Financial Development and Income Inequality: Is there any Financial Kuznets curve in Iran?

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Abstract

This deals with the investigation of the relationship between financial development and income inequality in case of Iran. In doing so, we have applied the ARDL bounds testing approach to examine the long-run relationship in the presence of structural break stemming in the series. The unit root properties have been tested by applying Zivot-Andrews (1992) and Clemente et al. (1998) structural break tests. The VECM Granger causality approach is used to detect the direction of causal relationship between financial development and income distribution. Moreover, Greenwood-Jovanovich (GJ) hypothesis has also been tested for Iranian economy.

Our results confirm the long run relationship between the variables. Furthermore, financial development reduces income inequality. Economic growth worsens income inequality, but inflation and globalization improve income distribution. Finally, GJ hypothesis is found as well as U-shaped relationship between globalization and income inequality in case of Iran. This study might provide new insights for policy makers to reduce income inequality by making economic growth more fruitful for poor segment of population and directing financial sector to provide access to financial resources of poor individuals at cheaper cost.

Keywords: Financial Development, Income Inequality, ARDL Bound Testing

Introduction

Higher economic growth with equal income distribution is a great matter of concern for all developing economics; those are trying to catch-up the growth path of developed countries, which is true for Iranian economy too. It has been verified by numerous empirical studies, for different countries, that for a developing country (in particular), which is trying to attain a high economic growth rate, that inequality on various grounds increases with the growth of an economy (Chambers et al. (2007 and, Baliscan and Fuwa, 2005). Our observation on the Gini coefficient and GDP per-capita (see figure 1 and figure 3 respectively) provides a clue for such a situation to exist in Iran too. We find from Figure-1 that the Gini coefficient was increased initially and thereafter it has shown fluctuating trends. The correlation between economic growth and income inequality is positive i.e. 0.2691 and negative i.e. 0998 between financial development and income inequality. By looking into trend of GDP per-capita we observe that it has initially increased, then decreased and now again has moved up word. Recognizing the problems associated with the increasing inequality, Iranian's government has taken various steps to combat with income inequality in order to mitigate negative consequences that might arise due to it. To combat with the inequality a prudential development of financial sector can be used as a big tool. Development and proper management of the financial sectors help in the faster and sustained economic growth. First, for example, easy access to financial resources boosts investment activities that directly increase the income of poor segments of population by generating employment opportunities. Second, easy access to financial resources provides various opportunities and enables the poor segments of population among other to increase human capital formation by investing in education, health and various aspects of socio-economic development of their children and family members. Third, financial development reduces income and wealth inequalities and mitigates various problems, which arises due to increasing inequality of such type and so on and so forth. Last but not least, development and proper management of the financial sector might also be helpful in protecting the indexed income of the elite class via easy access to financial resources during the instances of high inflations since inflation is very harmful for those who earn fixed income as high inflation reduces their purchasing power.

However, as Greenwood and Jovanovich, (1990) argued that initially financial development increases income inequality but declines income inequality once financial sector matures. This seems to be holding of inverted U-shaped hypothesis between financial development and income inequality. There is another mechanism through which financial sector may improve income distribution which is known as "trickle-down effect". According to "trickle-down effect", as economies expand, poverty is likely to be reduced but poverty reduction is likely to be adversely affected due to increased income inequality.

Income inequality is one of those problems that most of less developed countries have been facing for a long time. Slottje and Raj, (1998) showed that in South America and Asia, there is the worst income distribution while in Europe, income inequality is low. By a simple comparison between Iran and North Americas, Europe and Oceans in their study, it can be concluded that income inequality is high in Iran as compared to these regions. Over the years, it is observed that income inequality (Gini-coefficient) has fluctuated in Iran–(see Figure-1). It can be seen that from 1971 to 1975 Gini coefficient in Iran was increased. One of the most important reasons for this was increase in oil shock. After that and until 1978 it decreased slightly due to increase in import and subsidies. From 1979 to 1988 Iran had faced with revolution, war and economical restriction which affected income inequality. From 1985 to 1987 income inequality increased which could be the result of decreasing in oil income. After this period, war is terminated and Gini-coefficient diminished till 1992 but in 1993 Iran faced with the high inflation and again it started to rise. Improvements in income distribution have also been seen after 1997. Figure (2) belongs to real GDP per capita in Iran. This figure shows that most of the time real GDP per capita has an upward trend in Iran. But we didn't see a downward trend in Gini-coefficient and better income distribution was in this period.

Figure-1: Gini Coefficient in Iran

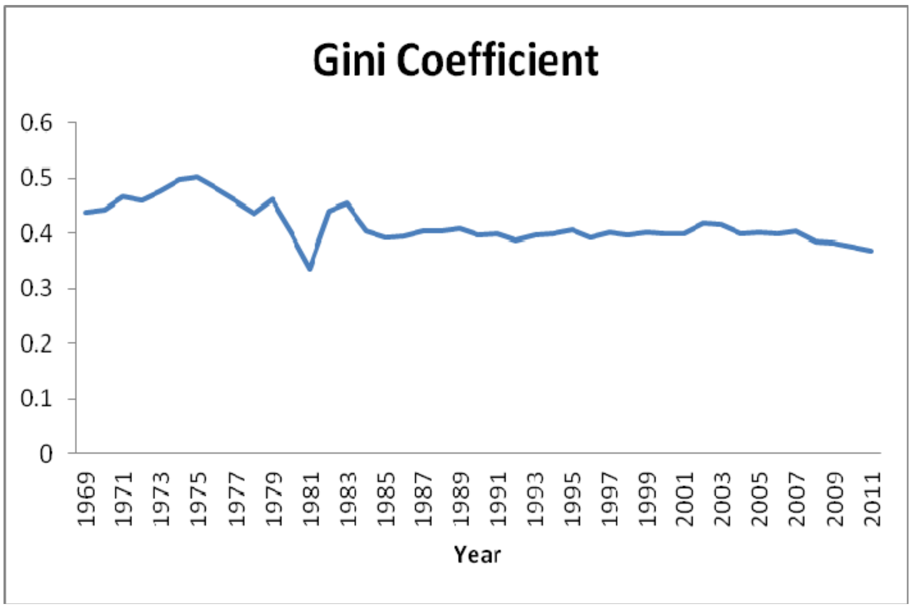


Figure-2: Real GDP Per Capita in Iran

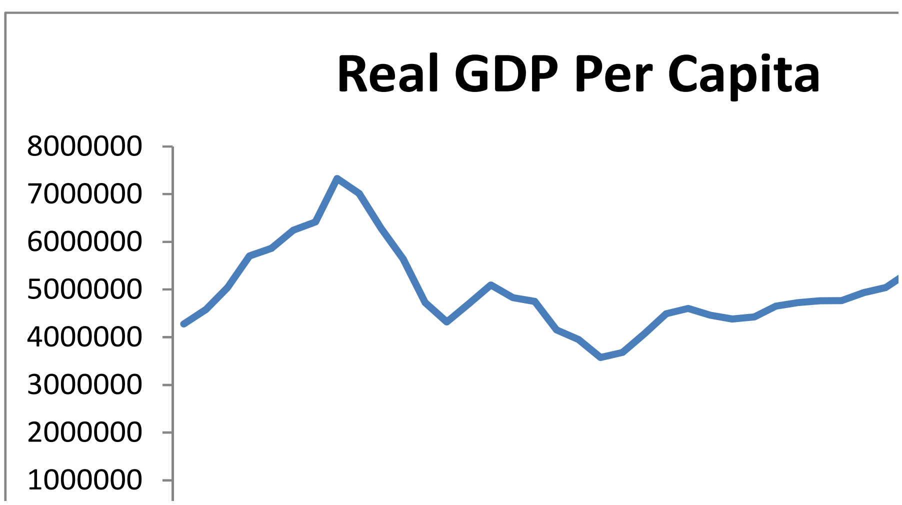
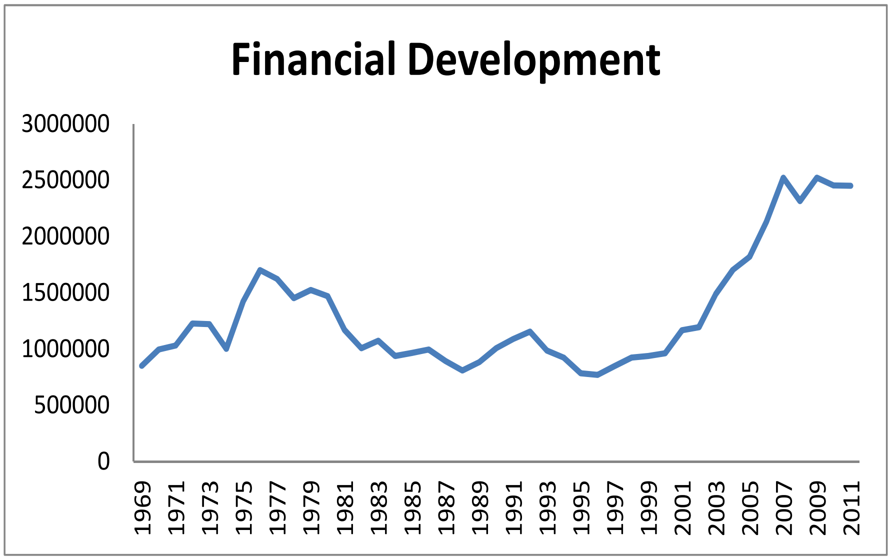


Figure-3: Financial Development in Iran



As it can be seen from figure (2), real GDP per capita rose before Iran's revolution, but after revolution it decreased. Revolution and war on the one hand and increasing in population on the other hand were the main factors for this decline. Increasing in production and diminishing in growth rate of population helped Iran's economy to increase its real GDP per capita in last decade of twentieth century and first decade of third millennium.

Figure (3) shows domestic credit to private sector per capita which is a proxy for financial development in Iran. Financial sector development began deteriorating after 1977 for a decade, remained relatively low in 1994 to1996 but gradually improved in subsequent years. Upward trend can be seen for this variable before the 1977, but after this time it started to decrease. This declining could be because of nationalizing and merging of banks. Moreover, increasing in invisible trade could be another reason. After war, Iran tried to develop his financial sector by launching 5 years development plan. From 1996 we can see an upward trend for this variable because at first, finance and credit institution, and then private banks started their job. Iran in its last 5 years development plan allowed the non-Iranian banks to open their branches to improve the efficiency of financial sector.

In the recent years there is increasing interest of researchers to analyze economic consequences of financial development on income inequality at national and cross-country levels. However, Iran has been departed from such research. The present study is intended to fill this gap. This paper contributes to existing literature by four folds: (i) the nexus between financial development and income inequality is investigated by using time series data in case of Iran, (ii), unit root properties of the variables have been examined by applying structural break unit root tests such as Zivot-Andrews (1992) and Clemente et al. (1998), (iii), in doing so, we have applied the structural break ARDL bounds testing approach to cointegration for long run relationship between the variables and, (iv) the VECM Granger causality is applied to test causal relation between the variables.

This paper is structured as follows. Section-II, presents a brief review of literature on relationship between financial development and income inequality. Modeling, methodological framework and data collection are presented in Section-III. Section-IV deals with results interpretation, and Section-V draws conclusion and policy recommendations.

II: Literature Review

Over the last three decades, there is growing interest of researchers on analyzing the financial development and economic growth (Pagano, (1993); Levine, (1997); Levine et al. (2000); Anderson and Tarp, (2003); Jalilian and Kirkpatrick, (2005)). Levine, (1997) confirms that long run economic growth has been experienced by those economies which have well developed banking system. However, theoretical concern is unclear in this aspect. But, Kirkpatrick, (2000) showed the role of well-functioning financial system in mobilization of savings, resource allocation, and facilitation of risk management which in turn provides support for capital accumulation, improves efficiency of investment and promotes innovations in technology and hence contributes to economic growth. Similarly; Goldsmith, (1969); Mckinnon, (1973); King and Levine, (1993); Pagano and Volpin, (2001); Christodoulou and Tsionas, (2004); Shan, (2005); Ma and Jalil, (2008) and Shahbaz et al. (2010) paid their attention to identify the degrees as well as effectiveness of financial development on sustained economic growth, physical capital accumulation and economic efficiency.

Our concern is to discuss the relationship between financial development and income inequality. There are various studies which have highlighted various aspect of association of financial development and income inequality. For example, Galor and Zeira (1993), and Banerjee and Newman (1993) have highlighted that financial markets particularly credit market improve income distribution. They suggested that the initial income gap would not be reduced unless financial markets are sound. Similarly, Canavire-Bacarreza and Rioja, (2009) document that "given their lack of collateral and scant credit histories, poor entrepreneurs may be the most affected by financial market imperfections such as information asymmetries, contract enforcement costs, and transactions costs".

There are some other ways also through which financial development may increase income inequality. For example, as Behrman et al. (2001); Dollar and Karaay, (2003); Beck et al. (2004) mentioned that in the early stage of financial development, financial sector may charge high set up cost against financial services during to gain advantages from the screening and risk pooling which is beyond the affordability of poor individuals. Hence, poor individuals are unable to come out from the circle of income inequality. Further, deficiencies in money markets in terms of asymmetric information, intermediation and transaction costs restrict the poor people to attain loans from financial institutions because they do not have collateral, credit records and political; and personal connections with high authorities of financial sector to get loans at reasonable interest rate. Hence, even if there is enough funds to be distributed at reasonable rate of interest among poor people then they are unable to avail benefit of such services. Claessense, (2006) and Perotti, (1996) provided another reason due to which poor people are unable to access the benefit of financial development. They argued that since poor individuals are not much educated and formal financial sector does not seem to prefer such un-educated or less-educated persons to offer loans and hence in many high income countries, financial sector has dualism in financial services.

Galor and Zeira, (1993) argued that access of poor entrepreneurs to financial resources enables them to start small to enhance their earnings. This not only reduces income inequality and hence declines poverty. On contrarily, Bourguignon and Verdier, (2000) noticed that since in almost cases, poor rely more on informal networks for credit hence, financial development would only benefit the rich class of the society and raises income inequality. Greenwood and Jovanovich, (1990) proposed a non-linear relationship between financial development and income inequality or what we may call as “inverted-U” hypothesis. They argued that initially financial development increases income inequality and improves income distribution once financial sector matures.

Furthermore; Westley, (2001) investigated the impact of financial markets on income distribution for Latin American countries in panel framework and reported that easy access to financial resources through micro finance policies reduces income inequality. Calderon and

Serven, (2003) disclosed that financial development worsens income distribution while education improves it. Similarly, Lopez, (2004) also found that better education and low prices seem to decrease income inequality. Financial development, international trade and government size hamper income distribution. Similarly; Honohan, (2004); Beck et al. (2004); Stijn and Perotti, (2007) noticed that financial development and income inequality is not only a correlation but also a causal relationship between both variables. For example, positive impact of financial development on economic growth may enable the poor segments of population to demand for loans from financial markets to increase their income levels as economy grows. However, Beck et al. (2007a) documented that strong relationship between finance and growth does not necessarily mean that financial development improves income distribution and hence reduces poverty. They claimed that financial development will help decline poverty if financial development increases average income of both rich and poor segments with of population. Financial development will help the poor if average income is higher achieved by rich class. On the other hand, Li et al. (1998) found that financial development lowers income inequality by raising the average income of bottom 20% population. Beck et al. (2007b) using cross-country data, found that financial development raises income of poor segment of population disproportionately and reduces income inequality. On contrary; Bonfiglioli, (2005) used cross-country data to examine the impact of financial development proxies by stock market development on income inequality and concluded that financial development has progressive effect on income inequality.

In case country studies; Liang, (2006) reported that financial development improves urban income distribution in post-reform China. In case of Malaysia; Law and Tan, (2009) examined the role of financial development in affecting income inequality. They used stock market capitalization and domestic credit to private sector proxy for financial development.

Their results supported favorable impact of financial development on income distribution while inflation raises income inequality. Shahbaz, (2009) used Pakistani data to examine the impact of financial development and financial instability on the income of bottom 20% population. The results indicted that financial development increases the income poor segment of population but this effect is nullified by financial instability. In case of India; Ang, (2010) investigated relationship between both variables and concluded that financial development helps reduce income inequality but financial liberalization deteriorates income distribution. Using Brazilian data, Bittencourt, (2010) investigated the impact of financial development on income inequality and found that financial development declines income inequality by increasing income bottom 20% population. Shahbaz and Islam, (2011) probed the relationship between financial development and income distribution in the presence of financial instability in case of Pakistan. Their results indicated that financial development declines income inequality while financial instability worsens income distribution.

Moreover; Wahid et al. (2011) found that financial development increases income inequality in case of Bangladesh. Furthermore, results revealed that economic growth improves income distribution suggesting that improvements in economic growth redistribute income and make the society more egalitarian. Using Bayesian structural autoregressive model (SVAR), Gimet and Lagoarde-Segot, (2011) reexamined the relationship between financial development and income inequality. They uncovered that financial development Granger causes income distribution. In case of China, Jalil and Feridun, (2011) questioned whether financial development improves income distribution or not. Their results accepted inequality narrowing hypothesis implying that financial development reduces income inequality. In case of Indian states, Arora, (2012) raised the issue of finance-inequality nexus for empirical investigation. The results showed that overall income inequality is deteriorated with financial development. Financial development improves inequality in rural but raises inequality in urban areas. Yu and Qin, (2011) also supported the fact that financial development helps to reduce rural-urban income gap in China. Similarly, Chun and Peng, (2011) reported favorable impact of financial development on income distribution. They suggested that government should loosen financial regulations, and lower market anticipation level to ensure the whole society can take advantage of economy development; open the financial market to higher degree, and promote the competition; accelerate interest rate marketization; build up a financial system which facilitates SMEs financing; develop micro-financial institutions and micro loans; develop technology and its application in financial areas, in order to lower financial cost; develop the financial industry support on human capital investment. Iyigun and Owen, (2012) found that financial development affects income inequality by controlling aggregate consumption variability. In low income countries, income inequality is linked with more consumption volatility and vice versa in high income countries. Hamori and Hashiguchi, (2012) documented that impact of financial development on income inequality depends on the choice of financial variables.

Various studies are available investigating GJ (1990) hypothesis between financial development and income inequality. For example; Li et al. (2008) investigated the relationship between financial development and income inequality and confirmed the existence of U-shaped Kuznets curve for East Asian countries while Rehman et al. (2008), while working on in similar line; reject inverted U-shaped relationship between financial development and income inequality. Sebastian and Sebastian, (2011) probed the relationship between financial development and income inequality by applying fixed effects model[1]. Their

[1] Albania, Algeria, Angola, Argentina, Armenia, Australia, Austria, Bahamas, Barbados, Belize, Bhutan, Bolivia, Belgium, Botswana, Brazil, Bulgaria, Canada, Chile, Colombia, Cameroon, Cape Verde, Cote d'Ivoire,Costa Rica, Croatia, Cyprus, Czech Republic, Dominica, Dominican Republic, Denmark, Ecuador, Egypt, Arab Rep. El Salvador, Estonia, Finland, Fiji, France, Gabon, Germany, Greece, Georgia, Guatemala,

results indicated that financial development worsens income inequality but could not find existence of GJ (1990) hypothesis between both the variables. Kim and Lin, (2011) noted that financial development improves income distribution if country achieves the threshold level of financial development and below this level financial development worsens income inequality i.e. GJ (1990) hypothesis and same inference is drawn by Rötheli, (2011). Shahbaz and Islam, (2011) also found U-shaped relationship between financial development and income inequality in Pakistan but it is statistically insignificant.

Batuo et al. (2012) investigated the empirical existence of GJ (1990) hypothesis using data of African countries applying dynamic panel estimation technique (GMM)[2]. They found that financial development has positive impact on income distribution but could not find evidence supporting the GJ (1990) hypothesis or inverted U-shaped relationship between financial development and income inequality. Nikoloski, (2012) investigated the linear and non-linear relationship between financial development and income inequality applying system generalized moments method (GMM)[3]. His empirical evidence proved the existence of inverted-shaped relationship between financial development and income equality i.e. GJ (1990) hypothesis. Tan and Law, (2012) investigated the dynamics of finance-inequality

Guyana, Grenada, Hong Kong, Hungary, Honduras, Iceland, India, Indonesia, Iran, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Korea, Rep. Latvia, Lithuania, Lesotho, Luxembourg, Malta, Macedonia, Malaysia, Mauritius, Mexico, Moldova, Mongolia, Morocco, Netherlands, New Zealand, Nigeria, Norway, Pakistan, Papua New Guinea, Paraguay, Philippines, Panama, Peru, Poland, Portugal, Romania, Russian Federation, Senegal, Sri Lanka, Swaziland, Serbia, Seychelles, Singapore, Slovak Republic, Slovenia, Spain, South Africa, St. Lucia, St. Vincent and the Gren, Suriname, Sweden, Switzerland, Turkey, Thailand, Tunisia, Trinidad a. Tobago, United Kingdom, United States, Uruguay, Venezuela RB, Vietnam, Yemen, Rep.

[2] Botswana, Ivory Coast, Cameroon, Egypt, Ethiopia, Ghana, Kenya, Lesotho, Morocco, Madagascar, Mauritania, Mauritius, Malawi, Nigeria, Senegal, Sierra Leone, South Africa, Tanzania, Tunisia, Uganda, Zambia and Zimbabwe.

[3] Argentina, Armenia, Australia, Austria, Barbados, Belarus, Belgium, Bolivia, Botswana, Bulgaria, Canada, Chile, China, Colombia, Costa Rica, Croatia, Cyprus, Czech Republic, Denmark, Dominican Republic, Ecuador, El Salvador, Estonia, Finland, France, Gabon, Georgia, Germany, Greece, Guatemala, Honduras, Hungary, Ireland, Israel, Italy, Japan, Kazakhstan, Kenya, Kyrgyzstan, Latvia, Lesotho, Lithuania, Luxembourg, Macedonia, Malta, Mauritius, Mexico, Moldova, Nepal, Netherlands, New Zealand, Nicaragua, Norway, Panama, Paraguay, Peru, Poland, Portugal, Romania, Russia, Serbia and Montenegro, Slovakia, Slovenia, South Africa, South Korea, Spain, Sweden, Switzerland, Tajikistan, Turkey, Ukraine, United Kingdom, United States, Uzbekistan, Venezuela and Zambia.

nexus using data of 35 countries[4]. Their results indicated U-shaped relationship between financial deepening and income distribution. This implies that financial markets are inefficient to improve income distribution in these countries. In case of China, Ling-zheng and Xia-hai, (2012) applied threshold model developed by Hansen, (1999) using provincial data to investigate the relationship between financial development and income inequality. Their results disclosed that financial development deteriorates income inequality and supported the existence of U-shaped relationship between both variables.

III- Modeling, Methodological Framework and Data Collection

The objective of this study is to examine the relationship between financial development and income inequality including economic growth, inflation and globalization are other potential determinates of income inequality in case of Iran. The general functional form of model is given below as following:

$$IE_t = f(Y_t, F_t, IN_t, G_t) \tag{1}$$

In this equation, IE_t is income inequality, Y_t shows economic growth, F_t illustrates financial development, IN_t represents inflation, and G_t is globalization. We have converted all the series into logarithm for consistent and reliable results. The log-linear specification provides better results because conversion of the series into logarithm reduces the sharpness in time series data (Shahbaz, 2010). The empirical equation is modeled as following:

$$\ln IE_t = \theta_1 + \theta_2 \ln Y_t + \theta_3 \ln F_t + \theta_4 \ln IN_t + \theta_5 \ln G_t + \varepsilon_i \tag{2}$$

[4] Algeria, Bangladesh, Bolivia, Cameroon, Chile, Colombia, Ecuador, Egypt, Ghana, Guatemala, India, Indonesia, Jamaica, Jordan, Kenya, Korea, Madagascar, Malawi, Malaysia, Mexico, Morocco, Pakistan, Peru, Philippines, Papua New Guinea, Senegal, South Africa, Sri Lanka, Syria, Thailand, Trinidad and Tobago, Turkey, Uruguay, Venezuela, Zimbabwe.

where, $\ln IE_t$, $\ln Y_t$, $\ln F_t$, $\ln IN_t$, $\ln G_t$ is natural log of income inequality proxies by Gini-coefficient, natural log of economic growth measured by real GDP per capita, natural log of financial development captured by real domestic credit to private sector per capita, natural log of inflation proxies by consumer price index, natural log of globalization measured by KOF globalization index (following Dreher, 2006). ε is residual term containing normal distribution with finite variance and zero mean. To test the GJ hypothesis following non-linear specification is considered:

$$\ln IE_t = \theta_{11} + \theta_{22} \ln Y_t + \theta_{33} \ln F_t + \theta_{44} \ln F_t^2 + \theta_{55} \ln IN_t + \theta_{66} \ln G_t + \varepsilon_t \quad (3)$$

Equation-3 envisages inequality reducing hypothesis if $\theta_{33} < 0$ keeping $\theta_{44} = 0$. Income inequality increases if $\theta_{33} = 0$ and $\theta_{44} > 0$. The GJ (1990) hypothesis would be confirmed if $\theta_{33} > 0$ and $\theta_{44} < 0$ otherwise U-shaped relationship between financial development and income inequality is accepted if $\theta_{33} < 0$ and $\theta_{44} > 0$. Similarly, nonlinear relationship between globalization and income inequality is investigated by including squared term of $\ln G_t$ i.e. $\ln G_t^2$. The empirical equation is modelled as following:

$$\ln IE_t = \beta_{11} + \beta_2 \ln Y_t + \beta_{33} \ln F_t + \beta_{44} \ln IN_t + \beta_{55} \ln G_t + \beta_{66} \ln G_t^2 + \varepsilon_t \quad (4)$$

The inverted-U shaped theory would be accepted if $\beta_{55} > 0$ and $\beta_{66} < 0$ otherwise U-shaped relationship between globalization and income inequality is accepted if $\beta_{55} < 0$ and $\beta_{66} > 0$.

Numerous unit root tests are available on applied economics to test the stationarity properties of the variables. These unit tests are ADF by Dickey and Fuller (1979), P-P by Philips and Perron (1988), KPSS by Kwiatkowski et al. (1992), DF-GLS by Elliott et al. (1996) and Ng-

Perron by Ng-Perron (2001). These tests provide biased and spurious results due to not having information about structural break points occurred in series. In doing so, Zivot-Andrews (1992) developed three models to test the stationarity properties of the variables in the presence of structural break point in the series: (i) this model allows a one-time change in variables at level form, (ii) this model permits a one-time change in the slope of the trend component i.e. function and (iii) model has one-time change both in intercept and trend function of the variables to be used for empirical propose. Zivot-Andrews (1992) followed three models to check the hypothesis of one-time structural break in the series as follows:

$$\Delta x_t = a + a x_{t-1} + bt + cDU_t + \sum_{j=1}^{k} d_j \Delta x_{t-j} + \mu_t \qquad (5)$$

$$\Delta x_t = b + b x_{t-1} + ct + bDT_t + \sum_{j=1}^{k} d_j \Delta x_{t-j} + \mu_t \qquad (6)$$

$$\Delta x_t = c + c x_{t-1} + ct + dDU_t + dDT_t + \sum_{j=1}^{k} d_j \Delta x_{t-j} + \mu_t \qquad (7)$$

Where dummy variable is indicated by DU_t showing mean shift occurred at each point with time break while trend shift variables is show by DT_t[5]. So,

$$DU_t = \begin{cases} 1 \ldots if \quad t > TB \\ 0 \ldots if \quad t < TB \end{cases} \text{ and } DU_t = \begin{cases} t - TB \ldots if \ t > TB \\ 0 \ldots if \ t < TB \end{cases}$$

The null hypothesis of unit root break date is $c = 0$ which indicates that series is not stationary with a drift not having information about structural break point while $c < 0$ hypothesis implies that the variable is found to be trend-stationary with one unknown time break. Zivot-Andrews unit root test fixes all points as potential for possible time break and does estimation through regression for all possible break points successively. Then, this unit

[5] We used model-4 for empirical estimations following Sen (2003)

root test selects that time break which decreases one-sided t-statistic to test $\hat{c}(=c-1)=1$. Zivot-Andrews intimates that in the presence of end points, asymptotic distribution of the statistics is diverged to infinity point. It is necessary to choose a region where end points of sample period are excluded. Further, Zivot-Andrews suggested the trimming regions i.e. (0.15T, 0.85T) are followed.

The Clemente et al. (1998) test is better suited when problems are due to structural break. This test has more power, compared to the Perron and Volgelsang (1992), Zivot-Andrews (1992), ADF, PP and Ng-Perron unit root tests. Perron and Volgelsang (1992) and Zivot-Andrews (1992) unit root tests are appropriate if the series has one potential structural break. Clemente et al. (1998) extended the Perron and Volgelsang (1992) method to allow for two structural breaks in the mean. The null hypothesis H_0 against alternate H_a is stated as follows:

$$H_0 : x_t = x_{t-1} + a_1 DTB_{1t} + a_2 DTB_{2t} + \mu_t \tag{8}$$

$$H_a : x_t = u + b_1 DU_{1t} + b_2 DTB_{2t} + \mu_t \tag{9}$$

In equation-8 and equation-9, DTB_{1t} is the pulse variable which equals 1 if $t = TB_i + 1$ and zero otherwise. Moreover, $DU_{it} = 1$ if $TB_i < t (i = 1,2)$ and zero otherwise. Modification of mean is represented by TB_1 and TB_2 time periods. To further simplify, we assume that $TB_i = \delta_i T (i = 1,2)$ where $1 > \delta_i > 0$ while $\delta_1 < \delta_2$ (see Clemente et al. 1998). If two structural breaks are contained by innovative outlier, then unit root hypothesis can be investigated by applying equation-8, as provided in the following model:

$$x_t = u + \rho x_{t-1} + d_1 DTB_{1t} + a_2 DTB_{2t} + d_3 DU_{1t} + d_4 DU_{2t} + \sum_{i=1}^{k} c_j \Delta x_{t-1} + \mu_t \qquad (10)$$

This equation helps us to estimate minimum value of t-ratio through simulations and the value of simulated t-ratio can be utilized to identify all break points if the value of autoregressive parameter is constrained to 1. For the derivation of the asymptotic distribution of the estimate, we assume that $\delta_2 > \delta_1 > 0, 1 > \delta_2 - 1 > \delta_0$ where, δ_1 and δ_2 obtain the values in interval i.e. $[(t+2)/T, (T-1)/T]$ by applying the largest window size. The assumption i.e. $\delta_1 < \delta_2 + 1$ is used to show that cases where break points exist in repeated periods are purged (see Clemente et al. 1998). Two steps approach is used to test the unit root hypothesis, if shifts can explain the additive outliers. In 1st step, we remove deterministic trend, following equation-8 for estimation as follows:

$$x_t = u + d_5 DU_{1t} + d_6 DU_{2t} + \hat{x} \qquad (11)$$

The second step involves search for the minimum t-ratio to test the hypothesis that $\rho = 1$, using the following equation:

$$\hat{x}_t = \sum_{i=1}^{k} \phi_{1i} DTB_{1t-1} + \sum_{i=1}^{k} \phi_{2i} DTB_{2t-1} + \rho \hat{x}_{t-1} + \sum_{i=1}^{k} c_i \Delta \hat{x}_{t-1} + \mu_t \qquad (12)$$

To make sure that the $\min t_{\rho\ t}^{IO}(\delta_1, \delta_2)$ congregates i.e. converges in distribution, we have included dummy variable in estimated equation for estimation: $\min t_{\rho\ t}^{IO}(\delta_1, \delta_2) \rightarrow \inf_{\gamma} = \wedge \dfrac{H}{[\delta_1(\delta_2 - \delta_1)]^{1/2} K^{1/2}}$

Avoiding traditional approaches to cointegration due to their demerits, we apply the structural break autoregressive distributed lag model or the ARDL bounds testing approach to cointegration in the presence of structural breaks in the series. The ARDL bounds testing approach to cointegration is preferred due to its certain advantages. For example, the ARDL bounds testing is flexible regarding the integrating order of the variables whether variables are found to be stationary at I(1) or I(0) or I(1) / I(0). The Monte Carlo investigation shows that this approach is superior and provides consistent results for small sample (Pesaran and Shin, 1999). Moreover, a dynamic unrestricted error correction model (UECM) can be derived from the ARDL bounds testing through a simple linear transformation. The UECM integrates the short run dynamics with the long run equilibrium without losing any information for long run. The empirical formulation of the ARDL bounds testing approach to cointegration is given below:

$$\begin{aligned}\Delta \ln IE_t = {} & \alpha_1 + \alpha_T T + \alpha_D D + \alpha_{IE} \ln IE_{t-1} + \alpha_Y \ln Y_{t-1} + \alpha_F \ln F_{t-1} + \alpha_{IN} \ln IN_{t-1} + \alpha_G \ln G_{t-1} + \sum_{i=1}^{p} \alpha_i \Delta \ln IE_{t-i} \\ & + \sum_{j=0}^{q} \alpha_j \Delta \ln Y_{t-j} + \sum_{k=0}^{r} \alpha_k \Delta \ln F_{t-k} + \sum_{l=0}^{s} \alpha_l \Delta \ln IN_{t-l} + \sum_{m=0}^{t} \alpha_m \Delta \ln G_{t-m} + \mu_t \end{aligned} \tag{13}$$

$$\begin{aligned}\Delta \ln Y_t = {} & \alpha_1 + \alpha_T T + \alpha_D D + \alpha_{IE} \ln IE_{t-1} + \alpha_Y \ln Y_{t-1} + \alpha_F \ln F_{t-1} + \alpha_{IN} \ln IN_{t-1} + \alpha_G \ln G_{t-1} + \sum_{i=1}^{p} \beta_i \Delta \ln Y_{t-i} \\ & + \sum_{j=0}^{q} \beta_j \Delta \ln IE_{t-j} + \sum_{k=0}^{r} \beta_k \Delta \ln F_{t-k} + \sum_{l=0}^{s} \beta_l \Delta \ln IN_{t-l} + \sum_{m=0}^{t} \beta_m \Delta \ln G_{t-l} + \mu_t \end{aligned} \tag{14}$$

$$\begin{aligned}\Delta \ln F_t = {} & \alpha_1 + \alpha_T T + \alpha_D D + \alpha_{IE} \ln IE_{t-1} + \alpha_Y \ln Y_{t-1} + \alpha_F \ln F_{t-1} + \alpha_{IN} \ln IN_{t-1} + \alpha_G \ln G_{t-1} + \sum_{i=1}^{p} \beta_i \Delta \ln F_{t-i} \\ & + \sum_{j=0}^{q} \beta_j \Delta \ln IE_{t-j} + \sum_{k=0}^{r} \beta_k \Delta \ln Y_{t-k} + \sum_{l=0}^{s} \beta_l \Delta \ln IN_{t-l} + \sum_{m=0}^{t} \beta_m \Delta \ln G_{t-m} + \mu_t \end{aligned} \tag{15}$$

$$\begin{aligned}\Delta \ln IN_t = {} & \alpha_1 + \alpha_T T + \alpha_D D + \alpha_{IE} \ln IE_{t-1} + \alpha_Y \ln Y_{t-1} + \alpha_F \ln F_{t-1} + \alpha_{IN} \ln IN_{t-1} + \alpha_G \ln G_{t-1} + \sum_{i=1}^{p} \vartheta_i \Delta \ln IN_{t-i} \\ & + \sum_{j=0}^{q} \vartheta_j \Delta \ln IE_{t-j} + \sum_{k=0}^{r} \vartheta_k \Delta \ln Y_{t-k} + \sum_{l=0}^{s} \vartheta_l \Delta \ln F_{t-l} + \sum_{m=0}^{t} \vartheta_m \Delta \ln G_{t-m} + \mu_t \end{aligned} \tag{16}$$

$$\Delta \ln G_t = \alpha_1 + \alpha_T T + \alpha_D D + \alpha_{IE} \ln IE_{t-1} + \alpha_Y \ln Y_{t-1} + \alpha_F \ln F_{t-1} + \alpha_{IN} \ln IN_{t-1} + \alpha_G \ln G_{t-1} + \sum_{i=1}^{p} \rho_i \Delta \ln G_{t-i}$$
$$+ \sum_{j=0}^{q} \rho_j \Delta \ln IE_{t-j} + \sum_{k=0}^{r} \rho_k \Delta \ln Y_{t-k} + \sum_{l=0}^{s} \rho_l \Delta \ln F_{t-l} + \sum_{m=0}^{t} \rho_m \Delta \ln IN_{t-m} + \mu_t \quad (17)$$

Where, Δ is difference operator, μ_s are residual terms and D is dummy variable to capture the structural breaks stemming in the series[6]. Here, we compute F-statistic to compare with critical bounds generated by Pesaran et al. (2001) to test whether cointegration exists or not. Pesaran et al. (2001) developed upper critical bound (UCB) and lower critical bound (LCB). We use F-test to examine the existence of long run relationship between the variables following null hypothesis i.e. $H_0 : \alpha_{IE} = \alpha_Y = \alpha_F = \alpha_{IN} = \alpha_G = 0$ against alternate hypothesis ($H_1 : \alpha_{IE} \neq \alpha_Y \neq \alpha_F \neq \alpha_{IN} \neq \alpha_G = 0$) of cointegration for equation-4. The F-test is non-standard and we may use LCB and UCB developed by Pesaran et al. (2001). Using Pesaran et al. (2001) critical bounds, there is cointegration between the variables if computed F-statistic is more than upper critical bound (UCB). The variables are not cointegrated for long run relationship if computed F-statistic does not exceed lower critical bound (LCB). If computed F-statistic falls between lower and upper critical bounds then decision regarding cointegration between the variables is uncertain. The critical bounds generated by Pesaran et al. (2001) may be inappropriate for small sample like ours case which has 43 observations in case of Iran. Therefore, we use lower and upper critical bounds developed by Narayan (2005). The stability tests, to scrutinize stability of the ARDL bounds testing estimates, have been applied i.e. CUSUM and CUSUMSQ (Brown et al. 1975).

The ARDL bounds testing approach can be used to estimate long run relationships between the variables. For instance, if there is cointegration in equation-4 where income inequality

[6] The structural breaks are based on Clemente et al. (1998)

(IE_t), financial development (F_t), inflation (IN_t) and globalization (G_t) are used as forcing variables then there is established long run relationship between the variables that can be molded in following equation given below:

$$\ln IE_t = \theta_0 + \theta_1 \ln Y_t + \theta_2 \ln F_t + \theta_3 \ln IN_t + \theta_4 \ln G_t + \mu_i \tag{18}$$

where $\theta_0 = -\alpha_1 / \alpha_{IE}, \theta_1 = -\alpha_Y / \alpha_1, \theta_2 = -\alpha_F / \alpha_1, \theta_3 = -\alpha_{IN} / \alpha_1, \theta_4 = -\alpha_G / \alpha_1$ and μ_t is the error term supposed to be normally distributed. These long run estimates are computed using the ARDL bounds testing approach to cointegration when income inequality (IE_t) treated dependent variables. This process can be enhanced by using other variables as dependent ones. Once, long run relationship is found between the variables, next is to test direction of causality between the variables following error correction representation given below:

$$(1-L)\begin{bmatrix} \ln IE_t \\ \ln F_t \\ \ln Y_t \\ \ln IN_t \\ \ln G_t \end{bmatrix} = \begin{bmatrix} a_1 \\ a_2 \\ a_3 \\ a_4 \\ a_5 \end{bmatrix} + \sum_{i=1}^{p}(1-L)\begin{bmatrix} b_{11i}b_{12i}b_{13i}b_{14i}b_{15i} \\ b_{21i}b_{22i}b_{23i}b_{24i}b_{25i} \\ b_{31i}b_{32i}b_{33i}b_{34i}b_{35i} \\ b_{41i}b_{42i}b_{43i}b_{44i}b_{45i} \\ b_{51i}b_{52i}b_{53i}b_{54i}b_{55i} \end{bmatrix} \times \begin{bmatrix} \ln IE_{t-1} \\ \ln F_{t-1} \\ \ln Y_{t-1} \\ \ln IN_{t-1} \\ \ln G_{t-1} \end{bmatrix} + \begin{bmatrix} \alpha \\ \beta \\ \delta \\ \phi \\ \varphi \end{bmatrix} ECT_{t-1} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \\ \varepsilon_{4t} \\ \varepsilon_{5t} \end{bmatrix} \tag{19}$$

Where difference operator is indicated by $(1-L)$ and ECT_{t-1} is lagged residual term generated from long run relationship while $\varepsilon_{1t}, \varepsilon_{2t}, \varepsilon_{3t}, \varepsilon_{4t}$, and ε_{5t} are error terms assumed to be normally distributed with mean zero and finite covariance matrix. The long run causality is indicated by the significance of t-statistic connecting to the coefficient of error correction term (ECT_{t-1}) and statistical significance of F-statistic in first differences of the variables shows the evidence of short run causality between variables of interest. Additionally, joint long-and-short runs causal relationship can be estimated by joint significance of both ECT_{t-1} and the estimate of lagged independent variables. For instance, $b_{12,i} \neq 0 \forall_i$ shows that financial development Granger-causes income inequality and causality is running from income inequality to financial development indicated by $b_{21,i} \neq 0 \forall_i$.

The study covers the period of 1965-2011. The data on real GDP, real domestic credit to private sector, Gini-coefficient (income inequality), consumer price index (inflation) has been sourced from world development indicators (CD-ROM, 2012). The KOF globalization index is borrowed from Dreher, (2006).

IV: Empirical Results and their Discussion

Stationary tests are among the most important tests to estimate regression with reliable coefficients. Stationary tests are also used to avoid spurious regression results. We have applied two tests for determining the stationarity properties of the variables. These tests are ADF developed by Dickey-Fuller (1981) and PP by Philips-Peron (1988). The null hypothesis of both tests reveals that there is unit root problem in the series. The results are of both are reported in Table-1. It can be concluded that all the variables have unit root in level, because the calculated statistics are not bigger than the critical values confirmed by probability values and the null hypothesis cannot be rejected. The null hypothesis of unit root problem is rejected at the first difference. This shows that variables are found to be stationary at 1st difference implying that variables are integrated at I(1).

Table-1: Unit Root Analysis

Variables	ADF Unit Root Test		P-P Unit Root Test	
	T-statistic	Prob. values	T-statistic	Prob. values#
$\ln IE_t$	-2.1195 (2)	0.5196	-2.2198(3)	0.2125
$\ln Y_t$	2.0100 (1)	0.5787	-1.4990 (3)	0.8142
$\ln F_t$	-1.1181 (1)	0.9134	-1.1529 (3)	0.9072
$\ln IN_t$	-2.9720 (2)	0.1520	-2.7361 (3)	0.2282
$\ln G_t$	-1.6859 (1)	0.7390	-1.617 (3)	0.7500
$\Delta \ln IE_t$	-8.1023 (1)*	0.0000	-8.0260 (3)*	0.0000
$\Delta \ln Y_t$	-3.5497 (1)*	0.0475	-3.5355 (3)*	0.0491
$\Delta \ln F_t$	-4.3091 (2)*	0.0077	-5.3795 (6)*	0.0004
$\Delta \ln IN_t$	-5.3421 (3)*	0.0005	-7.9863 (3)*	0.0000
$\Delta \ln G_t$	-4.6350 (0)*	0.0032	-4.6978 (3)*	0.0027

Note: * indicates significance at 1% level. Optimal lag order for ADF and bandwidth for PP unit root tests is determined by Schwert (1989) formula. The critical values of ADF and PP tests are -4.2191, -3.5330 and -3.1983 at 1%, 5% and 10% respectively. # MacKinnon (1996) one-sided p-values.

The results of ADF and PP unit root tests may be biased and inappropriate because both do not have information about structural break stemming in the series. This deficiency of ADF and PP tests has been covered by applying Zivot-Andrews, (1992) and Clemente et al. (1998) structural break unit root tests. Former contains information about one structural break and latter has information about two structural breaks stemming in the series. The results for Zivot and Andrew, (1992) unit root test are presented in Table-2. These results suggest that we cannot reject the null of unit root for these variables in level at 1% level but at 1st difference, it is possible to reject null hypothesis of unit root for all the variables.

Table-2: Zivot-Andrews Unit Root Test

Variable	At Level		At 1st Difference	
	T-statistic	Time Break	T-statistic	Time Break
$\ln IE_t$	-3.660(2)	1980	-12.304(1)*	1982
$\ln Y_t$	-4.298 (1)	1986	-6.410(2)*	1977
$\ln F_t$	-3.493 (0)	1993	-6.186 (0)*	1977
$\ln IN_t$	-4.011 (1)	1997	-7.492 (1)*	1986
$\ln G_t$	-3.238 (1)	1979	-5.940 (0)*	1981
Note: * represents significance at 1% level. Lag order is shown in parenthesis.				

To test the robustness of stationarity properties of the variables, Clemente et al. (1998) unit root test is also applied, which provides more consistent and reliable results as compared to Zivot-Andrews, (1992) unit root test. The main advantage of Clemente-Montanes-Reyes, (1998) unit root test is that it has information about two unknown structural breaks in the series by offering two models i.e. an additive outliers (AO) model informs about a sudden change in the mean of a series and an innovational outliers (IO) model indicates about the gradual shift in the mean of the series. The additive outlier model is more suitable for the variables having sudden structural changes as compared to gradual shifts.

Table-3 reports the results of Clemente et al. (1998) unit root test. The results reveal that all the variables have unit root at level but to found to be stationary at 1^{st} difference in the presence of various structural breaks. Unit root tests show that none of the variable is integrated at (2) or beyond that order of integration. The computation of the ARDL F-statistic for cointegration becomes unacceptable if any series is integrated at I(2) (Ouattara, 2004). The assumption of the ARDL bounds testing to cointegration is that integrating order of the variables should be I(1), or I(0) or I(1)/ I(0). Our results reveal that all the series are integrated at I(1). Because of the same integrating order of the variables, the ARDL bounds testing approach to cointegration must be applied to test whether cointegration exists among the series such as income inequality ($\ln IE_t$), financial development ($\ln F_t$), growth ($\ln Y_t$), inflation ($\ln IN_t$) and globalization ($\ln G_t$).

Table-3: Clemente-Montanes-Reyes Unit Root Test

Variable	**Innovative Outliers**			**Additive Outlier**		
	t-statistic	TB1	TB2	t-statistic	TB1	TB2
$\ln IE_t$	-3.995 (6)	1976	1978	-11.551 (3)*	1980	1984
$\ln Y_t$	-4.822 (3)	1975	2000	-8.316 (6)*	1975	1987
$\ln F_t$	-4.203 (3)	1979	2001	-5.997 (2)**	1977	1997
$\ln IN_t$	-4.813 (1)	1984	1998	-8.193 (4)*	1984	1989
$\ln G_t$	-4.528 (1)	1977	1996	-6.127 (2)*	1978	1988
Note: * and ** indicates significant at 1% and 5% level of significance. TB1 and TB2 show structural break point 1 and 2. Lag order is shown in small parenthesis.						

Once integrating order of the variables is confirmed, we chose an appropriate lag order of the variables to apply the ARDL bounds testing approach to cointegration. We use sequential modified LR test statistic (LR), Final Prediction Error (FPE); Akaike Information Criterion (AIC); Schwarz Information Criterion (SIC) and Hannan-Quinn Information criterion (HQ) to choose appropriate lag order but we prefer to take decision about appropriate lag after using AIC as it provides reliable and consistent information about lag order (Lütkepohl, 2006) in the presence of structural breaks stemming in the mentioned series.

Table-4 shows the results of the ARDL cointegration test. From these results, it is clear that our computed F-statistic exceeds critical bounds at 1% and 5% once we used income inequality ($\ln IE_t$), economic growth ($\ln Y_t$) and inflation ($\ln IN_t$) as dependent variables. The dummy for structural breaks in based on Clemente et al. (1998) unit root test. We have found three cointegrating vector confirming cointegration relationship between the variables. This implies that the long run relationship exists between income inequality, economic growth, financial development, inflation and globalization in case of Iran in the presence of structural breaks. Oil shock affects Iran's economy in 1975 and 1976 and made a wider gap between poor and rich in these years. Because, Iran reach to the higher oil revenue and it goes to industrial and services sectors, not agricultural sector. As a results, income of people who work in agricultural sector had a lower growth compared to others and income inequality increased. In 1984 government use a price adjustment and subsidies to decrease the income inequality that was have an upward trend because of war. Oil shock also affected economic growth in Iran.

Table-4: Results of the ARDL Cointegration Test

Estimated Model	$IE_t = f(Y_t, F_t, IN_t, G_t)$	$Y_t = f(IE_t, F_t, IN_t, G_t)$	$F_t = f(IE_t, Y_t, N_t, G_t)$	$IN_t = f(IE_t, Y_t, F_t, G_t)$	$G_t = f(IE_t, Y_t, F_t, IN_t)$
Lag order	(2, 2, 1, 1, 1)	(2, 2, 2, 2, 2)	(2, 2 , 2, 2, 1)	(2, 2, 2, 1, 2)	(2, 2, 2, 2, 2)
F-statistics	8.830*	10.004*	3.669	7.056**	2.957
Structural Break	1976	1975	1979	1984	1977
Critical values#	1 per cent level	5 per cent level	10 percent level		
Lower bounds	7.317	5.387	4.477		
Upper bounds	8.720	6.437	5.420		
Diagnostic tests					
R^2	0.8444	0.8701	0.8232	0.7659	0.8542
$Adj-R^2$	0.6889	0.6916	0.6969	0.4586	0.6076
F-statistics	5.4306*	8.8739*	6.5195*	2.4927**	3.4638

Durban Watson Test	2.1963	2.5606	1.7888	2.0940	2.1771
Note: * and ** show the significance at 1% and 5% level respectively. Critical bounds are generated by Narayan (2005).					

After finding cointegration between the variables, next round to investigate the impact of financial development, economic growth, inflation and globalization on income inequality.

The results of long-run relationship are reported in Table-5. Our findings based on the linear model show that economic growth has positive impact on income inequality and it is statistically significant at 1% level. It implies that economic growth hampers income distribution and less benefiting to the bottom 20 per cent population. All else is same, a 1 per cent increase in economic growth leads income inequality by 0.6615 per cent. These results are consistent with Shahbaz, (2010) in case of Pakistan but contradictory with Barro (2000) who reported negative impact of economic growth on income inequality in low income countries. The impact of financial development on income inequality is negative and it is statistically significant at 1% level. A 0.2529 per cent income distribution is improved by 1 per cent financial development i.e. disbursement of domestic credit to private sector by financial sector.

The inflation has inverse impact on income inequality and it is significant at 10 per cent level. Keeping other things constant, a 1 per cent increase in inflation is liked with 0.0248 per cent decline in income inequality. These findings are consistent with line of literature such as Shahbaz et al. (2010); Shahbaz and Islam, (2011) in Pakistan and Bittencourt, (2010) in Brazil. Globalization is inversely linked with income distribution and it is statistically significant at 1 per cent level of significance. This shows globalization improves income distribution by generating employment opportunities both for skilled and unskilled labour. A 1 per cent increase in globalization reduces income inequality by 0.1870 per cent, all else is same. Our results are contradictory with Dadgar and Nazari, (2011) who reported that globalization increases income inequality and, Mousavi and Taheri, (2008) found no significant relationship between globalization and rural-urban income distribution in case of Iran.

To test GJ (1990) hypothesis i.e. inverted U-shaped relationship between financial development and income inequality, we have included non-linear term of $\ln F_t$ in model-2. The coefficients of linear term and non-linear terms are positive and negative i.e. 5.989 and -0.2200 respectively. This implies that income inequality is increased with financial development and starts to decline once financial sector matures. Our results confirmed the empirical existence of an inverted U-shaped relationship between financial development and income inequality. Our results are consistent with the line of literature such as Clarke et al. (2003, 2007); Rehman et al. (2008); Kim and Lin, (2011); Rötheli, (2011); Batuo et al. (2012); Nikoloski, (2012). The U-shaped relationship between financial development and income inequality is also reported by Sebastian and Sebastian, (2011); Tan and Law, (2012); Ling-zheng and Xia-hai, (2012) etc.

There is a U-shaped relationship found between globalization and income inequality in case of Iran. In a third model in table (5) square term of $\ln G_t$ is included. Our finding shows that linear term is negative, non-linear is positive and both of them are significant. It indicates that globalization at low (high) levels tend to reduce (increase) income inequality. This result is against with the findings of Agenor (2003) which shows that there is an inverted U-shaped relationship between globalization and poverty, Lindert and Williamson (2001) and Heshmati (2004) which could not determine a U-shape relationship between inequality and globalization.

Table-5: Long Run Analysis

Dependent Variable = $\ln IE_t$						
Model	(1)		(2)		(3)	
Variables	Coefficient	T. Statistic	Coefficient	T. Statistic	Coefficient	T. Statistic
Constant	-6.8593*	-6.4535	-50.8719**	-2.7167	-41.0588**	-2.2136
$\ln Y_t$	0.6615*	6.2121	0.6503*	6.3934	0.8033*	6.7443
$\ln F_t$	-0.2529*	-4.6828	5.9890**	2.2346	5.8932**	2.4307
$\ln F_t^2$	….	….	-0.2200**	-2.3352	-0.2183**	-2.5723
$\ln IN_t$	-0.0248***	-1.7159	-0.0136	-0.8666	-0.0131	-0.7395
$\ln G_t$	-0.1870*	-2.8388	-0.2097*	-4.2305	-6.7423*	-2.9700
$\ln G_t^2$	….	….	….	….	0.9521*	2.8820
Diagnostic Tests						
R^2	0.5532	….	0.6279	….	0.6913	….
F-statistic	11.1433*	….	11.8151*	….	12.6942*	….
$\chi^2 NORMAL$	2.0170	(0.3647)	3.6200	(0.1636)	0.6240	(0.4687)
$\chi^2 SERIAL$	2.1456	(0.1132)	2.0182	(0.1489)	0.1552	(0.3277)
$\chi^2 ARCH$	0.3363	(0.5653)	0.0133	(0.9085)	0.9799	(0.3284)
$\chi^2 WHITE$	0.7034	(0.6861)	0.5167	(0.8510)	0.5589	(0.8434)
$\chi^2 RAMSEY$	1.8545	(0.1720)	3.3910	(0.1100)	0.4459	(0.5089)
Note: *, ** and *** denote the significant at 1%, 5% and 10% level respectively. $\chi^2 NORM$ is for normality test, $\chi^2 SERIAL$ for LM serial correlation test, $\chi^2 ARCH$ for autoregressive conditional heteroskedasticity, $\chi^2 WHITE$ for white heteroskedasticity and $\chi^2 REMSAY$ for Resay Reset test.						

Lower segment of Table-5 reports the results of diagnostic tests. Following these results, null hypothesis cannot be rejected. It is concluded that that residual term is normally distributed with constant variance and zero mean. There is no serial correlation and absence of autoregressive conditional heteroskedasticity is also found. There is no presence of white heteroskedasticity. Specification of model is well articulated confirmed by Ramsey test statistic.

After finding long run impacts of financial development, economic growth, inflation and globalization on income inequality, next round is to test their short-run dynamics using error correction method (ECM). Results of short run model are shown in Table-6. Economic growth is positively related with income inequality and it is significant at 1 per cent level. Financial development (lagged of financial development) and income inequality are inversely linked income inequality and it is significant at 5 (10) per cent level. Inflation has positive impact on income inequality and it is significant at 5 per cent level. Globalization improves income distribution as it is negatively linked with income inequality. It is statistically significant at 10 per cent level.

The coefficient of ECM_{t-1} indicates short run deviations towards long run equilibrium path. Our results postulates that the estimate of ECM_{t-1} is -0.5984. This implies that deviations in short run towards long run are corrected by 59 per cent per year. This would take 1 year and almost 7 months to attain full convergence process and it shows high speed of adjustment for Iranian economy in any shock to income inequality equation. The high significance of ECM_{t-1} with negative further confirms our established long run relationship between the variables.

Table-6: Short Run Analysis

Dependent Variable = $\Delta \ln IE_t$			
Variable	Coefficient	T-statistic	Prob. value
Constant	-0.0021	-0.2582	0.7978
$\Delta \ln Y_t$	0.6773*	3.7463	0.0007
$\Delta \ln F_t$	-0.0975**	-1.9957	0.0543
$\Delta \ln F_{t-1}$	-0.1296***	-1.9078	0.0651

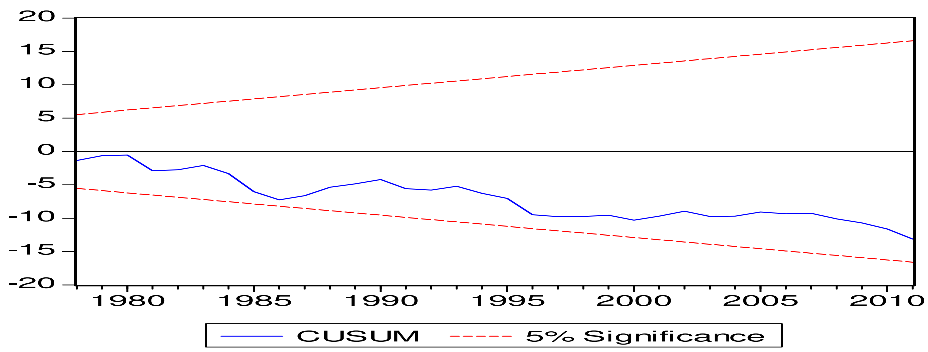
$\Delta \ln IN_t$	0.0275**	1.9989	0.0539
$\Delta \ln G_t$	-0.2297***	-1.7412	0.0910
ECM_{t-1}	-0.5984**	-2.7075	0.0107
R-Squared	0.5752		
F-statistic	7.4490*		
D. W Test	1.9921		
Diagnostic Tests			
Test	F-statistic	Prob. value	
$\chi^2 NORM$	0.9137	0.6332	
$\chi^2 SERIAL$	0.5282	0.5948	
$\chi^2 ARCH$	1.9551	0.1703	
$\chi^2 REMSAY$	2.0150	0.1920	
Note: *, ** and *** denote the significant at 1%, 5% and 10% level respectively. $\chi^2 NORM$ is for normality test, $\chi^2 SERIAL$ for LM serial correlation test, $\chi^2 ARCH$ for autoregressive conditional heteroskedasticity and $\chi^2 REMSAY$ for Resay Reset test.			

The results of diagnostics tests are reported in lower segment of Table-6. The results show that serial correlation and autoregressive conditional heteroskedasticity do not present between the variables used in short-run model. Residual term is normally distributed and model is well specified. Hansen, (1992) disclosed that potential biasedness and misspecification of the model should be avoided for testing the stability of long run parameters. Therefore, CUSUM and CUSUMsq tests are applied to examine the stability of the ARDL estimates. These tests are developed by Brown et al. (1975). Furthermore, Brown et al. (1975) indicated that recursive residuals are to be less affected by small or regular changes in parameters and these changes can be detected by using these residuals[7]. They

[7] The first of these involves a plot of the cumulative sum (CUSUM) of recursive residuals against the order variable and checking for deviations from the expected value of zero. The CUSUMSQs have expected values

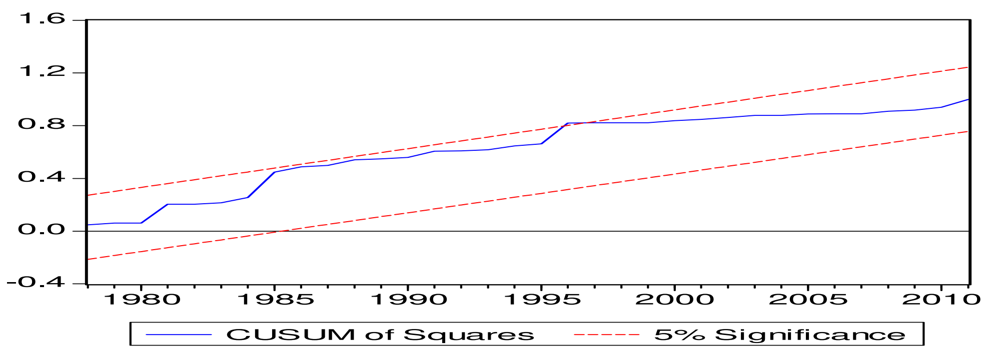
argue that if the null hypothesis of parameter constancy is correct, then the recursive residuals have an expected value of zero and if the parameters are not constant, then recursive residuals have non-zero expected values following the parameter change.

Figure-4: Plot of Cumulative Sum of Recursive Residuals



The straight lines represent critical bounds at 5% significance level.

Figure-4: Plot of Cumulative Sum of Squares of Recursive Residuals



The straight lines represent critical bounds at 5% significance level.

Figure 4 and 5 belongs to the results of CUSUM and CUSUMSQ tests, respectively. The graph of CUSUM test lies between the critical bounds (red lines) but graph of CUSUMsq test does cross red lines at 5 per cent confidence interval. This indicates the instability of the

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ranging in a linear fashion from zero at the first-ordered observation to one at the end of the sampling interval if the null hypothesis is correct. In both the CUSUM and CUSUMSQ tests, the points at which the plots cross the confidence lines give some in diction of value(s) of the ordering variable associated with parameter change.

ARDL estimates. Parameter instability is found around 1996-97 in CUSUMsq test at 5 per cent confidence interval. This structural break point is linked to efforts of Iranian government to control inflation. In 1994 and 1995 Iran faced very high inflation. So that in 1996-1997 government tried to control the liquidity by controlling the banks credit.

We have also applied Chow forecast test to validate the significance of structural break in Iran over the period of 1996-97. The results are reported in Table-7. It is pointed by Leow, (2004) that Chow forecast is preferred over graphs. Graphs often provide ambiguous results. The results in Table-7 indicate that forecast test accepts hypothesis of no structural change in our model.

Table-7: Chow Forecast Test

Forecast from 1996 to 2011			
F-statistic	0.2783	Probability	0.9923
Log likelihood ratio	8.1439	Probability	0.9178

The VECM Granger Causality Analysis

Once cointegration is found between the variables, we should apply the VECM Granger causality approach to examine causal relationship between income inequality, financial development, economic growth, inflation and globalization. It is also supported by Granger, (1969) to apply the VECM Granger approach if variables are found to stationary at same level. The direction of causal relationship between income inequality, financial development, economic growth, inflation and globalization would help policy makers to equalize income distribution by implementing appropriate economic and financial policies in Iran.

The results of the VECM Granger causality are reported in Table-8. It is found that the estimates of ECM_{t-1} have negative sign and statistically significant in all VECMs except in financial development ($\Delta \ln F_t$) and globalization ($\Delta \ln G_t$) equations. It implies that shock exposed by system converging to long run equilibrium path at a slow speed for income inequality equation (-0.5228) and economic growth equation (-0.4780) VECMs as compared to adjustment speed of inflation equation (-0.6477).

In long run, causal relationship reveals that feedback hypothesis is found between income inequality and economic growth. This finding is contradictory with Risso and Carrera, (2012) who reported unidirectional causality running from income inequality to economic growth in pre-reforms and neutral hypothesis is found between both variables in post reforms in China. But, Huang et al. (2011) reported that economic growth Granger causes regional income inequality. Financial development Granger causes income inequality. This finding is consistent with Gimet and Lagoarde-Segot, (2011) who reported that financial sector plays its vital in declining income inequality. The unidirectional causality running from financial development to economic growth confirms the existence of supply-side hypothesis in case of Iran. Our results have been supported by Shiva, (2001) who documented that financial development plays a vital role to lead economic growth. The feedback effect is found between inflation and income inequality. On contrary, Shahbaz et al. (2010) reported that inflation improves income distribution through redistributive policies. Globalization Granger causes income inequality. This view in contradictory to Mah (2002) who noted that globalization leads to deteriorate income inequality in Korea but Mousavi and Taheri, (2008) could not find a significant relationship between globalization and income distribution in case of Iran.

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Table-8: VECM Granger Causality Analysis

Dependent Variable	Type of causality					
	Short Run					Long Run
	$\sum \Delta \ln IE_{t-1}$	$\sum \Delta \ln Y_{t-1}$	$\sum \Delta \ln F_{t-1}$	$\sum \Delta \ln IN_{t-1}$	$\sum \Delta \ln G_{t-1}$	ECT_{t-1}
$\Delta \ln IE_t$	…	7.9826* [0.0017]	1.2436 [0.3032]	1.6248 [0.2144]	1.0938 [0.3484]	-0.5228** [-2.6066]
$\Delta \ln Y_t$	6.9088* [0.0035]	…	1.4883 [0.2492]	2.8118*** [0.0765]	3.1763*** [0.0566]	-0.4780* [-3.4499]
$\Delta \ln F_t$	0.7830 [0.4661]	2.8678*** [0.0725]	…	1.6603 [0.2071]	0.3132 [0.7735]	…
$\Delta \ln IN_t$	3.4047** [0.0470]	1.2088 [0.3132]	2.3398 [0.1147]	…	0.0171 [0.9831]	-0.6477* [-3.7251]
$\Delta \ln G_t$	0.6192 [0.5451]	2.8788*** [0.0718]	0.2836 [0.7550]	0.3811 [0.6863]	…	…
Note: *, ** and *** represent significance at 1%, 5% and 10% levels respectively.						

In short run, bidirectional causality exists between income inequality and economic growth. The feedback effect is found between economic growth and globalization. The unidirectional causal relationship is found running from income inequality to inflation. Economic growth Granger causes financial development.

V: Conclusion and Policy Implications

In this study long-run and short-run relationship between financial development and income inequality has been investigated in case of Iran. We have applied the ARDL bound testing approach for long run and error correction model for short run dynamics. The structural break unit root tests have applied to test the integrating order of all the variables. Greenwood-Jovanovich, (1990) hypothesis which illustrates an inverted-U shape relationship between financial development and income inequality is also tested.

Our results indicate that unique level of integration of the variables and presence of long run relationship between the series is validated. Furthermore, economic growth impedes income distribution. Financial development reduces income inequality. Inflation benefits income distribution. Globalization also improves income distribution. Our analysis has proved the empirical presence of GJ (1990) hypothesis between financial development and income inequality while U-shaped relationship between globalization and income inequality in case of Iran.

As a result, to have a better income distribution, financial sector in Iran must be developed. To reduce the gap between rich and poor, it is necessary to make it easy for entrepreneurs to reach the financial services. Expansion of capital market could be another remedy for Iran's economy. There can be numerous ways to expose the opportunities for better life to the poor's. Such as access to capital makes, the re-allocation of resources, technological innovation and proper human capital development, and last but not least giving proper attention to the financial sector. Access to capital market by poor and disadvantages might be helpful to them either by developing entrepreneurial skill and thus engaging them in productive activities and /or by allowing them to learn higher and quality education, particularly in the areas of science and engineering that would help human capital formation and innovation. Further, re-allocation of resources will help to increase income of the poor in the short run. The technological innovation and proper human capital development is very crucial for sustained long run growth path of an economy. Finally, proper attention of policy makers to the financial sector can prevent the mismanagement in the monetary and fiscal policy action and therefore save from a big disaster. Keeping the fact in mind that the main aim of public policy is to promote economic growth, create employment, and reduce poverty, the role of proper management of government policies should not be ignored. Policy makers need to pay a very great attention in initiating the reforms in the financial sector. It is expected that such reforms will surely have over all positive effects in the economic growth as well as development of society. Private players can also be given a great responsibilities and government should take steps which should allow private operates to operate without fear or undue political influence. Even if there is great practical relation between economics politics however, government should try not to take economic decisions based on political grounds but those should be taken on the basis of economic principles.

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