Economics Bulletin

Volume 32, Issue 3

The Impact of Oil Price Shocks on the Iranian Economy: New evidence

Katsuya Ito
Fukuoka University

Abstract

In a cointegrated VAR model we examined the relationship between oil price and macroeconomy in Iran, which is the third largest oil exporter in the world. The sample is quarterly data, ranging from 1994:1 to 2007:4. We find that an increase in real oil prices by 1% is associated with a 0.30% increase in real GDP in the long run, whereas the short-run impact is marginal.

Citation: Katsuya Ito, (2012) "The Impact of Oil Price Shocks on the Iranian Economy: New evidence", *Economics Bulletin*, Vol. 32 No. 3 pp. 2246-2254.
Contact: Katsuya Ito - masakidaden@gmail.com.
Submitted: February 04, 2012. **Published:** August 07, 2012.

1. Introduction

This paper is intended as an empirical investigation of the link between oil prices and macroeconomic variables in the Islamic Republic of Iran (hereafter Iran), which is the third largest oil exporter in the world (IMF, 2011). As is intuitively expected, rising oil prices enables the country to experience economic growth through income transfer from oil-importing countries. It is of interest to assess to what extent an oil price shock affects the economy. Although the linkage has been widely researched since the early 1980s, only a few studies have so far been carried out on this issue for the Iranian economy, and only focused on short-run analysis in a vector autoregression (VAR) system. Farzanegan and Markwardt (2009) using the data 1989:1-2006:4 with an unrestricted level-VAR model argued that a positive (real) oil-price shock contributes to real GDP growth, ranging from 0.001 to 0.007%, until next six quarters. Furthermore, based on the impulse-response functions they demonstrated that the shock raises inflation rate, reaching its peak in the second quarter, and real effective exchange rate reacts negatively to the shock except for the fifth and sixth quarters. Our analysis is different from previous studies in that a cointegration technique and generalized impulse-response functions are used. To the best of our knowledge, this paper is the first attempt in the literature to analyze the long-run relationship between oil price and the Iranian economy.

The rest of the paper is organized as follows. Section 2 and 3 provide the empirical framework and results, respectively. Section 4 concludes.

2. Methodology and Data

2.1 Econometric methodology

In our analysis, a cointegrated VAR model is employed. The model developed by Johansen (1988) has the following form:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \Gamma_2 \Delta Z_{t-2} + \cdots + \Gamma_{k-1} \Delta Z_{t-k-1} + \Pi Z_{t-1} + u_t \qquad [1]$$

where Δ is the difference operator, Z_t is an ($n \times 1$) vector of variables, k is the number of lags, u_t is an ($n \times 1$) vector of error terms for $t = 1,2,\cdots,T.$ Additionally, u_t is an independently and identically distributed with zero mean, *i.e.* $E(u_t) = 0$ and an ($n \times n$) symmetric variance-covariance matrix Σ, *i.e.* $E(u_t u_t') = \Sigma.$ Γ denotes an ($n \times n$) matrix of coefficients and contains information regarding the short-run relationships among the variables. The matrix Π contains information regarding the long-run relationships, and is an ($n \times n$) coefficient matrix decomposed as $\Pi = \alpha\beta'$, where α and β are ($n \times r$) adjustment and cointegration matrices, respectively.

To determine the number of cointegrating vectors the co-integration trace (λ_{trace})

test by Johansen (1988, 1991 and 1995) and Johansen and Juselius (1990) is used. The test is given by the following statistic:

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{n} \log(1-\lambda_{r+1}) \qquad [2]$$

where r is the number of cointegrating relations, and n is the number of variables. The null hypothesis is that the number of cointegrating vectors is less than or equal to r against the alternative hypothesis of $r > 0$.

2.2 Data

The data used are as follows: global oil consumption (*goc*); inflation rate (*ifr*) as measured by the percentage changes of consumer price index (*cpi*, 2005=100); net foreign assets/GDP ratio (*nfa/gdp*) (as a proxy for current account balance/GDP ratio or foreign reserves/GDP ratio); real effective exchange rate (*reer*, 2005=100); real GDP (*rgdp*, 2005=100) defined as the nominal GDP deflated by the *cpi*; and real oil price (*rop*) defined as Brent oil price in U.S. dollars converted (by the period average exchange rate) to Iranian rial per barrel deflated by the *cpi*. The data for the first variable are obtained from U.S. Energy Information Administration, *International Energy Statistics*, and the rest from International Monetary Fund, *International Financial Statistics*. The sample ranges from 1994:1 to 2007:4, for a total of 56 observations. The time series are transformed into logarithms, and were seasonally adjusted via the CensusX12-ARIMA (Autoregressive Integrated Moving Average) with the exception of the *ifr* and *reer*.

3. Empirical results

3.1. Unit root tests

To test for a unit root (non-stationarity) in our data series, the augmented Dickey-Fuller (ADF) test (1979) is adopted. Considering the low power of the ADF test in small sample size we also use the Phillips-Perron (PP) test (1988), which takes account of the serial correlation and heteroscedasticity, as an alternative test. The critical values for the ADF and PP *t*-statistics are obtained from the MacKinnon (1996) table. Table 1 shows results of unit root tests for six variables, and indicates (except for the *rgdp* with intercept and trend) that the null hypothesis is not rejected at the 5% level of significance when the variables are defined in log levels. The results of the ADF test suggest that the hypothesis of a unit root for $\Delta rgdp$ cannot be rejected at the 5% significance level, whereas those of the PP test indicate that the series are first-difference stationary at the preceding level. We adopt the results from the PP test, implying that all variables are integrated of the same order, *i.e.* $I(1)$.

Table 1. Unit root tests

	ADF		PP	
Variable	Intercept	Intercept and Trend	Intercept	Intercept and Trend
goc (log)	-1.122	-3.478	-1.305	-2.448
⊿*goc* (log)	-3.682**	-3.860*	-5.875**	-5.882**
ifr (log)	-2.373	-2.025	-1.738	-1.958
⊿*ifr* (log)	-4.099**	-4.112**	-5.896**	-5.918**
nfa/gdp (log)	-0.989	-1.603	-1.042	-1.650
⊿*nfa/gdp* (log)	-6.894**	-6.862**	-6.894**	-6.862**
reer (log)	-2.103	-2.949	-0.955	-2.763
⊿*reer* (log)	-3.659**	-3.697*	-6.150**	-6.080**
rgdp (log)	0.854	-2.066	-1.249	-5.171**
⊿*rgdp* (log)	-2.446	-2.794	-11.973**	-16.452**
rop (log)	-0.306	-1.914	-0.562	-1.955
⊿*rop* (log)	-3.067*	-4.152**	-4.557**	-4.899**

Notes: (1) ⊿ means 1st difference. (2) *and ** refer to the rejection of the null hypothesis at the 5% and 1% significant level, respectively. (2) The critical values at 1% level (at 5% level) are -3.56 (-2.92) with intercept and -4.14 (-3.50) with intercept and trend, respectively. The critical values are from MacKinnon (1996). The calculation of critical values is as follows: $\hat{u}(p) = \delta_0 + \delta_1\Phi^{-1}(p) + \delta_2 (\Phi^{-1}(p))^2 + \delta_3 (\Phi^{-1}(p))^3 + e^*_p$, where $\Phi^{-1}(p)$ denotes the inverse of the cumulative standard normal distribution function, evaluated at p with $0 < p < 1$. (3) The ADF test are given by: $\Delta y_t = \alpha_0 + \gamma y_{t-1} + \Sigma_{i=1}^{p} \beta_i \Delta y_{t-i} + u_t$ (4) The test regression for the PP test is the AR(1) process: $\Delta y_{t-1} = \alpha_0 + \gamma y_{t-1} + e_t$.

3.2. Cointegration tests

The results of the cointegration tests based on trace statistics are presented in Table 2. The number of lags was chosen to be 4 by the Akaike information criterion (AIC) (Akaike 1974). The critical values for the trace statistic are obtained from MacKinnon *et al.* (1999). In line with Pantula's (1989) principle we select the Model that the null hypothesis cannot be rejected for the first time. It is observed that Model 3 and 4 cannot reject the null hypothesis of $r = 3$ at the 5% level of significance. This provides a strong evidence of existing three cointegrating equilibrium relationships among the six variables. On the basis of the assumption that there are no long-run trends among the variables we apply Model 3 (with intercept and no trend in the long-run and the short-run models).

Table 2. Cointegration tests

No. of CE(s)		Model 2	Model 3	Model 4
H_0	H_1			
r=0	r=1	211.236*	165.741*	189.015*

		(103.847)	(95.753)	(117.708)
		[0.000]	[0.000]	[0.000]
r=1	r=2	134.776*	102.105*	125.292*
		(76.972)	(69.818)	(88.803)
		[0.000]	[0.000]	[0.000]
r=2	r=3	72.277*	59.042*	75.788*
		(54.079)	(47.856)	(63.876)
		[0.001]	[0.003]	[0.003]
r=3	r=4	20.4713*	29.445	41.216
		(35.192)	(29.797)	(42.915)
		[0.007]	[0.055]	[0.073]

Notes: (1) CE(s) refers to the cointegrating equation(s). (2) * denotes rejection of the hypothesis at the 5% significance level. (3) The lag length, which was determined by the AIC, was 4 lags. The AIC is defined as follows: $AIC = -2l/T + 2k/T$, where l is the log likelihood, k is the number of parameters and T is the number of observations. (4) Sample periods (adjusted) are from 1995:2 to 2007:4. (5) The values of round brackets and square brackets refer to critical values and p-values, respectively, based on MacKinnon et al. (1999). (6) Model 1: No intercept or trend in CE or VAR; Model 2: Intercept (no trend) in CE, and no intercept in VAR; Model 3: Intercept (no trend) in CE and VAR; Model 4: Intercept and trend in CE, and no trend in VAR; Model 5: Intercept and trend in CE, and linear trend in VAR. In general, the model 1 and model 5 are considered as rare cases. (7) The VAR model proposed by Sims (1980) can be written as follows:

$$Z_t = A_1Z_{t-1} + A_2Z_{t-2} + \cdots + A_pZ_{t-p} + u_t$$

where A is an (n x n) matrix of coefficients, p is the number of lags, u is an (n x 1) vector of error terms for t = 1, 2, …T.

We impose restrictions on each cointegrating vector. As a result, the hypothesis was not rejected with a p-value of 0.183 (Chi-square(4)=6.212), and the estimate of β' for Z_t = [*goc*, *ifr*, *nfa/gdp*, *reer*, *rgdp*, *rop*] is given by

$$\beta' = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & \underset{(0.006)}{-0.047} \\ 0 & \underset{(0.042)}{0.473} & \underset{(0.029)}{-0.133} & 1 & 0 & \underset{(0.042)}{0.088} \\ 0 & 0 & 0 & 0 & 1 & \underset{(0.019)}{-0.309} \end{bmatrix},$$

where the figures in parenthesis denote standard errors. The coefficients are statistically significant. The first cointegrating vector suggests that a 1% rise (fall) in real oil prices is positively (negatively) associated with global oil consumption by 0.04% in the long run. Given that global oil production is roughly equal to global oil consumption, this may be explained by the fact that rising oil prices induces an increase in oil production. The second vector represents that a 1% increase (decrease) in inflation rate and real oil prices leads to a depreciation (appreciation) of the real effective exchange rate, whereas that of the net foreign assets to GDP ratio appreciates (depreciates) the exchange rate. The third vector indicates that a 1% rise (drop) in the real price of oil contributes to the GDP growth (decline) by 0.30%.

Next, we conduct a test for misspecification in order to ascertain whether or not the model is appropriate. Thus we apply the Lagrange Multiplier (LM) test for autocorrelation presented by Breusch and Godfrey (1981). The results of the test are shown in Table 3, suggesting that there is no residual autocorrelation in the model since all p-values are greater than the 5% level of significance.

Table 3. LM test

Autocorrelation LM test	Lags	P-value
	1	0.28
	2	0.14
	3	0.65
	4	0.61

Notes: (1) Sample periods are from 1994:1 to 2007:4. (2) Probabilities are from chi-square with 36 degrees of freedom. (3) The LM test is defined as follows:

$\hat{u}_t = \alpha_0 + \alpha_1 X_{2t} \ldots \alpha_R X_{Rt} + \alpha_{R+1}\hat{u}_{t-1} \ldots \alpha_{R+p}\hat{u}_{t-p}$

LM statistic = $(n\text{-}p)R^2 \sim \chi^2_p$

where p is the number of lags used (: the degrees of freedom). If the LM statistic is larger than the χ^2_p critical value for a given significance level, the null of serial correlation can be rejected.

3.3 Generalized impulse-response functions

We employ impulse response functions (IRFs), which trace the impact of a one-standard-deviation shock in a variable on current and future values of the variables, in order to capture the short-run dynamics of the model. Considering that the IRFs based on a Cholesky decomposition is sensitive to the ordering of the variables, we apply generalized impulse-response functions (GIRFs) proposed by Pesaran and Shin (1998).

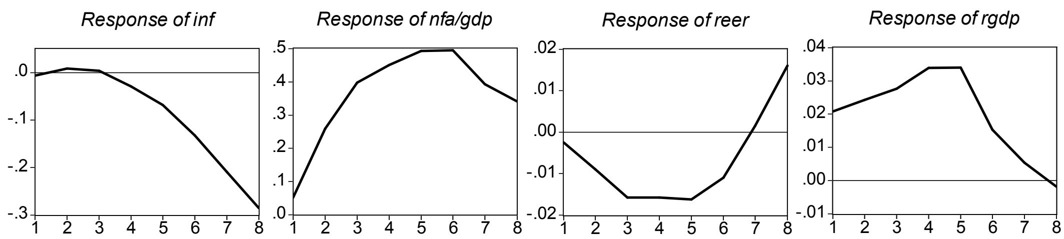
Table 4, Figure 1 and 2 show accumulated responses to a positive shock of oil price and net foreign assets to GDP ratio, respectively. These results suggest, as a whole, that rising oil prices are negatively associated with inflation rate in the short run. The response becomes positive in the second quarter, but remains negative from the fourth quarter. We observe that the oil shock has a positive impact on the ratio of net foreign assets to GDP. The response is estimated to be 0.34% over the next eight quarters. At the same time, despite a massive inflow of foreign currency earnings causing the appreciation of the Iranian rial, the oil shock leads to a decrease in real effective exchange rate until the sixth quarter. This phenomenon may be explained by the fact that the price level in Iran is relatively lower than that of its trading partners. Likewise, the response of real GDP to the shock is slightly positive and becomes negative after the fifth quarter, but not significant. The response is estimated to be marginally negative 0.001% in the eighth quarter. With respect to the net foreign assets/GDP shock, it is observed that inflation rate responds negatively, whereas the rest reacts positively. The response of real effective exchange rate to the shock would be interpreted as reflecting the long-run relationship between them. The magnitude of the inflation response is significant and estimated to be negative 0.59% over the entire horizon considered.

Table 4. Accumulated response to a positive shock of oil price and net foreign assets/GDP ratio

	Oil price shock				Net foreign assets/GDP shock			
Period	*ifr*	*nfa/gdp*	*reer*	*rgdp*	*ifr*	*nfa/gdp*	*reer*	*rgdp*
1	-0.0067	0.0518	-0.0024	0.0208	-0.0270	0.2393	0.0119	0.0152
2	0.0082	0.2594	-0.0090	0.0242	-0.0630	0.4629	0.0299	0.0101
3	0.0035	0.3978	-0.0157	0.0276	-0.1506	0.6295	0.0544	0.0241
4	-0.0289	0.4504	-0.0157	0.0338	-0.2484	0.7648	0.0918	0.0292
5	-0.0681	0.4928	-0.0161	0.0339	-0.3497	0.9998	0.1381	0.0399
6	-0.1323	0.4949	-0.0110	0.0153	-0.4507	1.1974	0.1898	0.0340
7	-0.2095	0.3927	0.0015	0.0053	-0.5305	1.3602	0.2430	0.0428
8	-0.2854	0.3404	0.0160	-0.0017	-0.5908	1.5062	0.2990	0.0343

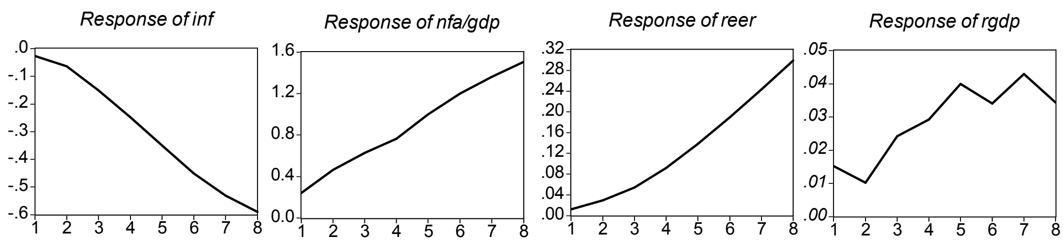
Notes: Sample periods are from 1994:1 to 2007:4 with 4 lags and three restricted co-integrating vectors. Accumulated responses to generalized one standard-deviation innovation for up to 8 quarters ahead are presented.

Figure.1 Accumulated response to a positive oil price shock



Notes: Sample periods are from 1994:1 to 2007:4 with 4 lags and three restricted co-integrating vectors. Accumulated responses to generalized one standard-deviation innovation for up to 8 quarters ahead are plotted. The horizontal axis is time (quarter), and the vertical axis is the magnitude of the response to the impulse (%).

Figure.2 Accumulated response to a positive shock of net foreign assets/GDP ratio



Note: Same as Figure 1.

4. Conclusions

In this paper, we have attempted to empirically investigate the oil price-macroeconomy nexus in Iran. Based on a cointegrated VAR framework we found that an oil price rise (drop) leads to the Iranian economic growth (decline) by 0.30% in the long run, whereas the contribution is insignificant in the short run. It was also found that the net foreign assets/GDP ratio is positively associated with the real effective exchange rate. At the same time, we see that in the short run a rise in the ratio of net foreign assets/GDP contributes to the decrease in the rate of inflation.

Overall, our findings suggest that net foreign assets (or current account surplus) have played a crucial role in the Iranian economy because it appreciates the real effective exchange rate, leading to a fall in the domestic price level through cheaper imports. Given the significance of oil to total exports, it seems reasonable to suppose that Iran does not suffer from slowdown in the manufacturing sector caused by real exchange rate appreciation (: Dutch disease). However, taking account of its sensitivity to oil price changes, Iran needs an industrial diversification away from natural resource-oriented economy over the long run.

Notwithstanding its data limitation, this paper contributes to the literature by offering some insight into oil price-the Iranian economy nexus.

References

Akaike, H. (1974) "A new look at statistical model identification" *IEEE Transactions on Automatic Control* **19**, 716-23.

Breusch, T. and L.G. Godfrey (1981) "A review of recent on testing for autocorrelation in dynamic simultaneous models" in *Macroeconomic analysis: essays in macroeconomics and econometrics* by D. Currie, R. Nobay and D. Peel, Eds., Croon Helm: London, 63-105

Dickey, D.A. and A. Fuller (1979) "Distribution of the estimators for autoregressive time series with a unit root" *Journal of American Statistical Association* **74**, 15-49.

Farzanegan, M.R. and G. Markwardt (2009) "The effects of oil price shocks on the Iranian economy" *Energy Economics* **31**, 134-51.

International Monetary Fund (2011) *World Economic Outlook September 2011*, IMF: Washington.

Johansen, S. (1988) "Statistical analysis of cointegrating vectors" *Journal of Economic Dynamics and Control* **12**, 231-54.

Johansen, S. (1991) "Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models" *Econometrica* **59**, 1551-80.

Johansen, S. (1995) *Likelihood-based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press: Oxford.

Johansen, S. and K. Juselius (1990) "The maximum likelihood estimation and inference on cointegration – with application to demand for money" *Oxford Bulletin of Economics and Statistics* **52**, 169-210.

MacKinnon, J. (1996) "Numerical distribution functions for unit root and cointegration tests" *Journal of Applied Econometrics* **11**, 601-18.

MacKinnon, J.G., M.A. Haug and L. Michelis (1999) "Numerical distribution functions of likelihood ratio tests for cointegration" *Journal of Applied Econometrics* **14**, 563-77.

Pantula, S.G. (1989) "Testing for unit roots in time series data" *Econometric Theory* **5**, 256-71.

Pesaran, H.H. and Y. Shin (1998) "Generalized impulse response analysis in linear multivariate models" *Economics Letters* **58**, 17-29.

Phillips, P.C.B. and P. Peron (1988) "Testing for a unit root in time series regression" *Biometrika* **75**, 335-46.

Sims, C.A. (1980) "Macroeconomics and Reality" *Econometrica* **48**, 1-48.