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The Impacts of Energy Sanctions on the Black-Market Premium: Evidence from Iran

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Abstract

This study extends the literature by analyzing the effects of energy sanctions on the black-market premium on the Iranian Rial – U.S. dollar exchange rate. Using a nonlinear autoregressive distributed lag (NARDL) model from 1959 to 2017, we find that falling oil revenues as a result of sanctions have an increasing impact on the black-market premium.

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1. Introduction

There is a growing interest in understanding the potential effects of (economic) sanctions on the political and socio-economic structures of the targeted country. The case study of Iran is particularly interesting due to the amplified conflict between the US and Iran following the election of Donald Trump as President of the United States in 2016. Iran has been under different forms of unilateral and multilateral political and economic sanctions since the 1979 Islamic revolution (see Dizaji and Farzanegan, 2019, and Farzanegan, 2013, for more details).

The foreign exchange revenues of Iran are highly dependent on oil exports (Farzanegan and Markwardt, 2009), which are significantly affected by US sanctions. Crude oil exports of Iran were around 2.4 million barrels per day (b/d) from 2000 to 2011; thereafter, oil exports dropped to about 1 million b/d in 2014 (IMF, 2020). Since the withdrawal of the US from the Joint Comprehensive Plan of Action and reactivation of economic and oil sanctions in 2018, oil exports are down 500,000 b/d to date (IMF, 2020). This significant drop in oil exports has major economic implications for Iran. Several studies have examined such implications (e.g., Dizaji and van Bergeijk, 2013; Farzanegan and Hayo, 2019; Farzanegan, 2019, and Khabbazan and Farzanegan, 2016). We follow these studies and focus on the effects of sanctions on the black-market premiums for the exchange rate between the US dollar and the Iranian Rial (BMPER).

An increasing BMPER has significant impacts on various aspects of the Iranian economy (e.g., on illicit trade as is shown by Farzanegan, 2009). Ishak and Farzanegan (2020) also show that the shadow economy which includes transactions in the black market of foreign exchange plays an important role in shaping the final effect of declining oil revenues on the taxation efforts of the government. The present analysis contributes to the literature, specifically on the black-market premium (e.g., Bahmani-Oskooee and Tankui, 2008), by shedding light on the effects of sanctions on the gap between formal and informal exchange rates for the US dollar in Iran. Using a nonlinear autoregressive distributed lag model, we examine the impacts of decreased oil revenue on the BMPER in the long- and short-term. Since international sanctions on Iran's energy sector will lead to negative shocks in oil revenue, we use negative changes in oil revenues as an empirical proxy of (energy) sanctions' impact (for a similar approach see Farzanegan, 2011; Farzanegan et al., 2016; Farzanegan and Raeisian Parvari, 2014, and Khabbazan and Farzanegan, 2016).

Oil revenue is the main source of foreign exchange for the Iranian economy. According to the Organization of the Petroleum Exporting Countries (OPEC) data, the average share of oil export revenues in total export revenues of Iran over the 1960 – 2019 period was approximately 80%. Declining oil income due to international sanctions not only prevents the government from accessing its foreign exchange assets abroad (because of banking sanctions and freezing of assets), but also leads to substantial decrement in oil production and revenue (Katzman, 2020). Consequently, the Iranian government faces scarcity of foreign currencies to meet the needs of importers, producers and individuals. To mitigate the negative externalities of the sanctions, normally governments allocate the reduced and limited foreign currencies at a formal (subsidized) rate to cover imports of strategic goods including agri-food products and medicines. The other needs for foreign exchange should then be covered at market rate (e.g., imports of machineries, equipment, consumer goods which are not classified as necessary, etc.). There will be an increasing pressure on demand for foreign exchange in the free market which increases the gap between formal and informal exchange rates. Farzanegan (2013) provides further descriptive information on how the Iranian government organized the rationing of hard currencies under the sanctions[1].

2. Data model and framework

We employ the most recent data obtained from the Central Bank of Iran, covering 59 annual observations for the period from 1959 to 2017. We follow Pesaran (1992), Kiguel and O'Connell (1995), Rodrick (2008), and Hebous (2011) for our model specification. The BMPER depends on oil export revenues (OILV), gross fixed capital formation (GFCF), government expenditure (GVEX), and the balance of trade (BOT). All variables in our analysis are measured in constant prices. Additionally, different dummy variables as deterministic components are introduced to control for the Iranian revolution, the Iran-Iraq war, and the major financial crises and policy reforms, respectively.

To explain the factors affecting BMPER, we should note that BMPER is composed of two variables; official and non-official exchange rates. The BMPER is defined as in Kiguel and O'Connell (1995):

$$\mathrm{BMERP_t} = \frac{(\mathrm{NOEX_t} - \mathrm{OEX_t})}{\mathrm{OEX_t}} \times 100 \quad (1)$$

where $\mathrm{OEX_t}$ and $\mathrm{NOEX_t}$ denote official and non-official exchange rates at time *t* respectively. The non-official exchange rate is the free market rate, while the official exchange rate is the government-subsidized rate, which is normally below the free market rate. We use both the official and non-official exchange rates for US dollar as published in the Economic Time Series Database of the Central Bank of Iran. The official exchange rate is the rate which the government uses to convert its foreign exchange revenues to Rial[2], and it is distributed by the Central Bank of Iran for specific purposes like importing strategic commodities. However, the non-official exchange rate is the rate which the private sector uses to purchase the foreign currencies from the open market. The non-official exchange rate is influenced by economic variables and market mechanism, while the official rate is directly regulated by the government. However, the macroeconomic conditions and external shocks may impose pressure on the government to revise its policies and hence the official rate is affected indirectly by the market. Oil export revenues are a fully government-owned source of foreign exchange supply (Farzanegan and Habibpour, 2017). Furthermore, we should also note that the foreign exchange needs of the government are usually met at the official rate, funded mainly by oil export revenues. Thus, we expect that a reduction in oil export revenue (because of sanctions) increases the pressure from the public sector on the supply of foreign exchange. A growing demand for hard currencies by households (to protect the value of their assets) and businesses (to finance their imports and investments) increases the gap between the formal and informal exchange rate, reflecting in a higher black-market premium in the market.

Gross fixed capital formation (GFCF) is the sum of public and private investments in the country. The foreign exchange needs for importing machinery related to public investments are met at the official rate. However, the private sector investments mainly use the non-official exchange rate. It means that GFCF potentially influences both the official and non-official rates. Hence, the final effects on BMPER depends on the share of public and private

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[1] We should also note that the falling oil revenues because of sanctions have also reduced the government's ability to support the fixed exchange rate, resulting to a variety of exchange rates for different purposes (for more information see Farzanegan, 2013).

[2] After the introduction of the *Shetab* (Interbank Information Transfer Network) system (an electronic banking clearance and automated payments system) in Iran in 2002, an equilibrium rate in *Shetab* is used as official exchange rate. For more detail on the history of official exchange rate in Iran please see Pesaran (1992).

sectors in total investments and their needs to get foreign exchange rates at official and non-official rates, respectively.

Moreover, neoclassical models predict increases in exchange rate and output following positive shock in the government expenditure. An increase in government expenditures (GVEX) may be associated with a positive effect on the demand for labor, which can increase the real wage. Therefore, the private consumption may expand which causes the free market exchange rate increases (Hebous 2011). Çebi and Çulha (2014) find that a positive shock to the government expenditure is associated with an increase in the free market exchange rate. This may lead to widening the gap between the free market and official exchange rates, which results in a higher BMPER.

The balance of trade (BOT) is basically affected by imports and exports volume. As imports raise (or the balance of trade decreases), the demand for foreign exchange currencies will grow and vice versa. However, the effects of BOT on the official and non-official rates individually depend on the government and private sectors trade activities. Direct government imports and exports are settled at the official exchange rate. The government (through the Central Bank) often provides lists of items that are allowed to be imported with the official rate (mainly for strategic agricultural commodities) and the exports that are allowed to offer their foreign exchange earnings in the market rate. Accordingly, the effects on BMPER depends on the interaction between imports and exports by both private and public sectors. Table 1 presents summary statistics of our variables. The average rate of the BMPER for the *post-revolution* period from 1978 to 2017 is 341%, which shows significant variation over time under different governments (CBI 2020).[3]

Table 1. Descriptive statistics (1959-2017)[4]

Variable	Abbreviation	Obs.	Mean	Min	Max
Black Market Premium (%)	$\mathrm{BMPER_t}$	59	231	0	2011
Oil revenue (billion IRR)	$\mathrm{OILV_t}$	59	315775	105337	641296
Balance of trade (billion IRR)	$\mathrm{BOT_t}$	59	-526930	-1612600	204337
Gross fixed capital formation (billion IRR)	$\mathrm{GFCF_t}$	59	303106	39343	699062
Government expenditure (billion IRR)	$\mathrm{GVEX_t}$	59	135511	14606	210805

Source: CBI (2020).

3. Empirical model

We use the nonlinear autoregressive distributed lag (NARDL) model offered by Shin et al. (2014) to investigate our hypothesis. The "distributed lag" implies that the model also includes lags of the explanatory variables. This technique combines the hidden cointegration concept with the standard linear ARDL models and allows the investigation of short and long-run asymmetries. In the present study, this approach is tended to capture the negative shocks of the oil revenue. As Shin et al. (2014), we specify the following level nonlinear ARDL (p, q) model,

$$\mathrm{BMPER_t} = \alpha_0 + \sum_{i=1}^{p} \emptyset_i \mathrm{BMPER_{t-i}} + \sum_{i=0}^{q} (\theta_i^{-\prime} \mathrm{X}_{t-i}^{-} + \theta_i^{+\prime} \mathrm{X}_{t-i}^{+}) + \varepsilon_t \quad (2)$$

where $\mathrm{BMPER_t}$ denote the black-market premiums for the exchange rate X_t is a $k \times 1$ vector of the independent variables (including oil revenue, balance of trade, government expenditure, gross fixed capital formation[5]), and α_0 is a constant. The superscripts + and – indicate the partial sum process of positive changes (e.g. $\mathrm{X}_t^{+} = \sum_{i=1}^{t} \Delta\mathrm{X}_i^{+} = \sum_{i=1}^{t} \max(\Delta\mathrm{X}_i, 0)$) and

[3] For a review of the development of the BMPER in Iran with reference to the sanctions, see Farzanegan (2013).
[4] Data and model codes are available on request.
[5] Dummy variables are added as constant.

the partial sum process of negative changes (e.g. $X_t^- = \sum_{i=1}^{t} \Delta X_i^- = \sum_{i=1}^{t} \min(\Delta X_i, 0)$) in the explanatory variables. p and q denote lag orders and Δ is a different operator. The NARDL model specification in equation (2) is estimated by ordinary least squares (OLS) (Pesaran and Shin 1998). Following Pesaran and Shin (1998) and Shin et al. (2014), we can re-express equation (2) in its equivalent Error Correction Model (ECM) representation as,

$$\Delta \mathrm{BMPER_t} = \beta_0 + \rho \mathrm{BMPER_{t-1}} + \theta^{-'} X_{t-1}^- + \theta^{+'} X_{t-1}^+ + \sum_{i=1}^{p-1} \gamma_i \Delta \mathrm{BMPER_{t-i}}$$
$$+ \sum_{i=0}^{q-1} (\omega_i^{-'} \Delta X_{t-i}^- + \omega_i^{+'} \Delta X_{t-i}^+) + \varepsilon_t$$
$$= \beta_0 + \rho \vartheta_{t-1} + \sum_{i=1}^{p-1} \gamma_i \Delta \mathrm{BMPER_{t-i}} + \sum_{i=0}^{q-1} (\omega_i^{-'} \Delta X_{t-i}^- + \omega_i^{+'} \Delta X_{t-i}^+) + \varepsilon_t \quad (3)$$

where $\gamma_i = -\sum_{j=i+1}^{p} \emptyset_i$ for i= 1, …, p, $\rho = -(1 - \sum_{i=1}^{p} \emptyset_i)$, $\omega_i^- = -\sum_{i=j+1}^{q} \theta_i^-$ and $\omega_i^+ = -\sum_{i=j+1}^{q} \theta_i^+$ for i=1,…,q-1, $\omega_0^- = \theta_0^-$, $\omega_0^+ = \theta_0^+$, and β_0 is a constant. The error correction form of equation 3 includes both the short- and long-run dynamics simultaneously. The first part of the equation is the long-run dynamics (i.e. $\rho \mathrm{BMPER_{t-1}} + \theta^{-'} X_{t-1}^- + \theta^{+'} X_{t-1}^+$) and the second part shows the short run dynamics (i.e. $\sum_{i=1}^{p-1} \gamma_i \Delta \mathrm{BMPER_{t-i}} + \sum_{i=0}^{q-1} (\omega_i^{-'} \Delta X_{t-i}^- + \omega_i^{+'} \Delta X_{t-i}^+)$). By equation 3, we can estimate the speed of adjustment of ρ and check whether our model is dynamically stable. This is done by substituting the lagged error term of ECM specification with its equivalent (i.e. $\vartheta_{t-1} = \mathrm{BMPER_{t-1}} - \frac{\theta^{-'}}{\rho} X_{t-1}^- - \frac{\theta^{+'}}{\rho} X_{t-1}^+$) (Shin et al., 2014; Bahmani-Oskooee and Bahmani, 2015). The sum over all short-run coefficients of equation 3 results the long-run effects, i.e. $\theta^+ = \sum_{i=0}^{q} \theta_i^+$, $\theta^- = \sum_{i=0}^{q} \theta_i^-$. Further, the asymmetric long-run coefficients of explanatory variables (i.e. X_t) are given by,

$$L_{X_t^-} = -\frac{\hat{\theta}^-}{\hat{\rho}} = -\frac{\sum_{i=0}^{q} \hat{\theta}_i^-}{\hat{\rho}}, \; L_{X_t^+} = -\frac{\hat{\theta}^+}{\hat{\rho}} = -\frac{\sum_{i=0}^{q} \hat{\theta}_i^+}{\hat{\rho}} \qquad (4)$$

where $L_{X_t^-}$ and $L_{X_t^+}$ respectively indicate the long-run effect of the negative and positive partial sums. To check for the presence of cointegration between $\mathrm{BMPER_t}$, $\mathrm{OILV_t}$, $\mathrm{BOT_t}$, $\mathrm{GFCF_t}$ and $\mathrm{GVEX_t}$, we apply the bounds testing approach proposed by Pesaran et al. (2001). This approach combines the Engle-Granger two steps into one, and it is applicable irrespective of whether the variables are integrated of order zero or one, I(0) or I(1) (Pesaran et al., 2001). Besides, it is more efficient in the case of small samples (Bahmani-Oskooee and Tankui, 2008; Pesaran et al., 2001).

The bounds testing technique should be performed on the EC representation of nonlinear (i.e. equation 3). Cointegration holds if the $F_{\_PSS}$ statistic obtained from equation (3) is compared with the critical value bounds proposed by Pesaran et al. (2001). There are two asymptotic bounds: the lower and upper bound critical values. If $F_{\_PSS}$ statistic exceeds the given upper bound, the null hypothesis of no cointegration, i.e. H_0: $\rho = \theta^+ = \theta^- = 0$, is rejected. By contrast, if the calculated $F_{\_PSS}$ statistic is lower than the lower bound it fails to reject the null hypothesis, and if the statistic falls between the critical bounds the test result is inconclusive. Additionally, we consider the t-test of the ECM, i.e. $t_{\_BDM}$, by Banerjee et al. (1998) which examines the lagged variable of the error correction term in an unrestricted ECM form. The null hypothesis of $t_{\_BDM}$ is H0: $\rho = 0$ against the alternative of H1: $\rho < 0$. If the null hypothesis is rejected, we assume cointegration.

According to the above explanations, the long-run relationship between the BMPER and the negative changes in oil revenue is expected to be positive (i.e. $L_{OILV_t^-} > 0$). We also expect BOT to have the same impact on the BMPER, in contrast to the government expenditure and investment which mainly influence the demand side of foreign exchange.

4. Estimation results

We first use the Akaike Information Criterion (AIC) to determine the appropriate NARDL-model. As it is presented in Appendix I, our specification has the lowest AIC. Although the NARDL model is appropriate with both I(1) and I(0) variables, we apply the unit root test to make sure that there is no I(2) variable involved. As shown in Table 2, the Phillips–Perron (PP) test rejects the null hypothesis of unit root for the first differences, but not for the levels. Thus, all variables can be I(1). With this assumption, we can test for cointegration among the variables in the next step.

Table 2. Results of Philips-Perron (PP) unit root test

Variables	Intercept	Intercept and Trend
$BMPER_t$	-2.501	-2.459
$OILV_t$	-2.447	-2.418
BOT_t	-2.150	-2.065
$GFCF_t$	-1.538	-2.152
$GVEX_t$	-1.744	-1.864
$\Delta BMPER_t$	-6.904^{***}	-6.851^{***}
$\Delta OILV_t$	-5.269^{***}	-5.235^{***}
ΔBOT_t	-6.760^{***}	-6.741^{***}
$\Delta GFCF_t$	-6.001^{***}	-5.975^{***}
$\Delta GVEX_t$	-5.230^{***}	-5.243^{***}

H_0: no stationary. *** p<.01, ** p<.05, * p<.1. Source: Own calculation.

The test statistics of $F_{\_PSS}$ and $t_{\_BDM}$ for both nonlinear and linear cointegration are presented in Table 3. For the asymmetric specification, bound test cointegration is beyond the critical values by Pesaran et al. (2001). We observed a similar result for the t-test on the lagged dependent variable. It means the null hypothesis of *no nonlinear cointegration* among variables is rejected. However, we do not find strong evidence of *symmetric cointegration* (see Table 3). This implies that *asymmetries* in variables can have permanent effects.

Table 3. Nonlinear Cointegration Tests

	Cointegration test	Test	Critical value		Conclusion
			I(0)	I(1)	
Nonlinear model	F-bound test	124.897	2.62	3.77	cointegration
	t-test on lagged dependent variable	-5.867	-2.58	-5.07	cointegration
Linear model	F-bound test	1.32	3.07	4.44	No cointegration
	t-test on lagged dependent variable	-2.37	-2.58	-4.23	No cointegration

Source: Own results. The critical values are from Pesaran et al. (2001).

We use different standard goodness-of-fit tests. For testing normality, we apply the test of Jarque-Bera (1980) which is asymptotically chi-squared distributed. As the test statistics in Table 4 show, the null hypothesis is not rejected which means the estimated residuals are normally distributed. Further, we perform the Ramsey's (1969) Regression Equation Specification Error Test (RESET) to determine general functional form misspecification. The idea behind RESET is to add quadratic independent variables and test if they significantly explain dependent variables. As shown, the null hypothesis of RESET is rejected indicating that our model is correctly specified. We further use a Lagrange multiplier test proposed by Breusch and Pagan (1979) to check for heteroscedasticity. The null hypothesis of Breusch-Pagan's test is no heteroscedasticity. According to the results (chi^2=0.874), heteroscedasticity is not present in our model. Finally, we use the Box–Pierce portmanteau test to examine the residuals for autocorrelation. According to test statistic (chi^2=18.170), the null hypothesis of serial correlation in the residuals is significantly rejected. In sum, all diagnosis tests confirm a good model fit.

Additionally, the Cumulative Sum of Recursive Residuals (CUSUM) and Cumulative Sum of Recursive Residual Square (CUSUM Q) statistics are applied to assess the stability of the model parameters. According to Figure 1, both plots are within the critical bounds at 5 percent significance level. Hence, we can conclude that our model is structurally stable.

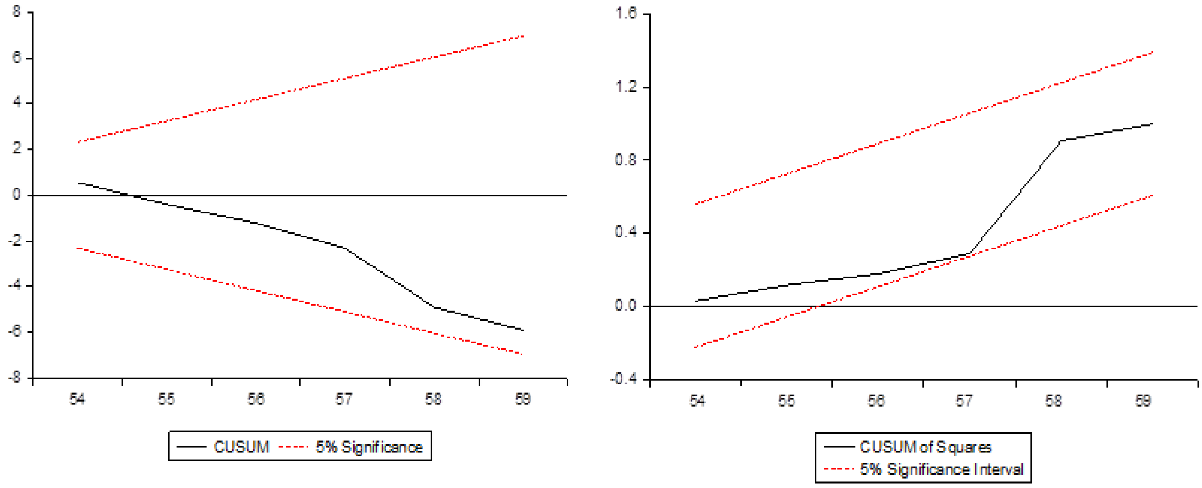

Figure 1. CUSUM and CUSUMS Plots

Source: Own calculation.

We estimate an error correction form of the unrestricted NARDL model to get the speed of adjustment. The model is dynamically stable as long as the speed of adjustment is negative and lower than one. By estimating the ECM representation, a significantly negative coefficient is obtained for the error term which supports the convergence of our estimates toward the long-run equilibrium (Table 4). Pesaran (1997; p 189) and Pesaran and Shin (1996; p 136) suggest that the long-run estimates are appropriate and sufficient for policy analysis when the error correction is fast. As shown, the speed of adjustment is significantly high. Thus, we only focus on interpreting the long-run equilibrium and the short-run effects are the lower order of importance. Nevertheless, the sum over all short-run coefficients of the transformed model results the long-run effects (see Appendix II for estimations of short-run coefficients).

Table 4 presents the parameter estimates for the long-run equilibrium of the asymmetric model. The long-run estimates show that increasing oil revenues reduce the BMPER. Oil revenues are always exchanged at official rates, but sometimes the Central Bank uses the US dollars gained from oil revenues (bought from the government) to meddle in the unofficial market (Valadkhani, 2004). Thus, there are two channels through which increasing oil revenues reduce pressure in the unofficial market. First, the list of priority items to receive foreign exchange in official rates expands, and that reduces the demand in the unofficial market. The second channel is the direct meddling in the unofficial market by the Central Bank[6].

Decreasing oil revenues have the opposite effect. The list of priority items to receive foreign exchange at the official rate gets smaller and the ability of the Central Bank to inject funds into the unofficial market diminishes. This leads to increased demand in the unofficial market and increases the gap between official and unofficial rates. The estimated long-run coefficients also imply that a reduction in oil revenues (e.g., due to oil sanctions) increases the BMPER[7]. Considering the estimated coefficients in Table 4, for the average negative change in oil revenues of 17,000 billion Rials[8], BMPER increases by 238%. For the same positive change of oil revenues, the BMPER decreases by 204%.

Besides, the BMPER is negatively affected by the BOT, which can be explained by the impacts of net exports on the source of foreign currencies. When BOT declines due to increasing imports, there is more demand for foreign exchange which in turn leads to higher BMPER. Additionally, the government tends to reduce the supply of foreign currencies at the official rate to mitigate the demand-side pressure following a positive change in imports and thereby narrows the gap between official and market rates. BOT is thus a complex combination of demand and supply of foreign exchange, in both official and market rates.

As explained above, the effect of an increase or decrease in GFCF on BMPER depends largely on the composition of public/private investment and also on how the government prioritizes official foreign exchange needs. On the one hand, the positive sign of the long-run negative shocks of the variable in the estimation means that a decrease in the investment reduces the demand pressures for foreign currencies. On the other hand, an increase in the GFCF often is associated with an economic boom that expands the need for foreign exchange. However, this pattern is often accompanied by a change in the types of investments that the government sponsors. In turn, this mechanism increases the demand for hard currencies at the official rate. This may lead to a marginal negative effect as presented in the estimation results. As we did not find a significant coefficient for positive variation in GFCF, we argue that the effects of GFCF on the BMPER are identified by the coefficient of negative changes, which is positive in this case. Decreasing investment reduces import-driven demand for foreign exchange and lowers the pressure on the unofficial market rate.

In line with Çebi and Çulha (2014) and Hebous (2011), the long-term estimate shows a positive relationship between BMPER and government expenditure. This is because government expenditures might crowd out the private sector's needs for foreign exchange from the official market, increasing pressure on the unofficial market by changing the prioritized items list. Following neoclassical models (Hebous 2011), an increase in government expenditure also means higher incomes for public employees and a higher demand for imported goods and outbound tourism.

As our findings in Table (4), the F statistics for the short-term asymmetry of oil revenues are significant while the F statistic for the long-term asymmetry is not. In line with Hufbauer et al. (2007) and Dizaji and van Bergeijk (2013), it shows that there might be a difference between the short and long-term impacts of positive and negative changes of oil revenues on BMPER.

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[6] This procedure is often a controversial action by the CBI and sometimes leads to legal problems.

[7] The negative signs of both estimated long-term effects are translated into an opposite correlation between dependent and independent variables and vice versa.

[8] It is equal to the average of negative shocks in oil revenue (i.e. OILV_t^-) from 1959 to 2017. According to the exchange rate of 2017, it is equal to 212 million US dollars.

Table 4. Estimation result of the Nonlinear ARDL model

long-run coefficient			
variable	coefficient	variable	Coefficient
$L_{OILV_t^-}$	-0.014*** (0.003)	$L_{OILV_t^+}$	-0.012*** (0.002)
$L_{BOT_t^-}$	-0.004** (0.001)	$L_{BOT_t^+}$	-0.003*** (0.001)
$L_{GFCF_t^-}$	0.018** (0.005)	$L_{GFCF_t^+}$	-0.004 (0.002)
$L_{GVEX_t^-}$	0.039 (0.024)	$L_{GVEX_t^+}$	0.049*** (0.048)

Short-run Asymmetry of oil revenue: F-test= 34.147***

Long-run Asymmetry of oil revenue: F-test= 2.202

Diagnostic Statistics:		
ECT_{t-1}	-0.928	[P-value < 0.01]
Portmanteau test [chi^2]	18.170	[P-value > 0.1]
Breusch-Pagan test [chi^2]	0.874	[P-value > 0.1]
Jarque-Bera test [chi^2]	2.754	[P-value > 0.1]
RESET test (F)	0.257	[P-value > 0.1]
lags (p, q_1,…, q_8)	(4,4,4,4,4,3,4,4,4)	
Adjusted R-squared	0.998	

Note: The dependent variable for the NARDL estimate is ΔBMPER_t. The superscripts ‘‘+’’ and ‘‘−’’ indicate positive and negative components of the variables. ECT stands for Error Correction Term which is estimated based on the conditional error correction specification. The standard errors are reported in parentheses. *** p<.01, ** p<.05, * p<.1. Source: Own calculation.

5. Conclusion

We use annual data from 1959 to 2017 for Iran to estimate the long- and short-term effects of increasing and decreasing oil revenues on the BMPER. An under-investigated aspect of sanctions is the response of BMPER which is closely associated with rent-seeking and corruption to sanction shocks. Our time-series analysis based on NARDL methodology is providing the first empirical evidence on the response of BMPER to negative (and positive) changes in the Iranian oil revenues. We find that falling oil revenues (as an empirical proxy for oil sanctions) cause the BMPER to significantly increase, controlling for other drivers of BMPER. Foreign exchange market restrictions and capital controls as a result of energy sanctions create parallel or black exchange markets and a premium of the parallel over the official exchange rate. A higher BMPER is an attractive incentive for traders to under-invoice exports or over-invoice imports. The unrecorded revenues are channeled in the black foreign exchange market to gain additional profits (see, e.g., Biswas and Marjit, 2007; Farzanegan, 2009; Buehn and Farzanegan, 2012). In short, energy sanctions distort the currency market and generate new rent-seeking opportunities in Iran.

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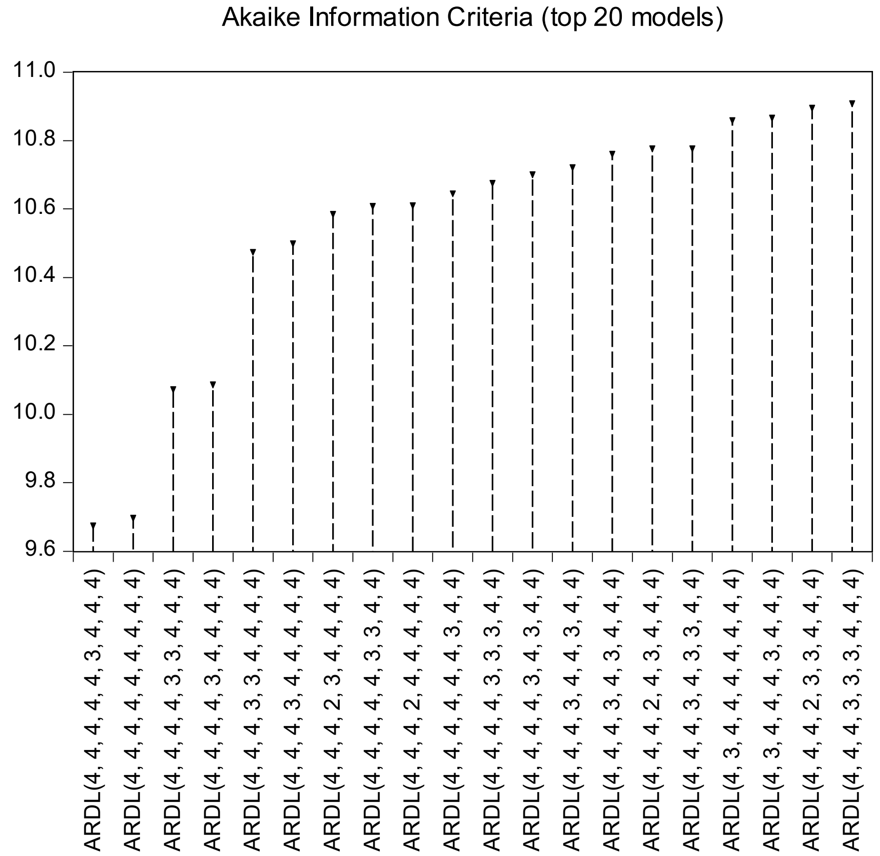
Appendix I: Akaike Information Criteria for top 20 models



Appendix II: Nonlinear ARDL(4,4,4,4,4,3,4,4,4) coefficients estimate

Dependent variable: $BMPER_t$

Variable	Coefficient	Variable	Coefficient
$OILV_t^-$	0.0022* (0.0009)	$OILV_t^+$	-0.0073*** (0.0015)
$OILV_{t-1}^-$	0.0003 (0.0006)	$OILV_{t-1}^+$	-0.0100*** (0.0013)
$OILV_{t-2}^-$	-0.0049** (0.0015)	$OILV_{t-2}^+$	0.0012 (0.0011)
$OILV_{t-3}^-$	-0.0089*** (0.0008)	$OILV_{t-3}^+$	-0.0020 (0.0012)
$OILV_{t-4}^-$	-0.0020 (0.0011)	$OILV_{t-4}^+$	0.0070** (0.0019)
BOT_t^-	-0.0011** (0.0003)	BOT_t^+	-0.0001 (0.0002)
BOT_{t-1}^-	-0.0017*** (0.0004)	BOT_{t-1}^+	-0.0002 (0.0002)
BOT_{t-2}^-	-0.0022*** (0.0003)	BOT_{t-2}^+	-0.0005* (0.0002)
BOT_{t-3}^-	-0.0016*** (0.0002)	BOT_{t-3}^+	-0.0003 (0.0003)
BOT_{t-4}^-	-0.0028*** (0.0004)	BOT_{t-4}^+	-0.0014*** (0.0003)
$GFCF_t^-$	-0.0012 (0.0010)	$GFCF_t^+$	0.0030** (0.0008)
$GFCF_{t-1}^-$	-0.0045*** (0.0012)	$GFCF_{t-1}^+$	-0.0042** (0.0015)
$GFCF_{t-2}^-$	0.0085*** (0.0002)	$GFCF_{t-2}^+$	0.0027** (0.0009)
$GFCF_{t-3}^-$	0.0108*** (0.0010)	$GFCF_{t-3}^+$	-0.0055*** (0.0010)
$GFCF_{t-4}^-$	0.0033** (0.0010)	$GVEX_t^+$	-0.0053 (0.0037)
$GVEX_t^-$	0.0096 (0.0091)	$GVEX_{t-1}^+$	-0.012** (0.0034)
$GVEX_{t-1}^-$	0.0380*** (0.0055)	$GVEX_{t-2}^+$	0.0166*** (0.0029)
$GVEX_{t-3}^-$	0.0166*** (0.0010)	$GVEX_{t-3}^+$	0.0139** (0.0046)
$GVEX_{t-4}^-$	-0.0300*** (0.0059)	$GVEX_{t-4}^+$	0.0319** (0.0048)
Const.	426.92*** (94.10)	$BMPER_{t-1}$	0.4666*** (0.0546)
DUM_{war}	-2421.58*** (230.60)	$BMPER_{t-2}$	-0.2133* (0.0943)
$DUM_{revolution}$	-1719.05*** (230.60)	$BMPER_{t-3}$	0.0281 (0.0566)
DUM_{policy_reform}	152.13 (85.90)	$BMPER_{t-4}$	-0.2090*** (0.0469)
$DUM_{financial\_crisis}$	235.28*** (38.80)		

Note: only significant coefficients are reported.

*** p<.01, ** p<.05, * p<.1