The Effects of Exchange Rate Volatility on Economic Growth in Iran

Sanginabadi, Bahram and Heidari, Hassan

Urmia University, Urmia University

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Bahram Sanginabadi[1], Hassan Heidari[2]

THE EFFECTS OF EXCHANGE RATE VOLATILITY ON ECONOMIC GROWTH OF IRAN

This paper investigates the effects of exchange rate volatilities on economic growth of Iran over the flexible exchange rate regime period (1988:Q1-2007:Q4). We use generalized autoregressive conditional heteroscedasticity (GARCH) family models to generate time-varying conditional variance of exchange rate as a standard measure of exchange rate volatility. We also use the autoregressive distributed lag (ARDL) bounds test approach to level relationship as proposed by Pesaran et al. (2001). Our results show a significant relationship between Iranian growth volume and real exchange rate volatility. The long-run results of ARDL model show that the effect of exchange rate volatility on economic growth is negative. ECM estimate shows that approximately 22% of disequilibria from the previous period's shocks converge back to the long-run equilibrium in the current period.

Keywords: *exchange rate volatility; economic growth; bounds test; Iran.*

JEL classification: F31; O40; O47.

Бахрам Сангінабаді, Хассан Хейдарі

ВПЛИВ ВОЛАТИЛЬНОСТІ ВАЛЮТНОГО КУРСУ НА ЕКОНОМІЧНЕ ЗРОСТАННЯ ІРАНУ

У статті досліджено вплив волатильності обмінного курсу на економічне зростання Ірану протягом періоду гнучкого режиму обміну (перший квартал 1988 — четвертий квартал 2007). Для вимірювання волатильності обмінного курсу використано родину моделей GARCH, побудовано часову дисперсію валютного курсу. Додатково використано авторегресивний розподілений лаг та метод граничних значень. Результати аналізу вказують на суттєвий взаємозв'язок між темпами зростання іранської економіки та волатильністю реального обмінного курсу. Результати довготермінового моделювання продемонстрували негативний вплив волатильності на економічне зростання. Приблизно 22% диспропорцій, згенерованих у попередньому періоді, повертаються у стан стабільності у чинному періоді.

Ключові слова: *волатильність обмінного курсу; економічне зростання; метод граничних значень; Іран.*

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Бахрам Сангинабади, Хассан Хейдари

ВЛИЯНИЕ ВОЛАТИЛЬНОСТИ ВАЛЮТНОГО КУРСА НА ЭКОНОМИЧЕСКИЙ РОСТ ИРАНА

В статье исследовано влияние волатильности обменного курса на экономический рост Ирана в течение периода гибкого обменного режима (первый квартал 1988 — четвертый квартал 2007). Для измерения волатильности обменного курса использовано семейство моделей GARCH и построена временная дисперсия валютного курса. Дополнительно использованы авторегрессивный распределенный лаг и метод граничных значений. Результаты анализа указывают на существенную взаимосвязь между темпами роста иранской экономики и волатильностью реального обменного курса. Результаты долгосрочного моделирования показали негативное влияние волатильности на

[1] Postgraduate Student, Urmia University, Iran.
[2] Corresponding Author, Ph. D., Assistant Professor, Department of Economics, Urmia University, Iran.

экономический рост. Примерно 22% диспропорций, сгенерированных в предыдущем периоде, возвращаются в состояние долгосрочной стабильности в текущем периоде.

Ключевые слова: *волатильность обменного курса; экономический рост, метод граничных значений; Иран.*

1. Introduction. Iran was moving between different exchange regimes in the recent decades. In a few last years, government intervenes in the currency market and tried to keep the exchange rate low. Shakeri (2008) argued that in recent decades we've been having foreign exchange regime in Iran. In some years we fixed exchange system and exchange in some years was floating. As the major currency supply is made by government due to oil export revenues, foreign exchange is always subject to exogenous shocks, and we do not have a competitive market with enough depth. In some years we had command exchange rate stabilization, which is incompatible with the fixed exchange rate system supported by a particular exchange rate with market forces. In recent years, floating exchange rate regime or a managed floating has been achieved in Iran, having little place of contemplation.

From the theoretical point of view, there is no clear consensus which exchange rate regime is more favorable for macroeconomic performance. Proponents of fixed exchange rate regime argue that exchange rate stability promotes economic performance through higher trade and enhanced macroeconomic stability, which could favor foreign investment and growth. This regime also affects investment and saving decisions (and therefore the current account balance) and financial development. In contrast, proponents of flexible exchange rate regimes emphasize the advantage of exchange rate flexibility to correct for domestic and external disequilibria in the face of real asymmetric shocks (see, e.g., Arratibel et al., 2011).

The success of currency depreciation in promoting trade balance and economic growth depends on demand side of economy as well as supply side to response the additional demand. This issue is discussed theoretically in the model by Meade (1951). If the Marshall-Lerner condition is not satisfied, currency depreciation could produce contraction. Hirschman (1949) pointed out that currency depreciation from an initial trade deficit reduces real national income and may lead to a fall in aggregate demand. Currency depreciation gives with one hand, by lowering export prices, while taking away with the other hand, by raising import prices. If trade is in balance and terms of trade are not changed these price changes offset each other. But if imports exceed exports, the net result is a reduction in real income within a country (Kandil, 2004).

Two opposite views in the literature exist when dealing with the effects of exchange rate volatility on growth. Bagella et al. (2006) defined first view as costs of volatility argument (CVA), and the second as advantage of flexibility argument (AFA).

The first view debates that exchange rate volatility decreases economic growth. "Although the associated costs have not been quantified rigorously, many economists believe that exchange rate volatility reduces international trade, discourages investment and compounds the problems people face in insuring their human capital in incomplete asset markets" (Obstfeld and Rogoff, 1995)[3].

[3] Note that the negative effects of exchange rate volatility on non-oil export of Iran are confirmed in the previous study by the authors (see Heidari et al., 2011).

The second view traces back to Meade's (1951) argument that, in countries with fixed exchange rates and inflexible money wages, adjustment in the equilibrium real exchange rate arising from external shocks will occur through domestic nominal prices and domestic wages. In such cases shock absorption would be easier under flexible exchange rate regimes (Bagella et al., 2006).

Yeyati and Sturzenegger (2003) findings confirm this view. "Terms of trade shocks are amplified in countries with more rigid exchange rate regimes and countries with flexible exchange rate regimes grow faster" (Edwards and Yeyati, 2003).

The review of recent empirical studies, however, shows that the results are mixed: absence of relation between exchange rate volatility and economic performance is reported by some researchers (see, e.g., Baxter and Stockman, 1989; Ghosh et al., 2003). According to some of them, the effect of exchange rate volatility on economic performance is negative, at the best (see, e.g., Bleaney and Greenaway, 2001; Bagella et al., 2004; Aghion et al., 2006; Bagella et al., 2006; Schnable, 2008; Aghion et al., 2009; Arratibel et al., 2011). And at least by Aliyu (2009) positive relation between exchange rate volatility and macroeconomic performance was reported.

This is a vacuum in the studies, as both theory and empirical studies do not have a clear conclusion on the relationship between exchange rate volatility and macroeconomic performance. Thus, the effect of exchange rate volatility on growth is ultimately an empirical issue. This paper contributes to this topic by analyzing the relation between exchange rate volatility and economic growth in Iran.

The rest of the paper is structured as follows: Section 2 provides the measurement of exchange rate volatility. Section 3 lays out the data and econometric methodology. Section 4 presents the empirical results, and finally section 5 concludes the paper.

2. The model. To motivate the empirical analysis, we draw on the theoretical predictions of the model in Kandil and Mirzaie (2002). According to this theory, the combination of demand and supply channels indicates that real output depends on movements in the exchange rate, money supply, and government spending. Bahmani-Oskooee and Kandil (2007) applied the real exchange rate, money supply, government spending, and exchange rate fluctuations proxy to study the relation between exchange rate fluctuations and output with Iranian data.

In the light of Kandil and Mirzaie (2002) theoretical predictions and Bahmani-Oskooee and Kandil (2007) empirical study we investigate the effects of exchange rate volatility on output, applying the following model:

$$\ln(Y_t) = \delta_1 \ln(G_t) + \delta_2 \ln(M_t) + \delta_3 \ln(RER_t) + \delta_4 \ln(V_t) + \varepsilon_t, \quad (1)$$

where $ln(Y_t)$ is the log of real GDP, $ln(G_t)$ is the log of real government consumption that measures fiscal policy, $ln(M_t)$ is the log of real money supply, approximates monetary policy, $ln(RER_t)$ is the log of real Exchange rate, $ln(V_t)$ is the exchange rate volatility in log, and ε_t is the error term.

3. Exchange rate volatility measure. In previous studies several different measures of exchange rate volatility (variability) were applied. However, following Heidari & Hashemi Pourvaladani (2011) we use generalized autoregressive conditional heteroscedasticity (GARCH) family models to generate time-varying conditional variance of exchange rate as a standard measure of exchange rate volatility. GARCH (p,q) model can be defined as follows:

$$X_t = \alpha + \beta \sum_{i=1}^{m} X_{t-i} + u_t \quad (2)$$

$$u_t \,|\, \Omega_t \sim iidN(0, h_t)$$

$$h_t = \gamma_0 + \sum_{i=1}^{p} \delta_i h_{t-i} + \sum_{j=1}^{q} \gamma_i u_{t-j}^2, \quad (3)$$

where Ω_t and h_t are the information set and the variance of residuals, respectively. m, p and q are set to one by using Schwarz Basyesian criterion (SBC). The estimated h_t in log (conditional variance) from the GARCH (p,q) model is peroxide volatility index. To get some insight into the trend, we plot the volatility of exchange rates in Figure 1.

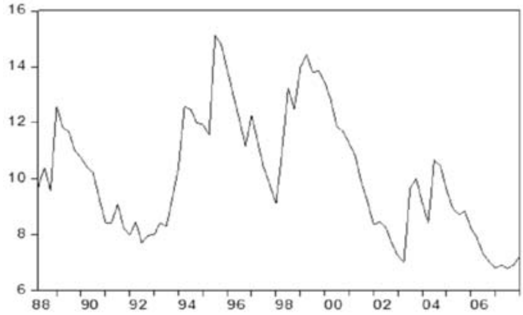

Fig. 1. **Volatility of exchange rates**

As Figure 1 shows, Iranian economy experienced even volatile exchange rate volatility during the period under study.

4. Data and Econometric Methodology

4.1. Data. This paper uses the quarterly data on Iranian economy over 1988:Q1-2007:Q4. All the data are gathered from Central Bank of Iran. Summary statistics for the series are given in Table 1.

Table 1. **Summary statistics for variables, 1988:q1-2007:q4**

	LnY_t	LnG_t	LnM_t	$LnRER_t$	LnV_t
Mean	11.23777	9.221208	11.67165	8.277429	10.15016
Median	11.21363	9.231935	11.74865	8.327340	9.830838
Maximum	11.75318	9.529085	14.20302	8.671704	15.11398
Minimum	10.69340	8.926252	9.436663	7.716356	6.793886
Std. Dev.	0.269100	0.162884	1.394912	0.226179	2.188084
Skewness	0.068812	-0.024484	0.041448	-0.537813	0.365123
Kurtosis	2.318658	2.036294	1.822940	2.550903	2.181930
Jarque-Bera	1.610556	3.103754	4.641143	4.528873	4.008328

Note: LnY_t is the log of real GDP, LnG_t is the log of real government consumption, LnM_t is the log of money supply, is the $LnRER_t$ log of real Exchange rate, LnV_t is the exchange rate volatility in log.

4.2. Unit root tests. In this paper, augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) are used to test for stationarity. Table 2 presents the ADF and PP test results at level and first difference. The test for the ADF and PP is applied with intercept, trend and intercept and non-intercept or trend. In this manner the results show that all series are stationary after first difference but are non-stationary in levels[4].

Table 2. **ADF and PP Unit root tests results**

	LnY_t	ΔLnY_t	LnG_t	ΔLnG_t	LnM_t	ΔLnM_t	$LnRER_t$	$\Delta LnRER_t$	LnV_t	ΔLnV_t
τ_μ(ADF)	-0.597	-11.159***	-0.932	-8.411***	0.937	-1.819	-1.025	-9.608***	-1.787	-8.525***
τ_T(ADF)	-2.760	-11.089***	-6.759***	-8.353**	-1.989	-2.076	-2.334	-9.568***	-2.080	-8.497***
τ(ADF)	3.362	-9.462***	1.955	-8.018***	2.116	0.530	-1.359	-9.430***	-0.648	-8.566***
τ_T(PP)	-0.439	-11.277***	-2.527	-33.37***	2.163	-14.848***	-0.832	-9.725***	-1.862	-8.520***
τ_μ(PP)	-2.489	-11.204***	-6.823***	-34.13***	-2.703	-15.588***	-2.313	-9.718***	-2.151	-8.491***
τ(PP)	5.227	-9.446***	2.722	-20.52***	19.958	-4.521***	-1.751*	-9.484***	-0.650	-8.564***

Note: Δ is the lag operator. τ_μ Represents the most general model with intercept, τ is the model with intercept and trend and τ is the model without intercept and trend. Both in ADF and PP tests, unit root tests were performed from the most general to the least specific model by eliminating trend and intercept across the models (See Enders, 2005: 181-199).*, ** and *** denote rejection of the null hypothesis at the 10%, 5% and 1% level respectively.

As Pahlavani et al. (2005) and Heidari and Parvin (2008) showed, Iranian economy was a subject to numerous shocks and regime shifts. On the other hand, Perron (1989) argued that failing to account for at least one structural break in the trend function may bias the usual unit root tests results towards their non-rejection of the null. In other words, standard unit root tests such as the ADF and PP tests may incorrectly indicate there is a unit root in a series, whereas in fact this series can be stationary around a one-time structural break (Zivot and Andrews, 1992). Moreover on the same issue, Ben-David et al. (2003) argued that just as failure to allow one break can cause non-rejection of the unit root null by the ADF test, failure to allow for two breaks, if they exist, can cause non-rejection of the unit root null by the tests which only incorporate one break.

Zivot and Andrews (1992) proposed a unit root test with one possible endogenous structural break and Lumsdaine and Papell (1997) proposed LM unit root test that allows for two unknown structural breaks under both the null and the alternative hypotheses. Table 3 presents Zivot and Andrews (ZA) and Lumsdaine and Papell (LP) unit root test's results.

The results of ZA and LP tests at 1% and 10% significance levels and reveal that while LnG_t is stationary, other variables are unit root. In other words, in the presence of possible structural breaks, the series under consideration are not in the same order of integrating.

[4] Note that according to ADF test LnMt is integrated of an order larger than one but from PP test LnMt is unit root. As the serial correlation does not affect the asymptotic distribution of the PP test statistic, we can say that LnMt is unit root (see Phillips and Perron, 1988).

Table 3. **ZA and LP unit root tests results**

Regressor	TB_{ZA}	*Lag*	τ_{ZA}	TB_{1LP}	TB_{2LP}	*lag*	τ_{LP}
LnY_t	1992Q3	0	-4.7562	1990Q3	1999Q3	2	-3.6943
LnG_t	1997Q2	4	-8.1015***	1993Q2	2000Q4	2	-6.5650*
LnM_t	2003Q2	4	-3.0192	1991Q1	2004Q2	2	-2.2664
$LnRER_t$	1998Q4	0	-4.2440	1992Q1	1998Q4	2	-4.8556
LnV_t	1993Q4	1	-4.0803	1990Q4	1993Q4	2	-4.0940

Notes: critical values for Zivot and Andrews unit root test at the 1, 5 and 10% levels are -5.57, -5.08 and -4.82, respectively (Zivot and Andrews, 1992). Critical values for Lumsdaine and Papell unit root test at the 1, 5 and 10% levels are -7.34, -6.82 and -6.49, respectively (Lumsdaine and Papell , 1997). ***, ** and * indicate that the corresponding null is rejected at the 1, 5 and 10% levels, respectively.

4.3. Econometric methodology. In the previous section, we conclude that the series under consideration are not in the same order of integration. As most of the cointegration tests such as Engel-Granger, and Johansen and Joselius (1990), are confident when the series are in the same order of integration, these tests cannot be suitable for our study. Thus, we use bounds test approach to level relationship, which can be applied irrespective of the order of integration of the series.

4.3.1. Bounds test approach to level relationship. This paper applies bounds test approach to level relationship within autoregressive distributed lag (ARDL) model proposed by Pesaran et al. (2001). This method has several advantages in comparison to other cointegration procedures: First, this approach yields consistent estimates of the long run coefficients that are asymptotically normal irrespective of whether the underlying regressors are I(1) or I(0) or fractionally integrated. Thus, the bounds test eliminates the volatility associated with pretesting the order of integration. Second, this technique generally provides unbiased estimates of the long-run model and valid t-statistics even when some of the regressors are endogenous. Third, it can be used in small sample sizes, whereas the Engle-Granger and the Johansen and Joselius procedures are not reliable for relatively small samples (Pesaran et al., 2001). We apply the bounds test procedure by modeling our regression (equation 3) as a general vector autoregressive (VAR) model of order p, in z:

$$z_t = C_0 + \beta_t + \sum_{i=1}^{p} \phi_i z_{t-i} + \varepsilon_t, t = 1,2,3,\dots,T, \tag{4}$$

where c_0 is a $(k+1)$ vector of intercepts and β denotesg a $(k+1)$ – vector of trend coefficients. Similar to Pesaran et al. (2001) our vector error correction model (VECM) is as follows:

$$\Delta Z_t = C_0 + \beta_t + \pi_t Z_{t-i} + \sum_{i=1}^{t} \Gamma_i \Delta Z_{t-i} + \varepsilon_t, t = 1,2,3,\dots,T, \tag{5}$$

where the $(k+1)\ x\ (k+1)$ – matrices, $\pi = I_{k+1} \sum_{i=1}^{p} \psi_i$ and, $\Gamma_i = -\sum_{j=i+1}^{p} \psi_i, i = 1,2,\dots,p-1$

contain the long-run multipliers and the short-run dynamic coefficients of VECM. z_i is the vector of variables y_t and x_t respectively. y_t is an $I(1)$ dependent variable defined as LnY_t and $x_t = [G_t, M_t, RER_t, V_t]$ is a vector of $I(0)$ and $I(1)$ regressors with a multivariate identically independently distributed zero mean error vector $\varepsilon_t = (\varepsilon_{1t}, \varepsilon'_{2t})'$, and a homoscedastic process. As Pesaran et al. (2001) and Katircioglu (2009a, 2009b) we consider 3 cases for VECM with regard to intercept and trends:

Case III: unrestricted intercepts; no trends and the ECM is

$$\Delta LnY_t = c_0 + \delta_1 LnY_{t-1} + \delta_2 LnG_{t-1} + \delta_3 LnM_{t-1} + \delta_4 LnRER_{t-1} + \delta_5 LnV_{t-1} + \\ + \sum_{i=1}^{p} \phi_i \Delta LnY_{t-i} + \sum_{l=1}^{q_1} \varphi_l \Delta LnG_{t-l} + \sum_{m=1}^{q_2} \eta_m \Delta LnM_{t-m} + \sum_{n=1}^{q_3} \theta_n \Delta LnRER_{t-n} \\ + \sum_{s=1}^{q_4} \varsigma_s \Delta LnV_{t-s} + \psi D_t + \varepsilon_t \quad (6)$$

Case IV: unrestricted intercepts; restricted trends and the ECM is

$$\Delta LnY_t = c_0 + \delta_1 (LnY_{t-1} - \gamma_y t) + \delta_2 (LnG_{t-1} - \gamma_x t) + \delta_3 (LnM_{t-1} - \gamma_x t) + \\ + \delta_4 (LnRER_{t-1} - \gamma_x t) + \sum_{i=1}^{p} \phi_i \Delta LnY_{t-i} + \delta_5 (LnV_{t-1} - \gamma_x t) + \sum_{l=1}^{q_1} \varphi_l \Delta LnG_{t-l} + \\ + \sum_{m=1}^{q_2} \eta_m \Delta LnM_{t-m} + \sum_{n=1}^{q_3} \theta_n \Delta LnRER_{t-n} + \sum_{s=1}^{q_4} \varsigma_s \Delta LnV_{t-s} + \psi D_t + \varepsilon_t \quad (7)$$

Case V: unrestricted intercepts; unrestricted trends and the ECM is

$$\Delta LnY_t = c_0 + \beta_t + \delta_1 LnY_{t-1} + \delta_2 LnG_{t-1} + \delta_3 LnM_{t-1} + \delta_4 LnRER_{t-1} + \\ + \delta_5 LnV_{t-1} + \sum_{i=1}^{p} \phi_i \Delta LnY_{t-i} + \sum_{l=1}^{q_1} \varphi_l \Delta LnG_{t-l} + \sum_{m=1}^{q_2} \eta_m \Delta LnM_{t-m} + \\ + \sum_{n=1}^{q_3} \theta_n \Delta LnRER_{t-n} + \sum_{s=1}^{q_4} \varsigma_s \Delta LnV_{t-s} + \psi D_t + \varepsilon_t \quad (8)$$

where δ_i are the long run multipliers, c_0 is the intercept, t is time trend and ε_t are white noise errors.

4.3.2. Bounds Testing Procedure. The first step in the ARDL bounds testing approach is to estimate equation (1) by ordinary least squares (OLS) in order to test for the existence of a long-run relationship among the variables by conducting an F-test for the joint significance of the coefficients of the lagged levels of the variables, i.e., $H_N : \delta_1 = \delta_2 = \delta_3 = \delta_4 = \delta_5 = 0$ against the alternative $HA : \delta_1 \neq \delta_2 \neq \delta_3 \neq \delta_4 \neq \delta_5 \neq 0$. We denote the test which normalized on X by $F_Y(Y|G, M, RER, V)$. Two asymptotic critical values bounds provide a test for cointegration when the independent variables are I(d) (where $0 \leq d \leq 1$): a lower value assuming the regressors are I(0), and an upper value assuming purely I(1) regressors. If the F-statistics is above the upper critical value, the null hypothesis of no long-run relationship can be rejected irrespective of the orders of integration for the time series. Conversely, if the test statistic falls below the lower critical value, the null hypothesis cannot be rejected. Finally, if the statistic falls between the lower and upper critical values, the result is inconclusive. The approximate critical values for the F- and t-tests are obtained from Pesaran et al. (2001).

In the second step, once cointegration is established, the conditional $ARDL(p_1, q_1, q_2, q_3, q_4)$ long-run model for Y can be estimated as follows:

$$LnY_t = c_0 + \beta t + \sum_{i=1}^{p} \delta_1 LnY_{t-i} + \sum_{i=0}^{q_1} \delta_2 LnG_{t-i} + \sum_{i=0}^{q_2} \delta_3 LnM_{t-i} + \\ + \sum_{i=0}^{q_3} \delta_4 LnRER_{t-i} + \sum_{i=0}^{q_4} \delta_5 LnV_{t-i} + \psi D_t + \varepsilon_t \quad (9)$$

where all variables are as previously defined. This involves selecting the orders of the $ARDL(p_1, q_1, q_2, q_3, q_4)$ model in 4 variables using Schwarz information criteria.

In the third and final step, we obtain the short-run dynamic parameters by estimating an ECM associated with the long-run estimates. This is specified as follows:

$$\Delta LnY_t = c_0 + \beta t + \sum_{i=1}^{p} \phi_i \Delta LnY_{t-i} + \sum_{l=1}^{q_1} \varphi_l \Delta LnG_{t-l} + \sum_{m=1}^{q_2} \eta_m \Delta LnM_{t-m} + \\ + \sum_{n=1}^{q_3} \theta_n \Delta LnRER_{t-n} + \sum_{s=1}^{q_4} \varsigma_s \Delta LnV_{t-s} + \vartheta ecm_{t-1} + \varepsilon_t \quad (10)$$

where ϕ, φ, η, θ and ζ are the short-run dynamic coefficients of the model's convergence to equilibrium, and ϑ is the speed of adjustment (Oteng-Abayie and Frimpong).

In the case of cointegration based on the bounds test, the Granger causality tests should be done under VECM when the variables under consideration are cointegrated. By doing so, the short-run deviation series from their long-run equilibrium path are also captured by including an error correction term (Katircioglu, 2009a, 2009b). Therefore, error correction models of cointegration can be specified as follow:

$$\begin{aligned}
\Delta LnY_t &= \alpha_0 + \varphi_{11}^p(L)\Delta LnY_t + \varphi_{12}^q(L)\Delta LnG_t + \varphi_{13}^r(L)\Delta LnM_t + \varphi_{14}^s(L)\Delta LnRER_t \\
&\quad + \varphi_{15}^v(L)\Delta LnV_t + \delta ECT_{t-1} + u_{1t} \\
\Delta LnG_t &= \alpha_0 + \varphi_{21}^p(L)\Delta LnG_t + \varphi_{22}^q(L)\Delta LnY_t + \varphi_{23}^r(L)\Delta LnM_t + \varphi_{24}^s(L)\Delta LnRER_t \\
&\quad + \varphi_{25}^v(L)\Delta LnV_t + \delta ECT_{t-1} + u_{2t} \\
\Delta LnM_t &= \alpha_0 + \varphi_{31}^p(L)\Delta LnM_t + \varphi_{32}^q(L)\Delta LnY_t + \varphi_{33}^r(L)\Delta LnG_t + \varphi_{34}^s(L)\Delta LnRER_t \\
&\quad + \varphi_{35}^v(L)\Delta LnV_t + \delta ECT_{t-1} + u_{3t} \\
\Delta LnRER_t &= \alpha_0 + \varphi_{41}^p(L)\Delta LnRER_t + \varphi_{42}^q(L)\Delta LnY_t + \varphi_{43}^r(L)\Delta LnG_t + \varphi_{44}^s(L)\Delta LnM_t \\
&\quad + \varphi_{45}^v(L)\Delta LnV_t + \delta ECT_{t-1} + u_{4t} \\
\Delta LnV_t &= \alpha_0 + \varphi_{51}^p(L)\Delta LnV_t + \varphi_{52}^q(L)\Delta LnY_t + \varphi_{53}^r(L)\Delta LnG_t + \varphi_{54}^s(L)\Delta LnM_t \\
&\quad + \varphi_{55}^v(L)\Delta LnRER_t + \delta ECT_{t-1} + u_{5t}
\end{aligned} \quad (11)$$

where

$$\varphi_{11}^p(L) = \sum_{i=1}^{p_{11}} \varphi_{11,i}^p L^i \varphi_{12}^p(L) = \sum_{i=0}^{p_{12}} \varphi_{12,i}^p L^i \varphi_{13}^p(L) = \sum_{i=0}^{p_{13}} \varphi_{13,i}^p L^i \ldots \quad (12)$$

$$\varphi_{21}^p(L) = \sum_{i=1}^{p_{21}} \varphi_{21,i}^p L^i \varphi_{22}^p(L) = \sum_{i=0}^{p_{22}} \varphi_{22,i}^p L^i \varphi_{23}^p(L) = \sum_{i=0}^{p_{13}} \varphi_{23,i}^p L^i \ldots \quad (13)$$

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where Δ denotes the difference operator and L denotes the lag operator, where $(L)(L)\Delta Lny_{t-1} = \Delta Lny_{t-1}ECT_{t-1}$ is the lagged error correction term derived from the long-run cointegration model. Finally, u_{1t} and u_{2t} are serially independent random errors with mean zero and finite covariance matrix. According to the VECM for causality tests, having statically significant F- and t-ratios for ECN_{t-1} in equation 11 confirms short-run and long-run causality relationship, respectively (Narayan and Smith, 2004).

5. Empirical results. In order to test for the existence of a long-run relationship between series under consideration, the bounds test approach to level relationship is used. Table 5 gives the results of the bounds test under 3 different scenarios as suggested by Pesaran et al. (2001), which are with restricted deterministic trend *(F_{IV})*, with unrestricted deterministic trend *(F_V)*, and without deterministic trend *(F_{III})*. Intercept in these scenarios are all unrestricted. Critical values for F-statistics are taken from Narayan (2005) and t-statistic from Pesaran et al. (2001), these critical values are presented in Table 4. The lag length *p* for this test is based on Schwarz Bayessian criterion (SBC). As can be seen from Table 5, F-statistic value confirms cointegration among series in F_{III}, F_{IV} and F_V at the 1% level of significance.

Table 4. **Critical values for ARDL modeling approach**

K=5	0.1		0.05		0.01	
	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
F_{III}	2.50	3.76	3.03	4.44	4.25	6.04
F_{IV}	2.83	3.87	3.35	4.50	4.84	6.51
F_V	3.08	4.27	3.67	5.00	5.09	6.77
t_{III}	-1.62	-3.49	-1.95	-3.83	-2.58	-4.44
t_V	-2.57	-3.86	-2.86	-4.19	-3.43	-4.79

Source: Narayan (2005): PP. 1988 to 1990 for F-statistics and Pesaran et al. (2001): pp. 300-301 for t ratio.
Note: F_{III} represents the F-statistic of the model with unrestricted intercept and no trend. F_{IV} represents the F-statistics of the model with unrestricted intercept and restricted trend, F_V represents the F-statistics of the model with unrestricted intercept and trend. t_V and t_{III} are the t ratios for testing $\delta_1 = 0$ in equation 6 to 9 with and without a deterministic linear trend.

Table 5. Bounds F- and t-statistics for the Existence of a Levels Relationship

		With Determintic Trends			Without Determintic Trends	
	Lag	F_{IV}	F_V	t_V	F_{III}	t_{III}
F_y ($\ln Y_t$|$\ln G_t$, LnM_t, $LnRER_t$, LnV_t)	1	9.245***	11.079***	-7.224***	6.632***	-5.549***

Table 6 presents the long-run coefficients of ARDL (2,0,0,1,0). As can be seen from Table 6 coefficient of the real exchange rate is not significant but the coefficient of exchange rate volatilities unable to be rejecting at the 5% significance level and its coefficient is negative and approximately equal to 0.021. Therefore, 5% increase in exchange rate volatility leads to approximately 2.3% decrease in economic growth. Thus we can conclude that exchange rate volatility reduces economic growth in the long run.

Table 6. **Estimated long-run coefficients using the ARDL approach**

Regressor	Coefficient	Std. Error	t-Statistics	Prob.
LnG_t	0.563490	0.256845	2.193893	0.0313
LnM_t	0.398144	0.330729	1.203839	0.2324
$LnRER_t$	0.429798	0.291918	1.472325	0.1451
LnV_t	-0.023385	0.011036	-2.118923	0.0374
c	-1.254286	6.655441	-0.188460	0.8510

Table 7 presents the ECM estimation results. As can be seen, the coefficient of ECMT(-1) is 0.217, significant at 1% level and negative as expected. Thus, approximately 22% of disequilibria from the previous period's shock in economic growth model converge back to the long-run equilibrium in the current period.

Table 7. **Error correction representation for the selected ARDL model**

Regressor	Coefficient	Std. Error	t-Statistic	Prob.
ΔLnY_{t-1}	-0.240052	0.095520	-2.513110	0.0143
ΔLnG_t	0.159947	0.032899	4.861703	0.0000
ΔLnM_t	0.014026	0.091418	0.153426	.8785
$\Delta LnRER_t$	-0.092425	0.044246	-2.088879	0.0404
$\Delta LnRER_{t-1}$	-0.103659	0.043211	-2.398890	0.0192
ΔLnV_t	-0.000895	0.003315	-0.269857	0.7881
C	0.001669	0.007382	0.226054	0.8218
ECMT(-1)	-0.217285	0.052210	-4.161725	0.0001
R2 = 0.444	S.E.R=0.027	F.st=6.906(0.000)	Shwarz.C=-4.000	
R2 = 0.380	RSS=0.050	D.W=2.074	Akaike.C=-4.272	

Table 8 presents the results of the Granger causality tests for the selected ARDL model. The results don't confirm the long-run causality between independent variable set and economic growth. The short-run causality from real exchange rate or exchange rate volatility to economic growth and vice versa does not exist.

Table 8. **ARDL(2,0,0,2,0) Model Granger Causality Tests: With Deterministic Trend**

	ΔLnG_t	ΔLnM_t	$\Delta LnRER_t$	ΔLnV_t	ΔLnY_t	ECM(t-1) t-stat
ΔLnG_t	--	0.258919 (0.7727)	1.235947 (0.2973)	1.503163 (0.2301)	8.630401 (0.0005)	1.41678 (0.16132)
ΔLnM_t	0.619962 (0.5411)	--	6.363770 (0.0030)	2.080833 (0.1331)	2.888586 (0.0628)	-0.04483 (0.96438)
$\Delta LnRER_t$	5.055710 (0.0091)	4.618798 (0.0133)	--	2.414319 (0.0974)	0.037489 (0.9632)	-0.44508 (0.65774)
ΔLnV_t	0.491173 (0.6142)	0.947693 (0.3929)	0.129457 (0.8788)	--	0.815341 (0.4470)	-2.42612 (0.01804)
ΔLnY_t	0.644220 (0.5284)	0.318501 (0.7284)	0.302349 (0.7401)	0.894486 (0.4138)	--	-2.62911 (0.01067)

To investigate the causality between exchange rate volatility and economic growth result of Granger causality test are between ΔLnV_t and ΔLnY_t is present in Table 9. As can be seen t-static fics in the second line is significant at the 1% level, therefore we can say that long-run causality from exchange rate volatility to economic growth exists. The t-statistics in the first line is significant at 10% level only, confirming the long-run causality from exchange rate volatility to economic growth at the 10% significance level.

Table 9. **Granger causality tests between exchange rate volatility and economic growth: With Deterministic Trend**

	Δ LnY_t	Δ LnV_t	ECM(t-1) -- t-stat
Δ LnV_t	--	0.306484 (0.7370)	-1.85911 (0.06715)
Δ LnY_t	0.812009 (0.4480)	--	-3.17213 (0.00224)

Table 10 presents Granger causality test results for real exchange rate and economic growth. As can be seen, t-statics in the second line is significant at 5% level, therefore we can say that the long-run causality from real exchange rate to economic growth exists. Moreover t-statistics in the first line is significant at 10% level. This result confirms the long-run causality from real exchange rate to economic growth at the 10% significance level.

Table 10. **Granger causality tests between real exchange rate and economic growth: With Deterministic Trend**

	Δ LnY_t	Δ LnRER$_t$	ECM(t-1) -- t-stat
Δ LnRER$_t$	--	0.093690 (0.9107)	-1.89059 (0.06276)
Δ LnY_t	1.440029 (0.2438)	--	-2.13056 (0.03659)

Table 11 shows the diagnostic tests results for the ARDL(2,0,0,2,0) model used in this paper. In this manner Breusch-Godfrey serial correlation LM test and heteroskedasticity ARCH test are used. LM test indicates that the residuals are not serially correlated and ARCH test shows that the residuals have no heteroskedasticity problem.

Table 11. **ARDL(2,0,0,2,0) model diagnostic tests**

Breusch-Godfrey Serial Correlation LM Test:			Heteroskedasticity Test: ARCH		
		probe			probe
F-statistics	0.575858	0.5651	F-statistics	1.021984	0.3153
Obs*R-squared	1.357998	0.5071	Obs*R-squared	1.035132	0.3090

6. Conclusion. This paper provides evidence on the impact of exchange rate volatilities on the economic growth of Iran. The conditional variance is estimated from a univariate GARCH(1, 1) model and used as a standard measure of exchange rate volatility. The paper contributes to the literature by employing the bounds test approach to level relationship as proposed by Pesaran et al. (2001). The results of bounds test approach confirm the existence of long-run relationship among the variables under consideration. The long-run results from ARDL estimate show that the coefficient of exchange rate volatility is negative and acceptable at the 5% significance level. ECM estimate shows that the coefficient of ECMT(-1) is 0.217, significant at the 1% level and negative as expected. Thus, approximately 22% of disequilibria from the previous period's shocks converge back to the long-run equilibrium in the current period. The results of Granger causality tests confirm causation from exchange rate volatility to growth. These results may imply that exchange rate volatility dampens economic growth of Iran.

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