Oil Exports and the Iranian Economy*

Hadi Salehi Esfahani[a], Kamiar Mohaddes[b†], and M. Hashem Pesaran[bc]

[a] Department of Economics, University of Illinois

[b] Faculty of Economics, University of Cambridge

[c] Department of Economics, University of Southern California

April 18, 2012

Abstract

This paper presents an error-correcting macroeconometric model for the Iranian economy estimated using a new quarterly data set over the period 1979Q1-2006Q4. It builds on a recent paper by the authors, Esfahani et al. (2012), which develops a theoretical long-run growth model for major oil exporting economies. The core variables included in this paper are real output, real money balances, inflation, exchange rate, oil exports, and foreign real output, although the role of investment and consumption are also analyzed in a sub-model. The paper finds clear evidence for the existence of two long-run relations: an output equation as predicted by the theory and a standard real money demand equation with inflation acting as a proxy for the (missing) market interest rate. The results show that real output in the long run is influenced by oil exports and foreign output. However, it is also found that inflation has a significant negative long-run effect on real GDP, which is suggestive of economic inefficiencies and is matched by a negative association between inflation and the investment-output ratio. Finally, the results of impulse responses show that the Iranian economy adjusts quite quickly to the shocks in foreign output and oil exports, which could be partly due to the relatively underdeveloped nature of Iran's financial markets.

JEL Classifications: C32, C53, E17, F43, F47, Q32.

Keywords: Growth models, long-run relations, oil exporters, Iranian economy, oil price and foreign output shocks, and error-correcting relations.

*We are grateful to conference participants at the University of Illinois at Urbana-Champaign, the University of Southern California, and the Dubai School of Government for constructive comments and suggestions. We would also like to thank Gary S. Becker, Massoud Karshenas, and Mehdi Raissi for most helpful suggestions and comments.

†Corresponding author. Email address: km418@cam.ac.uk.

1 Introduction

The first major oil field in Iran was discovered in 1908 with oil production flowing in sizeable amounts from 1912. Even after 100 years of exploration and production, Iran's current estimated reserve-to-extraction ratio suggests a further 87 years of oil production even in the absence of new oil field discoveries or major advances in oil exploration and extraction technologies. In addition, Iran has the second largest natural gas reserves after Russia, around 60% of which is yet to be developed.[1] Although, it is clear that Iran's oil and gas reserves will be exhausted eventually, this is likely to take place over a relatively long period. In fact over the past two decades the ratio of Iran's oil export revenues to GDP has fluctuated around 26%.[2] Therefore, there is little evidence to suggest that oil income will be diminishing any time soon for Iran. As such, rather than follow the approach in the 'Dutch disease' and 'resource curse' literature, which considers the revenues from the resource to be intrinsically temporary and focusses on the relatively short term implications of the resource discovery,[3] it makes more sense to view the income from such resources as permanent for the purpose of macroeconomic analysis over the medium term. Specifically, Esfahani et al. (2012) show that if the oil income to output ratio is expected to remain high and stable over a prolonged period, oil income will enter the long-run output equation with a coefficient which is equal to the share of capital if it is further assumed that the underlying production technology can be represented as a Cobb-Douglas production function. Esfahani et al. (2012) also provide empirical evidence in favour of such a long-run specification in the case of a number of major oil exporters using observations on real domestic and foreign outputs and revenues from oil exports.

In this paper we build on the theoretical results in Esfahani et al. (2012) and develop a small vector error-correcting model (VECX*) for Iran where we provide further evidence on the empirical validity of the long-run output equation. The core variables included in the model are real output, real money balances, inflation, exchange rate, oil exports, and foreign real output, although the role of investment and consumption are also analyzed in a sub-model. Interest rate is not included in our model because the domestic credit markets in Iran operate under tight controls and the interest rate is not market-determined. But assuming that the Fisher equation holds in the long run, the inflation rate can be used as a proxy for the interest rate. The foreign output variable is constructed as a weighted average of the log real output of Iran's trading partners with the weights based on the relative size of their trade with Iran (exports plus imports). For exchange rates we consider weighted averages of official and 'free' market exchange rates to capture the variety of the exchange rate regime that have been in place in Iran over the past three decades. The possible effects of Revolution and the eight-year war with Iraq are also analysed.

A number of models of Iran's macroeconomy have been developed in the past. The distinctive features of our model are: (1) a theory derived long-run model for oil exporting

[1] See, for example, Amuzegar (2008) and the British Petroleum Statistical Review of World Energy.

[2] Esfahani et al. (2012) show that most other members of the Organization of the Petroleum Exporting Countries (OPEC) such as Kuwait, Libya, Nigeria, Saudi Arabia, and Venezuela have also similar oil income GDP ratios that have remained relatively stable over time.

[3] See Corden and Neary (1982), Krugman (1987), Neary and van Wijnbergen (1986), Sachs and Warner (1995), and van der Ploeg and Venables (2009) for a recent survey.

countries in which the long-run role of oil export revenues for growth is explicitly modeled; (2) a careful and parsimonious modeling of the ways in which major external variables enter into the macroeconomic equations in Iran, taking into account the variety of channels through which the variables influence each other, including the implicit response of the government to macroeconomic developments; (3) parameterization of the model to allow for the measurement and testing of the macro-level impact of oil exports and global technological progress on the Iranian economy; (4) joint modeling and estimation of output, inflation, money supply, and the exchange rate, in contrast to models that focus on output or inflation alone, while treating the other variables as exogenous; and (5) use of quarterly data.

The maximum likelihood estimates of the VECX* model support the existence of two long-run relations, namely the real output and the real money demand equations, as predicted by the theory. Furthermore, it is not possible to reject the hypothesis that real output, real money balances, real oil income, and foreign output are co-trending. The evidence also supports the existence of a long-run relation between domestic output, foreign output, and real oil exports, although we also find that inflation has a statistically significant negative effect on real output. Once the effects of oil exports and inflation are taken into account, the estimates support output growth convergence between Iran and the rest of the world. These results seem to be reasonably robust regardless of how foreign output is constructed, what measure of the exchange rate is used, and whether a dummy variable for revolution and war (over the period 1979Q1-1988Q2) is included in the model.

From the estimates, several conclusions can be drawn. One key result is the economy's fast adjustment to shocks, when compared to the response rates of other economies, especially the developed ones. This seems to be due to the limitations of Iran's financial markets that restrict expenditure smoothing options and thereby cause the economy to move up and down quickly as external and internal conditions change. Second, we find that although Iran may lag behind its main trading partners in terms of its level of technology, it has experienced a similar rate of technological progress over the past three decades. Third, in the long run, oil exports contribute to real income through real capital accumulation. As a result, the elasticity of the aggregate real income with respect to real oil revenues (measured in term of domestic output units) is equal to the marginal product of capital. We confirm this result by showing that the nominal dollar value of oil revenues has the same impact on the real GDP as would be caused by a decline in the dollar value of one unit of domestic output. Fourth, our estimates suggest that in Iran, the output elasticity of capital (or the share of capital) is about 0.26, which is in line with the estimates obtained in recent studies for oil exporting economics, see for instance Cavalcanti et al. (2011a). Fifth, there is a statistically significant negative association between inflation and real GDP even in the long run, which is matched by a significant negative association between inflation and the investment-output ratio. The negative inflation effects found on real output and investment are in line with the theoretical literature and indicate certain inefficiencies in the Iranian economy where high and variable inflation seem to have led to lower investment and output. Sixth, in the long run, the elasticity of real money balances with respect to real output is around unity, and inflation (used as a proxy for interest rate) has a negative effect on real money balances.

The rest of the paper is set out as follows. Section 2 provides a review of the macroeconometric modelling literature for the Iranian economy to better place our contribution within the existing literature. Section 3 outlines the long-run relations considered for Iran, discusses the main macroeconomic trends in post-revolutionary Iran, and describes the VECX* econometric model that embodies the long-run relations. Section 4 presents the long-run estimates and the various tests of the long-run theory. Section 5 discusses the short-run dynamics and provides evidence on speed of convergence to equilibrium, impulse responses, and error correction estimates. Finally, Section 6 offers some concluding remarks.

2 Macroeconometric Models of the Iranian Economy

The origins of macroeconometric modeling in Iran date back to the early 1970's, when Habib-Agahi (1971)[4] pioneered the practice at the Iranian Plan and Budget Organization (PBO). Habib-Agahi's model contained 8 linear behavioral relations and 7 accounting identities, linking 3 categories of imports, aggregate output, real disposable income, private and government consumption and investment expenditures to the size of the development budget, oil and non-oil exports, and foreign loans for development. The model was estimated using annual time series data over the 1959-1970 period, and formed the basis of the first "official" macroeconometric model to be developed by the PBO. This was a modest exercise in macroeconometric modeling, largely reflecting the data and computational limitations prevailing in Iran at the time.

Considering other models subsequently developed at the PBO, a distinction needs to be made between the models that were constructed before the 1979 revolution and those constructed afterwards. The first model developed at the PBO before the revolution was a modification of Habib-Agahi's model and related non-oil exports to the value added in agriculture instead of treating it as exogenous. However, the value added in agriculture was now assumed to be exogenous. The second major macroeconometric model constructed at the PBO before the revolution was much more detailed, and represented important advances over the earlier one. It allowed for the effect of relative prices on imports and non-oil exports demands, contained equations for the determination of a number of key monetary aggregates and tax revenues, and used a Phillips type wage equation to close the model. The model was estimated over the period 1961-1975 and was the first serious empirical attempt at modeling the interactions of the monetary and real variables in the Iranian economy. However, as with the other models developed for the Iranian economy there is no systematic documentation of the model's short-run predictive performance or its long-run properties.

Perhaps not surprisingly, revolution and the subsequent eight-year war with Iraq halted any serious development of macroeconometric models both inside and outside of the PBO. But with the ending of the Iran-Iraq war in 1988, and the government's attempt at regeneration and reconstruction of the economy, once again the problem of economic planning and the development of appropriate macroeconometric models gained priority. But although the importance of macroeconometric models was recognized in the formulation of the First Five-Year Development Plan (1990-94), given the urgency surrounding the formulation of the First Plan and the limited time available to accomplish the task, serious attempts at macroeconometric modeling had to wait until after its approval and implementation.

[4] The dates in the 1300s are based on the Iranian calendar. The corresponding Gregorian dates are roughly equal to the Iranian date plus 621. Publications in Persian are cited with the Gregorian dates in the text but with both the Iranian and the Georgian calendar dates in the references.

According to official accounts, two different macroeconometric models were utilized in the formulation of the Second Five-Year Development Plan.[5] The first model, PBO1, contains 25 econometrically estimated behavioral, technical, and institutional relations and a number of accounting identities. With a few exceptions these relations are estimated by the least squares methods over the period 1974-1993, and are documented in PBO (1993).[6] This model is composed of a production and factor demand module, a Keynesian income-expenditure flow module, with investment expenditures disaggregated by 10 major production sectors, and an aggregate price equation. The second model developed seems to represent an extension of the first.

In addition to these models used for Iran's development plans, a number of other macroeconometric models have been developed for Iran since the contribution of Habib-Agahi in the beginning of the 1970's, both at the PBO and by other researchers around the world. Prominent examples of the latter category are the models developed by Baharie (1973), Vakil (1973), Shahshahani-Madani (1978), Heiat (1986) and Safai (1986). These models differ in the extent of detail and the level of disaggregation, but are very similar in their underlying structures. They are largely demand-determined Keynesian models, and with a few minor exceptions neglect the effect of relative prices and stock-flow relations on the economy's evolution. Also, very little is known about their short-run forecasting performance, or their long-term properties. In contrast to these papers, Noferesti and Arabmazar (1993) develop a model in which aggregate supply is not assumed to be perfectly elastic, while Valadkhani (1997), building on the work of the above-mentioned papers, develops a more comprehensive macroeconometric model for Iran.

There are also a number of simple planning/optimal control models developed notably by Motaman (1979), Razavi (1982), and Razavi (1983) for the analysis of the optimal rate of oil production in oil-based economies. These models are primarily concerned with the intertemporal optimization problem involved in oil production decisions (namely, whether to produce now or later), and are typically very simple as far as their main structural relations are concerned. They are not intended as forecasting or budgetary tools and, in view of the current constraints on Iran's capacity to produce oil, have limited relevance to Iran's economic policy problems.

An alternative strategy to the models developed in the above mentioned papers would be to estimate vector autoregressive (VAR) models in some of the main macroeconomic variables, such as output, price level, money supply, oil exports, consumption, and investment, along the lines originally developed by Sims (1980). In more recent papers, Mehrara and Oskoui (2007) make use of a structural VAR to determine whether oil price shocks are the main source of output fluctuation for Iran, while Elyasiani and Zhao (2008) make use of vector autoregression, generalized impulse response function and generalized variance decomposition techniques to determine the interdependencies of Iran with its major trading partners and the US. But the use of VAR models, without imposition of structural relations on their long-run solutions, will be limited to short-term forecasting and are unlikely to be relevant for medium term policy analysis. Thus, a long-run structural approach to VAR

[5] See PBO (1993). Initially the implementation of the Second Five Year Development Plan was intended to commence in March 1994, but due to delay in its approval by the Iranian Parliament was postponed by one year.

[6] The relations in the published version of model PBO1 were mainly estimated over the period 1974-1992.

modeling and its application to the Iranian economy, which we take up in this paper, are worth pursuing.

Other papers of interest dealing with economic growth in Iran are those of Valadkhani (2006), which looks at the determinants of the growing unemployment rate in Iran, Pahlavani et al. (2005), which tries to identify the short and long-run determinants of growth, taking into account the endogenously identified structural breaks in Iran, and Becker (1999), which looks at the development of several variables from pre- to post-revolution and the effect of monetary shocks on these variables. In addition, Bahmani-Oskooee (1995) and Kia (2006) explore the determinants of inflation in Iran, taking into account the role of external factors.

While all the papers discussed so far have used annual data, there are a few IMF working papers on Iran using quarterly observations. In particular, Bonato (2008) looks at the determinants of inflation in Iran, Celasun and Goswami (2002) develop an econometric model of short-run inflation and long-run money demand dynamics in Iran, and Liu and Adedeji (2000) construct a model to develop the determinants of inflation in Iran. However, all of the papers using quarterly data focus on a certain aspect of the Iranian economy, for instance the money demand relation or the determination of inflation, and as such do not consider the interconnection of the domestic variables with that of foreign variables. Neither do they explore the short-run and the long-run channels of growth. Part of our contribution is then to make use of quarterly data, while exploring the interconnection between the Iranian economy and the rest of the world and paying attention to both the short and long-run channels through which oil export revenues affect growth.

3 The Econometric Model and Methodology

3.1 A Long-Run Macroeconometric Model for Iran

Esfahani et al. (2012) develop a long-run growth model for a major oil exporting economy and derive conditions under which oil revenues are likely to have a lasting impact. They show that the possibility of a long-run impact from oil income to per capita output depends on the relative growth of oil income (g^o) relative to the combined growth of labour (n) and technology (g). In the case where $g^o < g + n$, the importance of oil income in the economy will tend towards zero in the limit and the standard growth model will become applicable. This is as to be expected since with oil income rising but at a slower pace than the growth of real output, the share of oil income in aggregate output eventually tends towards zero.

However, if $g^o \geq g + n$, oil income continues to exert an independent impact on the process of capital accumulation even in the long run. Under certain regularity conditions and assuming a Cobb-Douglas production function, it is shown that (log) oil exports enter the long-run output equation with a coefficient equal to the share of capital, α, or more specifically:

$$\ln(Y_t) - \theta \ln\left(Y_t^*\right) - (n - \theta n^*)t \sim I(0) \text{ , if } g^o < g + n, \tag{1}$$

and

$$\ln(Y_t) - \psi_1 \ln\left(Y_t^*\right) - \psi_2 \ln(E_t/P_t) - \psi_3 \ln(P_t^o X_t^o) - \gamma t \sim I(0) \text{ , if } g^o \geq g + n, \tag{2}$$

where Y_t (Y_t^*) is the real domestic (foreign) output, E_t is the nominal exchange rate in terms of US dollar, P_t is the consumer price index, P_t^o is the price of oil per barrel in US dollar, and X_t^o is the total number of barrels of oil exports. n and n^* are labour force growth rates of domestic and world economy, θ measures the extent to which foreign technology is diffused and adapted successfully by the domestic economy in the long run, and

$$\psi_1 = \theta(1 - \psi_2), \psi_2 = \psi_3 = \alpha, \text{ and } \gamma = (1 - \alpha)(n - \theta n^*). \tag{3}$$

Equation (2) is sufficiently general and covers both cases where $g^o < g+n$ and $g^o \geq g+n$. Under the former $\psi_1 = \theta$, $\psi_2 = \psi_3 = 0$, whilst under the latter $\psi_2 = \psi_3 \neq 0$. The above formulation also allows us to test other hypothesis of interest concerning θ and γ. The value of θ provides information on the long-run diffusion of technology to Iran. The diffusion of technology is at par with the rest of the world if $\theta = 1$, whilst a value of θ below unity suggests inefficiencies that prevents the adoption of best practice techniques, possibly due to rent-seeking activities. When $\theta = 1$ steady state per capita output growth in Iran can only exceed that of the rest of the world if oil income per capita is rising faster than the steady state per capita output in the rest of the world. The steady state output growth in Iran could be lower than the rest of the world per capita output growth if $\theta < 1$.

In what follows we estimate θ and the other parameters of the long-run output equation, (2), by embedding it within a vector error-correcting model of the Iranian economy estimated on quarterly observations over the past 28 years since the 1979 Revolution. To this end we first re-write the output equation as

$$y_t - \psi_1 y_t^* = \psi_2(e_t - p_t) + \psi_3 xo_t + c_y + \gamma_y t + \xi_{y,t} \tag{4}$$

where $y_t = \ln(Y_t)$, $y_t^* = \ln(Y_t^*)$, $e_t = \ln(E_t)$, $p_t = \ln(P_t)$, $ep_t = e_t - p_t$, $xo_t = \ln(X_t^o P_t^o)$, c_y is a fixed constant, and $\xi_{y,t}$ is a mean zero stationary process, which represents the error correction term of the long-run output equation. In addition to the output equation we also consider the real money demand equation (MD),

$$m_t - p_t = \phi_1 y_t + \phi_2(p_t - p_{t-1}) + c_{mp} + \gamma_{mp} t + \xi_{mp,t}, \tag{5}$$

where $mp_t = m_t - p_t$ is real money balances, c_{mp} is a fixed constant, and $\xi_{mp,t}$ is the stationary error-correcting term for the MD equation.

A number of other long-run relations considered in the literature, namely the purchasing power parity (PPP), the uncovered interest parity and the Fisher equation could also be included, see Garratt et al. (2006) for further details. But we have not been able to include these in our analysis as available data on interest rates are administratively determined, changed only at infrequent intervals, and as such do not reflect the market conditions. The money supply also comes to play an important role in the Iranian economy, since the capital markets are not developed in Iran. For the same reason we have used the inflation rate, $\pi_t = p_t - p_{t-1}$ rather than the interest rate in the MD equation specified above. The inflation rate could be a good proxy for the short-term interest rate assuming that the Fisher equation holds, at least in the long run. The analysis of PPP in Iran is also complicated by a prolonged period of black market in foreign exchange and the existence of multiple exchange rates, see Pesaran (1992). Also, to include a PPP relationship in the model, we need to introduce an effective exchange rate in addition to the US dollar rate, E_t. However, as a result of US sanctions only a very small fraction of Iran's trade is conducted with the US, and the use of US price level as a proxy for foreign prices will not be appropriate. Further work is clearly needed before a PPP relation can be added to the model in a satisfactory manner.

Our modelling strategy closely follows Garratt et al. (2003, 2006) and estimate a cointegrating VARX* model (which we also refer to as VECX*) with $\mathbf{x}_t = (y_t, mp_t, \pi_t, ep_t)'$ as the endogenous variables, and $\mathbf{x}_t^* = (y_t^*, xo_t)'$ as the exogenous variables. It is also possible to extend the model to include other macro variables such as consumption and investment. This is considered in Section 4.2.5. But before giving the details of the econometric model, we first discuss the data and the main economic trends of the Iranian economy over the period 1979Q1-2006Q4.

3.2 Macroeconomic Trends in Iran Since the 1979 Revolution

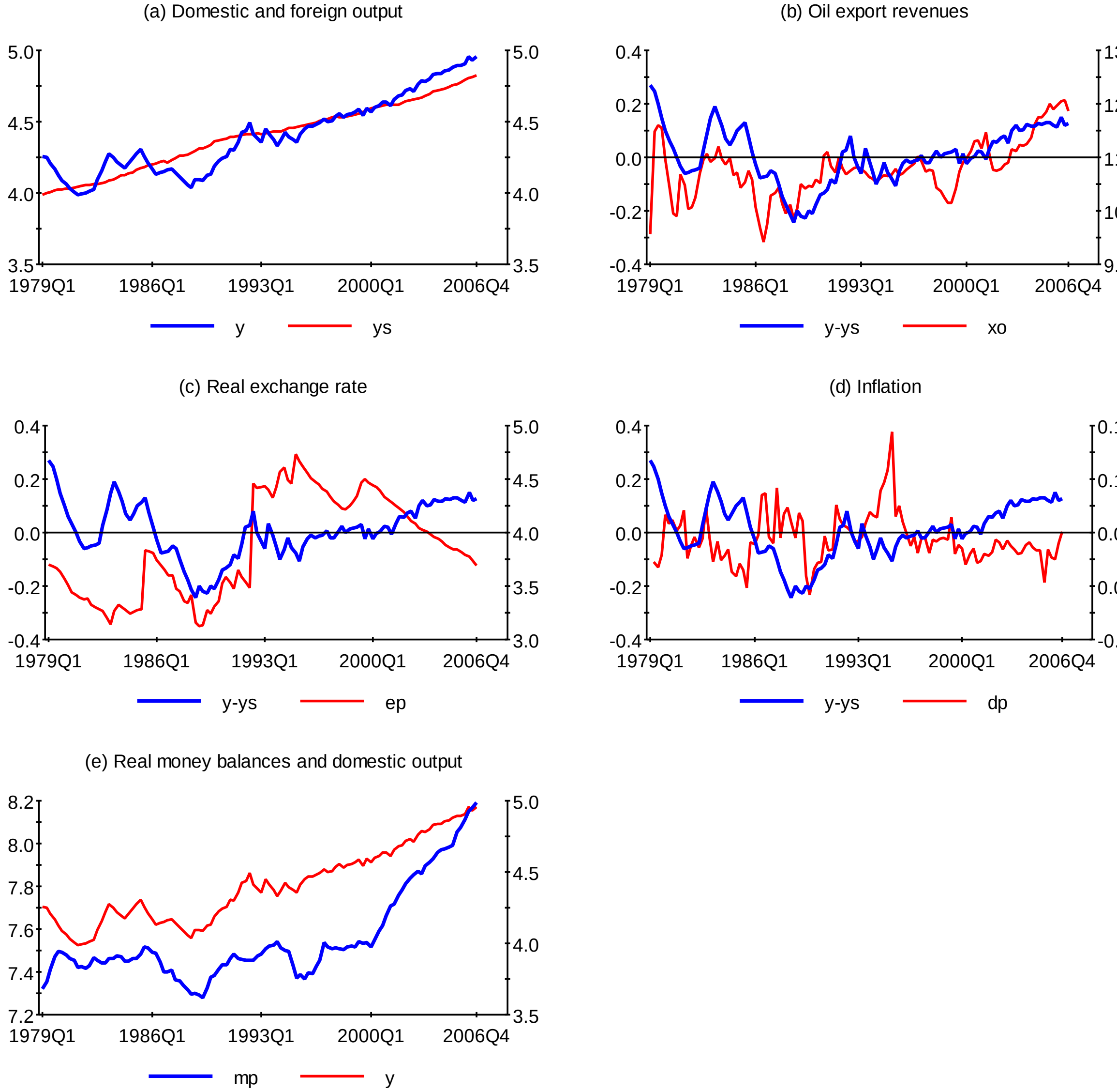
Iran's economy has gone through two major phases since the Islamic Revolution of 1979. The first phase was the aftermath of the Revolution and eight years of war with Iraq. Those years were characterized by mobilization of resources to deal with internal and external conflicts, massive extension of government controls over firms and markets, and efforts to define the institutions of the new political system, the Islamic Republic. The second phase started in 1989 with post-war reconstruction and a series of economic and institutional reforms. After a few years of market-oriented reforms, the government proceeded to liberalize the foreign exchange market and opened up the capital account in 1993. However, the process was not managed well and the country quickly accumulated a huge stock of short-term external debt, followed by a major balance of payments crisis in 1993-1994, see Pesaran (2000) and Esfahani and Pesaran (2009). The debt crisis put the reform program on hold and even reversed it in many areas, especially in the credit and foreign exchange markets. After the mid-1990s, a process of gradual change began in which the government tried to deal with the economy's problems in a more cautious manner.

The performance of real GDP since early 1979 is depicted in Figure 1a. Before the Revolution of February 1979, the Iranian economy was already on a downward trend. But, it went into a tailspin that lowered real GDP by almost a quarter of its 1979Q1 value in the two subsequent years. Part of the problem was the redistributive and political conflicts that undermined the production and investment incentives. The government quickly took over all large firms and all banks and financial companies, restricted trade and capital movements, and expropriated the properties of those believed to be associated with the Shah's regime. Property rights came into question more generally and the economy began to witness a major exodus of skilled labor. The costly war with Iraq during 1980-1988 also caused destruction of property and infrastructure and increasingly drained resources away from productive investment (Figure 1a).

A sharp drop in oil revenues between 1980 and 1982 must have also contributed to the decline in real GDP, see Figure 1b. Indeed, as oil revenues rose in 1982-1984 and then dropped again during 1984-1986, real GDP followed suit. Similar co-movements, especially long-term ones, can be seen after the end of the war in 1988 as well (Figure 1b). The rise of oil revenues during 1989-1991 helped the Iranian economy's quick recovery from the war and the decline of those revenues in 1993 triggered the balance of payments crisis that pushed

Iran's real GDP below its trend until the late 1990s. On the other hand, the recovery of oil prices in 2000 and especially after 2002 ushered in a period of relatively high growth that, so far, has lasted several years. As described in Section 3.1 we model this association between oil revenues and real GDP in the long run and confirm its existence and significance in our econometric exercise in Section 4.2.

Figure 1: Macroeconomic variables for Iran, in log level



Notes: The second variable in each of the figures (a) to (e) should be read using the right-hand scale. For the sources and construction of the data see Appendix A.

Oil revenues have also had an important impact on the exchange rate. The decline in oil revenues in the mid-1990s increased the purchasing power of the dollar in terms of domestic output, a process that has been reversed since the late 1990, see Figure 1c. Before the mid-1990s, the connection between the two variables was different because at that time the government controlled both foreign trade and the foreign exchange market much more tightly and tried to keep the real exchange rate of the dollar low by suppressing the demand for imports. These controls became tighter when oil revenues declined, inducing a positive correlation between foreign earnings and the real price of the dollar (Figure 1c). Such interventions must have had adverse effects on real GDP for a number of reasons. Besides causing inefficient allocation and discouraging exports, lower real value of the dollar meant that oil revenues could buy less domestic goods and resulted slower capital formation. Our econometric results are consistent with this claim.

Tightening of market controls in response to shocks was also a mean of controlling inflation. However, those measures could not work beyond the short or medium term and often resulted in high inflation in the long run. Institutional weaknesses in managing money supply, aggregate demand, and the operation of the markets in general also often manifested themselves in heightened inflation. As Figure 1d shows, the rate of inflation rose sharply in the early 1980s when the economy was grappling with internal political instability, external conflict, and declining oil revenues. The government managed to use monetary expansion and rationing of goods to keep up the real balances in those years (Figure 1e). In 1984 and 1985, the recovery of oil revenues helped lower inflation and raise output. But, the drop in oil prices in 1986 and the continuation of the war led to a sharp rise in inflation and the collapse of aggregate output and real balances until 1989 (Figure 1e).

End of war with Iraq and the start of reconstruction briefly lowered inflation and boosted real balances (Figures 1d and 1e). But, deregulation of many markets and a large depreciation of the rial (see Figure 1c) allowed prices to jump up in 1990. This was followed by a rapid expansion of credit and fiscal spending, which fueled inflation during the early 1990s. Increased imports and output growth were gradually lowering inflation when the balance of payments crisis of 1993-1994 broke out and led to shortage of imports and a significant depreciation of the rial. At the same time, the policy-makers decided to compensate those who owed foreign debt for their losses due to the depreciation. These developments jointly sent inflation soaring in 1995 and brought down real balances sharply (Figures 1d and 1e). In the following years, the government managed to bring down the rate of inflation to more moderate rates and stabilize the real balances, see Amuzegar (1997). Once the economy proved stable in the early 2000s, real balances took off and soon regained its position relative to the real GDP (Figure 1e). However, in recent years, as oil revenues have increased, the government's monetary and fiscal policies have become quite expansionary and have raised inflation to higher levels again.

3.3 A VECX* Model for Iran

In this section we begin by showing how the two long-run relations given by (4) and (5) can be embodied in a vector error-correcting model. We first note that the two long-run relations can be written compactly as deviations from equilibrium:

$$\boldsymbol{\xi}_t = \boldsymbol{\beta}'\mathbf{z}_t - \mathbf{c} - \boldsymbol{\gamma}t \tag{6}$$

where

$$\begin{aligned}\mathbf{z}_t &= (\mathbf{x}_t', \mathbf{x}_t^{*\prime})' = (y_t,\, mp_t,\, \pi_t,\, ep_t,\, y_t^*,\, xo_t)',\\ \mathbf{c} &= (c_y,\ c_{mp})',\ \boldsymbol{\gamma} = (\gamma_y,\ \gamma_{mp})',\ \boldsymbol{\xi}_t = \left(\xi_{yt},\, \xi_{mp,t}\right)'\end{aligned}$$

and

$$\boldsymbol{\beta}' = \begin{pmatrix} -1 & 0 & 0 & \psi_2 & \psi_1 & \psi_3 \\ \phi_1 & -1 & \phi_2 & 0 & 0 & 0 \end{pmatrix} \tag{7}$$

The long-run theory for oil exporting countries, as derived in Esfahani et al. (2012), require two further restrictions on the output equation (4) for Iran, namely $\psi_2 = \psi_3 = \alpha$ and $\psi_1 = \theta(1-\alpha)$, where we are interested in seeing whether in fact the coefficients of the real exchange rate variable, ep_t, and total oil revenues from oil exports are the same and equal to the share of capital in output (α) and whether technological progress in Iran is on par with that of the rest of the world, in other words whether $\theta = 1$, and as a result the coefficient of the foreign real output is equal to $(1-\alpha)$.

The VECX*(s, s^*) model that embodies $\boldsymbol{\xi}_t$ is constructed from a suitably restricted version of the VAR in $\mathbf{z}_t$. In the present application $\mathbf{z}_t = (\mathbf{x}_t', \mathbf{x}_t^{*\prime})'$ is partitioned into the 4×1 vector of endogenous variables, $\mathbf{x}_t = (y_t,\, mp_t,\, \pi_t,\, ep_t)$, and the 2×1 vector of the weakly exogenous variables, $\mathbf{x}_t^* = (y_t^*,\, xo_t)'$. Also as shown in Appendix B, the hypothesis that all the six variables are $I(1)$ cannot be rejected. Moreover, it is easily established that the two exogenous variables are not cointegrated. Under these conditions, following Pesaran et al. (2000), the VAR in $\mathbf{z}_t$ can be decomposed into the conditional model for the endogenous variables:

$$\Delta \mathbf{x}_t = -\boldsymbol{\Pi}_x \mathbf{z}_{t-1} + \sum_{i=1}^{s-1} \boldsymbol{\Psi}_i \Delta \mathbf{x}_{t-i} + \boldsymbol{\Lambda}_0 \Delta \mathbf{x}_t^* + \sum_{i=1}^{s^*-1} \boldsymbol{\Lambda}_i \Delta \mathbf{x}_{t-i}^* + \mathbf{a}_0 + \mathbf{a}_1 t + \boldsymbol{\upsilon}_t, \tag{8}$$

and the marginal model for the exogenous variables:

$$\Delta \mathbf{x}_t^* = \sum_{i=1}^{s-1} \boldsymbol{\Gamma}_i^* \Delta \mathbf{z}_{t-i} + \mathbf{b}_0 + \mathbf{u}_{x^*t}, \tag{9}$$

If the model includes an unrestricted linear trend, in general there will be quadratic trends in the level of the variables when the model contains unit roots. To avoid this, the trend coefficients are restricted such that $\mathbf{a}_1 = \boldsymbol{\Pi}_x \boldsymbol{\delta}$, where $\boldsymbol{\delta}$ is an 6×1 vector of free coefficients, see Pesaran et al. (2000) and Section 6.3 in Garratt et al. (2006). The nature of the restrictions on $\mathbf{a}_1$ depends on the rank of $\boldsymbol{\Pi}_x$. In the case where $\boldsymbol{\Pi}_x$ is full rank, $\mathbf{a}_1$ is unrestricted, whilst it is restricted to be equal to $\mathbf{0}$ when the rank of $\boldsymbol{\Pi}_x$ is zero. Under the restricted trend coefficients the conditional VECX*(s, s^*) model can be written as

$$\Delta \mathbf{x}_t = -\boldsymbol{\Pi}_x \left[\mathbf{z}_{t-1} - \boldsymbol{\delta}(t-1)\right] + \sum_{i=1}^{s-1} \boldsymbol{\Psi}_i \Delta \mathbf{x}_{t-i} + \boldsymbol{\Lambda}_0 \Delta \mathbf{x}_t^* + \sum_{i=1}^{s^*-1} \boldsymbol{\Lambda}_i \Delta \mathbf{x}_{t-i}^* + \tilde{\mathbf{a}}_0 + \boldsymbol{\upsilon}_t, \tag{10}$$

where $\tilde{\mathbf{a}}_0 = \mathbf{a}_0 + \boldsymbol{\Pi}_x \boldsymbol{\delta}$. We refer to this specification as the vector error-correcting model with weakly exogenous $I(1)$ variables, or VECX*(s, s^*) for short. Note that $\tilde{\mathbf{a}}_0$ remains unrestricted since $\mathbf{a}_0$ is not restricted. While for consistent and efficient estimation (and inference) we only require the conditional model as specified in (8), for impulse response analysis and forecasting we need the full system vector error correction model which also includes the marginal model; as such we need to specify the process driving the weakly exogenous variables, $\Delta\mathbf{x}_t^*$.

The long-run theory imposes a number of restrictions on $\mathbf{\Pi}_x$ and $\boldsymbol{\delta}$. First, for the conditional model to embody the equilibrium errors defined by, (6), we must have $\mathbf{\Pi}_x = \boldsymbol{\alpha}_x\boldsymbol{\beta}'$, which in turn implies that $rank(\mathbf{\Pi}_x) = 2$. Furthermore, the restrictions on the trend coefficients are given by

$$\mathbf{\Pi}_x\boldsymbol{\delta} = \boldsymbol{\alpha}_{\mathbf{x}}\boldsymbol{\beta}'\boldsymbol{\delta} = \boldsymbol{\gamma}.$$

Since under cointegration $\boldsymbol{\alpha}_x \neq \mathbf{0}$, it then follows that a trend will be absent from the long-run relations if one of the two elements of $\boldsymbol{\beta}'\boldsymbol{\delta}$ is equal to zero. These restrictions are known as co-trending restrictions, meaning that the linear trends in the various variables of the long-run relations gets cancelled out. This hypothesis is important in the analysis of output convergence between the domestic and the foreign variables, since without such a co-trending restriction the two output series will diverge even if they are shown to be cointegrated.

The theory also imposes a number of long-run over-identifying restrictions on the elements of $\boldsymbol{\beta}$. The total number of over-identifying restrictions is given by $12-4=8$, and there are 4 structural parameters to be estimated, α, θ, ϕ_1 and ϕ_2. This leaves us with 4 over-identifying restrictions to test.

4 Long-Run Estimates and Tests

4.1 Order Selection and Deterministic Components

We propose to use the VECX*(s, s^*) model defined by (10) to test the various long-run theory restrictions set out above. First we need to determine the lag orders s and s^* in the VECX*(s, s^*) model.[7] For this purpose we use the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC) applied to the underlying unrestricted VECX* model. The results are summarized in Table 1. SBC selects the lag orders $\hat{s} = \hat{s}^* = 1$, whilst, as to be expected, AIC selects a higher order lag for the endogenous variables, namely $\hat{s} = 2$ and $\hat{s}^* = 1$. We follow AIC and base our analysis on the VECX*(2,1), since under-estimating the lag orders is generally more serious than overestimating them.

As to the deterministic variables included in our model we make use of both a constant and a linear trend. As a trend may or may not be found in the long-run relations we also test for co-trending restrictions given by $\boldsymbol{\beta}'\boldsymbol{\delta} = \mathbf{0}$. We also experimented with including a war and revolution (WR) dummy amongst the deterministics. The WR dummy takes the value of 1 between 1979 quarter 1 and 1988 quarter 2 and zeros outside this period, and is intended to capture the joint effects of the 1979 Islamic Revolution and the war with Iraq which lasted from September 1980 until August 1988. The WR dummy could also pick up the effects of economic liberalisation that took place after the ending of the Iran-Iraq war.

[7] All estimations and test results are obtained using Microfit 5.0. For further technical details see Pesaran and Pesaran (2009), Section 22.10.

Table 1: Lag order selection criteria

Lag length		AIC	SBC
$s = 1$	$s^* = 1$	1455.32	1327.48
$s = 1$	$s^* = 2$	1445.35	1277.14
$s = 2$	$s^* = 1$	1459.09	1297.61
$s = 2$	$s^* = 2$	1451.63	1249.78

Notes: AIC refers to the Akaike Information Criterion and SBC refers to the Schwarz Bayesian Criterion.

But as we shall argue in Section 4.2.3 below, once xo_t, the oil exports variable, is included in the model the WR dummy ceases to be statistically significant.

4.2 Estimation and Testing of the Long-Run Relations

Having established the order of VECX* to be (2,1) we need to determine the number of cointegrating relations given by $r = rank(\mathbf{\Pi}_x)$, where $\mathbf{\Pi}_x$ is defined by (10). Cointegration tests with null hypothesis of no cointegration, one cointegrating relation, and so on are carried out using Johansen's maximum eigenvalue and trace statistics as developed in Pesaran et al. (2000) for models with weakly exogenous regressors. The test results are reported in Table 2. Both the maximal eigenvalue and the trace statistics suggest the presence of two cointegrating relations at the 5% level, which is the same as that suggested by economic theory, thus we set $r = 2$.

Table 2: Cointegration rank test statistics for the VECX*(2,1) model with endogenous variables (y, mp, dp, ep) and weakly exogenous variables (y*, xo)

H_0	H_1	Test statistic	95% Critical Values	90% Critical Values
(a) Maximal eigenvalue statistic				
$r = 0$	$r = 1$	55.84	41.93	38.29
$r \leq 1$	$r = 2$	40.31	33.79	31.23
$r \leq 2$	$r = 3$	24.66	26.26	23.93
$r \leq 3$	$r = 4$	6.30	17.73	16.08
(b) Trace statistic				
$r = 0$	$r = 1$	127.11	90.44	84.24
$r \leq 1$	$r = 2$	71.27	60.13	56.47
$r \leq 2$	$r = 3$	30.97	36.97	34.02
$r \leq 3$	$r = 4$	6.30	17.73	16.08

Notes: The underlying VECX* model is of order (2,1) and contains unrestricted intercept and restricted trend coefficients. y_t^* and xo_t are treated as weakly exogenous, non-cointegrated $I(1)$ variables. The test statistics refer to Johansen's log-likelihood-based maximum eigenvalue and trace statistics and are computed using 109 observations from 1979Q4 to 2006Q4.

In order to exactly identify the long-run relations, we must impose 4 restrictions, 2 restrictions on each of the 2 cointegration relations. The choice of the exactly identifying restrictions is econometrically innocuous and is best guided by economic theory. We proceed by taking the first cointegrating relation to be the output equation, defined by equation (4) and normalised on y_t, and the second one the money demand equation, defined by (5) and normalised on $mp_t = m_t - p_t$. Accordingly, we start with the following two exactly identified cointegrating vectors

$$\boldsymbol{\beta}'_{EX} = \begin{pmatrix} -1 & 0 & \beta_{13} & \beta_{14} & \beta_{15} & \beta_{16} \\ \beta_{21} & -1 & \beta_{23} & \beta_{24} & 0 & \beta_{26} \end{pmatrix}, \tag{11}$$

where the rows of $\boldsymbol{\beta}'_{EX}$ correspond to $\mathbf{z}_t = (y_t, mp_t, \pi_t, ep_t, y^*_t, xo_t)'$. Using this exactly identified specification we then test the co-trending restrictions, $\boldsymbol{\beta}'\boldsymbol{\delta} = \boldsymbol{\gamma} = (\gamma_y, \gamma_{mp})' = \mathbf{0}$. The log-likelihood ratio (LR) statistic for jointly testing the two co-trending restrictions takes the value 10.15, and is asymptotically distributed as a chi-squared variate with two degrees of freedom. Therefore, based on the asymptotic distribution the co-trending restrictions are rejected. But we are working with a relatively large dimensional VECX* model using a moderate number of time series observations. In such situations it is known that the LR tests could over-reject in small samples (see, for example, Gredenhoff and Jacobson (2001) as well as Gonzalo (1994), Haug (1996) and Abadir et al. (1999)). To deal with the small sample problem we computed bootstrapped critical values based on 1,000 replications of the LR statistic. Using the observed initial values of each variable, the estimated model, and a set of random innovations, an artificial data set is generated for each of the 1,000 replications under the assumption that the estimated version of the model is the true data-generating process. For each of the replicated data sets, we first estimate our VECX* model subject to the exact identifying restrictions in (11) and then subject to the two co-trending restrictions. Finally, the empirical distribution of the LR test statistic is derived using the 1,000 replications. Having applied this technique, the bootstrapped critical value for the joint test of the two co-trending restrictions is 10.20 at the 5% level, and 15.22 at the 1% level, as compared to the LR statistic of 10.15. Hence, based on the bootstrapped critical values the co-trending restrictions cannot be rejected at the conventional levels of significance, although the outcome of the test at the 5% level is rather marginal and is subject to the random variation of the bootstrapped critical values.

The results here contrasts with those of Esfahani et al. (2012), in which small quarterly models for six major oil economies (Iran, Kuwait, Libya, Nigeria, Saudi Arabia, and Venezuela) in $\mathbf{z}_t = (y_t, ep_t, y^*_t, xo_t)'$ is estimated, where the co-trending restriction on the long-run output equation cannot be rejected at the 1% level for all the major oil exporters under consideration with the exception of Iran. However, given the outcome of the test in this paper we shall impose the co-trending restrictions whilst considering the other theory restrictions, and return to them to see if they continue to be supported by the data once the other restrictions are imposed.

4.2.1 Testing Long-Run Theory Restrictions

We first consider the theory restrictions on the output equation whilst maintaining the exact identifying restrictions on the second long-run relation. Initially we impose the restriction that the coefficients of ep_t and xo_t are the same, namely that in (11) $\beta_{14} = \beta_{16} = \alpha$. We obtain the estimates

$$\hat{\psi}_1 = \underset{(0.2183)}{0.6931}, \; \hat{\psi}_2 = \hat{\psi}_3 = \hat{\alpha} = \underset{(0.1100)}{0.3140},$$

with the LR statistic of 10.52 for testing the three restrictions. The figures in brackets are asymptotic standard errors. The additional restriction has only marginally increased the LR statistic and is clearly not rejected. In fact the bootstrapped critical values for the test is now 11.91 at the 5% level and 17.12 at the 1% level. The implicit estimate of θ given by $0.6931/(1-0.3140) = 1.01$ is very close to unity and the null hypothesis that $\theta = 1$ cannot be rejected, thus implying that the technological growth in Iran is on par with that of the rest of the world. Under $\theta = 1$ we have $\beta_{15} + \beta_{14} = 1$, and imposing this additional restriction the LR statistic increases only marginally from 10.5181 to 10.5198. In addition, the coefficient of π_t in the long-run output equation is $\hat{\beta}_{13} = -14.72$ (5.91), which is statistically significant, implying that inflation has a negative effect on real output which is not supported by the long-run theory. This negative effect suggests inefficiencies in both the institutions and economic policies in Iran and shows the importance of controlling inflation for growth promotion in Iran.

While a short-run positive relationship between inflation and output growth has been widely documented in the literature investigating the empirical validity of the Phillips curve, there is also a strand in the literature that argues for a long-run negative association between the inflation rate and real output growth. This negative relationship is shown to hold even in advanced economies when inflation is above a certain threshold, see Fischer (1983), Fischer (1993), and López-Villavicencio and Mignon (2011) for a recent survey. In addition, inflation is often included as a control variable in the mainstream panel regressions in the growth literature, so as to capture the possible negative effects of price instability on economic growth (Aghion et al. (2009) and Cavalcanti et al. (2012)). It is argued, as we do, that the long-run negative impact of inflation on output comes about because (high) inflation reduces investment, and therefore capital accumulation and through that productivity growth. See also sub-section 4.2.5, where we consider the empirical evidence on the relationship between inflation and the investment-output ratio.

Consider now the second long-run equation. The theory restrictions in terms of the elements of $\boldsymbol{\beta}$ in (11) are

$$\beta_{24} = 0, \text{ and } \beta_{26} = 0.$$

Imposing these additional restrictions on $\boldsymbol{\beta}$ yields

$$\begin{aligned} \theta &= 1,\ \hat{\alpha} = \underset{(0.0465)}{0.2333},\ \hat{\beta}_{13} = \underset{(4.01)}{-13.06}, \\ \hat{\phi}_1 &= \underset{(0.1231)}{0.8277},\ \hat{\phi}_2 = \underset{(6.09)}{-14.53}\,. \end{aligned}$$

The long-run income elasticity of money demand is close to unity and the null hypothesis that it is equal to 1 cannot be rejected. The effect of inflation on real money balances is also negative and statistically significant. This is in line with our earlier discussion that inflation in the money demand equation acts as a proxy for the interest rate. In fact it would be a perfect proxy if it can be assumed that the Fisher parity holds in Iran. Imposing $\phi_1 = 1$ and re-estimating subject to all the seven over-identifying restrictions we obtain

$$\begin{aligned} \theta &= 1,\ \hat{\alpha} = \underset{(0.0489)}{0.2647},\ \hat{\beta}_{13} = \underset{(4.37)}{-13.84}, \\ \phi_1 &= 1,\ \hat{\phi}_2 = \underset{(6.79)}{-16.37}\,. \end{aligned}$$

The LR statistic for testing all 7 restrictions jointly is 23.34 which is to be compared to the bootstrapped critical values of 21.59 and 30.99 at the 5 and 1 percent significance levels, respectively. Therefore, the restrictions are rejected at 5% level, but not at the 1% level. The test outcome is inconclusive and further investigation seems in order. We considered relaxing some of the restrictions in the real money demand equation and found that the primary source of the rejection of the restrictions is the zero restriction imposed on the coefficient of the real exchange rate variable. Once this restriction is relaxed the following estimates are obtained

$$\begin{aligned}\theta &= 1,\ \hat{\alpha} = \underset{(0.0600)}{0.2467},\ \hat{\beta}_{13} = \underset{(3.36)}{-12.06}, \\ \phi_1 &= 1,\ \hat{\phi}_2 = \underset{(2.99)}{-1.91},\ \hat{\beta}_{24} = \underset{(0.0496)}{-0.2380}.\end{aligned}$$

There are now six over-identifying restrictions on the long-run relations, and the LR statistic for testing these restrictions is 13.37 as compared to the bootstrapped critical values of 16.29 and 19.34 at the 10 and 5 percent significance levels, respectively. Clearly, the restrictions are not rejected even at the 10% significance level. This is reassuring particularly as far as the long-run estimates of the output equation is concerned, since whether β_{24} is restricted or not seems to have little effects on the estimates of the output equation, which is the focus of the present investigation. However, relaxing $\beta_{24} = 0$ does significantly affect the inflation elasticity of the money demand which is reduced from -16.37 to -1.91 and is no longer statistically significant.

We are presented with a clear choice. Should we maintain the theory restrictions which are rejected at the 5% level, although not at the 1% level, or should we opt for the new specification of the real money demand equation that includes the $e_t - p_t$ variable which is difficult to justify in an economically meaningful sense. Given that we are primarily interested in the long-run effects of oil exports for real output, and the choice of the real money demand equation does not seem to play a central role for that issue, in the rest of the paper we shall maintain the theory consistent money demand equation since it is easier to interpret. Also, since the theory restrictions are not rejected at the 1% level, our adherence to a theory consistent real money demand equation is not without some empirical foundations.

Furthermore, the theory consistent specifications are robust to alternative measurements of foreign output and the exchange rate. For instance, estimating the VECX* model with foreign output computed using fixed weights based on the average of three consecutive years (2001-2003), yield similar outcomes

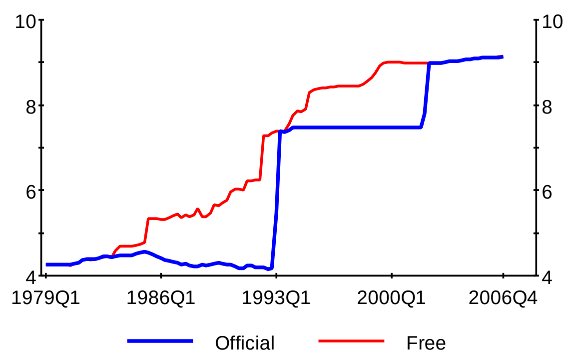
$$\begin{aligned}\theta &= 1,\ \hat{\alpha} = \underset{(0.0432)}{0.2311},\ \hat{\beta}_{13} = \underset{(5.08)}{-17.13}, \\ \phi_1 &= 1,\ \hat{\phi}_2 = \underset{(6.30)}{-16.06},\end{aligned}$$

to when we use foreign output based on time-varying weights (y_t^*), with the 7 over-identifying restriction now not being rejected at the 5% significance level.

4.2.2 Free and Official Exchange Rates

As noted earlier, a similar issue of measurement also arises with respect to the exchange rate. Since the 1979 Revolution the Iranian rial has depreciated significantly against the US dollar under a variety of exchange rate regimes from a fixed rate to multiple rates and back to a unified pegged managed rate. It has depreciated from 70 rials per US dollar in 1979 to 9170 rials in 2006, or around 131 fold increase, see Pesaran (1984) and Pesaran (2000). Figure 2 shows (in logs) the free rate (or black market in certain periods), e_t , and the official exchange rate , $e_{OF,t}$, over the period 1979Q1-2006Q4. The two rates are at par at the start of the Revolution but depart soon thereafter. They are, however, brought in line by two major jumps the last of which is associated with the successful unification of the exchange rates during Khatami's Presidency in 2002.

Figure 2: Free and official exchange rates



To investigate the robustness of our results to the choice of exchange rate we employ a geometrically weighted average of the free and the official rates, $e_{\omega,t} = \omega e_t + (1-\omega)e_{OF,t}$. The weights $\omega : (1-\omega)$ are intended to reflect the proportion of imports by public and private agencies that are traded at the two exchange rates, on average. There is little hard evidence on ω, although due to the gradual attempts at currency unification, it is reasonable to expect ω to have risen over time. Initially we set $\omega = 0.75$, but smaller values of $\omega = 0.70$ and 0.60 resulted in very similar estimates and test outcomes. Using $e_{\omega,t}$ with $\omega = 0.75$ we could not reject the 7 over-identifying restrictions even at the 10% level, since the LR statistic is 18.65 as compared to the bootstrapped critical values of 19.10 and 22.67 at the 10 and 5 percent significance levels, respectively. For $\omega = 0.75$ we obtained the following estimates:

$$\begin{aligned} \theta &= 1,\ \hat{\alpha} = \underset{(0..0308)}{0.1964},\ \hat{\beta}_{13} = \underset{(2.55)}{-8.97}, \\ \phi_1 &= 1,\ \hat{\phi}_2 = \underset{(6.84)}{-16.01}, \end{aligned}$$

which yield a smaller capital share of 0.1964 as compared to 0.2647, with the coefficient of inflation in the output equation still negative and statistically significant. However, the inflation elasticity of money demand, -16.01, is roughly the same as in the case when we use the floating exchange rate, e_t. Given that we do not know what these weights should be, for now we will proceed by only reporting the results when using the free exchange rate in our model, but we will return to this issue when looking at the short-run dynamics.[8]

[8]We also estimated the VECX* model with $e_{0.75,t}$ and the foreign output variable constructed using fixed

4.2.3 Including a War and Revolution Dummy

To see if the model captures the effects of the 1979 Islamic Revolution as well as the war with Iraq, which lasted from September 1980 until August 1988 and the economic liberalisation that followed after the war, we introduce a war and revolution (WR) dummy. This WR dummy takes the value of unity over the period 1979Q1 to 1988Q2, and zero otherwise. As before both the maximal eigenvalue and the trace statistics indicate the presence of two cointegrating relations at the 5% level.[9] Setting $r = 2$ and imposing the same over-identifying restrictions as in the above Sub-sections, namely:

$$\begin{aligned}
\boldsymbol{\beta}'\boldsymbol{\delta} &= \boldsymbol{\gamma} = \mathbf{0}, \\
\beta_{14} &= \beta_{16} = \alpha, \\
\beta_{15} + \beta_{14} &= 1 \Longrightarrow \theta = 1, \\
\beta_{21} &= 1 \Longrightarrow \phi_1 = 1, \\
\beta_{24} &= 0, \text{ and } \beta_{26} = 0,
\end{aligned}$$

and re-estimating subject to the seven over-identifying restrictions we obtain

$$\begin{aligned}
\theta &= 1,\ \hat{\alpha} = \underset{(0.0647)}{0.2870},\ \hat{\beta}_{13} = \underset{(12.57)}{-20.81}, \\
\phi_1 &= 1,\ \hat{\phi}_2 = \underset{(14.94)}{-18.48}.
\end{aligned}$$

The LR statistic for testing all 7 restrictions jointly is 24.02 which is to be compared to the bootstrapped critical values of 22.30 and 29.09 at 5% and 1% significance levels, respectively. Therefore, as before the restrictions are rejected at 5% level, but not at the 1% level. The estimates are fairly similar to the case when we do not include the WR dummy variable, with the long-run negative effects of inflation on real output still present, although now statistically less significant than previously. Table 3 reports the coefficient of the WR dummy variable in the error correction equations where we observe that the WR dummy is clearly insignificant at the 10% level in the real exchange rate and the inflation equations, while it is significant at the 10% level for the real money equation and at the 5% level in the output equation. These estimates suggest only a modest average decline in real output due to revolution and war, once the effects of the declines in real oil exports are taken into account.

This point is clearly illustrated using Figure 3 which shows the significant drop in oil exports in the aftermath of the revolution, which only begins to recover in a sustained manner after the end of the war with Iraq. In effect, the decline in oil exports, partly due to the economic disruptions, in turn puts further downward pressure on the real economy. Although the price of oil declined slightly and steadily between 1979 and 1986, this was not the case for Iranian revenues from oil exports which drop significantly after the revolution and again at the start of the Iran-Iraq war while on the other hand was at a higher level than the price of oil after the war. Thus, the negative effects of the war and revolution is largely picked up by the oil export variable, xo_t.

weights and obtained very similar estimates. These results are available upon request.

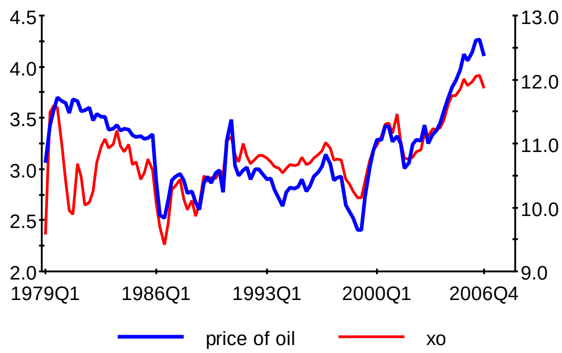
[9] The inclusion of the dummy variable changes the critical values of the test. The test statistics and the associated critical values are available on request.

Table 3: Reduced-form error correction equations of the VECX*

Equation	Δy_t	Δmp_t	$\Delta \pi_t$	Δep_t
WR Dummy	-0.0170^{*} (0.0063)	-0.0086^{**} (0.0047)	0.0046 (0.0039)	0.0079 (0.0266)

Notes: * denotes significance at the 5% level and ** denotes significance at the 10% level.

Figure 3: Price of oil and revenue from oil exports (xo)



However, if we had followed the literature and instead of total revenue from oil exports, xo_t, used the nominal price of oil, p_t^o, in our model, then the war and revolution dummy would have been necessary for modelling the disruptive effects of the revolution and the war on the real economy. In the light of these observations, we will work with the model with the xo_t variable included, but without the war and revolution dummy.

4.2.4 Import Weights as Opposed to Trade Weights

We also estimated our model with foreign output computed using import weights, both fixed and time-varying, rather than trade weights. The cointegration rank test statistics for the VECX* (2,1) model with the data vector defined by $\mathbf{z}_t = \left\{y_t, \, mp_t, \, \pi_t, \, ep_t, \, y^*_{t,IM}, \, xo_t\right\}$, where $y^*_{t,IM}$ is real foreign output using time-varying import weights, again suggest the presence of two long-run relations. Imposing the same 7 over-identifying restrictions as before and re-estimating we obtain

$$\begin{aligned}\theta &= 1, \; \hat{\alpha} = \underset{(0.0487)}{0.2702}, \; \hat{\beta}_{13} = \underset{(4.20)}{-13.79}, \\ \phi_1 &= 1, \; \hat{\phi}_2 = \underset{(6.47)}{-16.02}\,.\end{aligned}$$

The LR statistic for testing all 7 restrictions jointly is now 28.65 which is to be compared to the bootstrapped critical values of 21.79 and 30.43 at 5% and 1% significance levels, respectively. The results are very similar to the ones reported in the above Sub-sections, and shows that the choice of the weights in the construction of the foreign variable is of second order importance. However, given the important changes that have taken place in the geographical composition of the Iranian foreign trade since the revolution, gradually shifting Iran's trade from the West to the East, in what follows we use the time-varying trade weights as in Section 4.2.

4.2.5 The Role of Investment

As noted earlier it is relatively straightforward to augment the VECX* model with other aggregate variables such as log real consumption (c_t) and log real investment (i_t). But given the long-run focus of our analysis, the inclusion of these variables are unlikely to alter the long-run relationship that we have estimated between real output and oil income if c_t and i_t are cointegrated with y_t and xo_t. This is because any linear combination of cointegrating relations will also be cointegrated.

Here we focus on the role of investment in the interrelation of oil income, real output, and inflation since it is generally believed that changes in oil income affect real output primarily through changes in investment. Real consumption is also quite stable and does not seem to respond significantly to short-run changes in oil income.

Figure 4: Real domestic output (y) and investment (i)

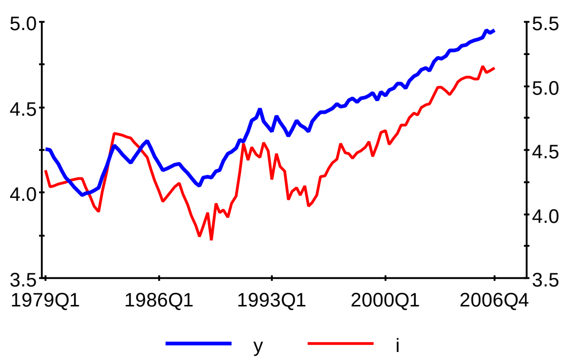


Figure 4 shows the evolution of log real output and investment over the period 1979q1-2006q4. For comparability output and investment data are first transformed into indices with 2000 as the base year before taking logarithms. It is clear that the two series move quite closely, although investment tend to be much more volatile than output. To check their cointegrating properties we estimated an exactly identified cointegrating VAR(2) in i_t and y_t with unrestricted intercepts and a restricted trend. The cointegration rank test statistics for this model is given in Table 4. The test results strongly support the existence of cointegration between y_t and i_t. But the co-trending restriction (that real output and investment have the same deterministic trend components) is rejected. The cointegrating relationship between output and investment is given by

$$y_t = \underset{(0.0406)}{0.3179} i_t + \underset{(0.0004)}{0.0059} t + \xi_{yi,t}, \text{ where } \xi_{yi,t} \sim I(0). \tag{12}$$

The long-run impact of investment on real output is significantly different from unity, as implied by the standard neoclassical growth model without oil income. In contrast the exactly identified cointegrating relation between log real output and consumption is given by $c_t - y_t$ which does satisfy the standard long-run theory restriction.

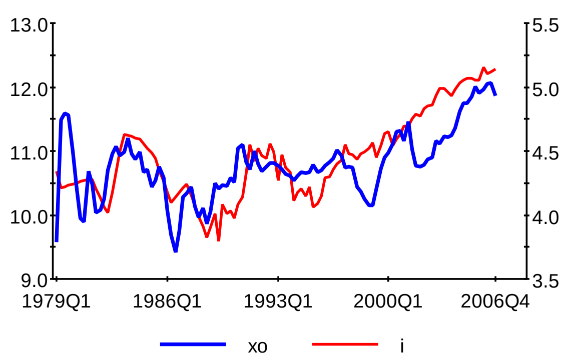
Table 4: Cointegration rank test statistics for the VAR(2) model with y and i

H_0	H_1	Test statistic	95% Critical Values	90% Critical Values
(a) Maximal eigenvalue statistic				
$r = 0$	$r = 1$	29.59	19.22	17.18
$r \leq 1$	$r = 2$	3.94	12.39	10.55
(b) Trace statistic				
$r = 0$	$r = 1$	33.53	25.77	23.08
$r \leq 1$	$r = 2$	3.94	12.39	10.55

Notes: The test statistics refer to Johansen's log-likelihood-based maximum eigenvalue and trace statistics and are computed using 110 observations from 1979Q3 to 2006Q4.

The fact that output is not responsive to investment on a one-to-one basis, even in the long run, might be indicative of some inefficiencies in the way oil income has been utilized in the Iranian economy. To see this in Figure 5 we show the evolution of (log) oil export revenues and log investment. It is clear that both variables share the same trend over the long run, with some important short-run deviations. Estimating a cointegrating VAR(2) model for investment and oil export revenues, the cointegration rank test statistics in Table 5 suggest that there is cointegration relation between investment and oil export revenues. It is also interesting that in the case of these variables the co-trending restriction is not rejected, and the hypothesis that the long-run elasticity of investment to real oil income is unity cannot be rejected either, and as a result: $i_t = xo_t + \xi_{xoi,t}$, where $\xi_{xoi,t} \sim I(0)$.

Figure 5: Oil export revenues (xo) and investment (i)



Therefore, oil export revenues represent an excellent proxy for investment in the Iranian economy, providing further justification for our modelling strategy of using oil exports as one

Table 5: Cointegration rank test statistics for the VAR(2) model with i and xo

H_0	H_1	Test statistic	95% Critical Values	90% Critical Values
(a) Maximal eigenvalue statistic				
$r = 0$	$r = 1$	30.57	19.22	17.18
$r \leq 1$	$r = 2$	6.33	12.39	10.55
(b) Trace statistic				
$r = 0$	$r = 1$	36.90	25.77	23.08
$r \leq 1$	$r = 2$	6.33	12.39	10.55

Notes: The test statistics refer to Johansen's log-likelihood-based maximum eigenvalue and trace statistics and are computed using 110 observations from 1979Q3 to 2006Q4.

of the main long-run drivers of real output. The above results also show that from a long-run perspective only one of the two variables (investment or oil exports) need to be included in the cointegrating model. Our decision of including oil exports rather than investment is justified on the ground that xo_t is likely to be exogenous to the Iranian economy whilst the same cannot be said of i_t.

Table 6: Cointegration rank test statistics for the VAR(2) model with i, y, and dp

H_0	H_1	Test statistic	95% Critical Values	90% Critical Values
(a) Maximal eigenvalue statistic				
$r = 0$	$r = 1$	35.52	25.42	23.10
$r \leq 1$	$r = 2$	16.96	19.22	17.18
(b) Trace statistic				
$r = 0$	$r = 1$	56.12	42.34	39.34
$r \leq 1$	$r = 2$	20.60	25.77	23.08

Notes: The test statistics refer to Johansen's log-likelihood-based maximum eigenvalue and trace statistics and are computed using 109 observations from 1979Q4 to 2006Q4.

It is also worth noting that our analysis is compatible with the traditional view that changes in oil income primarily affect output through investment. But the relatively low long-run impact of investment on output, estimated to be around 0.32, and the rejection of the co-trending restriction in the long-run investment-output equation could also be due to the high levels of inflation experienced in Iran over the past three decades, and the negative effects that such high rates of inflation can have on investment, as argued above. To test this viewpoint we re-estimated the investment sub-model including inflation as an additional variable. We found that investment, real output and inflation are cointegrating (see Table 6), but with inflation included in the sub-model the trend term becomes less statistically significant and the joint hypothesis that the cointegrating relation is co-trending and the long-run elasticity of investment to output is unity is now not rejected. The log-likelihood ratio statistic for testing the two restrictions is 13.81 as compared to the bootstrapped critical value of 14.87 at the 1% significance level. More specifically, we obtain the following estimate

$$i_t - y_t = \underset{(5.10)}{-13.89}\ \pi_t + \xi_{yi,t},\ \xi_{yi,t} \sim I(0), \tag{13}$$

which confirms a statistically significant negative association between inflation and the (log) investment-output ratio. Clearly, further research is required on the adverse effects of high inflation on the Iranian economy. Understanding the nature of these inefficiencies is beyond the scope of the present paper and requires more detailed disaggregated analysis.

5 Short-Run Dynamics

We now return to the full model and use it to examine the dynamic responses of the Iranian economy to shocks to oil exports and foreign output. Initially, we consider the effects of system-wide shocks on the cointegrating relations using the persistence profiles, developed by Lee and Pesaran (1993) and Pesaran and Shin (1996). On impact the persistence profiles (PP) are normalized to take the value of unity, but the rate at which they tend to zero provide information on the speed with which equilibrium correction takes place in response to shocks. The PP could initially over-shoot, thus exceeding unity, but must eventually tend to zero if the long-run relationship under consideration is cointegrating. To investigate the effects of variable specific shocks on the Iranian economy we make use of the Generalized Impulse Response Functions (GIRFs), developed in Koop et al. (1996) and Pesaran and Shin (1998). Unlike the orthogonalized impulse responses popularized in macroeconomics by Sims (1980), the GIRFs are invariant to the ordering of the variables in the VECX* model.

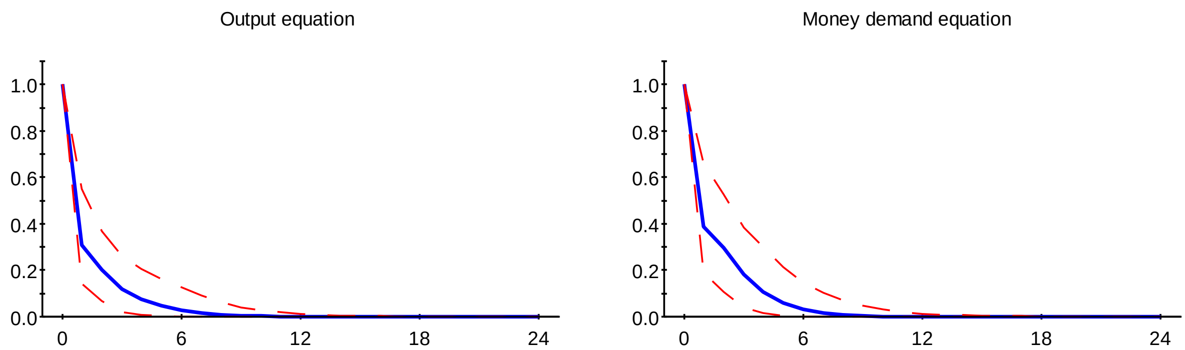
5.1 Persistence Profiles

Figure 6 depicts of the effect of a system-wide shock to the cointegrating relations with 95% bootstrapped confidence bounds. The speed of convergence to equilibrium for the two cointegrating relations are quite fast as compared, for example, with the UK (Garratt et al. (2006)) and Switzerland (Assenmacher-Wesche and Pesaran (2009)). The half life of the shock is less than one quarter and the life of the shock is generally less than eight quarters. Thus the effect of shocks tend to disappear rather quickly. This could be due to lack of access to capital markets and an absence of a developed domestic capital and money markets, which allows little possibility for shock absorptions. The recently created National Development Fund could, in principle, if used appropriately act as a shock absorber which might lead to a more sluggish response of the economy to shocks.

5.2 Generalized Impulse Responses

Generalized Impulse Response Functions (GIRFs) can be computed for shocks to any of the variables in the model, but they are more straightforward to interpret in the case of shocks to the exogenous variables, namely oil exports and foreign output. Consider first the GIRFs of a positive unit shock (equal to one standard error) to oil exports given in Figure 7. These figures clearly show that the shock to oil exports significantly increases inflation, strengthens

Figure 6: The persistence profiles of the effect of a system-wide shock to the cointegrating relations with 95 percent bootstrapped confidence bounds



the exchange rate variable $(e_t - p_t)$, increases real output, but its effect on real money balances whilst positive is not statistically significant. These results are as to be expected, but also show that the effects of the shock work themselves through the economy rather rapidly. Note also that these effects tend to be permanent, due to the presence of unit roots in the underlying variables. Quantitatively, the positive oil export shock increases inflation by 0.8% per annum, real output by 3.2% and results in an exchange rate appreciation (relative to domestic prices) of around 7.6%. The rise in the exchange rate variable in the aftermath of the positive shock to oil exports can also be viewed as supporting the Dutch disease, although here the rise in the exchange rate relative to domestic prices is in fact accompanied with a rise in real output which does not sit comfortably with those that view the Dutch disease as a resource curse.[10]

The GIRFs of a unit shock to foreign output are given in Figure 8. By comparison to the oil export shock these effects are muted and generally statistically insignificant. By far the most important effect of the foreign output shock is on the real exchange rate variable $e_t - p_t$, which appreciates by 2 per cent and is statistically significant for the first 3-4 quarters after the shock.

5.3 Error-Correcting Equations

Using the estimates of the conditional model, (10), the error-correcting property of the model can also be seen in the size and significance of the coefficients of the error-correcting terms, $\boldsymbol{\xi}_t = (\xi_{t,y}, \xi_{t,mp})'$, defined by (6). The estimates of the reduced form error correction equations are given in Table 7, from which we can see that $\hat{\xi}_{t-1,y}$ and $\hat{\xi}_{t-1,mp}$ are both statistically significant in the output and real exchange rate equations but not in the real money and inflation equations. There seems to be a dichotomy between the real and the financial sides of the economy as far as their responses to disequilibria are concerned with

[10] For a short-run macroeconomic analysis where a rise in oil exports induces a rise in real output see Pesaran (1984). See also Cavalcanti et al. (2011b) and Cavalcanti et al. (2012) for recent panel studies.

Figure 7: Generalized Impulse Responses of a positive unit shock to oil export revenues (with 95 percent bootstrapped confidence bounds)

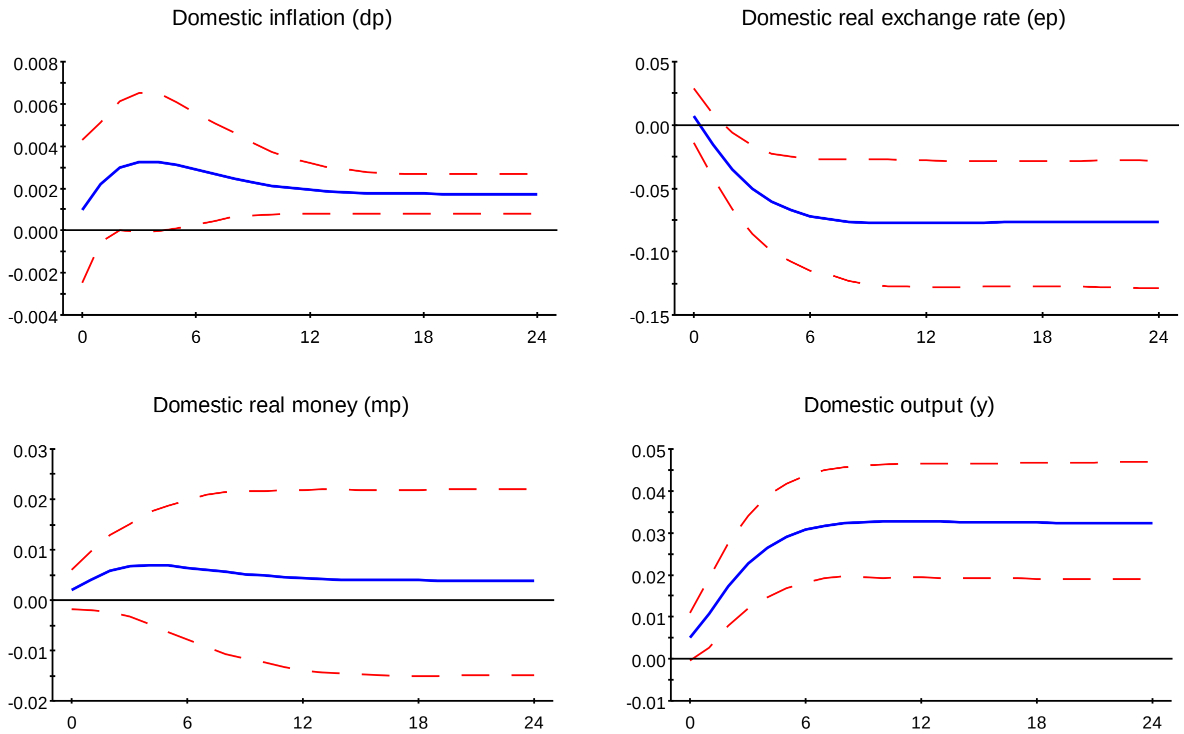
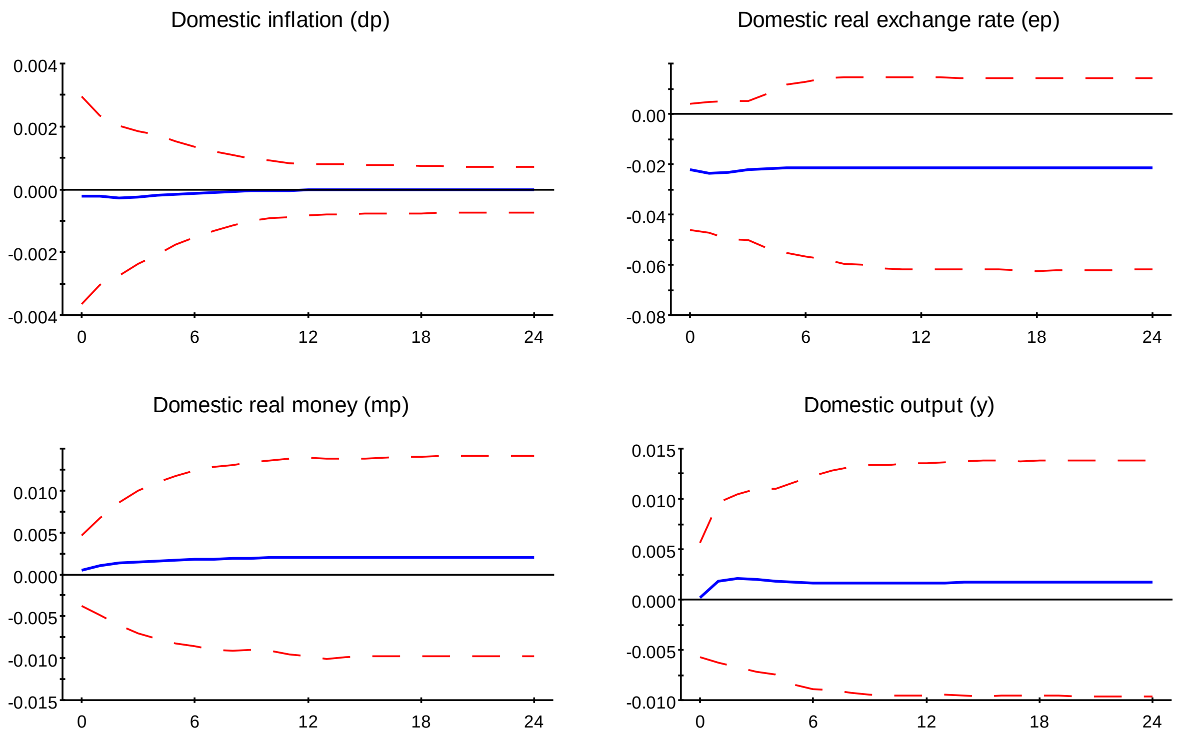


Figure 8: Generalized Impulse Responses of a positive unit shock to foreign output (with 95 percent bootstrapped confidence bounds)



the real output and exchange rate adjusting most to shocks.

Table 7: Reduced-form error correction equations of the VECX*

Equation	Δy_t	Δmp_t	$\Delta \pi_t$	Δep_t
$\hat{\xi}_{y,t-1}$	0.089* (0.029)	0.026 (0.021)	0.028 (0.018)	−0.401* (0.118)
$\hat{\xi}_{mp,t-1}$	−0.047* (0.023)	0.011 (0.017)	−0.008 (0.014)	0.332* (0.093)
Δy_{t-1}	0.325* (0.094)	0.086 (0.069)	−0.038 (0.058)	−0.184 (0.386)
Δmp_{t-1}	−0.515* (0.174)	0.052 (0.128)	0.061 (0.106)	0.542 (0.711)
$\Delta \pi_{t-1}$	−0.021 (0.167)	0.232** (0.123)	−0.143 (0.102)	0.335 (0.681)
Δep_{t-1}	−0.082* (0.023)	−0.016 (0.017)	−0.001 (0.014)	−0.042 (0.094)
Δy_t^*	0.073 (0.547)	0.111 (0.402)	−0.036 (0.334)	−4.162** (2.235)
Δxo_t	0.023** (0.014)	0.009 (0.010)	0.004 (0.008)	0.027 (0.056)
intercept	−0.363* (0.144)	−0.005 (0.106)	−0.089 (0.088)	2.157* (0.591)
$\bar{R}^2$	0.189	0.226	0.138	0.085
$\bar{R}^2$-$AR(p)$	0.054 (p=1)	0.158 (p=2)	0.141 (p=2)	0.00 (p=1)
SC: $\chi^2(4)$	0.71	7.95	11.74	8.17
FF: $\chi^2(1)$	3.49	0.57	2.18	11.27
N: $\chi^2(2)$	1.97	2.96	9.56	3354.6
HS: $\chi^2(1)$	0.22	3.45	9.55	19.91

Notes: The two error correction terms are given by:

$$\begin{aligned} \xi_{y,t} &= y_t + \underset{(4.37)}{13.84}\pi_t - \underset{(0.0489)}{0.2647}ep_t - \underset{(0.0489)}{0.7353}y_t^* - \underset{(0.0489)}{0.2647}xo_t \\ \xi_{mp,t} &= mp_t - y_t + \underset{(6.79)}{16.37}\pi_t \end{aligned}$$

*denotes significance at the 5% level and ** denotes significance at the 10% level. SC is a test for serial correlation, FF a test for functional form, N a test for normality of the errors, and HS a test for heteroscedasticity. Critical values are 3.84 for $\chi^2(1)$, 5.99 for $\chi^2(2)$ and 9.49 for $\chi^2(4)$. $\bar{R}^2$ is the adjusted squared multiple correlation coefficient, and $\bar{R}^2$-AR(p) refers to the $\bar{R}^2$ of a univariate autoregressive equation. The sample period is 1979Q1 to 2006Q4.

Turning to the fit of the error-correcting equations, the inflation and the real money balances equations seem to be the least satisfactory. In the case of the inflation equation none of the regression coefficients are statistically significant and the error-correction regression suffers from statistically significant residual serial correlation.[11] In the case of the real money balances the only significant coefficient is that of the lagged inflation which is significant at the 10% level. The fit of the exchange rate equation seems reasonable, considering the general unpredictably of exchange rates documented in the literature. By contrast, the

[11]The inflation equation also seems to suffer from multicollinearity since despite the fact that none of its coefficients are statistically significant the overall fit of the equation is highly significant.

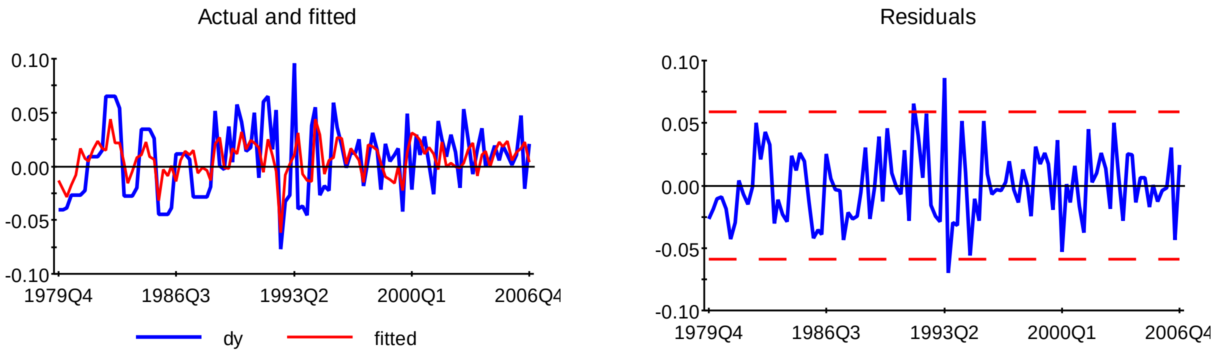
output equation provides a reasonable explanation, particularly considering the significant disruptions experienced by the Iranian economy over the period under study and the fact that no dummy variables are included in the regressions.

To evaluate the importance of the error correction terms we also estimated univariate autoregressive moving average (ARMA) time series equations for the four endogenous variables in the VECX* model and concluded that an $AR(1)$ specification fits best for the real output growth (Δy_t), and the exchange rate changes (Δep_t), and an $AR(2)$ specification for changes in inflation ($\Delta \pi_t$) and real money balances (Δmp_t) . The adjusted squared multiple correlation coefficient of these univariate equations are denoted by $\bar{R}^2$-$AR(p)$, which needs to be compared to the $\bar{R}^2$ of the error correction equations also presented in Table 7. It is clear that the fit of the ECM equation for output at 19% is substantially better than the fit of the associated univariate AR(1) equation of only 5.4%. The ECM equations of the exchange rate variable (at 8.5%) also fits much better than the univariate equation (at 0%). By contrast the ECM equations for inflation and the real money balances are either worse or not that much better than the univariate alternatives. This seems to be largely due to the fact that the univariate specifications point to a higher order dynamics for these variables. Unfortunately the available data does not allow us to experiment with a VECX*(3,1) or VECX*(3,2) specifications that might be needed to accommodate such higher order dynamics.

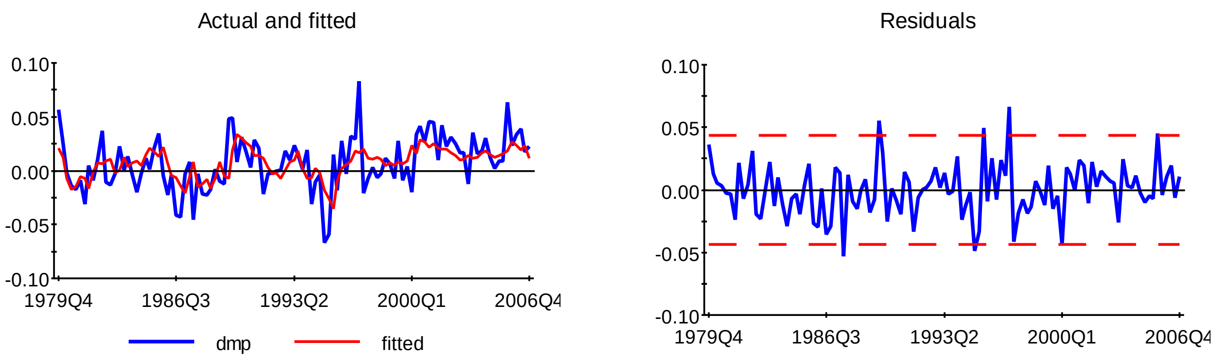
The actual and fitted values for each of the four equations together with the associated residuals are displayed in Figure 9. We observe that while there are some large outliers, especially for the exchange rate equation in the mid 1980's and the beginning of the 1990's and for output and real money in the early 1990's, the fitted values seem to track the main movements of the dependent variables reasonably well. The presence of large outliers are reflected in the massive rejection of the normality of the errors in the case of the real exchange rate equation.

Figure 9: Actual, fitted, and residuals for the core equations

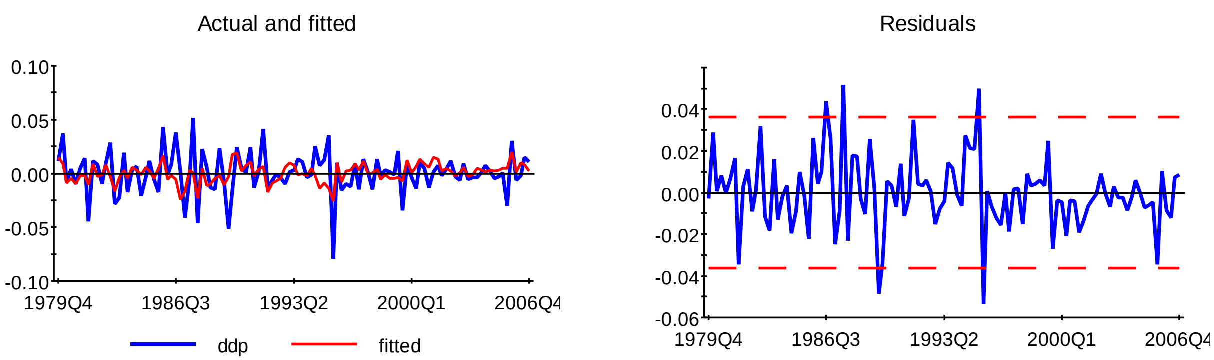
(a) Output equation



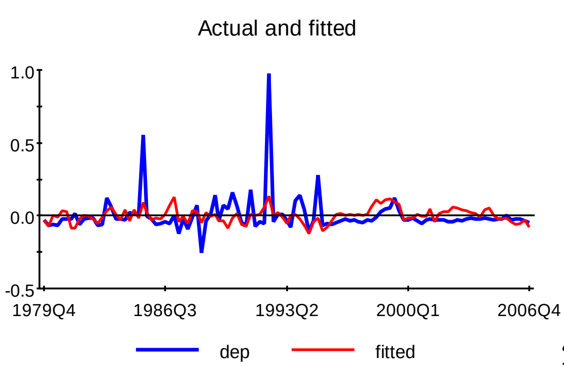
(b) Real money demand equation

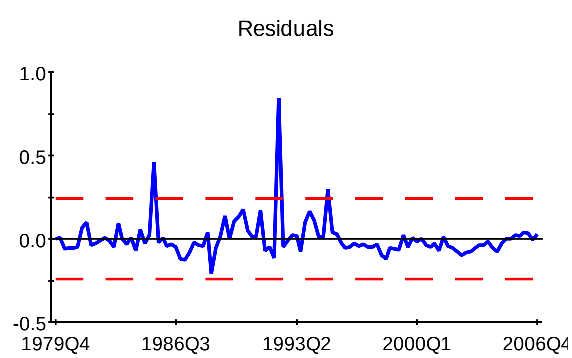


(c) Inflation equation



(d) Real exchange rate equation





6 Concluding Remarks

This paper, using a new quarterly data set on the Iranian economy over the period 1979Q1-2006Q4, provides a small quarterly model of the Iranian economy, where the long-run implications of oil exports for real output, inflation, real money balances, and the real exchange rate are tested. The results are generally supportive of the long-run theory developed in Esfahani et al. (2012) for a major oil exporting economy, although they also point to certain inefficiencies in the demand management of the economy that manifest themselves as negative long-run effects of inflation on real output and investment.

The estimates also suggest a rather rapid response of the economy to shocks, which could be due to the relatively underdeveloped nature of the money and capital markets in Iran. Such markets tend to act as shock absorbers in developed economies during normal conditions, although as we have seen recently, they can also act as shock magnifiers during crisis periods. The recently created National Development Fund could, in principle, if used appropriately act as a shock absorber which might lead to a more sluggish response of the economy to shocks.

The research in this paper can be extended in a number of directions. The current VECX* model is connected to the rest of the world through oil exports and foreign real output. Although these are clearly the most important channels of the transmission of shocks to the Iranian economy, there could be others. It would be interesting to see if the model can be linked to the global model recently developed in Dees et al. (2007), where the differential effects of supply and demand shocks and different regional shocks on the Iranian economy could be investigated.

Appendix A: Sources and Construction of the Data

Domestic and Foreign Data Series

Our data set contains quarterly observations on Iran and another 33 countries, from the first quarter of 1979 to the fourth quarter of 2006. The domestic variables included are (log) real output, y_t, (log) real money supply, mp_t, (log) price level, p_t, the rate of inflation, $\pi_t = p_t - p_{t-1}$, and (log) nominal exchange rate, e_t. Specifically

$$\begin{aligned} y_t &= \ln(GDP_t/CPI_t), \quad mp_t = \log(M_t/CPI_t), \\ e_t &= \ln(E_t), \qquad\qquad\quad p_t = \ln(CPI_t), \end{aligned} \tag{14}$$

where GDP_t is the nominal Gross Domestic Product, M_t is a broad liquidity measure that includes M1 and Quasi Money, CPI_t is the consumer price index, and E_t is the number of domestic currency (rials) per one US dollar exchanged on 'free' markets.

The two exogenous variables in the model are foreign output, y_t^*, and oil income in US dollars defined as $xo_t = \ln(P_t^o X_t^o)$, where P_t^o is the nominal price of oil per barrel in US dollars, and X_t^o is the domestic oil export in thousands of barrels per day. Foreign output was computed as the trade weighted average of log real output indices (y_{jt}) of Iran's trading partners:

$$\begin{aligned} y_t^* &= \sum_{j=1}^{33} \bar{\omega}_{j,t-1} y_{jt}, \text{ time varying weights,} \\ y_{t,FW}^* &= \sum_{j=1}^{33} \omega_{j,2001-03} y_{jt}, \text{ fixed weights,} \end{aligned}$$

where $\omega_{j,2001-03}$ and $\bar{\omega}_{j,t-1}$ are defined below by (15) and (16). The countries included in these weighted averages are: Argentina, Australia, Austria, Belgium, Brazil, Canada, China, Chile, Finland, France, Germany, India, Indonesia, Italy, Japan, Korea, Malaysia, Mexico, Netherlands, Norway, New Zealand, Peru, Philippines, South Africa, Saudi Arabia, Singapore, Spain, Sweden, Switzerland, Thailand, Turkey, United Kingdom, and United States.

The trade weights are computed based on the IMF Direction of Trade Statistics between 1980 and 2006. The bilateral trade of Iran with country j during a given year t, denoted by T_{jt}, is calculated as the average of exports and imports of Iran with that country. Trade data for Belgium is only available from 1997, and so the trade shares for Belgium between 1980 and 1996 was calculated by using the data on Belgium-Luxembourg and multiplying it by 0.93 (this procedure was also adopted in Dees et al. (2007)). In addition, trade data between South Africa and partner countries are only available from 1998, and so the data of all trading partners with South Africa was used to construct the South African trade shares with partners between 1980 and 1997.

The fixed trade weights were computed over the period 2001-2003 and are given by

$$\omega_{j,2001-03} = \frac{T_{j,2001} + T_{j,2002} + T_{j,2003}}{T_{2001} + T_{2002} + T_{2003}}, \tag{15}$$

where $T_t = \sum_{j=1}^{N} T_{jt}$, for $t = 2001, 2002, 2003$. The time varying trade weights are computed as

$$\bar{\omega}_{jt} = \frac{T_{j,t} + T_{j,t-1} + T_{j,t-2}}{T_t + T_{t-1} + T_{t-2}} \tag{16}$$

We also considered three year moving averages of the the annual trade shares, $\omega_{jt} = T_{jt}/T_t$ and obtained very similar results.

The most important trading partner for Iran is Japan, which accounts for between 15 and 20 percent of the total Iranian trade. More than 40% of the Iranian trade originates in or is destined for the euro area economies with Germany, Italy and France being Iran's most important trading partners in Europe.[12] Trade with China has increased significantly over the past two decades, emphasising the shift in the Iranian trade from the west to the east. Other countries in our data set with whom Iran's total trade is more than 5% are UK, Korea, and Turkey.

Data Sources

Real Output

The main source of data on Iran's real output is the Central Bank of the Islamic Republic of Iran (CBI) online database: Economic Time Series Database (http://tsd.cbi.ir/). Quarterly observations are available from 1988Q2 while annual data is available from 1959. We seasonally adjust the quarterly data using the U.S. Census Bureau's X-12 ARIMA seasonal adjustment program.[13] Quarterly series were interpolated (backwards) linearly from the annual series using the same method as that applied by Dees et al. (2007) to data for a number of the 33 countries in their data set. This data source is also updated to the end of 2006 and used for the computation of the foreign output variable described above. For a description of the interpolation procedure see Dees et al. (2007) Section 1.1 of Supplement A.

Consumer Price Indices

The CBI online database contains annual CPI data from 1959 and quarterly data from 1990Q2. To complete the quarterly data series we make use of several volumes of the CBI's Economic Report and Balance Sheets. We first use the 1981, 1987, and the 1989 Economic Report and Balance Sheets to compute quarterly data between 1976 and 1989 from the monthly data available in these reports. We then obtain quarterly CPI series by splicing the three series such that our quarterly CPI data stretches from 1976Q2 to 2007Q1, setting the average value of the index for 2000 equal to 100. Finally, we seasonally adjust the quarterly data using the U.S. Census Bureau's X-12 ARIMA seasonal adjustment program.

[12] When computing the trade weights, and thus the foreign variables, we aggregate Austria, Belgium, Finland, France, Germany, Italy, Netherlands, and Spain as the euro countries and so use their combined trade weight and output.

[13] For further information see U.S. Census Bureau (2007): X-12-ARIMA Reference Manual at http://www.census.gov/srd/www/x12a/

Exchange Rates

We obtain the official exchange rate series from the CBI online database. This data is available from 1959Q2. The nominal 'market' or 'free' exchange rate series used is from the IMF INS database and is available from 1979Q1.

Money and Quasi Money

The data on money and quasi money supply are from the IMF IFS series 34 and 35 and are 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. Quasi money data was missing for 1960Q4, 1978Q4, and between 1985Q2 to 1986Q2. Again we filled in for the missing data by splicing the IFS and the CBI data, but as CBI data was only available from 1974Q1, the complete series for quasi money is available only from 1961Q1.

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.

Oil Exports and Prices

Annual and quarterly oil export series (thousand barrels per day) are available, from 1973 and 1978Q2 respectively, from the CBI online database. Quarterly crude oil production data is also available from the CBI online database. Quarterly nominal oil prices were obtained from monthly averages of the Brent crude series from Datastream.

Data on value added of oil group, Gross Domestic Product at Basic Prices, and Non-Oil Gross Domestic Product at Basic Prices are available annually from 1959 and quarterly from 1988Q2. We first seasonally adjust the quarterly data and then obtain quarterly series from 1959Q2 by linearly interpolating (backwards) the 'missing' quarterly series from the annual series.

Population

Annual data on population was obtained from the IMF IFS series 99. This data was available from 1948. As quarterly data on population were not available, quarterly series were interpolated linearly from the annual series using the same method used to generate quarterly output series described above.

Conversion from Iranian to Gregorian Years

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. More complex ways of calculating this, such as taking into account exact number of days in the Iranian Quarter and converting the data was also investigated, but there were essentially no differences in the series.

Appendix B: Unit Root Test Results

For interpretation of the long-run relations and also to ensure that we do not work with a mixture of $I(1)$ and $I(2)$ variables we need to consider the unit root properties of the core variables in our model $(y_t, mp_t, \pi_t, ep_t, y_t^*, xo_t)$. Table 8 reports the standard Augmented Dickey-Fuller (ADF) test. But as the power of unit root tests are often low we also report the generalized least squares version of the Dickey-Fuller test (ADF-GLS) proposed by Elliott et al. (1996), and the weighted symmetric ADF test (ADF-WS) of Park and Fuller (1995), as they both have been shown to have better power properties than the ADF test.

Table 8: Unit root test statistics (based on AIC order selection)

Unit root test statistics for the levels									
	y_t	p_t	e_t	ep_t	mp_t	y_t^*	xo_t	CV	$CV\ T$
ADF	-2.67	-2.01	-0.79	-1.43	-0.19	-1.80	-2.50	-2.89	-3.45
ADF-GLS	-1.29	-2.01	-1.08	-1.37	-0.68	-2.02	-1.10	-2.14	-3.03
ADF-WS	-1.48	-2.06	-1.02	-1.72	-0.59	-2.12	-3.09	-2.55	-3.24
Unit root test statistics for the first differences									
	Δy_t	Δp_t	Δe_t	Δep_t	Δmp_t	Δy_t^*	Δxo_t	CV	$CV\ T$
ADF	-8.42	-3.82	-10.35	-10.45	-4.31	-3.37	-8.36	-2.89	-3.45
ADF-GLS	-7.66	-2.96	-9.79	-10.29	-2.94	-1.94	-1.09	-2.14	-3.03
ADF-WS	-8.16	-4.04	-10.63	-10.73	-4.29	-3.64	-6.08	-2.55	-3.24
Unit root test statistics for the second differences									
	$\Delta^2 y_t$	$\Delta^2 p_t$	$\Delta^2 e_t$	$\Delta^2 ep_t$	$\Delta^2 mp_t$	$\Delta^2 y_t^*$	$\Delta^2 xo_t$	CV	$CV\ T$
ADF	-9.07	-5.95	-7.78	-7.55	-6.90	-8.79	-10.12	-2.89	-3.45
ADF-GLS	-3.72	-5.43	-7.82	-7.55	-2.10	-1.50	-0.28	-2.14	-3.03
ADF-WS	-9.38	-5.84	-8.08	-7.82	-6.77	-8.69	-6.33	-2.55	-3.24

Notes: ADF denotes the Augmented Dickey-Fuller Test, ADF-GLS the generalized least squares version of the ADF test, and ADF-WS the weighted least squares ADF test. The sample period runs from 1979Q1 to 2006Q4. CV T gives the 95% simulated critical values for the test with intercept and trend, while C is the 95% simulated critical values for the test including an intercept only.

It is clear from Figures 1a to 1e that most of the core variables are trended and so we will include a linear trend and an intercept in the ADF regressions for all the variables except for the ep_t and xo_t series which do not seem to have a trend. When testing for the presence of unit roots in the first and second differences of the core variables only an intercept is included in the ADF regressions. As can be seen from Table 8, both the ADF and ADF-WS tests provide strong support that y_t, mp_t, π_t, ep_t, y_t^*, and xo_t are all $I(1)$, as the unit root hypothesis is clearly rejected when applied to the first differences of these variables, but not when the tests are applied to the (log) levels. The ADF-GLS test results are not as conclusive. But overall, as a first order approximation, the available evidence supports our treatment of the core variables as being $I(1)$.

References

Abadir, K., K. Hadri, and E. Tzavalis (1999). The Influence of VAR Dimensions on Estimator Biases. *Econometrica 67*, 163–181.

Aghion, P., P. Bacchetta, R. Rancière, and K. Rogoff (2009). Exchange Rate Volatility and Productivity Growth: The Role of Financial Development. *Journal of Monetary Economics 56*(4), 494–513.

Amuzegar, J. (1997). *Iran's Economy under the Islamic Republic*. I.B. Tauris.

Amuzegar, J. (2008). Iran's Oil as a Blessing and a Curse. *The Brown Journal of World Affairs 15*, 46–61.

Assenmacher-Wesche, K. and M. H. Pesaran (2009). A VECX* Model of the Swiss Economy. *Swiss National Bank Economic Studies No. 6*.

Baharie, N. (1973). *Economic Policy and Development Strategy in Iran: A Macro-Simulation Approach*. Ph. D. thesis, University of Illinois, Illinois.

Bahmani-Oskooee, M. (1995). Source of Inflation in Post-Revolutionary Iran. *International Economic Journal 9(2)*, 61–72.

Becker, T. (1999). Common Trends and Structural Change: A Dynamic Macro Model for the Pre- and Postrevolution Islamic Republic of Iran. *IMF Working Paper WP/99/82*.

Bonato, L. (2008). Money and Inflation in the Islamic Republic of Iran. *Review of Middle East Economics and Finance 4*(1), Article 3.

Cavalcanti, T. V. d. V., K. Mohaddes, and M. Raissi (2011a). Growth, Development and Natural Resources: New Evidence Using a Heterogeneous Panel Analysis. *The Quarterly Review of Economics and Finance 51*, 305–318.

Cavalcanti, T. V. d. V., K. Mohaddes, and M. Raissi (2011b). Does Oil Abundance Harm Growth? *Applied Economics Letters 18*(12), 1181–1184.

Cavalcanti, T. V. d. V., K. Mohaddes, and M. Raissi (2012). Commodity Price Volatility and the Sources of Growth. *IMF Working Paper WP/12/12*.

Celasun, O. and M. Goswami (2002). An Analysis of Money Demand and Inflation in the Islamic Republic of Iran. *IMF Working Paper WP/02/205*.

Corden, W. M. and J. P. Neary (1982). Booming Sector and De-Industrialisation in a Small Open Economy. *The Economic Journal 92*(368), 825–848.

Dees, S., F. di Mauro, M. H. Pesaran, and L. V. Smith (2007). Exploring the International Linkages of the Euro Area: A Global VAR Analysis. *Journal of Applied Econometrics 22*, 1–38.

Elliott, G., T. J. Rothenberg, and J. H. Stock (1996). Efficient Tests for an Autoregressive Unit Root. *Econometrica 64*, 813–836.

Elyasiani, E. and W. Zhao (2008). International Interdependence of an Emerging Market: The Case of Iran. *Applied Economics 40*, 395 – 412.

Esfahani, H. S., K. Mohaddes, and M. H. Pesaran (2012). An Empirical Growth Model for Major Oil Exporters. *CESifo Working Paper No 3780*.

Esfahani, H. S. and M. H. Pesaran (2009). The Iranian Economy in the Twentieth Century: A Global Perspective. *Iranian Studies 42*, 177 – 211.

Fischer, S. (1983). Inflation and Growth. *National Bureau of Economic Research Working Paper No. 1235*.

Fischer, S. (1993). The Role of Macroeconomic Factors in Growth. *Journal of Monetary Economics 32*(3), 485 – 512.

Garratt, T., K. Lee, M. H. Pesaran, and Y. Shin (2006). *Global and National Macroeconometric Modelling: A Long Run Structural Approach.* Oxford University Press, Oxford.

Gonzalo, J. (1994). Five Alternative Methods of Estimating Long-run Equilibrium Relationships. *Journal of Econometrics 60*(1-2), 203 – 233.

Gredenhoff, M. and T. Jacobson (2001). Bootstrap Testing Linear Restrictions on Cointegrating Vectors. *Journal of Business and Economic Statistics 19*, 63–72.

Habib-Agahi, H. (1971). A Macro Model for Iranian Economy: 1338-1349. *Plan and Budget Organization, Tehran, Iran*.

Haug, A. (1996). Tests for Cointegration A Monte Carlo Comparison. *Journal of Econometrics 71*, 89–115.

Heiat, A. (1986). *An Econometric Study of an Oil-Exporting Country: The Case of Iran.* Ph. D. thesis, Portland State University, Portland, Oregon.

Kia, A. (2006). Deficits, Debt Financing, Monetary Policy and Inflation in Developing Countries: Internal or External Factors?: Evidence from Iran. *Journal of Asian Economics 17*(5), 879 – 903.

Koop, G., M. H. Pesaran, and S. M. Potter (1996). Impulse Response Analysis in Nonlinear Multivariate Models. *Journal of Econometrics 74*, 119–147.

Krugman, P. (1987). The Narrow Moving Band, the Dutch Disease, and the Competitive Consequences of Mrs. Thatcher: Notes on Trade in the Presence of Dynamic Scale Economies. *Journal of Development Economics 27*(1-2), 41–55.

Lee, K. and M. H. Pesaran (1993). Persistence Profiles and Business Cycle Fluctuations in a Disaggregated Model of UK Output Growth. *Ricerche Economiche 47*, 293–322.

Liu, O. and O. Adedeji (2000). Determinants of Inflation in the Islamic Republic of Iran - A Macroeconomic Analysis. *IMF Working Paper No. 00/127*.

López-Villavicencio, A. and V. Mignon (2011). On the Impact of Inflation on Output Growth: Does the Level of Inflation Matter? *Journal of Macroeconomics 33*(3), 455 – 464.

Mehrara, M. and K. N. Oskoui (2007). The Sources of Macroeconomic Fluctuations in Oil Exporting Countries: A Comparative Study. *Economic Modelling 24*(3), 365 – 379.

Motaman, H. (1979). *Expenditure of Oil Revenue: An Optimal Control Approach with Application to the Iranian Economy*. Frances Pinter.

Neary, J. P. and S. J. G. van Wijnbergen (1986). *Natural Resources and the Macroeconomy*. MIT Press, Cambridge, MA.

Noferesti, M. and A. Arabmazar (1374, 1993). A Macroeconomic Model for the Iranian Economy (in Persian). *Journal of Research and Economic Policies 2(1)*, 5–39.

Pahlavani, M., E. Wilson, and A. Worthington (2005). Trade-GDP Nexus in Iran: An Application of the Autoregressive Distributed Lag (ARDL) Model. *Faculty of Commerce Papers 144, University of Wollongong*.

Park, H. and W. Fuller (1995). Alternative Estimators and Unit Root Tests for the Autoregressive Process. *Journal of Time Series Analysis 16*, 415–429.

PBO (1372, 1993). Documentations of the Second Economic, Social and Cultural Development Plan of the Islamic Republic of Iran (1373-1377), Volume 5, (Macroeconomic Models and Models for Prediction of Population). Technical report, Plan and Budget Organization, Iran.

Pesaran, M. H. (1984). Macroeconomic Policy in an Oil-Exporting Economy with Foreign Exchange Controls. *Economica 51*(203), 253–270.

Pesaran, M. H. (1992). The Iranian Foreign Exchange Policy and the Black Market for Dollars. *International Journal of Middle East Studies 24*(1), 101–125.

Pesaran, M. H. (2000). *Economic Trends and Macroeconomic Policies in Post-Revolutionary Iran in The Economy of Iran: Dilemmas of an Islamic State*. I.B. Tauris, London.

Pesaran, M. H. and B. Pesaran (2009). *Microfit 5.0: An Interactive Econometric Software Package*. Oxford University Press, Oxford.

Pesaran, M. H. and Y. Shin (1996). Cointegration and Speed of Convergence to Equilibrium. *Journal of Econometrics 71*, 117–143.

Pesaran, M. H. and Y. Shin (1998). Generalised Impulse Response Analysis in Linear Multivariate Models. *Economics Letters 58*, 17–29.

Pesaran, M. H., Y. Shin, and R. J. Smith (2000). Structural Analysis of Vector Error Correction Models with Exogenous I(1) Variables. *Journal of Econometrics 97*, 293–343.

Razavi, H. (1982). Optimal Rate of Oil Production for OPEC-member Countries. *Resources and Energy 4*(3), 291 – 305.

Razavi, H. (1983). An Analysis of Iran's Oil Production Policy: A Welfare Maximization Approach. *Applied Economics 15*(2), 243–254.

Sachs, J. D. and A. M. Warner (1995). Natural Resource Abundance and Economic Growth. *National Bureau of Economic Research Working Paper 5398.*

Safai, J. (1365, 1986). Development and Estimation of a Macro-econometric Model for the Iranian Economy. M.Sc. Thesis, Shiraz University (in Persian).

Shahshahani-Madani, A. (1357, 1978). Econometric Models for Iran and their Applications. *Publication of Tehran University (in Persian).*

Sims, C. (1980). Macroeconomics and Reality. *Econometrica 48*, 1–48.

Vakil, F. (1973). An Econometric Model for Iran - Estimated Structural Equations. *Bank Markazi of Iran Bulletin March-April Issue.*

Valadkhani, A. (1997). *A Macroeconometric Model for the Iranian Economy.* Ph. D. thesis, University of Queensland, Brisbane.

Valadkhani, A. (2006). Unemployment Conundrum in Iran. *University of Wollongong Economics Working Paper Series WP 06-15.*

van der Ploeg, F. and T. Venables (2009). Symposium on Resource Rich Economies Introduction. *Oxford Economic Papers 61*(4), 625–627.