beta' &= \begin{pmatrix} 0 & -1 & 1 & 0 & \beta_{15} & \beta_{16} & 0 \\ -1 & 0 & 0 & 1 & \beta_{25} & 0 & 0 \end{pmatrix} \quad (\text{Case I}) \\ \beta' &= \begin{pmatrix} 0 & -1 & 1 & 0 & \beta_{15} & \beta_{16} & 0 \\ \beta_{21} & 0 & 0 & 1 & \beta_{25} & 0 & 0 \end{pmatrix} \quad (\text{Case II})\end{aligned}$$

β' is the over-identifying matrix, imposing all the restrictions as suggested by the two long run relations: PPP-UIP and MD.

Table 3 reports the long run restrictions that correspond to the Iranian economy where the output coefficient in the money demand equation is: (I) restricted to unity, (II) left unrestricted. We calculate the simulated critical values for the LR statistics by applying the non-parametric bootstrap method with 3000 replications. The results show that we cannot reject the null hypothesis of the 11 (case I) and 10 (case II) overidentifying restrictions. They also suggest that the presence of structural breaks in the data, which are taken into account by the inclusion of the dummy variable WR, does not affect the long run relationships. The absence of a time trend term and a dummy variable in the long-run relationships suggests that the variables in the two cointegrating vectors (based on Iran data) are not only co-trending but also co-breaking. In addition, we check whether the coefficients in the long run relations have the correct sign and plausible magnitude as

theory suggests. We proceed with case (II) for two reasons: 1) it is subject to fewer restrictions 2) the LR statistic for the joint test of the over-identifying restrictions, 23.73, is lower than its bootstrapped critical values at both 5 and 10 percent significance levels: 30.21 and 27.10, respectively.

Table 3. Over identified long run relations for Iran (1979Q1-2007Q4)

I- Restricted output Coefficient in money demand		LR(df)	CV
PPP-UIP	$(p_t - p_t^*) - e_t = b_{10} + 22r_t^{s*} - 140.84r_t^s + \zeta_{1,t}$ (16.48) (32.23)	$\chi^2[11]=36.30$	$\chi^2[11]=38.62$ 95% Critical Values
MD	$m_t - y_t = b_{20} - 8.76r_t^s + \zeta_{2,t}$ (5.51)		$\chi^2[11]=34.95$ 90% Critical Values
II- Unrestricted output Coefficient in money demand		LR(df)	CV
PPP-UIP	$(p_t - p_t^*) - e_t = b_{10} + 35.42r_t^{s*} - 121.34r_t^s + \zeta_{1,t}$ (14.51) (27.74)	$\chi^2[10]=23.73$	$\chi^2[10]=30.21$ 95% Critical Values
MD	$m_t = b_{20} + 1.39y_t - 17.52r_t^s + \zeta_{2,t}$ (0.13) (5.38)		$\chi^2[10]=27.10$ 90% Critical Values

Notes: Number in brackets refers to the number of over identifying restrictions =total number of restrictions- number of just identifying restrictions. Standard errors are given in parenthesis.

The coefficients on the price levels (domestic and foreign) are restricted symmetrically in the PPP-UIP relationship, while interest rates are unrestricted and allowed to affect the real exchange rate non-symmetrically. As shown in Table (3), the real exchange rate is positively related to the nominal interest semi-differential. It tends to depreciate when the domestic interest rate is higher than the foreign interest rate, reflecting expectations about future depreciations. The second cointegrating vector (money demand equation) shows that demand for real money balances increases with a rise in real income and fall in interest rate. The long run income and interest rate elasticities of demand are 1.39 and -17.52, respectively. Overall, an increase in domestic interest rate causes a fall in money demand and depreciation of the real exchange rate.

4.3. Error Correction Specifications

Short run dynamics of the model are characterised by the five error correction specifications given in Table (4). The error correcting terms are of significant importance as they provide a complex set of interactions and feedbacks across real and financial markets. The results show that $\hat{\xi}_{1,t-1}$ is statistically significant in the output and interest rate equations, while $\hat{\xi}_{2,t-1}$ is significant at the 5% level in the output and real money equations.

Table 4. Reduced form error correction specification for the Iran Model

Equation	Δy_t	Δe_t	Δr_t	$\Delta(p_t - p_t^*)$	Δm_t
Δr_t^*	-0.132 (0.522)	-0.097 (1.327)	-0.037** (0.007)	2.518** (1.218)	0.105 (0.473)
$\Delta Poil$	0.009 (0.036)	0.168* (0.093)	-0.001 (0.005)	0.047 (0.086)	-0.003 (0.033)
$\hat{\xi}_{1,t-1}$	0.015* (0.008)	-0.028 (0.021)	-0.006** (0.001)	0.007 (0.020)	-0.006 (0.008)
$\hat{\xi}_{2,t-1}$	0.111** (0.043)	-0.080 (0.109)	-0.009 (0.006)	0.151 (0.101)	-0.164** (0.039)
C	-0.320** (0.124)	0.406 (0.315)	0.006** (0.002)	-0.312 (0.289)	0.367** (0.112)
\bar{R}^2	0.028	0.044	0.404	0.038	0.116
$\chi_{SC}^2[4]$	11.180	1.708	8.911	2.514	39.034
$\chi_{FF}^2[1]$	1.713	1.618	35.580	7.131	3.540
$\chi_N^2[2]$	15.374	4838.3	454.119	1671.4	17.384
$\chi_H^2[1]$	1.219	2.060	33.294	5.790	0.009

Notes: The error correction terms, ζ_i , are defined on page 13. The two cointegrating relations are output gap and money demand equation respectively. Standard errors are given in parenthesis. An asterisk denotes significance at the 10 percent level and ‘***’ denotes significance at the 5 percent level. The diagnostics are chi-squared for serial correlation (SC), functional form (FF), normality (N), and heteroscedasticity (H). Critical values are 3.84 for $\chi^2(1)$, 5.99 for $\chi^2(2)$ and 9.49 for $\chi^2(4)$.

The diagnostic statistics indicate that serial correlation is present in the output and real money equations. This could be improved by including further lags. However, the size of the system would increase considerably as a result of adding lags, rendering this solution

unattractive. The test hypothesis for functional form cannot be rejected for the output, the exchange rate and the real money equations. The hypothesis of homoskedasticity of errors is rejected for the interest rate and the price differential equations. The assumption of normally distributed errors is rejected in all the error correction equations, which are understandable if we consider the three major hikes in oil prices experienced during the estimation period. Overall, the equations of Table (4) appear to capture well the time series properties of the main macroeconomic aggregates in Iran over the period 1979Q1-2007Q4.

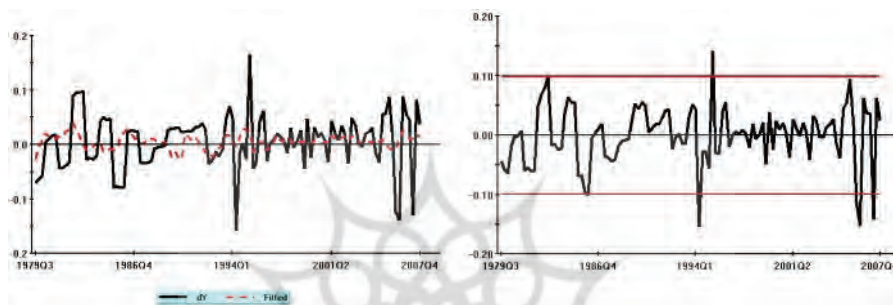


Figure 4. Real output equation

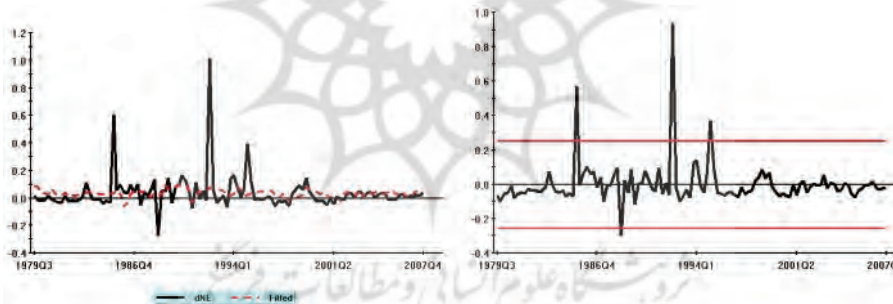


Figure 5. Nominal effective exchange rate equation

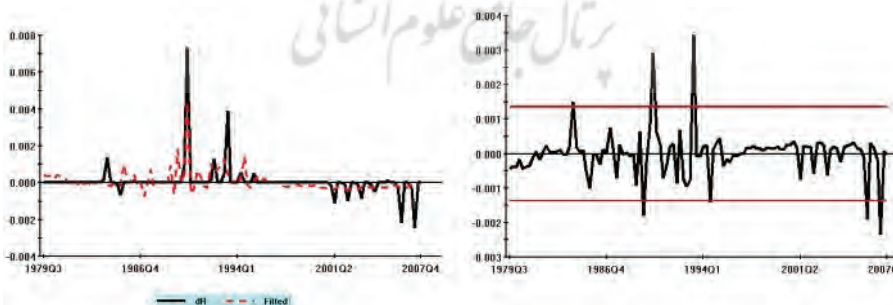


Figure 6. Interest rate equation

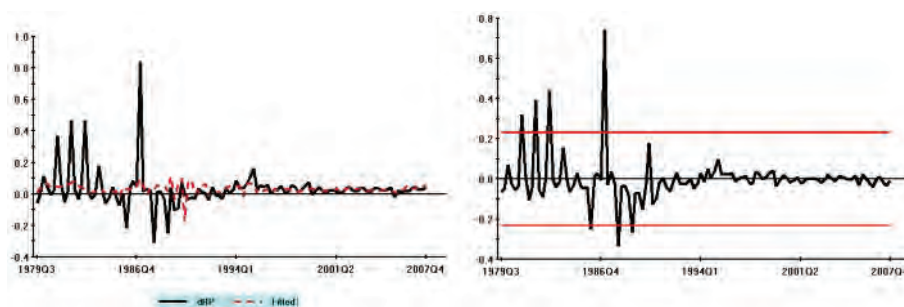


Figure 7. Relative price level equation

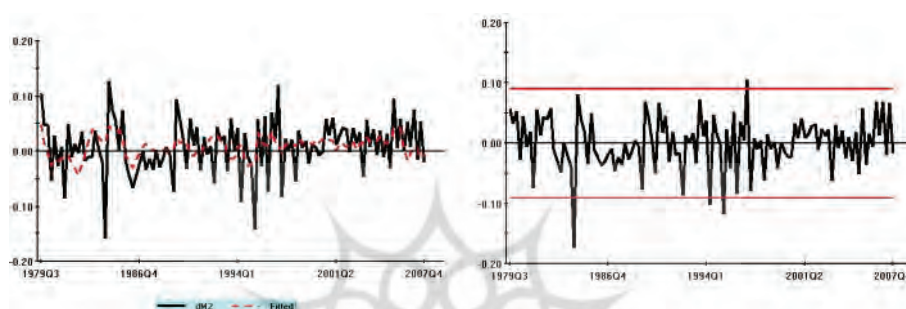


Figure 8. Real Money equation

Figures (4) to (8) plot the actual and fitted values of the reduced form error correction equations together with the corresponding residuals. We observe some large outliers especially in the mid 1990s for real output, in the mid 1980s and the beginning of the 1990s for the exchange rate, in the early 1990s for the interest rate, and at the beginning of the sample for the price differential. However, these outliers mainly reflect departures from normality, which are unlikely to have significant impacts on our main findings in our view.

4.4. Simulation Results

We analyze the model dynamics using impulse responses and persistence profiles. The former studies the evolution of the conditional means of the target variables in response to different types of shocks. In direct contrast to impulse responses, which focus on the impact of variable-specific shocks, persistence profiles are used to map out the dynamic response paths of system-wide shocks or composite shocks to the cointegrating relationships. In this section, we begin by presenting the persistence profiles to measure the speed of convergence of the cointegrating relations to equilibrium. We then

present the generalized impulse responses of the endogenous variables to oil price and foreign interest rate shocks. In all figures we plot the point estimates as well as the 95% simulated confidence intervals obtained from a non-parametric bootstrap method using 3000 replications. Also, we show the generalized impulse responses of the endogenous variables in the system to a one standard error shock to the various observables in the model in appendix C.

4.4.1. Persistence Profiles

This section uses persistence profiles to examine the effects of system wide shocks on the long run relationships. The value of these profiles is unity on impact. If the vector under investigation is indeed an equilibrium relationship, it should tend to zero as the time horizon goes to infinity. In addition, the persistence profiles provide useful information about the speed with which the different cointegrating relations in the model, once shocked, will return to their long-run equilibria.

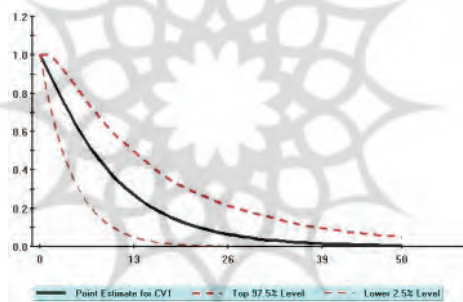


Figure 9. Persistence Profile of the effect of a system-wide shock to PPP-UIP

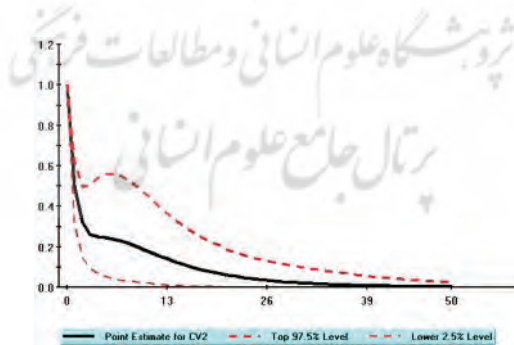
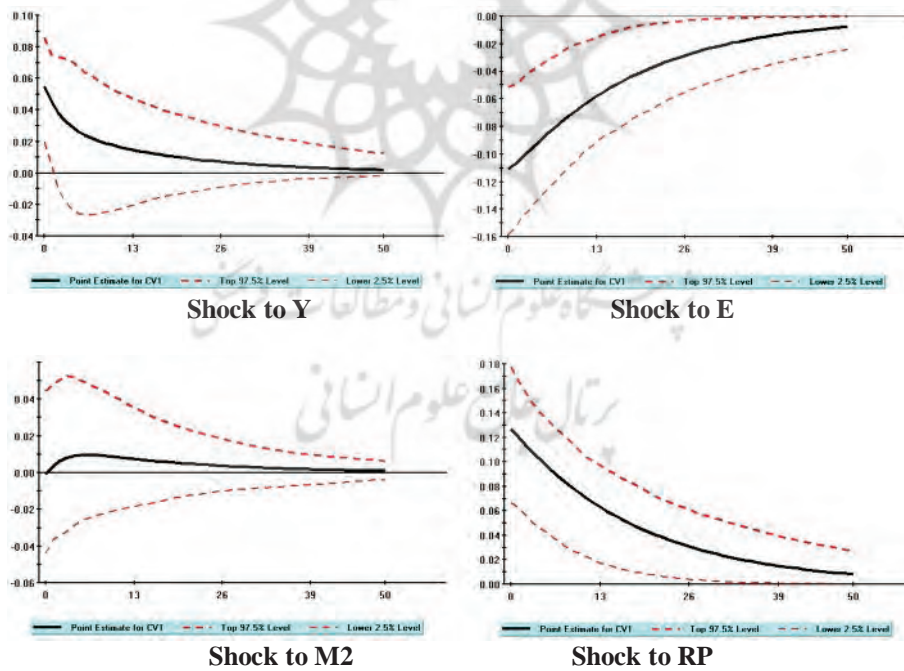


Figure 10. Persistence Profile of the effect of a system-wide shock to money demand

Figures (9) and (10) show that following a system-wide shock both vectors converge towards zero, thus confirming the cointegrating properties of the long-run relationships. As shown in Figure (9), convergence to equilibrium in the PPP-UIP relation is sluggish. The real money demand relation returns to equilibrium after five years following the initial shock (Figure 10), while real exchange rate will be misaligned for quite a while. The real exchange rate misalignment is a key variable in policy circles and its calculation is one of the most controversial issues in Open Economy Macroeconomics. Macdonald (2007) argues that the PPP-UIP relation can be used to calculate the equilibrium value of the real exchange rate and thus any possible misalignments. In general, misalignments are used as a tool to predict future exchange rate shifts among floaters and to evaluate the need to adjust the exchange rate among countries with less flexible regimes. It has been argued that sustained real exchange rate (RER) overvaluations constitute an early warning indicator of possible currency crashes and they also have led to a drastic adjustment of relative prices and to a decline in the aggregate growth rate of the economy.



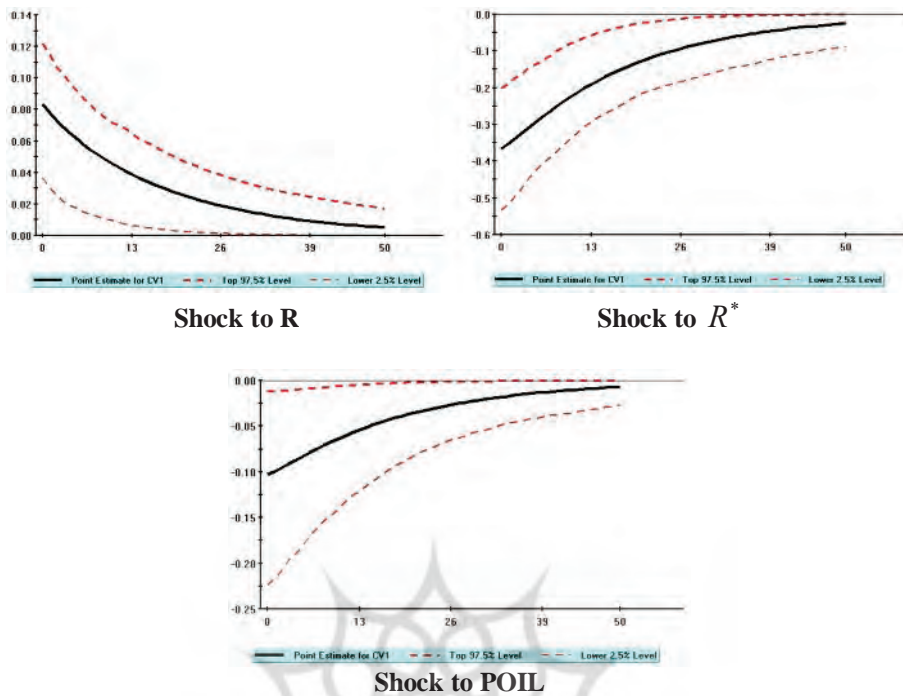
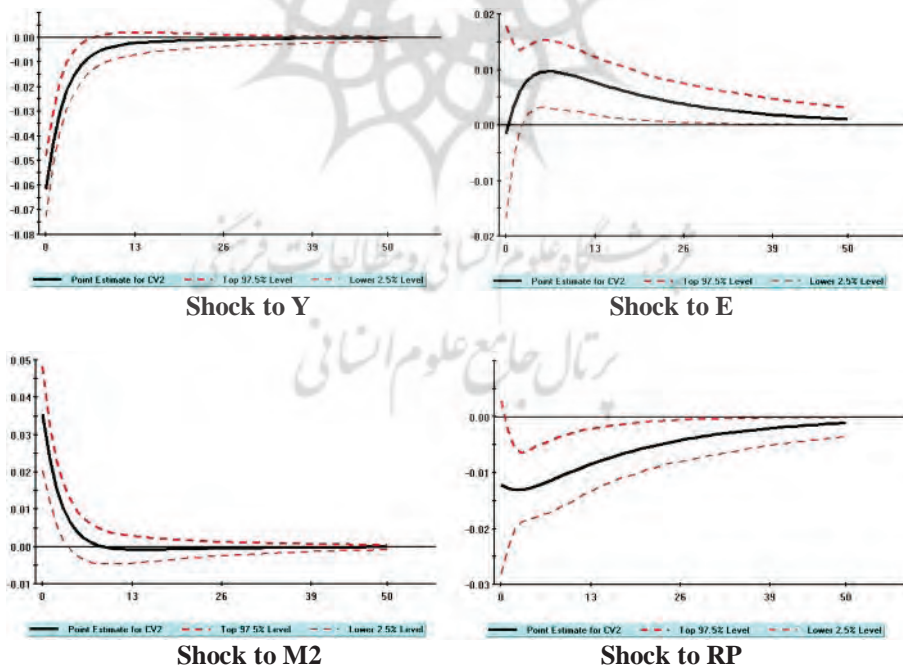


Figure 11. Persistence profile for PPP-UIP equation with 95 % bootstrapped confidence bounds



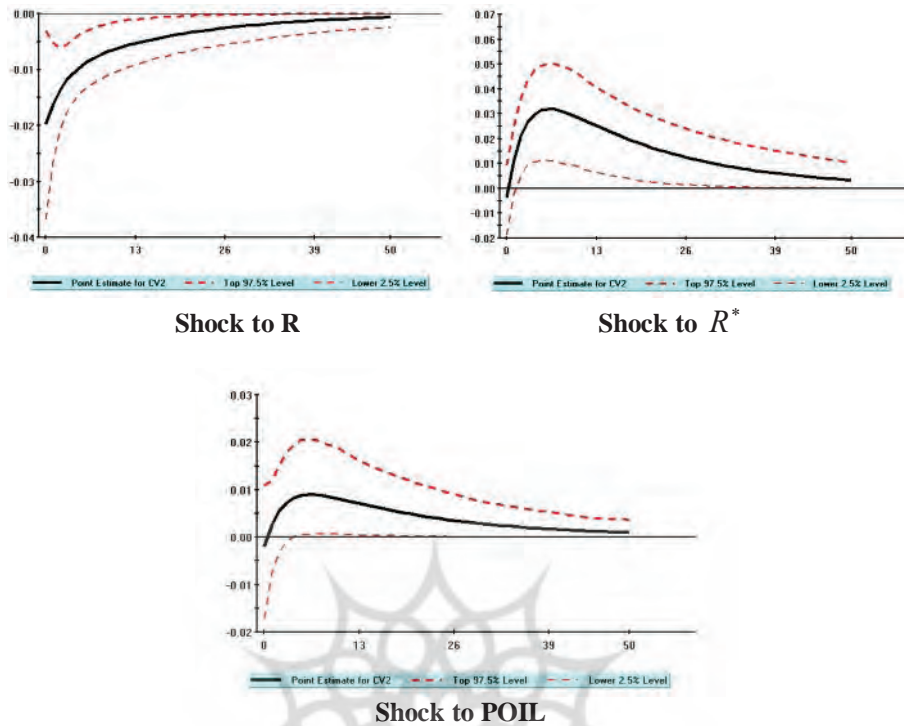


Figure 12. Persistence profile for MD relation with 95 % bootstrapped confidence bounds

We also show, in figures (11) and (12), the persistence profile of the two cointegrating relations following shocks to individual variables. It seems that shocks to Y and $M2$ die out quickly for the Money demand equation, while they have a relatively large and long-lasting impact on PPP-UIP equation. For the Iranian economy, the PPP-UIP relation, in comparison to money demand, shows slower rates of adjustments in the aftermath of shocks, reflecting misalignment in the real exchange rate in Iran. In addition, the oil price and foreign interest rate shocks on PPP-UIP equation have a negative and relatively long-lasting influence whereas their effects on money demand are positive and take two years to die out.

4.4.2. Impulse Response Analysis

Impulse response functions are of great importance in the analysis of dynamic systems. They characterize the response of the system over time after an impulse to one of the model variables. Generalized Impulse Response Function (GIRF) was proposed in Koop *et al.*

(1996), and developed further in Pesaran and Shin (1998) for vector error-correcting models¹. The GIRF is an alternative to the Orthogonalized Impulse Responses (OIR) of Sims (1980).

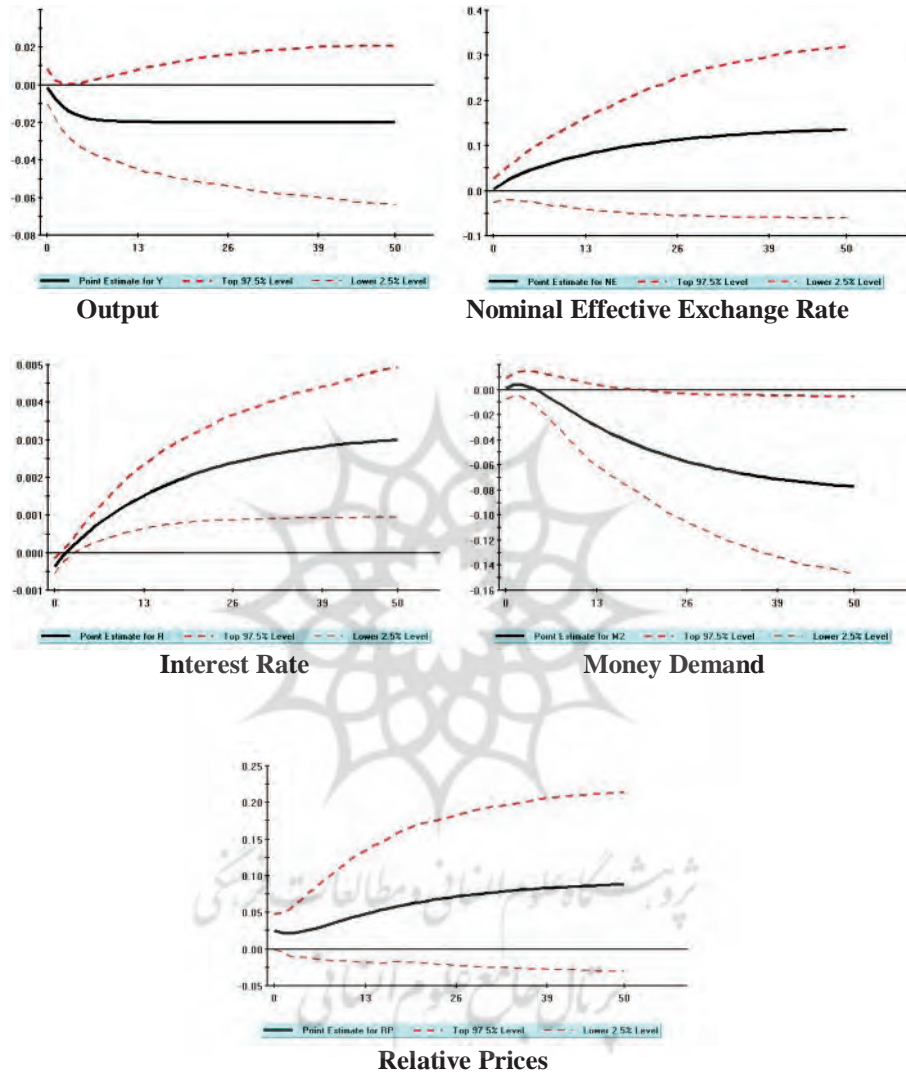


Figure 13. Generalized impulse responses of a positive unit shock to the foreign interest rate equation with 95% bootstrapped confidence bounds

1. For an account of the GIRF applied to VARX, see Garratt *et al.* (2006, Chapters 6 and 10)

The OIR approach requires the impulse responses to be computed with respect to a set of orthogonalized shocks, while the GIRF considers shocks to individual errors and integrates out the effects of the other impulses using the observed distribution of all shocks without any orthogonalization. Unlike the OIR approach, the GIRF is invariant to the ordering of the variables, which is clearly an important advantage.

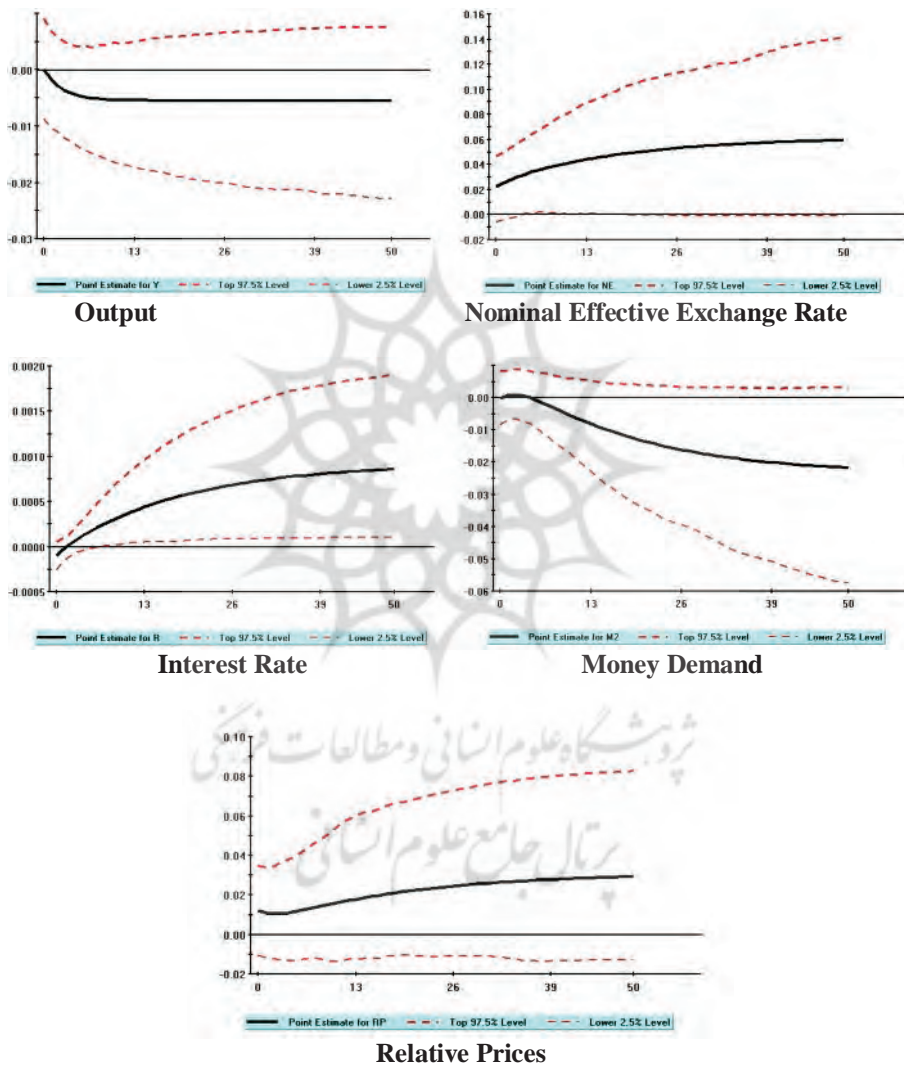


Figure 14. Generalized impulse responses of a positive unit shock to the oil price equation with 95% bootstrapped confidence bounds

Figures (13) and (14) show the impulse responses of the five endogenous variables to foreign interest rate and oil price shocks. These shocks have permanent effects on the level of the individual series, reflecting their unit root properties. A shock to foreign interest rate leads to an increase in domestic interest rates by approximately 0.3% and reduces output by 1% below its baseline value. However, the latter effect is not statistically significant based on the simulated confidence intervals around the point estimates. The effects of the shock to the foreign interest rate equation on real money balances are negative. The shock to the foreign interest rate equation depreciates the nominal exchange rate by 4% in the short term (the effects are not significant though).

Figure (14) shows that a positive oil price shock decreases real output by approximately 0.5% below its baseline level after 2.5 years; a pattern that is consistent with the resource curse paradox. According to this paradox, resource rich countries perform poorly when compared to countries, which are not endowed with oil, natural gas, and minerals. Therefore, resource abundance is believed to be an important determinant of economic failure, which implies that oil abundance is a curse and not a blessing. The nominal effective exchange rate increases in the aftermath of oil price shock; a pattern inconsistent with the “Dutch disease” view. Higher oil prices would imply a rise in export revenues in Iran. This leads to an increase in spending on all products, which in turn increases the relative prices (domestic relative to foreign prices). The oil price shock is accompanied by an increase in the domestic short-term interest rate, suggesting a possible tightening of the monetary policy in response to the rise in oil prices. The shock affects the real money balances both directly and indirectly through its impact on interest rates. The overall outcome is a fall in real money balances by around 2% in the long run. These two shocks illustrate the importance of including foreign variables in macroeconomic models for Iran in order to get a comprehensive overview of the Iranian macroeconomy.

5. Conclusion

We have provided a cointegrating VAR analysis of the Iranian economy in a global context. The specified VECX* model combines the implications of economic theory for identification of the long-run relationships and a data-driven approach to modeling the short-run

dynamics. We have managed to identify two long-run relationships amongst the variables considered. Those are: i) the money demand equation and ii) the PPP-UIP relationship, which is a systematic interaction between the real exchange rate, domestic and foreign interest rates. The latter cointegrating vector indicates that the nominal exchange rate is not only determined by the domestic and foreign goods market forces, but also it needs the short run adjustments coming from the interest rate semi-differentials to achieve its long-run balance. Furthermore, the likelihood ratio tests did not reject the over-identifying restrictions suggested by economic theory. We believe that the estimated model in this paper is both theory- and data-consistent. In addition, the well-behaved persistence profiles of the long run relations support the validity of these restrictions. However, the rate of adjustment of the real exchange rate to equilibrium following a system-wide shock is slow, reflecting the misaligned real exchange rate in Iran. We also have analyzed the short-term and long-term implications of external and internal shocks for the Iranian economy. We have provided the time profiles of the effects of external/internal shocks on the core macro variables of Iran using generalized impulse responses together with the corresponding bootstrapped standard errors.

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Appendix A. sources and construction of the data

This appendix provides the definitions and data sources of the variables used in constructing a VARX* model for Iran. The variables used in this paper are Y : Real GDP; CPI : Consumer price index; E : Exchange rate; $M2$: Broad real money; R^s : Short term interest rate, and P_t^0 : Oil price index.

A.1. Real GDP (y_t)

The source of all 29 countries is the IMF's International Financial Statistics (IFS) GDP series in 2000 constant prices. We seasonally adjusted the quarterly data using the U.S. Census Bureau's X-12 ARIMA seasonal adjustment program, for further details see U.S. Census Bureau 2007. Seasonally adjusted and rebased into to 2000=100 index. As quarterly data were not available [i.e. for Argentina (1979Q1-1989Q4), Iran (1979Q1-1994Q4), India (1979Q1-2006Q4), Malaysia (1979Q1-1987Q4), Pakistan (1979Q1-2007Q4), Philippines (1979Q1-1980Q4), Sri Lanka (1979Q1-2007Q4), Thailand (1979Q1-1992Q4) and Turkey (1979Q1-1986Q4)], quarterly series were interpolated (backwards) linearly from the annual series using the same method as that applied by Dees, di Mauro, Pesaran, and Smith (2007) to data for a number of the 33 countries in their data set and also included in ours for the computation of the foreign (star) variables. For a description of the Interpolation Procedure, please see Dees, di Mauro, Pesaran, and Smith (2007) Section 1.1 of Supplement A.

A.2. Consumer Price Indices (p_t)

The CPI data source for all countries except China is the IFS Consumer Price Index series 64 zf. In case of China we use 64 xzf. The natural logarithm of the foreign price index is computed as

$$p_t^* = \sum_{j=0}^N w_{jt} p_{jt}$$

where p_{jt} is the price index of foreign country j , where $j = 1, 2, \dots, N$, seasonally adjusted and rebased into a 2000=100 index. w_{jt} is the

trade share of country j in total trade of Iran with all countries in Table (A.2.1).

Table A.2.1. Countries in the VARX* Model for Iran

Argentina	Malaysia
Australia	Netherlands
Austria	Pakistan
Brazil	Philippines
Canada	Singapore
China	Spain
Denmark	South Africa
France	Sri Lanka
Germany	Sweden
Greece	Switzerland
India	Thailand
Italy	Turkey
Japan	United Kingdom
Korea	United States

A.3. Exchange rates (e_t)

The natural logarithm of Iran's nominal effective exchange rate is computed as $-\ln(E_t)$ where E_t is the IMF INS nominal effective exchange rate (code NECZF). E_t is defined such that a rise represents Rial's appreciation. We take minus the log of E_t and rename it as e_t to obtain a convenient measure of exchange rate defined as domestic price of foreign currency.

A.4. Money and Quasi Money Supply (m_{2t})

The data on Money and Quasi Money Supply is from the IMF IFS series 34 and 35 and is available from 1957Q1. As Money Supply data between 1984Q2 and 1986Q2 is missing in the IFS series, we obtained the complete series by splicing the IFS and CBI data on Money Supply. For Quasi Money data was missing between 1985Q2 to 1986Q2. We filled in for the missing data by splicing the IFS and the CBI data.

We seasonally adjust the quarterly data on Money and Quasi Money Supply using the U.S. Census Bureau's X-12 ARIMA seasonal adjustment program.

A.5. Interest Rate (r_t^s)

The domestic quarterly nominal interest rate is computed as:

$$0.25 * Ln(1 + R_t/100)$$

where R_t is the average of expected rates of return (on exports, trade, services and miscellaneous, constructions and housing, manufacturing and mining, and agriculture), short term investment deposit rate, and one-year special deposit rate and is denoted in percent per annum. The data is obtained from the Central Bank of the Islamic Republic of Iran (CBI) online database: Economic Time Series Database (<http://tsd.cbi.ir/>).

The natural logarithm of the foreign short-term interest rate is computed as

$$r_t^* = \sum_{j=0}^N w_{jt} r_{jt}$$

where r_{jt} is short term interest rate of foreign country j , where $j = 1, 2, \dots, N$, and w_{jt} is the trade share of country j in total trade of Iran.

The Iranian year generally starts on the 21st of March, as such the Iranian quarter 1 contains 10 days of the Gregorian quarter 1 and 80 days of Gregorian quarter 2. To convert the data from Iranian to Gregorian calendar we simply adopt the following rule:

$G(Q) = \frac{8}{9} Iran(Q-1) + \frac{1}{9} Iran(Q)$, where $G(Q)$ is the Gregorian quarter Q and $Iran(Q)$ is the Iranian quarter Q .

A.6. Oil Prices (p_t^0)

The natural logarithm of the oil prices is computed as

$$p_t^0 = Ln(POIL_t)$$

where $POIL_t$ is the nominal crude oil prices in US Dollar per barrel. The data source is IFS, code 00176 AAZZF which is rebased into a 2000=100 index.

Appendix B. Unit Root Test Results

Table B.1. Augmented Dickey-Fuller and Philips - Perron unit root tests (1979Q2-2007Q4)

Variable	ADF(0)	ADF(1)	ADF(2)	ADF(3)	ADF(4)
Levels					
y_t	-2.56	-3.16 ^a	-2.81	-2.38	-2.10
p_t	-0.73	-1.25	-0.74	-1.63	-2.85 ^a
e_t	-1.05 ^a	-1.11	-1.13	-1.01	-1.17
m_{2t}	-0.95	-0.33	-0.69	-0.03 ^a	-1.20
r_t	-0.32 ^a	-0.10	-0.24	-0.25	-0.10
r_t^*	-2.91	-2.02 ^a	-1.29	-1.44	-1.21
p_t^*	-2.57 ^a	-2.58	-2.60	-2.62	-2.60
$p_t - p_t^*$	-2.94 ^a	-3.02	-2.95	-3.08	-3.23
$Poil$	-1.06	-1.36	-0.84 ^a	-1.08	-0.79
Variable	PP(0)	PP(5)	PP(10)	PP(15)	PP(20)
Levels					
y_t	-3.02	-3.18	-3.32	-3.23	-3.09
p_t	-1.50	-1.38	-1.50	-1.51	-1.51
e_t	-1.62	-1.28	-1.39	-1.52	-1.60
m_{2t}	-0.33	-0.52	-0.45	-0.32	-0.31
r_t	-0.29	-0.14	-0.06	-0.15	-0.27
r_t^*	-3.60	-2.59	-2.70	-2.63	-2.52
p_t^*	-2.40	-2.29	-2.28	-2.31	-2.38
$p_t - p_t^*$	-2.47	-2.62	-2.54	-2.46	-2.46
$Poil$	-0.12	-0.66	-0.31	-0.16	-0.07
Variable	ADF(0)	ADF(1)	ADF(2)	ADF(3)	ADF(4)
First differences					
Δy_t	-8.74	-7.72	-7.47	-7.16 ^a	-5.75
Δp_t	-7.98	-8.28	-4.19	-2.54 ^a	-3.26
Δe_t	-10.10 ^a	-7.19	-6.30	-4.94	-4.76
$\Delta^2 p_t$	-12.85	-16.85	-14.73 ^a	-6.92	-5.35
Δm_{2t}	-13.23	-6.54	-6.99	-3.55	-4.50 ^a

Δr_t	-9.21 ^a	-7.00	-5.60	-4.10	-3.89
Δr_t^*	-14.89	-12.52 ^a	-7.33	-6.92	-5.72
Δp_t^*	-3.09 ^a	-2.20 ^a	-1.95	-2.08	-1.78
$\Delta^2 p_t^*$	-14.95 ^a	-9.55	-6.64	-6.46	-6.28
$\Delta(p_t - p_t^*)$	-9.95 ^a	-7.52	-5.61	-4.61	-4.27
$\Delta Poil$	-8.80	-8.39	-5.50	-5.35	-5.87 ^a
Variable	PP(0)	PP(5)	PP(10)	PP(15)	PP(20)
First differences					
Δy_t	-9.16	-8.45	-8.46	-8.65	-8.97
Δp_t	-5.04	-5.20	-5.30	-5.19	-5.11
Δe_t	-10.34	-10.25	-10.26	-10.30	-10.33
$\Delta^2 p_t$	-26.02	-15.81	-20.43	-25.12	-25.32
Δm_{2t}	-12.94	-12.97	-12.98	-12.97	-12.95
Δr_t	-10.83	-10.68	-10.66	-10.71	-10.80
Δr_t^*	-19.29	-17.62	-19.03	-19.45	-19.42
Δp_t^*	-3.63	-2.82	-3.07	-3.29	-3.54
$\Delta^2 p_t^*$	-20.01	-16.35	-17.78	-19.60	-20.06
$\Delta(p_t - p_t^*)$	-10.11	-10.13	-10.11	-10.11	-10.11
$\Delta Poil$	-9.24	-9.22	-9.28	-9.28	-9.27

Note: ADF and $PP(\ell)$ represent augmented Dickey-Fuller (1979, ADF) test statistics, and Phillips and Perron (1988) unit root statistic based on the Bartlett window of size ℓ respectively. The lower 5% critical values for the tests are -2.88 for the first differences and -3.45 for the levels. Regarding the levels, ADF and PP regressions with an intercept and a linear time trend are estimated, for the first differences, ADF and PP regressions with an intercept and no time trend are estimated. The symbol 'a' denotes the order of augmentation in the Dickey-Fuller regressions chosen using the Akaike information Criterion, with a maximum lag order of four.

Appendix C. Generalized impulse responses of the endogenous variables

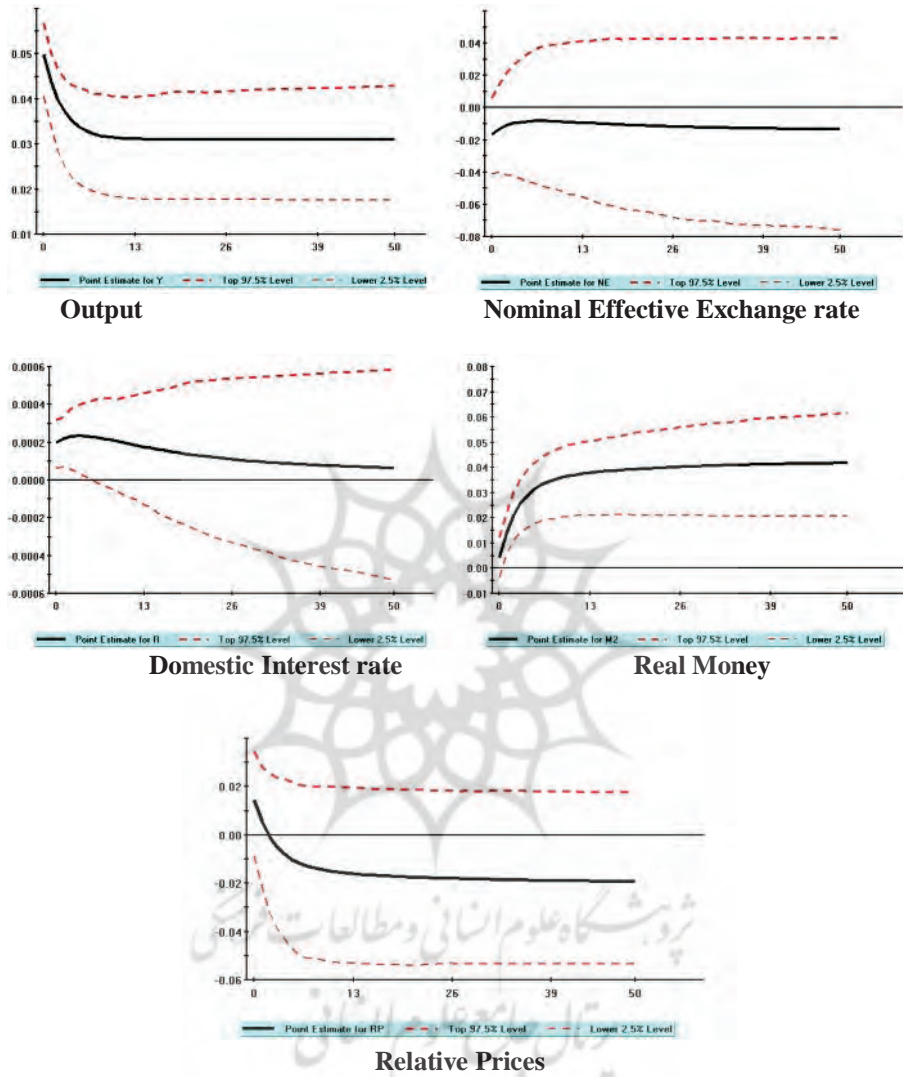


Figure C.1. Generalized impulse responses of a positive unit shock to the output equation with 95% bootstrapped confidence bounds

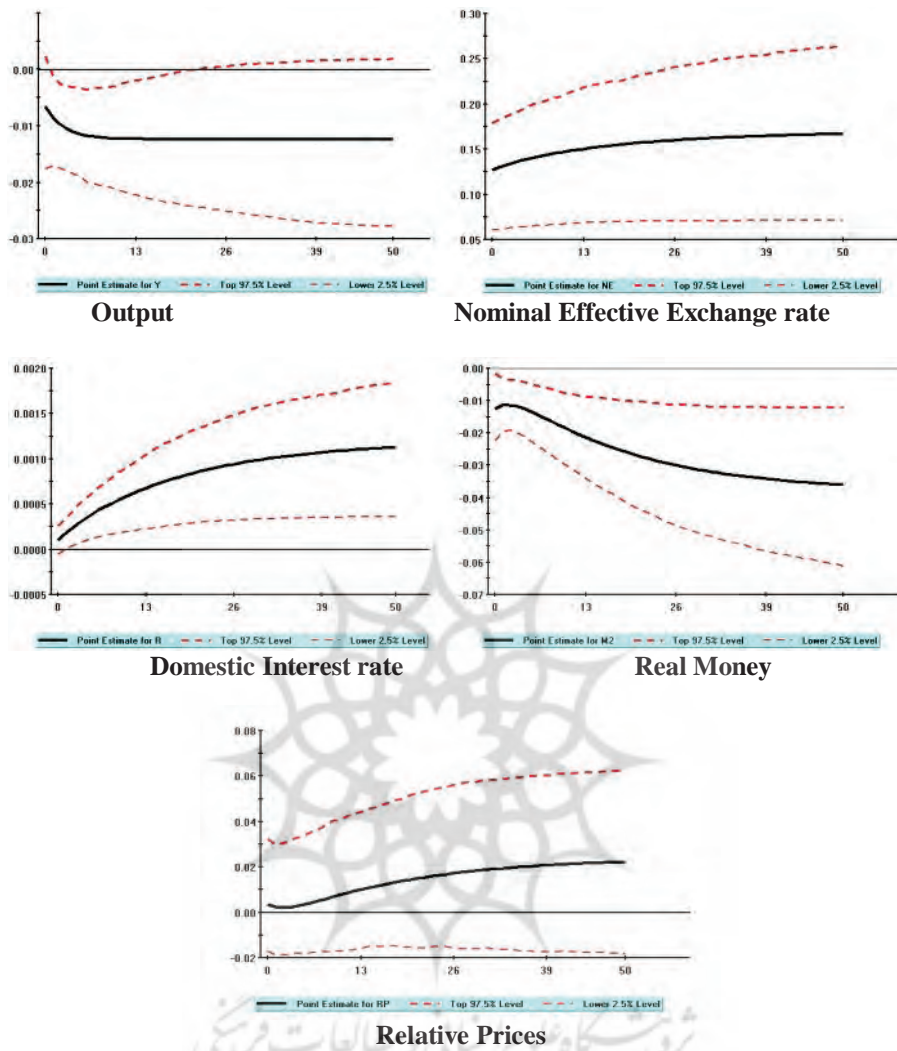
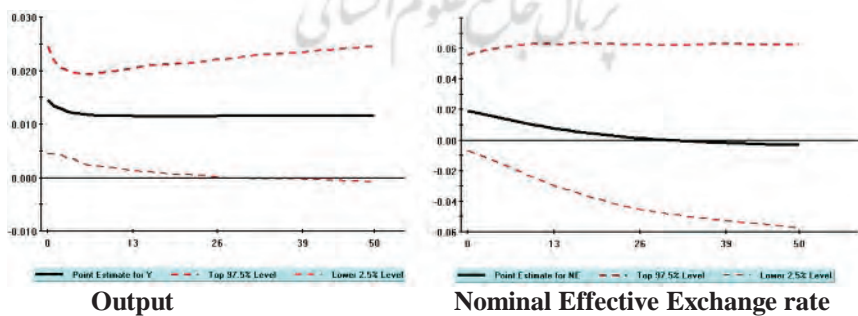


Figure C.2. Generalized impulse responses of a positive unit shock to the exchange rate with 95% bootstrapped confidence bounds



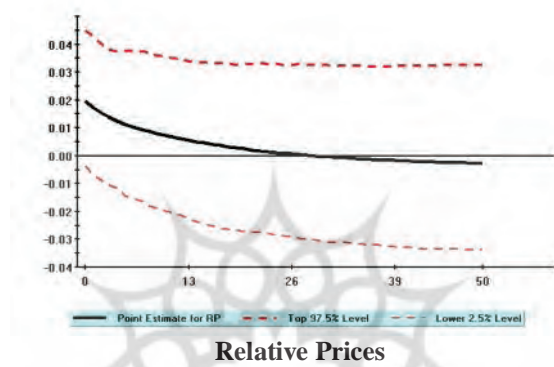
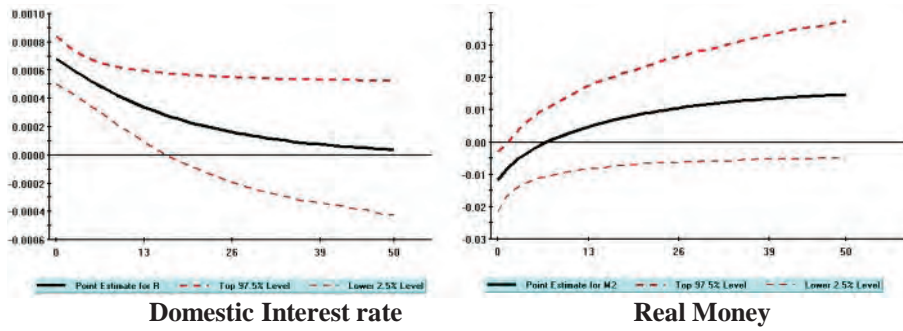
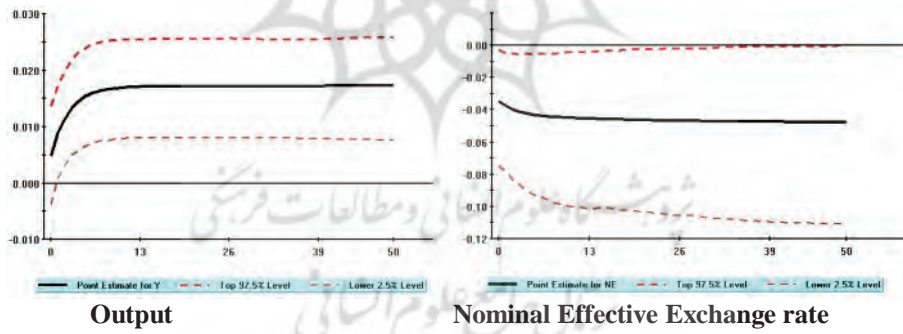


Figure C.3. Generalized impulse responses of a positive unit shock to the domestic interest rate with 95% bootstrapped confidence bounds



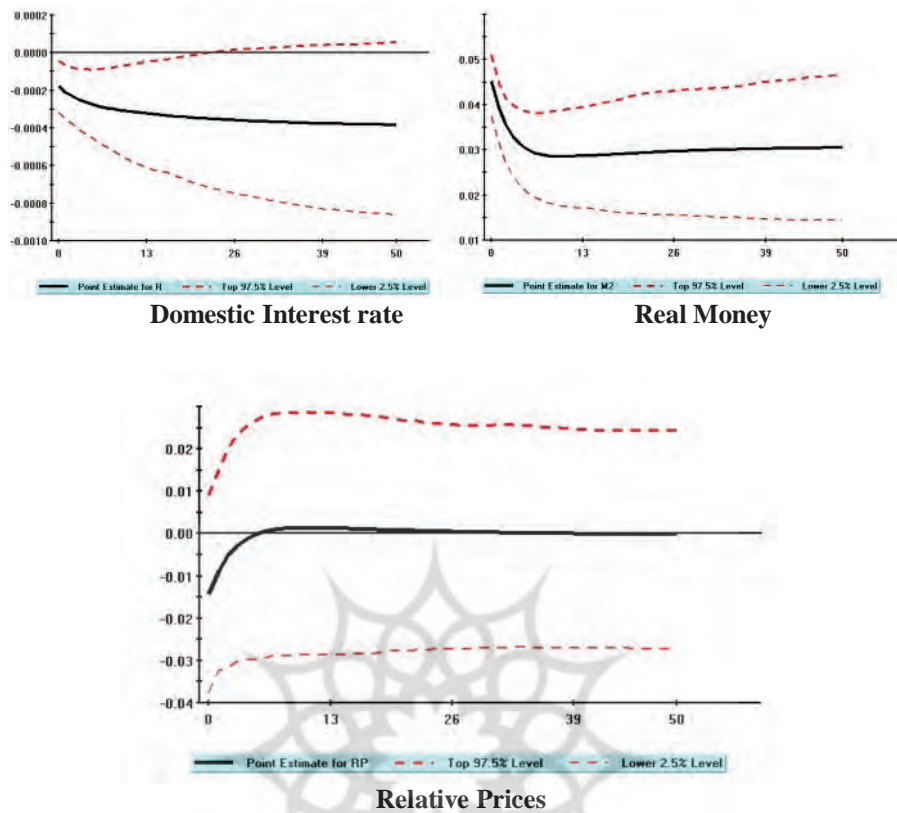
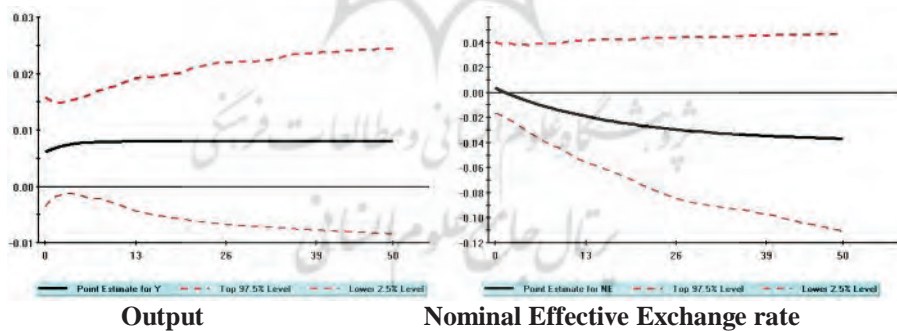


Figure C.4. Generalized impulse responses of a positive unit shock to the real money with 95% bootstrapped confidence bounds



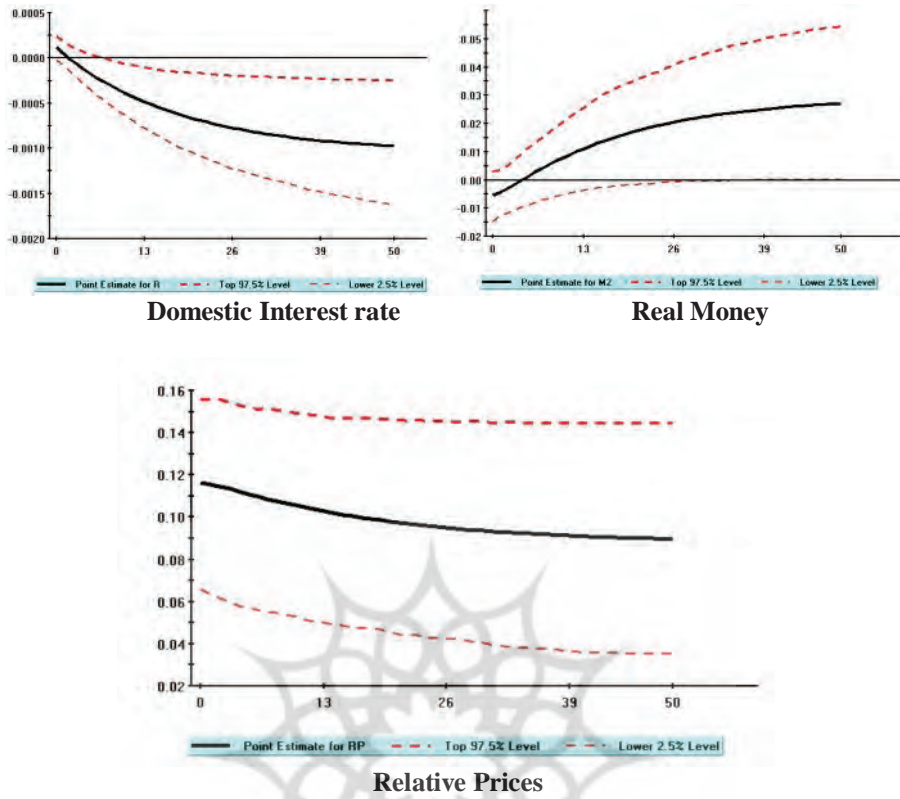


Figure C.5. Generalized impulse responses of a positive unit shock to the relative price level with 95% bootstrapped confidence bounds

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