Bangladesh: Income Inequality and Globalization

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ABSTRACT

Global anxiety about the effect of globalization is increasing. Throughout recent years the influence of globalization on income allocation has been fiercely discussed. This research carries out a time series review to examine the effect of globalization on income inequality in Bangladesh between 1975 and 2018. Study results indicate that globalization variables – exports, imports, foreign aid, foreign direct investment (FDI), and remittance inflows – have a significant long-term impact on income inequality in Bangladesh. Long-term foreign aid and imports are improving, while exports, FDI, and remittance inflows are deteriorating income distribution in Bangladesh during the study period. Nevertheless, in *the* short term, exports, imports, FDI have little to no impact in the model and a change in foreign aid and remittances will have a significant conservative force attempting to resolve the system.

Key words: Globalization, Income Inequality, Co-integration

INTRODUCTION

Globalization is defined as a mechanism of growing obstacles to a liberalized movement of money, finance, products, services, resources, and labor across national borders to create an integrated global economy (Bhensdadia & Dana, 2004; Bowles, 2005; Jermsittiparsert, Sriyakul, & Rodboonsong, 2013; Lal, 2000). This globalization includes worldwide export, import, and capital flow, remittances, foreign direct investment, migration, foreign aid, and all types of obstacles to trade (Arif et al., 2015; Sifat & Israt, 2011). In recent years, the socio-economic development of each nation has been closely related to globalization (Haseeb, Suryanto, Hartani, & Jermsittiparsert, 2020). Globalization can have a positive effect on expanded production and innovation, emerging technology, promoting financial expansion cheaper imports; nonetheless, it may harm the factors allude above, which could abuse equal benefits and widen income gaps (Hussain, 2011) (Wu, 2012). Since the advantages of increasing incomes are not spread relatively among all segments of the population in the state of income inequality, which is likely to impede sustainable economic growth (Arif et al., 2015; Mohanty, 2017). Increasing inequalities in both developing and industrialized countries could deepen conflicts and delay economic and social growth (World Social Report, 2020). According to this report, more than two-thirds of the world's population is still residing in countries where inequality has risen, and the disparity is increasing again in Brazil, Argentina, Mexico, and several other states that have seen inequality decreased in recent decades. The report presents evidence that technical advancement, environmental degradation, industrialization, and global migration have had an impact on inequality patterns. The evidence thus demonstrates that the impact of globalization on income disparity is intertwined. The causative impacts of globalization on income inequality are a major issue of academic concern.

Ali and Isse (2007); Aradhyula, Rahman, and Seenivasan (2007); Beckfield (2006); Bensidoun, Jean, and Sztulman (2011); Cassette, Fleury, and Petit (2012); Demir, Ju, and Zhou (2012); Felbermayr (2005); Hepenstrick and Tarasov (2015); Kai and Hamori (2009; Lu and Cai (2011); Meschi and Vivarelli (2009); Munira, Kianib, Khanc, and Jamald (2012); Rodríguez-Pose (2012); Rudra (2004); Silva & Leichenko, 2004; Wagle (2007) , etc. claim that globalization is rising income inequality in various regions at different periods. Stiglitz (2002) argued that globalization has negative consequences on, especially weak economies. Fischer (2003); Hurrell and Woods (1995); Stiglitz (1998) show that globalization contributes to a rise in inequalities as trade raises differentials in exchange for schooling and expertise. It marginalizes other classes of citizens or geographical areas, and liberalization is not complemented by the establishment of appropriate organizations and governance. Babones and Zhang (2008); Ben-David (1993); Bhagwati (2004); Chakrabarti (2000); Georgantopoulos and Tsamis (2011); Heshmati (2007); Round and Whalley (2006); Silva (2007); Tian, Wang, and Dayanandan (2008); Zhou, Biswas, Bowles, and Saunders (2011) , etc. showed that globalization is contributing to reduce income inequality and deprivation and to generate jobs. Throughout their research, Dollar & Kraay (2002); Edwards (1997); Li, Squire, & Zou (1998); Lindert & Williamson (2003); Mah (2003); Mahler, Jesuit, & Roscoe (1999) noticed no vital connection between globalization and income inequality. Felbermayr (2005); Jaumotte, Lall, and Papageorgiou (2013); Milanovic (2005) identified the mixed impact of globalization on income inequality. Jaumotte et al. (2013) reported that globalization contributes to a reduction in the distribution of poor countries' income, but that rich countries benefit from it. Felbermayr (2005) observed a favorable trade-income link and consider little proof that trade decreases income disparities over two separate periods.

Therefore, the above issue sparks debate in the literature for researchers to reinvestigate the connection between globalization and income inequality. Knowing the roots of disparities is key to the implementation of policy initiatives. This analysis aims to provide empirical proof of the effect of globalization on income disparity in Bangladesh. Five variables are used to denote globalization, and the internationally agreed Gini Coefficient is used as a metric of income inequality in this regard.

LITERATURE REVIEW

Connecting globalization with income inequality is an important topic of discussion in economic literature. It is, therefore, essential to take into account the existence and ties between globalization and income disparity. Studies have used foreign direct investment, economic integration, foreign aid, remittance inflows, etc. as the indicator of globalization. Relatively few researches have been performed on the consequences of globalization on income disparity in the world. Stolper and Samuelson (1941) points out that income inequality has risen in developed countries due to globalization, and the reverse has occurred in emerging countries [their finding is known as Stolper-Samuelson hypothesis]. Some scholars argue that economic globalization inevitably results in a decrease in income inequality in less developing countries and a rise in advanced industrialized countries, in support of the Stolper and Samuelson. Mundell (1957) has used FDI as an indicator of globalization and shown that FDI contributes to a general rise of capital, since it moves primarily from industrialized countries to developing countries, raising the total physical productivity of labour. As a consequence, the earnings of the laborer are increasing, and income inequality in developing countries is declining. Therefore an inverse relation is explored between globalization and income inequality. Yet there are disputes over Mundell's results. Cornia (2003) points out that globalization raise the difference in income levels of people in various countries. More than a few research, Anderson (2005); Atif, Srivastav, Sauytbekova, and Arachchige (2012); Haseeb et al. (2020); Meschi and Vivarelli (2009); Munira et al. (2012); Ogunyomi, Daisi, and Oluwashikemi (2013), etc. looked at the impact of international trade on income discrimination in developing nations. Haseeb et al. (2020) found that globalization, in Indonesia's economy, increases income inequality. Figini and Görg (2006) argue the enhanced penetration of FDI deepens the gap between inequalities in developing countries. Contrairement to Figini and Görg, Anderson (2005) claims that greater openness decreases inequality in developing countries. Meschi and Vivarelli (2009) estimate the effect of trade on income inequality of developing countries (DCs) through various forms of countries. The findings indicate that interaction with high-income countries worsens the distribution of income in the DCs. The relationship of the sector-wise employment, trading system, and income disparity in 55 developing countries were investigated by Demir et al. (2012). The study shows that, if the job share of the manufacturing sector crosses a threshold, the rise in the percentage of manufacturing exports decreases income disparity within the nation. The relation between trade openness and income disparity in Pakistan has been studied by Munira et al. (2012). Results have shown that interest rate, remittances, trade, and urbanization raise inequality, while FDI minimized it. Using annual time series data for the duration of 1986 to 2010, Ogunyomi et al. (2013) observed that globalization continues to increase wealth disparities and decrease economic development in Nigeria's economy. Sylwester (2005), looked at the impacts of foreign direct investment on economic growth and allocation of income in less developed countries (LDCs). FDI has a favorable association with economic development, but there is no proof that FDI is rising income disparity within the community of LDCs.

Bensidoun et al. (2011) looked at the relationship of income inequality and foreign trade by introducing the trading trend of 41 nations. The analysis concluded that foreign exchange leads substantially to rising income gaps in developing economies. Atif et al. (2012) examine the effect of globalization on income disparities by evaluating static and dynamic models for panel data from 68 developed countries across the period 1990-2010. Their research suggests a favourable association between globalization and income inequalities, indicating that growing globalization will contribute to a deterioration of income distribution. In the developing economies, there is no statistically relevant association between FDI inflow and income disparities (Mah, 2003; Mahler et al., 1999).

Beckfield (2006); Cassette et al. (2012); Chintrakarn, Herzer, and Nunnenkamp (2012); Herzer and Nunnenkamp (2013); Lu and Cai (2011); Silva and Leichenko (2004); Tian et al. (2008) have explored the effect of trade on income inequality in developed countries. Silva and Leichenko (2004) looked at the consequences of international trade on income disparities in various U.S. states utilizing panel statistics between 1972 and 1994. The findings of their analysis indicated that costly imports and inexpensive exports exacerbate the income gap in the different United Nations states. Beckfield (2006) also found the same phenomenon in 12 European countries from 1973 to 1997. His work has shown that enhanced regional economic integration across European countries increases inequality in incomes. However, Tian et al. (2008) have been demonstrated that FDI, trade, and government spending continue to boost the income allocation in China. Economic disparity is thus not induced by trade liberalization, but by other factors. The previous findings are not confirmed by Lu and Cai (2011), who finds that increased trade openness leads to rising income disparities in China. They looked at the relationship between globalization and allocation of individual income for four Chinese provinces from 1997 to 2005. Chintrakarn, Herzer, and Nunnenkamp (2012) and Herzer and Nunnenkamp (2013) noted that the impacts of FDI on income disparity are negligible, or weakly significant, and negative in the short term. Still, in the long run, FDI has a significant and detrimental influence on income inequality in the United States and Europe. Cassette et al. (2012) differentiated the short-term and long-term effects of foreign trade in products and services on income inequality in 10 developed nations. The results of the report found that trade in services had a short-term impact only, whereas trade in products influenced income disparities in both the short and long term. Their research found that global foreign trade raises the income gap.

Heshmati (2007); Spilimbergo, Londoño, and Székely (1999); Zhou et al. (2011)) etc. investigated the convergence of globalization and income disparities in diverse countries. Spilimbergo et al. (1999) examined that trade openness induces a reduction in income disparity in capital-abundant countries, yet worsens the distribution of income in skill-abundant countries. For the year 1985, Chakrabarti (2000) examined the effect of intra-national income allocation and foreign trade on low-income, lower-middle-class, upper-middle-income, and high-income 73 nations. The study found that income disparity is minimized by rising foreign trade participation and production. Using two indices Heshmati (2007) he studied the correlation with income inequality and globalization, and noticed that industrialized nations had a more balanced income distribution than emerging economies. Milanovic (2005) examines the relationship between openness and income inequality in 1993 for 113 poor, wealthy, and middle-class nations. The study found that the gains of foreign trade have largely been earned by the wealthy and that the share of income for the poor is smaller. Rudra (2004) studied the interaction between openness and income distribution across 11 developed and 35 less developed countries and noticed that trade for developed countries is significantly stronger than less developed ones. Aradhyula et al. (2007) studied the effect of trade openness on income per capita and income inequalities in 60 emerging and industrialized economies using balanced and unbalanced panel data. The study concluded that although globalization raises disparity, the extent of variation is less severe in developing countries. Zhou et al. (2011) examine the effect of globalization on the distribution of income disparity in 60 industrialized, transitional and emerging countries, and the study argues that globalization tends to minimize differences in income distribution within countries. Rodríguez-Pose (2012) researched that rising economic globalization has a beneficial effect on global disparities, and trade contracts had a more substantial influence on income disparity in middle and low-income economies than high-income countries. Hepenstrick and Tarasov (2015) explored how fluctuations in trade openness lead to income gaps across the world. The study found that there would be no disparity owing to trade openness if the states were utterly symmetric. As with the counterfactual world where nations compete in terms of wealth, demographic growth and rising trading prices, income disparities will rise as a consequence of globalization.

Two native papers on this issue are reviewed for this study. Sifat and Israt (2011) utilize foreign direct investment and trade openness as the factors of globalization and monitor their effect on Bangladesh's income distribution. This article has demonstrated that globalization is exacerbating inequality, and increased openness is contributing to greater inequalities. (Arif et al., 2015) described globalization as trade and foreign direct investment (FDI) flows, foreign aid and remittance inflows, and tried to demonstrate their effect on income disparities in Bangladesh. Empirical findings suggest that the rise in trade increases inequalities, but the rise in FDI and remittance inflows reducing inequality.

DATA REVIEW

To analyze the effect of globalization on income inequality, this analysis uses export, import, foreign aid, remittance, and foreign direct investment to be the proxies of globalization the getting idea from the researches of Arif et al. (2015); Sifat and Israt (2011). Such statistics are obtained from World Bank Development Indicators Database. There are a variety of income inequality indicators around the globe. Among such coefficients, Gini is the most common, but this data is not available in the database of the World Bank. The Standardized World Income Inequality Database (SWIID) has assembled a chart on Gini coefficient. This analysis utilized Gini

Coefficient as the measure of income inequality obtained from that source. The accompanying econometric model has been employed to explore the effect of globalization on income inequality.

$$G_i = \alpha_0 + \alpha_1 EX_i + \alpha_2 IM_i + \alpha_3 FA_i + \alpha_4 RE_i + \alpha_5 FDI_i + U_i \quad (1)$$

Where,

G_i = Gini Coefficient
$\alpha_i^{\prime} s$ = Coefficients of globalization variables
EX_i = Export
IM_i = Import
FA_i = Foreign Aid
RE_i = Remittance
FDI_i = Foreign Direct Investment
U_i = Random Disturbance Term

If the sign of either of the α_i is positive, the resulting corresponding globalization indicator decreases the level of income inequality and vice versa. There are various opinions on the connection between FDI and income inequalities. FDI may decrease Mundell (1957) or raise Feenstra and Hanson (1997) income inequality. Hence, the sign of α_5 can be either positive or negative. But according to Mohanty (2017), FDI raises disparities in income at all levels of economic growth.

Remittance inflow can decrease/ increase the difference in the distribution of income of a nation; thus, the sign of α_4 can be either positive or negative. Foreign aid will ultimately aim to reduce income inequality in the case that the sign of α_3 is negative (Arif et al., 2015).

Methodology and Result Analysis

Because the investigation deals with time-series data, we need to analyze the stationary properties of variables. If the series is not stationary, we then perform a co-integration test to decide if there is a long-term relationship between dependent and explanatory variables.

Unit Root and Co-integration Test

Time-series data is known to be stationary if its value fails to revert to its long-term average value and the data set features are not just affected by the time change (Shrestha & Bhatta, 2018). If the time series is non-stationary, it is presumed to have a root unit, and, in econometrics, the time series stationary is tested by the root unit test (Shrestha & Bhatta, 2018). This analysis used Dickey-Fuller (DF), Augmented Dicey-Fuller (ADF) and Phillips-Perron (PP) unit root test units to verify whether or not the time series is stationary.

Table 1: Unit root test (Null Hypothesis: No Unit root)

		Gini	Export	Import	Foreign Aid	Remittance	FDI
DF	I(0)	-2.4090	-1.5467	-1.4012	-1.39155	-1.20586	-2.7883
	I(1)	-3.0412*	-3.6772***	-4.6431***	-4.97274***	-4.26661***	-7.7898***
ADF	I(0)	-2.4774	-0.7620	-1.7243	-0.22394	1.12087	-2.9143
	I(1)	-3.3748*	-3.6451**	-4.6280***	-3.59164**	-4.39745***	-3.5915**
PP	I(0)	-1.2761	-1.0983	-1.3825	-2.11623	-1.04618	-2.8575
	I(1)	-2.9730**	-3.6773**	-4.4097***	-11.4183***	-4.49088***	-8.4450***
Decision		I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
Author calculation using E-Views 9							

The findings of the unit-root test are displayed in **Table 1**. The results demonstrate that the null hypothesis of a unit root cannot be discarded at all significance levels for all variables in the level, i.e., there is a unit root issue with the data set. But this issue does not appear in the first difference for these variables, i.e., all the sequences are I(1). Under this situation, the sequence must be cointegrated, if not, using ordinary least square or other equivalent approaches for non-stationary time series will yield spurious results (Shrestha & Bhatta, 2018). This paper carries out the Johansen cointegration study by deploying logG, logEX, logIM, logFA, logRE and logFDI. The study documents trace statistics as well as maximum eigenvalue statistics displayed in **Table-2**. Trace statistics checks on the null hypothesis of k cointeg rating ties against the hypothesis of $k-1$. On the other side, the maximum eigenvalue statistics check the r cointegrating ties with the alternative $r+1$. For both approaches, we continue sequentially from $r=0$ to $r = k-1$ before the null hypothesis is not dismissed. Throughout this method, the unrestricted cointegration rank test focused on trace statistics and maximum eigenvalue all suggest that there are two cointegration relationships at 5% level of significance. The analysis noticed a long-term association among the variables.

Table 2: Johansen cointegration test results

Unrestricted Cointegration rank test (Trace)					Unrestricted Cointegration rank test (Maximum eigenvalue)				
Hypothesized No. of CE(s)	Eigenvalue	Trace statistic	0.05 Critical value	Prob.**	Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.897843	151.5350	95.75366	0.0000	None *	0.897843	70.71870	40.07757	0.0000
At most 1 *	0.789341	80.81634	69.81889	0.0051	At most 1 *	0.789341	48.28299	33.87687	0.0005
At most 2	0.444970	32.53335	47.85613	0.5827	At most 2	0.444970	18.25070	27.58434	0.4744
At most 3	0.297244	14.28265	29.79707	0.8244	At most 3	0.297244	10.93512	21.13162	0.6538
At most 4	0.093748	3.347524	15.49471	0.9489	At most 4	0.093748	3.051573	14.26460	0.9432
At most 5	0.009501	0.295951	3.841466	0.5864	At most 5	0.009501	0.295951	3.841466	0.5864
* denotes rejection of the hypothesis at the 0.05 level									
Author calculation using EViews 9									

Evidence of cointegration can also be evaluated by the residual test. As all series contain the root unit and I(1), we can test the unit root of residual u_i from equation (1). If we have u_i is stationary, i.e. I(0), stationary, i.e. I(0), the variables can be said to be co-integrated (Damodar, 2004). **Table 3** reveals that the residuals in this model do not have the root unit at all level of significance**.** The Augmented Dickey Fuller test has been carried out in this respect. The ADF statistic in **Table 3** is -3.038363 , and probability value is 0.0040. We should deny the null hypothesis that the least residual squares are non-stationary, which implies that the variables are still cointegrated (Xu, 2012).

Table 3: Unit root test on the residuals (Null Hypothesis: Residual has a unit root)

		t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic		-3.038363	0.0040
Test critical values:	1% level	-2.664853	
	5% level	-1.955681	
	10% level	-1.608793	
Author calculation using EViews 9			

OLS Process for Long-run Estimates

As all the variables are cointegrated the long-run coefficient(s) of the equation (1) can be calculated by the OLS process (Moniruzzaman, Toy, & Hassan, 2011). Table 4 displays the OLS estimates, and the approximate equation is as follows:

$$G_i = 3.58 + 0.036885\,EX_i - 0.019025IM_i - 0.025229FA_i + 0.004122RE_i + 0.002877FDI_i$$

Table 4: Long-run Estimates of the Model (Using OLS process)

Variable	Coefficient	Std. Error	t-Statistic	Probability
Log of Foreign_AID	-0.025229	0.004745	-5.316995	0.0000
Log of EXPORT	0.036885	0.009165	4.024410	0.0003
Log of FDI	0.002877	0.001370	2.100674	0.0432
Log of IMPORT	-0.019025	0.009133	-2.083158	0.0448
Log of REMIT	0.004122	0.002405	1.714218	0.0956
C	3.583340	0.110475	32.43561	0.0000
R-squared	0.970333	Mean dependent var		3.508066
Adjusted R-squared	0.965970	S.D. dependent var		0.045673
S.E. of regression	0.008425	Akaike info criterion		-6.577661
Sum squared resid	0.002414	Schwarz criterion		-6.324329
Log likelihood	137.5532	Hannan-Quinn criter.		-6.486064
F-statistic	222.4072	Durbin-Watson stat		0.964401
Prob(F-statistic)	0.000000	Dependent Variable		Log of GINI
Author calculation using EViews 9				

All derived coefficients are found to be important at the 5 percent level, except for the coefficient of remittances (coefficient of remittances are significant at 10%). The R^2 and Adjusted R^2 values are 97% and 96%, respectively, which indicates the goodness of fit is very tight. It implies that globalization can be described around 97 percent variation in income inequalities. The Durbin-Watson statistics is reported to be 0.964401, which is less than 1 and that is a question regarding autocorrelation (Field, 2009). Nevertheless, the autocorrelation will be tested later by the Breusch-Godfrey Serial Correlation LM Test. The overall significance of the model is high because the F-statistics is highly significant. The export, remittance and FDI coefficients are shown to be positive, meaning income inequality will increase if exports, remittances and FDI are increased; these results bolster the study of (Mohanty, 2017). A percentage increase in export, remittances and FDI level contributes to an increase in income inequality by 3.6%, 0.2877% and 0.4122%, respectively. In the other side, it is observed that the coefficients of foreign aid and import variables are negative and the percentage change in foreign aid and imports would reduce income inequality by 2.5 percent and 1.9 percent, respectively.

Diagnostic Test for Long-run Estimates

The outcome of the Breusch-Godfrey Serial Correlation LM test is stated in Table 5 to analyze the residuals under OLS regression if there is autocorrelation in the long-term model. The probability value is stated to be 0.0466, which is less than 5% but more than 1%. Therefore, the error terms are not autocorrelated at the 1% level.

Table 5: Breusch-Godfrey Serial Correlation LM Test (Null Hypothesis: No serial correlation)

F-statistic	2.373249	Prob. F(1,33)	0.0638
Obs*R-squared	11.61472	Prob. Chi-Square(1)	0.0405

Heteroskedasticity test result for BPG is stated in Table 6. The probability value calculated to be 0.0634 is more than 5%. Thus we cannot deny the null hypothesis that the residuals are homoskedastic.

Table 6: Breusch-Pagan-Godfrey Heteroskedasticity Test (Null Hypothesis: Homoscedasticity)

F-statistic	2.404807	Prob. F(1,33)	0.0570
Obs*R-squared	10.45022	Prob. Chi-Square(1)	0.0634
Scaled explained SS	6.770371	Prob. Chi-Square(1)	0.2383

The Jarque – Bera test is a fitness test to check on how sample results follow a normal distribution with the skewness and kurtosis. Figure-1 indicates that the distribution's skewness is 0.287477 and that the Kurtosis is 2.793408, so it is skewed to the right and platykurtic. The Jarque-Bera value is 0.622087 with a probability value of 0.732682 which is more than 5%, so we cannot dismiss the null hypothesis which implies that the residual distribution is naturally distributed.

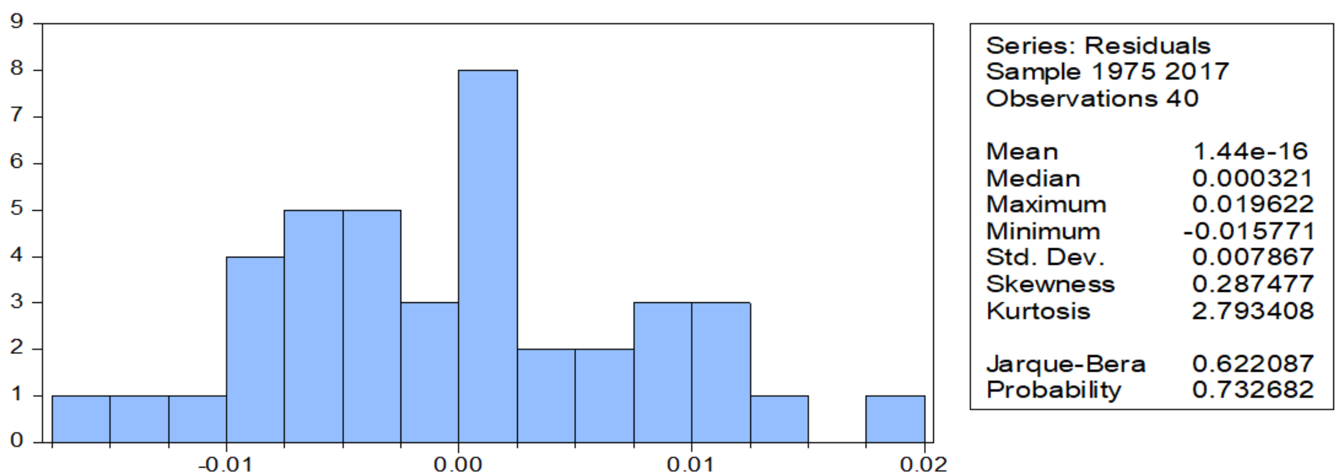

Figure 1: Normality Test

Ramsey's RESET is recorded in Table-7 to check whether the model is correctly identified. The calculated result suggests that the significance of F-statistic is 4.726475 with the p-value 0.0670, and we cannot dismiss the null hypothesis (H_0: The model is correctly specified) (Ramsey, 1969; Wooldridge, 2016).

Table 7: Ramsey RESET Test

(H_0: The model is correctly specified)

Specification: Log(GINI) Log(FDI) Log(F_AID) Log(EXPORT) Log(IMPORT) Log(REMIT) C			
Omitted Variables: Squares of fitted values			
	Value	Degrees of freedom	Probability Value
t-statistic	2.174046	33	0.0670
F-statistic	4.726475	(1, 33)	0.0670
Likelihood ratio	5.354182	1	0.0207

The CUSUM check determines whether the regression model is undergoing any structural break or sudden variation. In Figure-2, the red lines are higher and lower limits at 5% level of significance and the blue line is CUSUM line. As in our case, the CUSUM line is between the two red lines, and our model has passed the CUSUM test, there is no structural beak in the model.

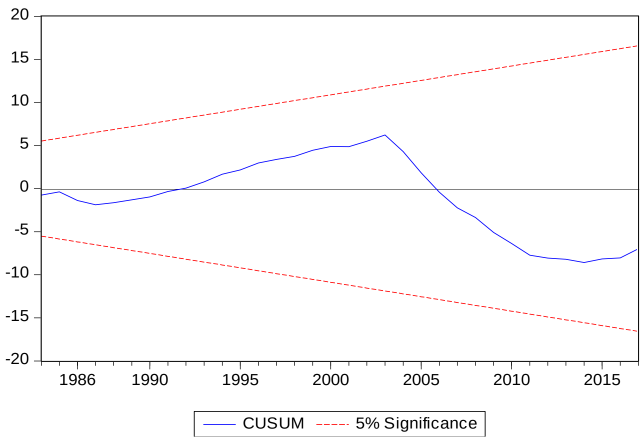

Figure 2: CUSUM Test

Vector Error Correction Mechanism

Depending on the outcome of two cointegrated relations from the Johansen test, we may estimate an error correction model (ECM). To determine the number of optimal lags, we can run standard unrestricted VAR, and the optimal lag length in this study is two (2) as indicated by FRE, AIC & SC criterion seen in **Table 8.**

Table 8: VAR Lag Order Selection Criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	22.10945	NA	1.56e-08	-0.947614	-0.678257	-0.855756
1	246.3343	356.1218*	2.52e-13	-12.01966	-8.744681	-11.37665*
2	286.1876	49.23061	2.47e-13*	-12.24633*	-10.13416*	-11.05217

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan-Quinn information criterion

To determine the short-term dynamics of a long-run equilibrium relationship, we perform vector error correction after the cointegration test. The error correction term calculates the velocity of adjustment toward long-term balance. The approximate VECM coefficients for the globalization function are seen in **Table 9.** The error correction term is seen as a predicted negative symbol, which is accepted at the 10% level. As a result, approximately 19 percent of the model's disequilibrium is dissipated, and 81 percent persists until the next period (*it is often of interest to estimate how long it will take for an existing disequilibrium to be reduced by* 50%. *For our* -0.19 *this is given by* n *in the solution of*, $0.81^n = 0.5 \rightarrow n\log 0.81 = \log 0.5 \rightarrow n = \frac{\log 0.5}{\log 0.81} = 3.28$ *Years*). The short-run elasticity of foreign aid is -0.005875 at lag one and insignificant, which is -0.009353 at lag two and significant at 5%. The elasticity of export, import, FDI is insignificant in both lag 1 and 2. The elasticity of remittances is -0.007744 and significant at the 10% level. Such findings indicate that, whenever there is instability in the whole structure, a shift in foreign aid and remittances would have a major conservative influence that tries to pull the mechanism back into line if it goes too fast.

Table 9: Vector Error Correction Model (VECM) Results

Variable	Coefficient	Std. Error	t-Statistic	Probability Value
D(Log of Foreign_AID(-1))	-0.005875	0.003411	-1.722181	0.0997
D(Log of EXPORT(-1))	0.012302	0.009295	1.323438	0.1999
D(Log of FDI (-1))	0.000671	0.000566	1.186452	0.2487
D(Log of IMPORT (-1))	-0.010309	0.007038	-1.464611	0.1578
D(Log of REMIT (-1))	0.006681	0.006944	0.962089	0.3470
D(Log of Foreign_AID(-2))	-0.009353	0.003218	-2.906314	0.0084
D(Log of EXPORT (-2))	0.004860	0.008772	0.554013	0.5854
D(Log of FDI (-2))	0.000605	0.000637	0.949459	0.3532
D(Log of IMPORT (-2))	-0.005518	0.007023	-0.785676	0.4408
D(Log of REMIT (-2))	-0.007744	0.004173	-1.855896	0.0776
ECT(-1)	-0.188187	0.093668	2.009094	0.0576
C	0.002400	0.000888	2.702955	0.0133
R-squared	0.534922	Mean dependent var		0.002904
Adjusted R-squared	0.291309	S.D. dependent var		0.003830
S.E. of regression	0.003224	Akaike info criterion		-8.361187
Sum squared resid	0.000218	Schwarz criterion		-7.817003
Log likelihood	149.9596	Hannan-Quinn criter.		-8.178086
F-statistic	2.195790	Durbin-Watson stat		1.207562
Prob(F-statistic)	0.005835	Dependent Variable:		D(GINI_LL)

Diagnostic Test for VECM

From Breusch-Godfrey Serial Correlation LM test (reported in **Table-10**) we can accept the null hypothesis at 1% that there is no autocorrelation and cannot reject the null hypothesis equal variance of error terms at 5% from Breusch-Pagan-Godfrey Heteroskedasticity test (reported in **Table-11).** Therefore VECM is free from autocorrelation and Heteroskedasticity. The probability value of Jarque Bera is more than 5%, i.e. the model is normally distributed **(Figure-3).** If we observe Ramsey's RESET test, find that the probability value is more than all level of significance. Therefore we cannot reject the null hypothesis that the model is correctly specified (shown in **Table-12**). Since the blue line remains in between both red lines in **Figure-4** and **Figure-5** the plot of CUSUM's stays within the 5% critical value. Thus, according to CUSUM and CUSUM of square coefficients are stable in the long run.

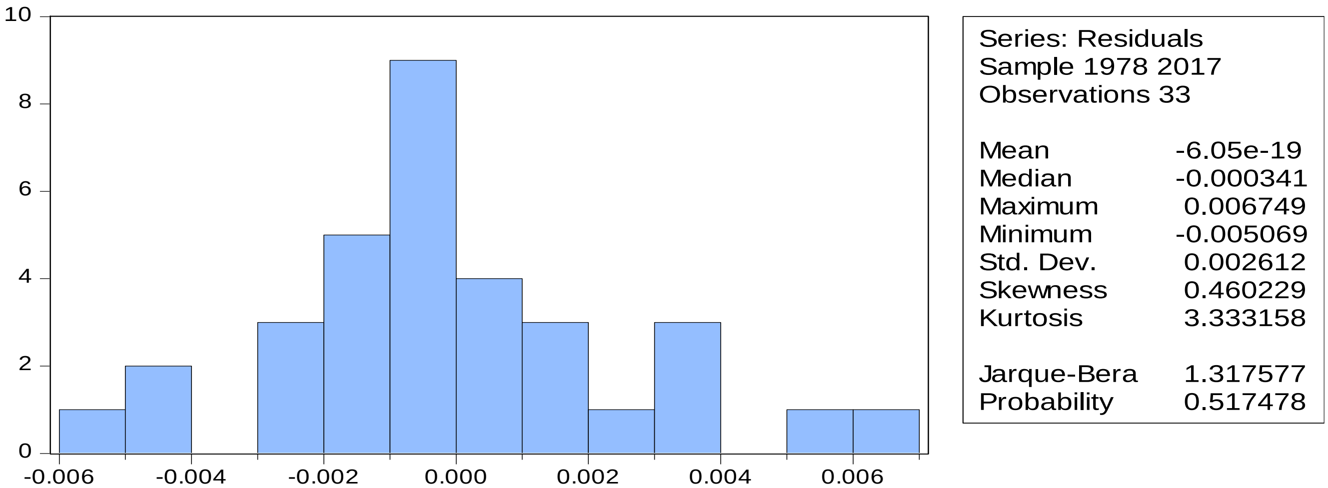

Figure 3: Normality Test under VECM

Table 10: Breusch-Godfrey Serial Correlation LM Test Under VECM

F-statistic	2.168383	Prob. F(2,19)	0.1418
Obs*R-squared	6.132525	Prob. Chi-Square(2)	0.0466

Table 11: Breusch-Pagan-Godfrey Heteroskedasticity Test Under VECM

F-statistic	0.860808	Prob. F(11,21)	0.5881
Obs*R-squared	10.25549	Prob. Chi-Square(11)	0.5076
Scaled explained SS	4.844859	Prob. Chi-Square(11)	0.9384

Table 12: Ramsey RESET Test Under VEM

	Value	df	Probability
t-statistic	0.700477	20	0.4917
F-statistic	0.490668	(1, 20)	0.4917
Likelihood ratio	0.799830	1	0.3711

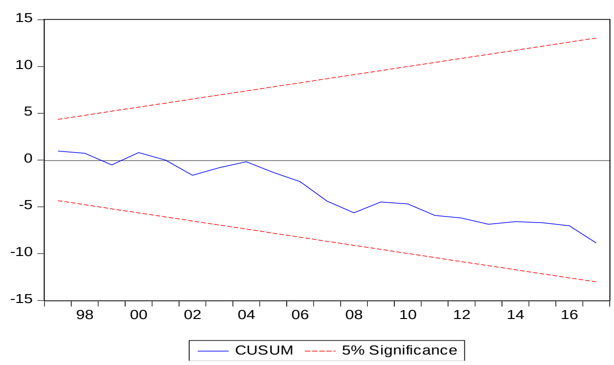

Figure 4: CUSUM test

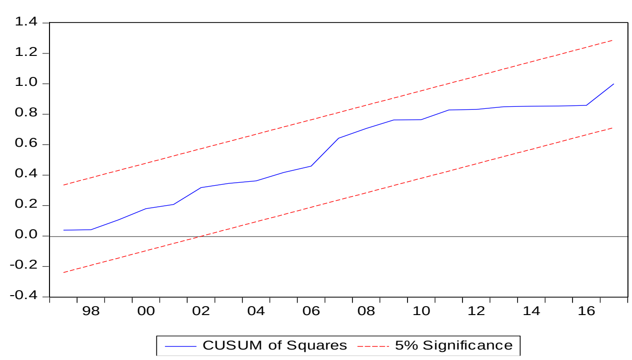

Figure 5: CUSUM Squares Test

CONCLUSION

It is commonly agreed that growth in globalization is correlated with increasing income disparity in the world. The analysis utilizes imports, exports, foreign aid, remittances, FDI as the globalization metric and monitors its effects on income disparities in Bangladesh between 1975 and 2018. DF, ADF and PP tests show that series are stationary at level 1, i.e., they are I (1). Johansen cointegration test and residuals test expose that the variables are cointegrated and two (2) cointegration equations are observed. The Ordinary Least Squares approach indicates that long-term foreign aid and imports reduced income disparity. In contrast, long-term exports, FDI, and remittances will rise in income inequality that is incompatible with Arif et al. (2015) but close to the Sifat and Israt's (2015) analysis. The Vector error correction mechanism demonstrates that, in short-run, FDI, export, and imports have no impact on the dependent variable but foreign aid and remittances may help to minimize the income inequality. Nevertheless, the professional policymakers in Bangladesh will take note of the findings while implementing income disparity mitigation policies. Policymakers should take into account the position of international aid and imports in reducing income disparity. Through their sophisticated initiatives, prospective researchers will concurrently test the validity of the present analysis.

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