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Openness and Inflation in Iran

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Abstract-**The purpose of this paper is to test the hypothesis first documented by [1], that inflation is lower in more open economies. According to this hypothesis, central banks have a smaller incentive to engineer surprise inflations in more-open economies because the Phillips curve is steeper.We utilized the ARDL Bounds test approach to level relationship proposed by [2] for Iranian annual data over the period 1973-2007. Results from Bounds test approach confirm existence of long-run relationship among the variables under consideration. The results show that openness has negative and significant effect on inflation in short-run but its effect in long-run is not significant.**

Keywords-Openness; Inflation; Iran; ARDL Bounds test approach.

I. INTRODUCTION

One of the most striking events of the past two decades has been the remarkable decline in inflation around the world [3]. Inflation has always been a concern for the policy makers as it creates uncertainty in the economy that may adversely affect economic growth. Therefore, maintaining noninflationary stable economic growth has been at the core of macroeconomic policies in developing countries. The concern with inflation stems not only from the need to maintain overall macroeconomic stability, but also from the fact that inflation hurts the poor particularly hard as they do not possess effective inflation hedges [4].

Reference [1] argues that more-open economies will have steeper Phillips curves. The reason for this is that a monetary expansion in an open economy will be accompanied by a real depreciation of the currency, raising costs for households and businesses. The larger share of imported goods, the greater the increase in inflation. Reference [1] also argues that the relative weight on stabilizing output is smaller in more-open economies, again because of the real depreciation induced by the monetary shock.

Reference [5] follows a different line of reasoning to explain the negative correlation and argues that the effects of openness are straightforward: As the country becomes more open, the nontrade sector becomes less important than the traded goods sector. Therefore, the monetary authorities stand to gain less by creating surprise inflation in a more open economy [3].

Most studies have focused on the estimation of cross-country averages of many different levels of economies. However, these studies cannot identify country-specific differences. Little works have been done the impact of openness on inflation at a country level. The literature on the openness-inflation association in Iran is scarce and this study tries to fill this gap to some extent. Thus, the present article tests the hypothesis that inflation is lower in more open economies for Iran economy during 1973-2007. This paper is organized in five sections. After the introduction in the first section, section 2 provides a theoretical background and reviews empirical research. Section 3 presents model specification and data description. Section 4 considers the empirical results and finally a conclusion will be provided in section 5.

II. REVIEW OF LITERATURES AND EMPIRICAL RESEARCH

The theoretical reasoning for why more open economies tend to have less inflation follows from Reference's [6] model, which shows that such economies gain less from surprise inflation. Surprise monetary expansions cause the real exchange rate to depreciate, leading to a negative terms-of-trade effect. The more open the economy the more the real exchange depreciates, thus reducing incentives to undertake expansion [7].

Reference [1] proposes an explanation of this relationship. Because unanticipated monetary expansion causes real exchange rate depreciation, and because the harms of real depreciation are greater in more open economies, the benefits of surprise expansion are a decreasing function of the degree of openness. Thus, if the monetary authorities' temptation to expand is an important determinant of inflation-that is, if the absence of binding pre commitment is important to monetary policy-monetary authorities in more open economies will on average expand less, and the result will be lower average rates of inflation [1].

The relationship between inflation and openness has been a subject of research, theoretical as well as empirical. However, the literature on the subject is relatively scant. According to 'new growth theory', openness is likely to affect inflation through its likely effect on output [8]. This link could be operating through: a) increased efficiency which is likely to reduce cost through changes in composition of inputs procured domestically and internationally, b) better allocation of resources, c) increased capacity utilization, d) rise in foreign investment which can stimulate output growth and ease pressures on prices [9]. Reference [10] postulates that the shocks to the domestic price level due to domestic output fluctuation are likely to ease as the economy opens up [4].

Reference [11] documents that in Small open countries prices of traded goods converge across counties because of free trade; therefore, theory suggests a lower degree of price distortions in outward-looking countries. Moreover, in highly open countries conversion of domestic currency into foreign currency is very easy. Therefore, the inflation rate –a kind of tax on domestic currency –will be low in more open countries [12].

In early empirical literature [13] tested the hypothesis that openness leads to cheaper availability of goods that are costly in the country otherwise and confirmed that more open economies tended to experience lower inflation in 5 countries in the European Economic Community.

Reference [14] used a sample of 33 less developed countries and analysed the relationship for both yearly and 5 year average data from 1960-61 to 1964-65. A negative relationship between openness and inflation emerged when [14] related inflation and openness in a bivariate framework using method of ordinary least squares. However, when the analysis was extended to a multivariate exercise, the results were not unambiguous. Though the openness variable was not always significant, it always had a negative sign [4].

Reference [1] used a Barro- Gorden type of model for a cross section of 114 countries and argued that inflation is lower in small and open economies even in the absence of an independent central bank with pre-commitment to price stability [1].

Reference [5] proposes that it is existence of imperfect competition and the presence of rigid nominal prices in the non-tradable sector that leads inverse relationship between openness and inflation. According to new growth theory, openness reduces inflation through its positive influence on output, mainly through increased efficiency, better allocation of resources, improved capacity utilization, and increased foreign investment [8].

Reference [7] using panel of 148 countries finds that openness does not seem to play a role in the short run in restricting inflation, but a fixed exchange-rate regime plays a significant role [7].

Reference [4] tested the hypothesis that inflation is lower in small and open economies for Pakistan economy using annual time series data for the period 1973-2005.They found that the openness variable such as growth in 'overall trade to GDP ratio' also has significant negative impact on the domestic price growth in Pakistan.

Reference [3] showed that there is a robust negative relationship across countries between a country's openness to trade and its long-run inflation rate in the United States. Also, Reference [15] find support for Reference's [1] argument concerning the relation between monetary policy and economic openness. Their study links economic openness to the slopes of aggregate supply and aggregate demand to explain why the openness-inflation relation can be ambiguous. Their empirical results from 15 developed countries support the recent empirical failure in finding the negative openness-inflation relation. Reference [16] checked the validity of Reference's [1] main result and also tested the reference's [17] criticism that the negative relationship between openness and inflation is due to severely indebted countries in the debt crisis period. Their analysis showed that the principal result of [1] still holds in the 1990s, however, Terra's criticism fails to hold in the 1990s as the negative relationship between inflation and openness remains unrestrictive to a subset of countries or specific time period.

Reference [18] has been used multivariate cointegration and a vector error correction model in Pakistan during 1960 to 2007. The empirical findings under the cointegration test show that there is a significant negative long-run relationship between inflation and trade openness, which confirms the existence of Romer's hypothesis in Pakistan.

In turn, opponents (cost push hypothesis) argue that trade openness does not necessarily reduce inflation; rather it increases inflation. Inflowing there are researches that have shown that openness increases inflation:

Reference [19] argues that the positive effect of openness on inflation is driven by the fact that the monetary authority enjoys a degree of monopoly power in international markets as foreign consumers have some degree of inelasticity in their demand for goods produced in the home country. The decision of the monetary authority is then to balance the benefits of increased money growth that come from the open economy setting with the well-known consumption tax costs of inflation. Further, it is also possible for an open economy to import inflation from the rest of the world via the prices of manufactured imports or raw material imports. Moreover, as the economy opens up, the fiscal and monetary authorities tend to lose their ability to control inflation through fiscal and monetary policies [19].

Reference [20] developed a two-country general equilibrium model to analyze the optimal rate of inflation under discretion. He shows that when agents' welfare is the sole policy objective it is possible to show that openness and inflation no longer have a simple inverse relationship. Because the terms of trade are related to monopoly markups, a greater degree of openness may lead the policymaker to exploit the short-run Phillips curve more aggressively, even if it involves a smaller short-run benefit. Then inflation can be higher in a more open economy [20].

Reference [12] examined relationship between trade openness and inflation in Pakistan using annual time-series data for the period 1947 to 2007. The empirical analysis shows that a positive relation holds between trade openness and inflation in Pakistan.

III. MODEL, DATA AND EMPIRICAL RESULTS

We have used following model as:

$$INF_t = \delta_1 M_t + \delta_2 GS_t + \delta_3 OPN_t + \delta_4 Y_t + \varepsilon_t \tag{1}$$

Where INF_t is the inflation rate calculated by CPI, M_t is the money growth, GS_t is the government size, OPN_t is the openness criteria, Y_t is the GDP per capita and ε_t is the error term.

Empirical studies have shown that money is neutral in the long-run, so that the level of the money supply at any time has no influence on real magnitudes, money could still be non-super neutral: the growth rate of the money supply could affect real variables. A rise in the monetary growth rate leads to a new dynamic equilibrium with an equally increased inflation rate. Reference [21] tested the neutrality and super neutrality of money in Iran during 1959-2002. They have shown that money is neutral but the results of the super neutrality tests suggest that inflation driven by money growth. Thus, we have employed money growth in the model.

A. *Data*

This paper uses annual data of the Iranian economy during 1973-2007 taken from WDI. Summary statistics for the series are given in Table I.

TABLE I SUMMARY STATISTICS FOR VARIABLES

	M	INF	GS	Y	OPN
Mean	0.207514	18.48353	16.44441	1618.818	42.75751
Median	0.229982	17.21305	15.50931	1527.982	41.14782
Maximum	0.566546	49.65599	25.77113	2270.596	76.77430
Minimum	-1.000000	4.389341	11.01334	1122.060	13.77244
Std. Dev.	0.238034	8.513558	4.324425	295.9608	14.62797
Skewness	-3.750304	1.435188	0.716250	0.464852	0.266211
Kurtosis	20.31585	6.294721	2.333364	2.337362	2.844922
Jarque-Bera	519.3093	27.84578	3.640670	1.900848	0.448470
Probability	0.000000	0.000001	0.161972	0.386577	0.799127
Sum	7.262975	646.9236	575.5545	56658.62	1496.513
Sum Sq. Dev.	1.926442	2464.343	635.8223	2978155.	7275.237
Observations	35	35	35	35	35

B. *Unit Root Tests*

In this paper, Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests have been used to test for stationarity. Table II 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 results show that Inflation (INF), Government Size (GS) and GDP per capita (Y) are stationary after first difference but are non-stationary in levels.

According to ADF and PP openness (OPN) is non stationary in level, and stationary after first difference according to PP but it is not according to ADF. Reference [22] argues that an important advantage of PP test is that the serial correlation does not affect the asymptotic distribution of the PP test statistic. Therefore we can say that OPN is unit root. Money growth (M) unit root test results are different. M is stationary in level as well as first difference, therefore it is I(0).

TABLE II ADF AND PP UNIT ROOT TEST RESULTS

	IN	ΔIN	M	ΔM	GS	ΔG	OPN	ΔOPN	Y	ΔY
τ_μ	-3.813	-6.50	-4.372	-9.125**	-1.122	-6.406	-2.476	-1.634	-2.083	-3.207
τ_T	-3.742*	-6.45	-4.302	-9.005**	-2.996	-6.301	-1.240	-2.353	-1.656	-7.240
τ	-1.168	-6.61	-2.998	-9.270**	-0.883	-6.341	-0.440	-1.675	-0.121	-3.252
τ	-3.467*	-12.5	-4.384	-11.41**	-1.076	-6.406	-1.699	-4.661*	-0.989	-3.209
τ	-3.383*	-13.7	-4.312	-11.40**	-3.203	-6.301	-1.446	-5.053*	-0.540	-4.151
τ	-0.811	-12.7	-2.917	-11.62**	-0.883	-6.510	-0.542	-4.718*	0.1	-3.250

Note: Δ is the lag operator. τ_μ Represents the most general model with intercept, τ_T 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 [23, pp; 181-199].*, ** and ***denotes rejection of the null hypothesis at the 10%, 5% and 1% level.

C. *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 [24] and [25], 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 variables.

1) *ARDL Model Specification*

This paper applies Bounds test approach to level relationship with in Autoregressive Distributed Lag (ARDL) model proposed by [2] 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 pre-testing 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 procedures are not reliable for relatively small samples [2]. We apply the Bounds test procedure by modeling our regression (equation 3) as a general vector autoregressive (VAR) model of order p, in z:

$$Z_t = c_0 + \beta t + \sum_{i=1}^{t} \phi_i Z_{t-i} + \varepsilon_t \qquad (2)$$
$$t = 1,2,3,...,T$$

Where c_0 is a (k+1) vector of intercepts and β denoting a (k+1)-vector of trend coefficients. Similar to [2] our Vector Error Correction Model (VECM) is as follows:

$$\Delta Z_t = c_0 + \beta t + \pi Z_{t-i} + \sum_{i=1}^{t} \Gamma_i \Delta Z_{t-i} + \varepsilon_t$$
$$t = 1,2,3,...,T \qquad (3)$$

Where the (k+1) x (k+1)- matrices, $\pi = I_{k+1} \sum_{i=1}^{p} \psi_i$ and $\Gamma_i = -\sum_{j=i+1}^{p} \psi_i, i = 1,2,..., p-1$, contain the long-run multipliers and short-run dynamic coefficients of the VECM. z_i is the vector of variables y_t and x_t respectively. y_t is an I(1) dependent variable defined as INF_t and $x_t = [M_t, OPN_t, GS_t, Y_t]$ 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. We consider two cases for VECM with regard to intercept and trends:

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

$$\Delta INF_t = c_0 + \delta_1 INF_{t-1} + \delta_2 M_{t-1} + \delta_3 OPN_{t-1} + \delta_4 GS_{t-1} + \delta_5 Y_{t-1} + \sum_{i=1}^{p} \phi_i \Delta INF_{t-i} + \sum_{l=1}^{q_1} \varphi_l M_{t-l} + \sum_{m=1}^{q_2} \eta_m \Delta OPN_{t-m} + \sum_{n=1}^{q_3} \theta_n \Delta GS_{t-n} + \sum_{s=1}^{q_4} \varsigma_s \Delta Y_{t-s} + \psi D_t + \varepsilon_t \qquad (4)$$

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

$$\Delta INF_t = c_0 + \beta t + \delta_1 INF_{t-1} + \delta_2 M_{t-1} + \delta_3 OPN_{t-1} + \delta_4 GS_{t-1} + \delta_5 Y_{t-1} + \sum_{i=1}^{p} \phi_i \Delta INF_{t-i} + \sum_{l=1}^{q_1} \varphi_l M_{t-l} + \sum_{m=1}^{q_2} \eta_m \Delta OPN_{t-m} + \sum_{n=1}^{q_3} \theta_n \Delta GS_{t-n} + \sum_{s=1}^{q_4} \varsigma_s \Delta Y_{t-s} + \psi D_t + \varepsilon_t \qquad (5)$$

Where δ_i are the long run multipliers, c_0 is the intercept, t is time trend and ε_t are white noise errors [26].

2) Bounds Testing Procedure

The first step in the ARDL Bounds testing approach is estimate of equation (5) 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 $H_A : \delta_1 \neq \delta_2 \neq \delta_3 \neq \delta_4 \neq \delta_5 \neq 0$. We denote the test which normalized on INF by $F_{INF}(INF_t | M_t, OPN_t, GS_t, Y_t)$. Two asymptotic critical values bounds provide a test for cointegration when the independent variables are I(d) (where 0≤d≤1): a lower value assuming the regressors are I(0), and an upper value assuming purely I(1) regressors. If the F-statistic 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 were obtained from [2].

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

$$INF_t = c_0 + \beta t + \sum_{i=1}^{p} \delta_1 INF_{t-i} + \sum_{i=0}^{q_1} \delta_2 M_{t-i} + \sum_{i=0}^{q_2} \delta_3 OPN_{t-i} + \sum_{i=0}^{q_3} \delta_4 GS_{t-i} + \sum_{i=0}^{q_4} \delta_5 Y_{t-i} + \psi D_t + \varepsilon_t \qquad (6)$$

This involves selecting the orders of the $ARDL(p_1, q_1, q_2, q_3, q_4)$ model in the four variables using Schwarz information criteria.

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

$$\Delta INF_t = c_0 + \beta t + \sum_{i=1}^{p} \phi_i \Delta INF_{t-i} + \sum_{l=1}^{q_1} \varphi_l \Delta M_{t-l} + \sum_{m=1}^{q_2} \eta_m \Delta OPN_{t-m} + \sum_{n=1}^{q_3} \theta_n \Delta GS_{t-n} + \sum_{s=1}^{q_4} \varsigma_s \Delta Y_{t-s} + \vartheta ecm_{t-1} + \varepsilon_t \quad (7)$$

Where $\phi, \varphi, \eta, \theta$ and ς are the short-run dynamic coefficients of the model's, and ϑ is the speed of adjustment [26].

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 sort-run deviation series from their long-run equilibrium path are also captured by including an error correction term [27-28]. Therefore error correction models of cointegration can be specified as follows:

$$\Delta INF_t = \alpha_0 + \varphi_{11}^p(L)\Delta INF_t + \varphi_{12}^q(L)\Delta M_t + \varphi_{13}^r(L)\Delta OPN_t + \varphi_{14}^s(L)\Delta GS_t + \varphi_{15}^v(L)\Delta Y_t + \delta ECT_{t-1} + u_{1t}$$

$$\Delta M_t = \alpha_0 + \varphi_{21}^p(L)\Delta M_t + \varphi_{22}^q(L)\Delta INF_t + \varphi_{23}^r(L)\Delta OPN_t + \varphi_{24}^s(L)\Delta GS_t + \varphi_{25}^v(L)\Delta Y_t + \delta ECT_{t-1} + u_{2t}$$

$$\Delta OPN_t = \alpha_0 + \varphi_{31}^p(L)\Delta OPN_t + \varphi_{32}^q(L)\Delta INF_t + \varphi_{33}^r(L)\Delta M_t + \varphi_{34}^s(L)\Delta GS_t + \varphi_{35}^v(L)\Delta Y_t + \delta ECT_{t-1} + u_{3t}$$

$$\Delta GS_t = \alpha_0 + \varphi_{41}^p(L)\Delta GS_t + \varphi_{42}^q(L)\Delta INF_t + \varphi_{43}^r(L)\Delta M_t + \varphi_{44}^s(L)\Delta OPN_t + \varphi_{45}^v(L)\Delta Y_t + \delta ECT_{t-1} + u_{4t}$$

$$\Delta Y_t = \alpha_0 + \varphi_{51}^p(L)\Delta Y_t + \varphi_{52}^q(L)\Delta INF_t + \varphi_{53}^r(L)\Delta M_t + \varphi_{54}^s(L)\Delta OPN_t + \varphi_{55}^v(L)\Delta GS_t + \delta ECT_{t-1} + u_{5t} \quad (8)$$

Where

$$\varphi_{11}^p(L) = \sum_{i=1}^{p_{11}} \varphi_{11}^p L^i \qquad \varphi_{12}^p(L) = \sum_{i=0}^{p_{12}} \varphi_{12}^p L^i \,.. \quad (9)$$

$$\varphi_{21}^p(L) = \sum_{i=1}^{p_{21}} \varphi_{21}^p L^i \qquad \varphi_{22}^p(L) = \sum_{i=0}^{p_{22}} \varphi_{22}^p L^i \,.. \quad (10)$$

Where Δ denotes the difference operator and L denotes the lag operator where $(L)\Delta x_t = \Delta x_{t-1}.ECT_{t-1}$ 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 ECT_{t-1} in equations 8 confirms short-run and long-run causality relationship respectively [26-28].

IV. 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 4 gives the results of the bounds test under two different scenarios as suggested by [2], which with unrestricted deterministic trend (F_V) and without deterministic trend (F_{III}). Intercept in these scenarios are all unrestricted. Critical values for F-statistic are taken from [30] and t-statistic from [2], and presented in Table III. The lag length p for this test is based on Schwarz-Bayessian criterion (SBC). As can be seen from Table IV, F-statistic value confirms cointegration among series in F_{III} and F_V at %1 level significance.

TABLE III 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_V	3.08	4.27	3.67	5.00	5.09	6.77

Source: Reference [30, pp. 1988-1990] for F-statistics and [2, pp.300-301] for *t* ratio.

TABLE IV Bounds F- and t-statistics for the Existence of a Levels Relationship

		With Determintic Trends	**Without Determintic Trends**
	Lag	F_V	F_{III}
$F_{INF}(INF_t \mid M_t, OPN_t, GS_t, Y_t)$	2	9.042***	7.586***

Table V presents long-run coefficients of ARDL(2,0,1,1,2). The coefficients of money growth (M) and GDP per capita (Y) unable to be rejecting at 1%significance level and their coefficients are positive and negative respectively. The coefficient of M is large, therefore it seems that money growth has been hardly increased inflation in Iran, and economic growth reduces inflation. The coefficient of GS is positive and significant at 5% level; therefore we can say that inflation increases with increasing government size in long-run. As can be seen from Table V coefficient of OPN is not significant, and then openness does not have a significant effect on inflation in long run.

TABLE VESTIMATED LONG RUN COEFFICIENTS USING THE ARDL APPROACH

Regressor	Coefficient	Std. Error	t-Statistic	Prob.
M	15.33019	4.381650	3.498725	0.0015
OPN	0.257275	0.163057	1.577820	0.1251
GS	1.964510	0.775745	2.532417	0.0168
Y	-0.043358	0.008311	-5.216879	0.0000
C	26.44367	12.17522	2.171925	0.0379

Table VI presents ECM results. As can be seen, except GS, all variables have significant effect on inflation in short-run. The most important among others is the negative impact of openness on inflation. The coefficient of ECMT (-1), is 1.17, significant at 1% level and negative as be expected, thus approximately all of disequilibria from the previous year's shocks in our model converge back to the long-run equilibrium in less than a year.

TABLE VI ERROR CORRECTION REPRESENTATION FOR THE SELECTED ARDL MODEL

Regressor	Coefficient	Std. Error	t-Statistic	Prob.
DINF(-1)	0.543814	0.125130	4.345982	0.0002
DM	16.92782	3.891225	4.350255	0.0002
DOPN	-0.387091	0.161904	-2.390874	0.0247
DGS	-0.412848	0.534535	-0.772350	0.4471
DY	-0.064066	0.011470	-5.585414	0.0000
DY(-1)	0.064576	0.011586	5.573527	0.0000
C	0.983775	0.882133	1.115223	0.2754
ECMT(-1)	-1.168353	0.134878	-8.662317	0.0000

$R^2 = 0.770$	S.E.R=4.891	F.St=12.00(0.000)	Schwarz.C=6.583
$\overline{R}^2 = 0.706$	RSS=598.245	D.W=2.376	Akaike.C=6.220

TABLE VIIGRANGER CAUSALITY TESTS RESULT

Y / X	M	OPN	GS	Y	INF	ECM(t-1) -- t-stat
Without Deterministic Trend						
M	--	0.174257 (0.8413)	1.162754 (0.3329)	1.275815 (0.3010)	0.309546 (0.7372)	-0.20146 (0.84238)
OPN	0.826908 (0.4518)	--	0.432511 (0.6548)	0.272433 (0.7643)	0.627570 (0.5441)	-0.98643 (0.33571)
GS	1.050469 (0.3683)	0.411041 (0.6684)	--	4.696394 (0.0213)	1.663200 (0.2147)	-1.27602 (0.21657)
Y	1.790722 (0.1926)	5.131920 (0.0159)	2.704849 (0.0913)	--	1.096695 (0.3532)	1.45504 (0.16118)
INF	0.289136 (0.7520)	4.065727 (0.0330)	2.714679 (0.0906)	1.315982 (0.2905)	--	-3.11032 (0.00551)
With Deterministic Trend						
M	--	0.253749 (0.7783)	1.174655 (0.3294)	1.371916 (0.2765)	0.387480 (0.6838)	0.24091 (0.81207)
OPN	0.512721 (0.6065)	--	1.065852 (0.3632)	0.052759 (0.9487)	0.731848 (0.4935)	-1.06337 (0.30028)
GS	0.601755 (0.5575)	0.726844 (0.4958)	--	1.714333 (0.2055)	0.447855 (0.6453)	0.24653 (0.80778)
Y	4.301250 (0.0279)	1.373687 (0.2760)	1.890213 (0.1771)	--	0.082920 (0.9207)	-0.49427 (0.62650)
INF	0.451951 (0.6427)	1.572742 (0.2321)	0.415690 (0.6655)	0.577446 (0.5704)	--	-1.67416 (0.10967)

Table VII presents results of Granger causality tests for the selected ARDL model without deterministic trend and with deterministic trend respectively. As can be seen results confirm long-run causation between independent variable set and inflation in the model without deterministic trend. In this case granger causality from openness to inflation cannot reject in short-run. In the second case results don’t confirm short-run and long-run causation between independent variable set and inflation.

To investigation causality between openness and inflation in Table VIII result of Granger causality tests between OPN and INF has showed. As can be seen results from Without Deterministic Trend and with Deterministic Trend are same totally. T-static in second lines are significant at 5% level, therefore we can say that long-run causality from openness to inflation is exist .t-statistic in first lines are not significant, Thus long-run causality from inflation to openness is not verifiable.

TABLE VIII GRANGER CAUSALITY TESTS BETWEEN OPENNESS AND INFLATION

Y / X	OPN	INF	ECM(t-1) -- t-stat
	Without Deterministic Trend		
OPEN	--	0.192792 (0.8258)	-0.40341 (0.68995)
INF	0.063531 (0.9386)	--	-2.46831 (0.02047)
	With Deterministic Trend		
OPEN	--	0.179248 (0.8369)	-0.35838 (0.72295)
INF	0.060287 (0.9416)	--	-2.47227 (0.02029)

Table IX shows diagnostic tests for ARDL(2,0,1,1,2) model that used in this paper. In this manner Breusch-Godfrey serial correlation LM test and Heteroskedasticity ARCH test are used. LM test indicate that the residuals are not serially correlated and ARCH test shows that the residuals have not Heteroskedasticity problem. The cumulative sum (CUSUM) and cumulative sum of squares (CUSUMQ) plots (Fig. 1) from a recursive estimation of the model also indicate stability in the coefficients over the sample period

TABLE IXARDL(2,0,1,1,2) MODEL DIAGNOSTIC TESTS

Breusch-Godfrey Serial Correlation LM Test:			**Heteroskedasticity Test: ARCH**		
		probe			probe
F-statistic	0. 713894	0. 5024	F-statistic	0.380114	0. 5422
Obs*R-squared	2.306514	0. 3156	Obs*R-squared	0.400382	0. 5269

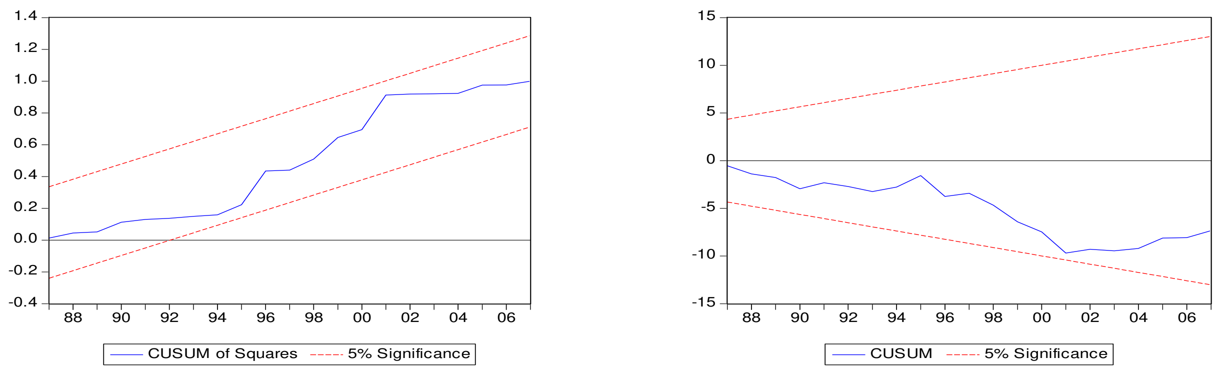

Fig. 1: CUSUM and CUSUMQ tests for coefficients stability of ARDL(2,0,1,1,2) model

V. CONCLUSION

This paper provides evidence on the impact of openness on the inflation in Iran. We apply Bounds test approach to level relationship with in Autoregressive Distributed Lag (ARDL) model proposed by [2] The Results from Bounds test approach confirm existence of long-run relationship among the variables under consideration. The results show that openness has negative and significant effect on inflation in short-run but its effect in long-run is not significant. The coefficient of ECMT(-1) is 1.17, significant at 1% level and negative as be expected, thus approximately all of disequilibria from the previous year's shocks in our model converge back to the long-run equilibrium in less than a year.

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