# Income Inequality, Economic Growth and Carbon Dioxide Emissions Nexus: Empirical Evidence from Ethiopia

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**Research Article**

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## Abstract

The relationship between income inequality, economic growth and $CO_2$ emissions is ambiguous both theoretically and empirically. Hence, this study examines the link between income inequality and $CO_2$ emissions in Ethiopia for time span covering 1979–2014 using ARDL bounds test and DOLS approach to cointegration. The Zivot-Andrews unit root test and Clemente-Montanes-Reyes unit root test reveal that some of the variables under consideration are stationary at level while others become stationary after first differencing. Both ARDL and DOLS approaches confirm that there is a long-run relationship among the series during the study period. The long-run empirical results show that a 1% increase in economic growth accounts for a 1.05% increase in $CO_2$ emissions while a 1% increase in economic growth squared reduces CO2 emissions by 0.11%. The U-test result reveals that the relationship between $CO_2$ emissions and economic growth confirms existence of the Environmental Kuznets Curve hypothesis. The effect of income inequality on $CO_2$ is not robust to alternative estimation techniques; it is statistically insignificant under the ARDL estimation, but DOLS estimates show that a 1% increase in income inequality increases $CO_2$ emissions by 0.21% in the long-run during the study period. In the long-run a 1% rise in urbanization, population size, energy intensity and industrialization each positively contribute to environmental degradation in Ethiopia by 0.38%, 0.22%, 0.07% and 0.11% respectively. Results from the Toda-Yamamoto Granger causality show a bidirectional causal relationship between $CO_2$ emissions and all other variables except economic growth. $CO_2$ emissions granger causes economic growth with no feedback effect. Results suggested important policy implications in the light of achieving its 2030 targets of low-carbon economy for Ethiopia.

## 1. Introduction

The world has been facing serious global social and environmental crises (M. A. Baloch, Danish, et al., 2020; Berthe & Elie, 2015; Bhattacharya, 2019; Knight et al., 2017). Rising economic inequality is among social crisis that has been facing most countries in the world while the environmental crisis has revealed in the rapid upsurge in environmental pressures (M. A. Baloch, Danish, et al., 2020). In response to addressing these and related global issues, the United Nations (UN) implemented the 2030 sustainable development goals (SDGs) in 2015 with the aim of comprehensively coordinating economic development, social progress, and environmental protection. The goals focus on eradicating poverty, improving sustainable economic growth, promoting social progress, promoting equality, protecting the eco-environment and addressing climate change (He et al., 2019). Developing countries face more severe challenges in achieving the goals of SDGs and national determined contributions (NDCs) than developed countries (M. A. Baloch, Danish, et al., 2020; He et al., 2019).

Developing countries are trapped in multidimensional problems. They need to develop their economies to alleviate poverty, advance the process of industrialization and modernization, and improve the living standards of their people (He et al., 2019). However, sustained economic growth takes place at the expense of environmental quality (M. A. Baloch, Danish, et al., 2020; He et al., 2019; Ravallion et al., 2000). This is because energy consumption increases with rising economic activities, which increases $CO_2$ emissions (A. Baloch et al., 2017; He et al., 2019; Padhan et al., 2019; Zhu et al., 2018) Furthermore, the current economic growth is responsible for income inequality in developing countries like Sub-Saharan Africa (Demir et al., 2019). This clearly indicates that there is a trade-off between economic growth, income inequality and environmental quality which are among the key goals of the SDGs.

Ethiopia has emerged as one of the fastest-growing economies in Africa in the early twenty-first century (Mondal et al., 2018; Oqubay, 2018; Taka et al., 2020; Usama et al., 2020). The country is witnessing rapid urbanization and industrialization along with the growing economy which leads to higher energy consumption and high levels of $CO_2$ emissions. Ethiopia is the third largest $CO_2$ emitter in East Africa (Taka et al., 2020). Even though Ethiopia's contribution to GHG emissions is very low on a global scale, it has been rapidly rising and it is projected to increase from 150 Mt CO2e to 400 Mt CO2e in 2030 under business as usual scenario (FDRE, 2011). While Ethiopia still has a low level of income inequality, its Gini coefficient has shown the upward trend since 2002 (Araya & Woldehana, 2019). For example, Ethiopia's Gini coefficient increased from 0.2867 in 2002 to 0.3370 in 2014 which was about 17.5% increment.

Ethiopia has exerted concerted efforts to bring sustained economic growth, eradicate poverty, reduce income inequality and mitigate climate change. The country has been facing a trade-off in increasing its economic growth sustainably; reducing income inequality and controlling climate change. For instance, the country has a vision of achieving middle-income status by 2025 while developing green economy as stated in its green economy strategy known as Climate Resilient Green Economy (CRGE). Following the Paris Agreement, Ethiopia pledged to reduce its GHG emissions by 64% below the business-as-usual levels by 2030. The SDGs also urges Ethiopia to reduce within country income inequality in order to reinforce the economic growth, improve social as well as environmental crises. These three goals are very interconnected with trade-offs. The $CO_2$ reduction is not possible unless the country follows evidence-based green economy strategy while reducing income inequality. However, empirical evidence shows that the inequality-growth-$CO_2$ emissions nexus has been a debatable topic ( A. Baloch et al., 2017; Berthe & Elie, 2015; Grunewald et al., 2017; Hailemariam et al., 2019; Hübler, 2017; Jorgenson et al., 2017; Knight et al., 2017). Therefore, successful reduction of the high $CO_2$ emission requires further investigation of impact of economic growth and income inequality on $CO_2$ emissions in order to devise informed policy decisions.

Several empirical studies have examined impact of economic growth on environmental quality in both developed and developing countries using different approaches. Some researchers follow linear relationship framework while others apply an N-shaped relation (Allard et al., 2017). However, the well-known framework under which the impact of economic growth on environmental quality has been studied is the Environmental Kuznets Curve (EKC) (Adedoyin et al., 2020; M. A. Baloch, Ozturk, et al., 2020; Kusumawardani & Dewi, 2020; Pata & Aydin, 2020). The EKC hypothesis postulates the inverted U-shaped relationship between economic growth and $CO_2$ emissions. The hypothesis shows that at early stage of economic development process economic growth harms the environment, and it improves the environment at the later stage when a certain income threshold is reached (M. A. Baloch, Ozturk, et al., 2020; Dogan & Aslan, 2018; Dogan & Seker, 2016).

Empirical studies that try to test the EKC hypothesis show inconclusive findings regarding the connection between economic growth and $CO_2$ emissions (A. Baloch et al., 2017; M. A. Baloch, Ozturk, et al., 2020; Bhattacharya, 2019; Pata & Aydin, 2020; Sarkodie & Ozturk, 2020). Several empirical studies lend support

to the EKC hypothesis (M. A. Baloch, Ozturk, et al., 2020; Dogan & Inglesi-Lotz, 2020; Kusumawardani & Dewi, 2020; Pata & Aydin, 2020; Sarkodie & Ozturk, 2020; Sharif et al., 2020), while others fail to confirm the hypothesis (Dogan & Ozturk, 2017; Dogan & Turkekul, 2016). Some of the studies that do not support the EKC hypothesis argue that the relationship between economic growth and environmental quality is linear (Gill et al., 2017). Other findings corroborates with the U-shaped relationship between the two variables (Dogan & Ozturk, 2017; Dogan & Turkekul, 2016; Sohag et al., 2019). Some empirical evidences, on the other hand, lend support to an N-shaped relationship (Allard et al., 2017; Caravaggio, 2020; Lorente & Álvarez-herranz, 2016) and M-shaped (Bousquet et al., 2005; Terrell & Terrell, 2020) between economic growth and environmental quality. Allard et al. (2018) argued that the negative relationship between economic growth and environmental quality that may occur when income reaches a turning point may reverse at higher income levels, leading to N-shape. According to the N-shaped relationship between economic growth and environmental quality, the original EKC hypothesis will not hold in the long-run. Instead, increased income may lead to higher $CO_2$ emissions beyond a certain income level. Scholars attribute the existence of the M-shaped relationship between economic growth and income inequality to different factors. According to Bousquet et al. (2005) inclusion of income inequality is responsible for such type of relationship between the two variables while Egli and Steger (2007) associated it to a policy-induced discontinuity.

Different scholars associate the inconclusive results on the effect of economic growth on environmental quality to different factors. For instance, Padhan et al. (2019) claimed that the variations in the association between economic growth and environmental quality is because of changing economic parameters, natural disasters, energy and environmental strategies, and regulatory and technological innovations. Some researchers argue that the EKC hypothesis may depend on the way economic growth variable is measured (Dogan & Inglesi-Lotz, 2020). According to Narayan and Narayan (2010) and Dinda (2004) most of the EKC literatures employed weak econometric approaches which are responsible for inconclusiveness of results.

Scholars argue that environmental quality is affected both by the income level and the way the income is distributed (Hübler, 2017; Jorgenson et al., 2017; X. Wang et al., 2020; Wolde-rufael & Idowu, 2017; Yang et al., 2020). Excluding income distribution in the test of the EKC hypothesis results in omitted variable bias (M. A. Baloch, Ozturk, et al., 2020; Dogan & Turkekul, 2015). Mcgee and Greiner (2018) claimed that impact of economic growth on environmental quality is influenced by level of income inequality. According to them in countries with the highest income inequality, economic growth harms environmental quality while the opposite is true in regions having the lowest income inequality.

However, many empirical studies on determinants of $CO_2$ emissions have failed to incorporate income inequality as a potential determinant (Baek & Gweisah, 2013; Bhattacharya, 2019; Chen et al., 2020; Ravallion et al., 2000; Yang et al., 2020; You et al., 2019). The existing evidences regarding the impact of income inequality on environmental quality are quite mixed (Berthe & Elie, 2015; Hailemariam et al., 2019; Knight et al., 2017). Findings of these studies vary depending on economic status of countries (Grunewald et al., 2017; Jorgenson et al., 2016; Kusumawardani & Dewi, 2020; Liobikienė, 2020; Magnani, 2000; Mittmann & de Mattos, 2020), time periods (Bhattacharya, 2019; Uddin et al., 2020), and modelling techniques used for the analysis (Hübler, 2017; Uddin et al., 2020). Jorgenson et al. (2017) argued that the mixed findings are attributed to several channels through which income inequality might affect emissions. According to Mushtaq et al. (2020), Jorgenson et al. (2015, 2016, 2017) and Liobikienė (2020) income inequality affects $CO_2$ emissions through three main channels.

The first pathway is explained through the political economy approach (PEA). According to the PEA, the impact of income inequality on environmental degradation is determined by the characteristics of winners and losers in economic projects instigating environmental degradation (Boyce, 1994; Knight et al., 2017; Torras & Boyce, 1998; Uzar & Eyuboglu, 2019). In this regard, Boyce (1994) argued that a significant gap between power and income detrimentally affect environmental quality because high income inequality favors political elites to influence approval of environmentally unfriendly projects that harm the environment. High income inequality strengthens the power of the rich to impose environmental costs on the poor and compromises the ability of the society to achieve optimal solutions to environmental problems (Boyce, 1994; Wolde-rufael & Idowu, 2017; Yang et al., 2020). Magnani (2000) stressed this by arguing that income equality increases demand for environmental quality by creating environmental awareness. However, this idea is criticized by Scruggs (1998) who argued that income equality is not a guarantee for improvement of environmental quality. According to Scruggs (1998) a rise in income may be associated with a lower level of preference for environmental quality. There is a situation under which higher income inequality is conducive to lower environmental degradation. Liu et al. (2018) also argued that only optimal income inequality reduces CO2 emissions. Recent empirical evidences also show that high income inequality rises working hours which contributes to high energy consumption and $CO_2$ emissions (Fitzgerald et al., 2018; Jorgenson et al., 2017, 2019). According to Liobikienė (2020), the PEA explanation mostly explains the impact of income inequality on consumption-based $CO_2$ emissions.

The second explanation focuses on the consumption behaviors of households and marginal propensity to emit (MPE). According to this theory, individuals' propensities to consume more polluting goods differ due to the change in the patterns of consumption which is determined by the level of income and income inequality (Berthe & Elie, 2015; Hao et al., 2016; Heerink et al., 2001; Hübler, 2017; Jorgenson et al., 2017; Ravallion et al., 2000). Following the Keynesian marginal propensity to consume (MPC) concept poor households' marginal propensity to consume energy is higher than that of rich households. This indicates that energy consumption level determines the MPE. The MPE suggests that increasing income of the poor to catch up with the rich means a higher marginal propensity to consume energy, and therefore a higher $CO_2$ emissions. Accordingly, if MPE of the rich group is greater than that of the poor, an increase in income distribution imbalance improves environmental quality (Heerink et al., 2001; Hübler, 2017; C. Liu et al., 2019; Q. Liu, Wang, Zhang, & Li, 2019; Ravallion et al., 2000). However, some authors argue that the poorest people cannot afford to buy environmentally friendly products since they are too expensive (see, e.g., A. Baloch et al., 2017; Hao et al., 2016; Magnani, 2000) which indicates that poor-households have higher MPE than the rich (Chen et al., 2017). As a result, increasing income inequality tends to improve environmental quality. Moreover, rich people are assumed to be more concerned about environmental quality since they are more important for them compared to the poor (Heerink et al., 2001; Scruggs, 1998; Wolde-rufael & Idowu, 2017) which implies that income inequality and environmental quality are positively related.

The third approach which is known as Veblen's (1899) Emulation Theory (ET) relates impact of income inequality on environmental quality to individual's economic behavior. Veblen (2009) shows that people desire to show their financial resources and social status through consumption. Bowles and Park (2005)

 time allocation between work and leisure. As a result, increased income inequality encourages longer

working time and consumption. Chen et al. (2017) states that longer working hours and higher consumption induced by the emulation behavior leads to increased energy consumption. Longer working hours and increased consumption both deteriorate environmental (Grunewald et al., 2017; Jorgenson et al., 2017; Liobikienė, 2020; Wolde-rufael & Idowu, 2017).

The results of empirical researches conducted so far to examine the impact of income inequality on environmental quality are quite mixed and inconsistent. Some findings support the PEA theory (Bae, 2018; Baek & Gweisah, 2013; A. Baloch et al., 2017; M. A. Baloch, Danish, et al., 2020; Boyce, 1994; Demir et al., 2018; Hailemariam et al., 2019; Hao et al., 2016; Knight et al., 2017; Q. Liu, Wang, Zhang, Li, et al., 2019; Morse, 2018; Mushtaq et al., 2020b; Padhan et al. 2019; Torras & Boyce, 1998; Uzar & Eyuboglu, 2019; Yang et al., 2020; Zhu et al., 2018) while others are consistent with the MPE theory (Bhattacharya, 2019; Heerink et al., 2001; Hübler, 2017; Jorgenson et al., 2015; Ravallion et al., 2000; Sager, 2019). On the other hand, some authors argue that it is difficult to attach the relationship between income inequality and environmental quality exclusively to any of the three theories. This group claims that income level and country specific macroeconomic characteristics matter a lot (Grunewald et al., 2017; Jorgenson et al., 2016; Kusumawardani & Dewi, 2020; Mittmann & de Mattos, 2020). Jorgenson et al. (2017) found that the impact of income inequality is statistically insignificant while Uddin et al. (2020) and Bhattacharya (2019) argued that the impact can be positive, negative or insignificant based on the study period. Some scholars like Chen et al. (2020), Grunewald et al. (2017) and Y. Liu et al. (2020) associated the interplay between income inequality and environmental quality with the level of economic development. According to them equal income distribution reduces $CO_2$ emissions in developing countries, but income inequality does not affect $CO_2$ emissions in most developed countries. The authors reasoned that income inequality in developing countries harms innovation, which reduces green technology.

Methodologically, most studies applied either panel data (M. A. Baloch et al., 2020; Hailemariam et al., 2019; Hübler, 2017; Ravallion et al., 2000; Uddin et al., 2020; Yang et al., 2020; Zhang & Zhao, 2014; Zhu et al., 2018) or cross-sectional data (Heerink et al., 2001; Jorgenson et al., 2017; Knight et al., 2017) for a large sample of countries covering different income levels and time periods. Others use either panel (Hao et al., 2016; Kasuga & Takaya, 2017; Zhou & Liu, 2016) or time series data (Baek & Gweisah, 2013; A. Baloch et al., 2017; Demir et al., 2018; Uzar & Eyuboglu, 2019; Wolde-rufael & Idowu, 2017). In this regard, Uddin et al. (2020) argued that results of studies on the impact of income inequality on environmental degradation are sensitive to choice of econometric techniques, sample countries and the measure of income inequality.

The sensitivity of results of the relationship between income inequality and environmental degradation to choice of time, level of income and inequality, country-specific macroeconomic contexts and econometric techniques suggests that individual country-specific case studies may be more appropriate for assessing the relationship between income distribution and environmental degradation (Baek & Gweisah, 2013; Jorgenson et al., 2016; Wolde-rufael & Idowu, 2017).

With this backdrop, this study contributes to the existing literature in the following ways. First, to the best knowledge of the author, there is no empirical study that has attempted to investigate the interrelationship between income inequality, economic growth and environmental quality in Ethiopia using the EKC hypothesis framework. Related previous studies in Ethiopia overlooked income inequality in modelling environmental quality which results in specification bias. This study has incorporated income inequality as a determining factor of environmental quality to overcome the specification bias. Furthermore, incorporating income inequality factor helps better understand the links between income inequality, economic growth and environmental quality in Ethiopia to improve the design and implementation of sound policies that can simultaneously address sustainable economic growth, environmental problems and socioeconomic inequalities.

Second, the present study contributes to existing literature in methodological aspects by applying recent econometric techniques. Most of the previous studies applied conventional unit root tests that may result in biased and spurious estimates because they ignore the presence of structural breaks which puts uncertainties on cointegration and Granger causality tests. This is due to the fact that both cointegration and Granger causality tests depend on order of integration of series under consideration. The novel contribution of this paper to the existing of literature is that it applies recently developed unit root tests that consider the presence of structural breaks. Endogeneity problem, small sample bias, autocorrelation problems and being dependent on order of integration are shortcomings of the conventional cointegration techniques. This study addresses these problems by employing innovative cointegration approaches that yield more precise and reliable estimates. Moreover, novel Granger causality testing technique that is applicable regardless of cointegration status and order of integration of the series, which also avoids loss of information since it can be applied to VAR at level is employed in the current study. Unlike the previous related studies conducted in Ethiopia, the current study employed exact U-test, which provides both necessary and sufficient condition for the presence inverted U-shape relationship between economic growth and environmental quality in Ethiopia. Almost all previous studies employ the quadratic test for the EKC hypothesis which only provides the necessary condition for inverted U-shape or U-shape relationship leading to misleading results because of high correlation between the quadratic and linear variables. With this regard, findings of this study are robust, reliable and accurate since it employs superior econometric techniques.

Findings of this study reveal that the effect of economic growth on environmental quality depends on the level of economic development. That is, economic growth has a detrimental effect on environmental quality at the early stage of economic development. It turns to improve environmental quality when economic development is sufficiently high. At a sufficiently high level of economic development industrial and other sectors become cleaner, people desire cleaner environment, and regulatory framework becomes more efficient, which supports the hypothesis that environmental quality is superior good when overall economic conditions of the counties are improved. The neutrality effect of income inequality on environmental quality is inconsistent with the marginal propensity to emit approach.

The rest of this study is structured as follows. Section 2 describes the variables, methodology, and sources of data. Section 3 details the empirical results and discussion. Section 4 gives concluding remarks and policy recommendations.

## 2. Materials And Methods

## 2.1. Data and Variable Description

The data used in this study entirely depends on World Bank's Development Indicators (WDI) and Standardized World Income Inequality Database. Annual data covering time span of 1979–2014 is used while the selection of the time period is dictated by data availability. Variables of interest are $CO_2$ emissions, real gross domestic product (GDP), real GDP square, Gini index, urbanization, industrialization, energy intensity and population.

**Table 1**:: Variables, Notation and Descriptive Statistics

| ariable | Description | Source |
|---|---|---|
| $O_2$ | Carbon dioxide emissions (metric tons per capita) | World Bank WDI |
| | Real Gross domestic product (GDP) | World Bank WDI |
| $)^2$ | GDP squared | World Bank WDI |
| ini | Gini coefficient of income inequality | Standardized World Income Inequality Database |
| R | Percentage of urban population | World Bank WDI |
| d | Ratio of industrial sector GDP to total GDP | World Bank WDI |
| I | Ratio of energy use to GDP | World Bank WDI |
| | Population size | World Bank WDI |

## 2.2. Model Specification

The major aim of this study is to econometrically examine the impact of income inequality, economic growth, industrialization, urbanization, population and energy intensity on CO2 emissions for Ethiopia over the period of 1979–2014. In order to achieve this objective, this study employed the extended stochastic impacts by regression on population, affluence, and technology (STIRPAT) model developed by Dietz and Rosa (1994) and refined by York et al. (2003). The advantage of the STIRPAT model is that it allows for a more precise specification of the sensitivity of environmental impacts to the forces driving them (York et al., 2003), in the EKC hypothesis framework following Mcgee and Greiner (2018), Shahbaz et al. (2016) and York et al. (2003).

The STIRPAT model has been successfully applied to analyze the effects of driving forces on a variety of environmental impacts (Dietz & Rosa, 1994; Fan et al., 2006; Fu et al., 2015; Mcgee & Greiner, 2018; Poumanyvong & Kaneko, 2010; Shahbaz et al., 2016; Yang et al., 2020; York et al., 2003).

The STIRPAT model is specified as:

$$I_t = \alpha P_t^b A_t^c T_t^d \varepsilon_t \ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots (1)$$

Transforming the above equation into natural logarithm gives it the following form:

$$\ln(I_t) = \alpha + b\ln(P_t) + c\ln(A_t) + d\ln(T_t) + \varepsilon_t \ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots\ldots (2)$$

where the subscript t stands for time; P stands for population; A denotes GDP; T is technology; I represents environmental impact and $\varepsilon$ is disturbance term.

The main advantage of the STIRPAT model is that it allows other factors to be added to examine their impact on environmental parameters because P and A (Dietz & Rosa, 1994) and T (York et al., 2003) are decomposable. For instance, Zhou and Liu (2016) decomposed P into household size, population size, urbanization level and working-age population; and Zhou and Liu (2016), Poumanyvong and Kaneko (2010), Fan et al. (2006) and S. Wang et al.(2017) used energy intensity and industrialization as a proxy for T. Following previous empirical studies (see for instance, Fan et al., 2006; Fu et al., 2015; Zhou & Liu, 2016), therefore, this study added urbanization and energy intensity into the STIRPAT model.

Empirical evidence shows that the interaction between economic growth and environmental quality follows the EKC hypothesis (Bae, 2018; Baek & Gweisah, 2013; Hao et al., 2016; Hundie, 2018; Kusumawardani & Dewi, 2020; Padhan et al., 2019; Uzar & Eyuboglu, 2019; Zhu et al., 2018). Therefore, economic growth squared is added to the STIRPAT model to capture the EKC hypothesis. Moreover, several scholars argue that income disparity affects environmental quality (Bae, 2018; A. Baloch et al., 2017; M. A. Baloch, Danish, et al., 2020; Bhattacharya, 2019; Demir et al., 2018; Hübler, 2017; Kusumawardani & Dewi, 2020; Q. Liu, Wang, Zhang, Li, et al., 2019; Masud et al., 2018; Mittmann & de Mattos, 2020; Yang et al., 2020). This firmly suggests the need of including the variable into the STIRPAT model. The modified and extended STIRPAT model is then specified as follows:

$$\ln CO_{2t} = \alpha + \beta_1 \ln Gini_t + \beta_2 \ln Y_t + \beta_3 (\ln Y_t)^2 + \beta_4 \ln P_t + \beta_5 UR_t + \beta_6 EI_t + \beta_7 Ind_t + \varepsilon_t \ldots\ldots\ldots\ldots (3)$$

## 2.3. Estimation Techniques

## 2.3.1. Unit Root Tests

Even though several conventional unit root tests are available to test the stationarity properties of the variables, they were not used in this study because they provide biased and spurious results because they ignore structural break in the series (Baum, 2001, 2005; Clemente et al., 1998; Katircioglu, 2014; Muhammad

et al., 2013). To this end, this study employed Zivot and Andrews (1992) (ZA hereafter) and Clemente, Monatne and Reyes (1998) (CMR hereafter) unit root tests which consider structural break/s in the series.

Zivot and Andrews (1992) (ZA hereafter) test the stationarity properties of the variables in the presence of single structural break point in the series with three options: a one-time change in variables at level form, a one-time change in the slope of the trend component and a one-time change both in intercept and the trend function of the variables. It can be captured by the model given in Equations a-d under Appendix part.

It is common for macroeconomic variables to exhibit the presence of multiple breaks. In this case, the unit root test method proposed by Clemente, Monatne and Reyes (1998) (CMR hereafter) which takes two break dates was used in this study in addition to ZA method. The model has two forms: the additive outliers (the AO model) and the innovative outliers (the IO model). See the appendix part for the detailed mathematical presentation.

## 2.3.2. Cointegration: ARDL Bounds Test Approach

Different conventional approaches to cointegration tests are available but with several critical drawbacks. The ARDL bounds testing approach to cointegration developed by Pesaran et al. (2001) outperforms these conventional approaches to cointegration (Chindo et al., 2014; c & Ozturk, 2015; Halicioglu & Ketenci, 2016; Hundie, 2018; Hundie & Daksa, 2019; Jalil & Mahmud, 2009; Pesaran et al., 2001; Shahbaz et al., 2013, 2015; Wolde-rufael & Idowu, 2017) both in terms of estimation precision and reliability of statistical inferences (Panopoulou & Pittis, 2004). The ARDL approach to cointegration is applicable regardless of the order of integration of the series under consideration, therefore, it avoids the uncertainties created by pre-testing for unit root (Hundie, 2014, 2018; Jalil & Mahmud, 2009; Wolde-rufael & Idowu, 2016). Unlike other conventional tests for cointegration, the ARDL approach can be applied to studies that have small sample size (Hundie, 2014; Jalil & Mahmud, 2009; Menyah & Wolde-rufael, 2010; M. H. Pesaran et al., 2001; Wolde-rufael & Idowu, 2016). Under the ARDL approach, both short-run and long-run relationships are simultaneously estimated (Hundie, 2014; Jalil & Mahmud, 2009; Wolde-rufael & Idowu, 2017). Another advantage of the ARDL approach is that it efficiently corrects for a possible endogeneity of explanatory variables (Menyah & Wolde-rufael, 2010; Panopoulou & Pittis, 2004) and mitigates problems related to omitted variables and autocorrelation (A. Baloch et al., 2017).

The unrestricted error correction model (UECM) version of the ARDL model for Eq. (3) is specified as follows:

$$\Delta \ln CO_{2t} = \alpha_1 + \sum_{i=1}^{p} \beta_{1i} \Delta \ln CO_{2t-i} + \sum_{i=0}^{q_1} \eta_{1i} \Delta \ln Y_{t-i} + \sum_{i=0}^{q_2} \gamma_{1i} \Delta (\ln Y)^2_{t-i} + \sum_{i=0}^{q_3} \theta_{1i} \Delta \ln UR_{t-i} + \sum_{i=0}^{q_4} \pi_{1i} \Delta \ln Ind_{t-i}$$
$$+ \sum_{i=0}^{q_5} \phi_{1i} \Delta \ln Gini_{t-i} + \sum_{i=0}^{q_6} \omega_{1i} \Delta \ln P_{t-i} + \sum_{i=0}^{q_7} \psi_{1i} \Delta \ln EI_{t-i} + \delta_1 \ln CO_{2t-1} + \delta_2 \ln Y_{t-1} + \delta_3 \left(\ln Y_{t-1}\right)^2$$
$$+ \delta_4 \ln UR_{t-1} + \delta_5 \ln InD_{t-1} + \delta_6 \ln Gini_{t-1} + \delta_7 \ln P_{t-1} + \delta_8 \ln EI_{t-1} + \varepsilon_{1t} \quad \ldots (4)$$

The parameters $\delta_i$ $(i = 1, 2, 3, 4, 5, 6, 7, 8)$ are the corresponding long-run multipliers, while the parameters $\beta_i$, $\eta_i$, $\gamma_i$, $\theta_i$, $\pi_i$, $\phi_i$, $\omega_i$, $\psi_i$ are the short-run dynamic coefficients of the underlying ARDL model.

Testing the presence of long-run relationship among variables in Eq. (4) is the first step of the ARDL approach. To that effect, an F-test is used to test the null hypothesis of $(H_0: \delta_1 = \delta_2 = \delta_3 = \delta_4 = \delta_5 = \delta_6 = \delta_7 = \delta_8 = 0)$. Since the F-statistic is much more sensitive to lag order selection investigating the presence of long-run relationship among the variables in Eq. (9) using Fisher (F) or Wald (W) statistics is the first step in the ARDL bounds testing approach to cointegration. Shahbaz et al. (2015) contended that the F-statistic is much more sensitive to lag order selection. In this study, the Schwartz Bayesian Criterion (SBC) was used to select appropriate lag length. Then, F-statistic is compared to the critical bounds generated by Narayan (2005) because it better fits small sample observations (Narayan, 2004, 2005; Narayan & Narayan, 2004; Uzar & Eyuboglu, 2019). If the computed F-statistic falls below (above) the lower (upper) critical value, the null hypothesis cannot (can) be rejected, indicating the non-existence (existence) of the long-run relationship (Baek & Gweisah, 2013; Pesaran et al., 2001). However, conclusive decision cannot be made if the F-statistic lies between the lower and upper critical bounds unless we know the order of integration of series under consideration (Pesaran et al., 2001; Wolde-rufael & Idowu, 2017). If the F-statistic falls between the lower and upper critical bounds; the result is inconclusive and alternative cointegration methods should be applied (Uzar & Eyuboglu, 2019).

Once the long-run relationships (cointegration) among the variables is established, the second step is to estimate the following long-run and short-run models that are represented in Equations (8) and (9) respectively.

$$\ln CO_{2t} = \alpha_2 + \sum_{i=1}^{p} \beta_{2i} \ln CO_{2t-i} + \sum_{i=0}^{q_1} \eta_{2i} \ln GDP_{t-i} + \sum_{i=0}^{q_2} \gamma_{2i} \left(\ln GDP_{t-i}\right)^2 + \sum_{i=0}^{q_3} \theta_{2i} \ln UR_{t-i} + \sum_{i=0}^{q4} \pi_{2i} \ln Ind_{t-i}$$
$$+ \sum_{i=0}^{q_5} \phi_{2i} \ln Gini_{1i} + \sum_{i=0}^{q_6} \omega_{2i} \ln P_{t-i} + \sum_{i=0}^{q_7} \psi_{2i} \ln EI_{t-i} + \varepsilon_{2t} \quad \ldots (5)$$

$$\Delta \ln CO_{2t} = \alpha_3 + \sum_{i=1}^{p} \beta_{3i} \Delta \ln CO_{2t-i} + \sum_{i=0}^{q_1} \eta_{3i} \Delta \ln GDP_{t-i} + \sum_{i=0}^{q_2} \gamma_{3i} \Delta \left(\ln GDP_{t-i}\right)^2 + \sum_{i=0}^{q_3} \theta_{3i} \Delta \ln UR_{t-i}$$
$$+ \sum_{i=0}^{q_4} \pi_{3i} \Delta \ln Ind_{t-i} + \sum_{i=0}^{q_5} \phi_{3i} \Delta Gini_{t-i} + \sum_{i=0}^{q_6} \omega_{3i} \Delta \ln P_{t-i} + \sum_{i=0}^{q_7} \psi_{3i} \Delta \ln EI_{t-i} + \mu ECT_{t-1} + \varepsilon_{3t} \quad \ldots (6)$$

Where, $\mu$ is the speed of adjustment parameter and $ECT_{t-1}$ is the lagged residuals that are obtained from the estimated cointegration model.

In order to confirm the robustness and stability of the estimated model this study applied a diagnostic test including investigating normality, no autocorrelation and homoscedasticity using the Jarque-Bera, Breusch-Godfrey LM and Breusch-Pagan-Godfrey tests, respectively. The decision rule for all assumptions is not to reject the null hypothesis ($H_0$) if the probability is greater than 0.05. Moreover, CUSUM and CUSUM Square (CUSUMQ) tests were also employed to determine the stability of the model.

### 2.3.3. Toda-Yamamoto (TY) Approach to Granger Causality

The existence and direction of causal relationship among variables under consideration is tested the model developed by Toda and Yamamoto (1995) called Toda-Yamamoto (TY hereafter) causality test. The TY method has numerous statistical advantages over conventional Granger causality testing methods (Chindo et al., 2014; Gokmenoglu & Taspinar, 2018; Toda & Yamamoto, 1995; Uzar & Eyuboglu, 2019). The TY method can be applied regardless of cointegration status and order of integration of series under consideration and, therefore, avoids the uncertainties created by pre-testing for cointegration and unit root (Uzar & Eyuboglu, 2019). Moreover, there is no loss of information because the TY method is applicable to VAR at level (Toda & Yamamoto, 1995; Uzar & Eyuboglu, 2019).

The basic idea behind TY method is estimating a $(k+d_{\max})^{th}$-order VAR where k is the correct lag length of the VAR model and $d_{\max}$is the maximal order of integration. The TY representation of $CO_2$ emissions is given as follows:

$$\ln CO_{2t} = \beta_{1_0} + \sum_{i=1}^{k} \theta_{1i} \ln CO_{2t-i} + \sum_{i=p+1}^{k+d_{\max}} \Omega_{1_i} \ln CO_{2t-i} + \sum_{i=1}^{k} \delta_{1_i} \ln Gini_{t-i} + \sum_{i=p+1}^{k+d_{\max}} \phi_{1_i} \ln Gini_{t-i} + \sum_{i=1}^{k} \gamma_{1i} \ln Y_{t-i}$$
$$+ \sum_{i=p+1}^{k+d_{\max}} \psi_{1_i} \ln Y_{t-i} + \sum_{i=1}^{k} \mu_{1i} \ln P_{t-i} + \sum_{i=p+1}^{k+d_{\max}} \eta_{1i} \ln P_{t-i} + \sum_{i=1}^{k} \vartheta_{1i} \ln UR_{t-i} + \sum_{i=p+1}^{k+d_{\max}} \omega_{1i} \ln UR_{t-i}$$
$$+ \sum_{i=1}^{k} \varphi_{1i} \ln IE_{t-i} + \sum_{i=p+1}^{k+d_{\max}} \mu_{1i} \ln EI_{t-i} + \sum_{i=1}^{k} \tau_{1i} \ln Ind_{t-i} + \sum_{i=p+1}^{k+d_{\max}} \tau_{1i} \ln Ind_{t-i} + \varepsilon_{1t} \quad (7)$$

Equations for other series can be constructed in the same fashion.

The modified Wald (MWald) test is used to test the direction of causal relationship among the variables under study.

## 3. Results And Discussion

## 3.1. Unit Root Test Results

Even though the ARDL and DOLS approaches to cointegration and TY Granger causality method are applicable regardless of the order of series under consideration, conducting unit root test is needed for two purposes: to identify that no variable is I(2) under which the ARDL is not valid and to know the maximal lag length to augment the VAR (p) for the purpose of conducting TY Granger causality test. Clemente-Montanes-Reyes unit root and Zivot-Andrews unit root tests are employed to test the order of integration of variables under consideration.

The estimated unit root test results are summarized in Table 2. The result shows that income inequality (lnGini) and population (lnP) are stationary at a level under both unit root test approaches. Moreover, urbanization (lnUR) is stationary at a level under CMR unit root test. All variables become stationary after first differencing [I(1)]. This implies that the variables are a combination of I(0) and I(1), thus the ARDL approach to cointegration is the most appropriate for this study.

Table 2
Unit Root Tests with Structural Breaks

| Clemente-Montanes-Reyes Unit-Root Test with Double Mean Shifts | | | | | | | Zivot-Andrews Unit Root test allowing for a single break in intercept and/or trend | | |
|---|---|---|---|---|---|---|---|---|---|
| **At levels** | Innovative outliers | | | Additive outliers | | | t-statistic | | Break date |
| | t-statistic | TB1 | TB2 | t-statistic | TB1 | TB2 | | | |
| lnY | -0.572 (3) | 1990 | | -3.153 (0) | 1991*** | 2006*** | -2.983 (2) | 2001 | |
| $(lnY)^2$ | -2.892 (4) | 1990*** | 2004*** | -3.364 (1) | 2009*** | | -4.090(0) | 2001 | |
| lnGini | -6.660 (5)** | 1989*** | 1997*** | -5.090 (1) | 1989*** | 2001*** | -6.433 (1)*** | 1999 | |
| lnP | -3.168 (4) | 1992 | | -4.012 (10)** | 1989*** | | -5.134 (2)** | 1994 | |
| lnUR | -6.468 (1)** | 1984*** | 2007*** | -8.048 (7)** | 1992*** | 2006*** | -3.378 (1) | 2010 | |
| lnInd | -1.633 (1) | 2001* | 2008** | 0.209 (8) | 2000*** | 2008*** | -2.642 (1) | 1994 | |
| lnEI | 3.067 (10) | 1992** | 2002 | -3.262 (0) | 1995*** | 2006*** | -1.938 (1) | 1991 | |
| **First Difference** | | | | | | | | | |
| lnY | -5.260 (1)** | 2001*** | | -7.866 (1) | 1991 | 2003 | -4.472 (2) | | 1992 |
| $(lnY)^2$ | -5.309 (1)** | 2003*** | 2001** | -5.520 (1)** | 2004*** | | -3.892 (2) | | 1992 |
| lnGini | -7.828 (1)** | 1997*** | 2004*** | -7.828 (1)** | 1997*** | 2004*** | -4.971 (1)** | | 1997 |
| lnP | -5.613 (10)** | 1990 | | -5.093 (12)** | 1985*** | | -12.071 (1)*** | | 1992 |
| lnUR | -33.914 (0)** | 1995*** | 2007*** | -5.147 (1) | 1995*** | 2010*** | -6.450 (2)*** | | 2009 |
| lnInd | -5.324 (0) | 1986*** | 1991*** | -0.401 (3) | 1989 | 2008*** | -7.490 (0)*** | | 1993 |
| lnEI | -12.028 (0)** | 1992** | 2001 | -11.213 (0)** | 1991** | | -12.130 (0)*** | | 1983 |

## 3.2. Cointegration Test Results

The first step in conducting the long-run and short-run estimations using ARDL approach is ensuring the existence of cointegration among the variables of interest through a bounds test. Since bounds test is lag length sensitive (A. Baloch et al., 2017), Schwartz Bayesian Criterion (SBC) was employed to choose the lag length due to its parsimonious specification (Pesaran, M. H., Shin, 1999). The determined lags length for the model in Eq. 9 is given in Table 3.

Table 3 below presents the ARDL bounds testing approach to cointegration. The calculated F-statistic is 45.75 which is greater than the upper critical bound (3.91) at 1% level of significance. This confirms that a long-run relationship exists between the variables of interest.

**Table 3: ARDL Bounds Cointegration Test Result**

| Model Specification | Selected Model | F-Stat. | Result |
|---|---|---|---|
| $F_{lnCO2}(lnCO_2 \mid lnY, ln(Y)^2, lnGini, lnP, lnUR, lnInd, lnES)$ | (2,3,3,3,3,3,3,3) | 45.75*** | Cointegration |
| Bounds Test Critical Values | CV | I(0) | I(1) |
| | 1% | 2.54 | 3.91 |
| | 5% | 1.97 | 3.18 |

One the existence of cointegration among the variables of interest is established the next step is to estimate the long-run and short-run relationships based on Eq. (3). Long-run and short-run relationships are estimated based on the selected ARDL model. The empirical results of long-run and short-run are presented in Table 4 and Table 5.

Table 4
Long-run Coefficients

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| lnY | 1.054495 | 0.117389 | 8.982943 | 0.0029 |
| $(lnY)^2$ | -0.105909 | 0.012416 | -8.530360 | 0.0034 |
| lnGini | 0.073499 | 0.032163 | 2.285197 | 0.1064 |
| lnP | 0.228519 | 0.048479 | 4.713812 | 0.0181 |
| lnUR | 0.381021 | 0.055748 | 6.834637 | 0.0064 |
| lnInd | 0.118412 | 0.008197 | 14.445068 | 0.0007 |
| lnEI | 0.072886 | 0.023646 | 3.082393 | 0.0540 |
| $R^2$ | 0.99*** | | | |
| Diagnostic Tests | | | | |
| Test | F-Statistic | P-value | DW Stat | 3.737018 |
| $\chi^2 Normal$ | 0.0277 | 0.9862 | | |
| $\chi^2 Serial$ | 19.2997 | 0.1589 | | |
| $\chi^2 RESET$ | 0.0499 | 0.8440 | | |
| $\chi^2 Hetero$ | 6.3516 | 0.1450 | | |

The long-run relationship results given in Table 4 show that the coefficient of economic growth (lnY) is positive and statistically significant at 1% level of significance. More specifically, a 1% increase in GDP leads to 1.05% rise in $CO_2$ in Ethiopia during the study period, indicating that economic growth is the major cause of environmental degradation. This finding is supported by Taka et al.( 2020) who found that about 52% of $CO_2$ emissions in Ethiopia have been strongly driven by the economic effects during 1970–2017.

The positive effect of economic growth on $CO_2$ emissions is attributable to the fact that at the early development stage of the economy, developing countries often ignore the possible negative effects of economic growth on the environment (Khan, 2019). Moreover, at an early stage of economic development process, industrialization increases and developing countries have weak environmental regulations to curb emissions from the increasing industrialization. The positive association between environmental quality and economic growth is related to the scale effect (Dogan & Inglesi-Lotz, 2020; Lorente & Alvarez-Herranz, 2016). According to the scale effect, economic activities rise proportionally as economy grows, therefore, environmental quality diminishes with economic growth. For instance, Usama et al. ( 2020) argued that expansion in economic activities like farming, construction, mining, settlement and other economic activities have led to deforestation, a factor responsible for $CO_2$ emissions, in Ethiopia. Nathaniel and Khan (2020) asserted that the growing economy demands more energy to run the economic activities. However, increasing consumption of non-renewable energy, which is a case of developing countries, deteriorates environmental quality. For the growing economy to contribute to the environmental quality, Khan et al. (2020), Khan, Yu, Belhadi, et al. (2020) and Dogan and Ozturk (2017) suggested that renewable energy should replace non-renewable energy sources.

The coefficient of economic growth squared is negative and statistically significant. A 1% increase in economic growth squared will decrease $CO_2$ emissions in Ethiopia, keeping other things constant. It indicates that countries are concerned to reduce environmental degradation after they achieve a certain development level. Dasgupta et al.(2002) argued that industrial and other sectors become cleaner, people give more value to cleaner environment, and regulatory framework becomes more efficient at later stage of economic development. As the economy climbs along the development ladder, it shifts from production of energy-intensive industrial products to services (composition effect) and resorts to energy-efficient techniques of production (technical effect). The results from ARDL, DOLS and U-test confirm that effect of economic growth on environmental quality follows the EKC hypothesis. This is consistent with the finds of Hundie (2018) and Usama et al. (2020) for Ethiopia, Kusumawardani and Dewi (2020) for Indonesia, Uzar and Eyuboglu (2019) for Turkey, Zhu et al. (2018) for BRICS, Sarkodie & Ozturk (2020) for Kenya, M. A. Baloch, Ozturk, et al. (2020) for the OECD countries, Hailemariam & Dzhumashev (2019) for OECD countries and Dogan & Inglesi-Lotz (2020) for European countries. But this finding is not consistent with results of Dogan and Turkekul (2016), Dogan and Ozturk (2017) and for USA and Sohag et al. (2019) for the OECD countries which found U-shape relationship and Gill et al. (2017) for Malaysia which found a linear relationship. The inconsistence in the findings regarding the relationship between environmental quality and economic growth may be associated with factors like selection of economic growth indicator. For instance, Dogan and Inglesi-Lotz (2020) found that the EKC hypothesis holds in African countries when aggregate GDP is used to measure economic growth. However, the hypothesis is not confirmed when industrial share to GDP is used. The contradictions in findings on the impact of economic growth on environmental quality can also be associated with various pollution reduction strategies as contended by M. A. Baloch, Ozturk, et al. (2020).

Considering the extreme point given in Table 9, the turning point of real GDP at which economic growth decouples from $CO_2$ emissions is $276.4 billion, and real GDP during 2014 is $44.01 billion. Average growth rate during the study period is 5.3%. Based on this, it was estimated that it would take 35.5 years to reach the turning point. This shows that decoupling will occur after a year 2050.

Coefficient of income inequality variable is positive in sign, but statistically insignificant showing that policies that affect income distribution do not affect $CO_2$ emissions in Ethiopia. This empirical finding is supported by Wolde-rufael and Idowu (2016) for China and India, Jorgenson et al. (2017) for USA, Uddin et al. (2020) for G7, Bhattacharya (2019) for India, Wu and Xie (2020) for low income non-OECD countries and A. Baloch et al. (2017) for Pakistan. The neutrality of effect of income inequality on environmental quality implies that income redistribution to reduce inequality induces both propensities to consume and emit which rises and then falls. This is because poor people's consumption increases as they transit to the middle class. According to the PEA, improvement in income inequality reduces environmental pollution by the wealthy class while the poor class positively contribute to environmental degradation. The two forces may offset each other and make income inequality neutral in determining environmental quality. Finding of this study contradicts with the findings of Padhan et al. (2019) for the Next-Eleven countries and Hailemariam et al. (2019) for the OECD countries who found a statistically significant impact of income inequality on environmental quality. The variations in result regarding the interplay between income inequality and environmental quality is due to different factors like how one captures income inequality in a particular study and magnitude of income inequality. With this regard, Padhan et al. (2019) found that wealth disparity decreased $CO_2$ emissions when income inequality was more than 45% while income inequality less than 40% resulted in increased $CO_2$ emissions. Unlike previous studies, Hailemariam et al. (2019) argued that income share of the top 10% as a measure of income inequality rather than Gini coefficient is more appropriate in examining the effect of income inequality on environmental quality. Therefore, variations in measurement of income inequality may be responsible for the inconsistent findings regarding the interplay between income inequality and environmental quality.

Another long-run predictor of environmental degradation in Ethiopia during the study period is population size. The coefficient of population size (lnP) is positive and statistically significant at 1% level of significance. It implies that population growth has a detrimental effect on environment in Ethiopia. One channel through which population size affects environmental quality is age structure of the total population. Cao et al.(2020) contended that countries with a higher proportion of young people more degrades the environment because the young produce more pollutants than the old. According to United Nations Department of Economic and Social Affairs Population Division [UNDESA] (2019) the share of working age population (15–64) was 35% about in 2020 while that of youth (15–24) was 21.5%. Therefore, the positive effect of population size on environmental quality may be attributed to the growing share of young population in Ethiopia. Besides, the higher the population size, the higher energy and food consumption which increases $CO_2$ emissions. The population has been striving to fulfill the increasing demand for energy and food by expanding agricultural lands that has led to deforestation. According to the Global Forest Watch (2020) Ethiopia lost about 25.6kha of tree cover which is equivalent to 8.56Mt of CO□ of emissions. This finding is consistent with results reported in Hundie (2018), Chen et al. (2020), Ravallion et al. (2000), Yang et al. (2020), Zhou and Liu (2016) and Li et al.(2019); but in contradiction with M. A. Baloch et al. (2020) who found a statistically significant negative relationship between $CO_2$ emissions and population in Sub-Saharan Africa. On the other hand, finding reported in Wu and Xie (2020) reveals that population size has no significant effect on environmental quality in OECD and non-OECD countries. Evidences reveals that the impact of population size is much stronger in developing countries than the developed ones ( see Poumanyvong and Kaneko (2010). The reason is that developing countries' primary source of energy is biomass, even though it has been decreasing slightly in countries like Ethiopia. For instance, the share of biomass in energy consumed was 92.6% in 2012 and 87.7% in 2020 out of which the lion's share goes to firewood according to Mondal et al. (2018). Moreover, the growing population size in Ethiopia resulted in increasing demand for food. Scarcity of land and low productivity are among the major bottlenecks to satisfy the increasing demand for food, which increases the risk of deforestation as a result of agricultural land expansion and agricultural land degradation due to overutilization. The situation is common in rural Ethiopia, though the government has recently started putting a concerted effort in order to reverse it through afforestation and soil protection programs.

Urbanization has a statistically significant positive impact on CO2 emissions, which indicates that it has intensify environmental degradation during the study period. This is due to the fact that urban population puts pressure on environment through their consumption because consumption pattern of urban population is different; they consume more food and durable goods, energy and water compared to rural population. Wu and Xie (2020) also argued that urbanization leads to increase in usage of and demand for commodity and public services, and expansion of manufacturing industry. Moreover, mode of transport in urban areas is different where vehicles are crowded for road transport and responsible for air pollution. For example, energy demand by the transport sector constitutes the second largest (7.8%) share out of the total energy demand in Ethiopia during 2020, as indicated in Mondal et al. (2018). This result stands with the finding of Shahbaz et al. (2015) for Malaysia, (Zhou & Liu, 2016), Li et al. (2019) for China, Wu and Xie (2020) for the OECD countries and Cao et al. (2020) for China. But it is in contradiction with the finding of Chen et al. (2020) find that accelerated urbanization reduces $CO_2$ emissions in G20 countries while Wu and Xie (2020) find insignificant interplay between urbanization and environmental quality in OECD and Non-OECD countries.

Industrialization has a positive impact on environmental degradation in Ethiopia during the study period; and this finding corroborates with the finding of A. Baloch et al. (2017) for Pakistan, Ullah et al. (2020), for Pakistan, M. A. Baloch et al. (2020) for Sub-Saharan Africa, Yang et al.(2020) for developing countries and Cao et al. (2020) for China, but contradicts with A. Baloch et al. (2017) for Pakistan. Ethiopia's rise in GDP has been largely driven by an increase in industrial activity, which is energy intensive relative to other sectors. According to Mondal et al. (2018) industrial sector's share of demand for electricity is projected to be the highest (41.2%) in Ethiopia during 2020. The increment in energy consumption leads to rise in $CO_2$ emissions. Moreover, expanding industrial sector positively contributes to environmental degradation through polluting rivers, water and environment.

The coefficient of energy intensity is positive and statistically significant at 10% level of significance, indicating that higher energy intensity increases $CO_2$ emissions in Ethiopia during the study period. This finding corroborates with the findings of Hundie (2018) for Ethiopia, Shahbaz (2014) for Bangladesh, Yang et al. (2020) for developing countries and Padhan et al. (2019) for the next eleven countries. The positive interaction between $CO_2$ emissions and energy intensity in Ethiopia is related to high energy intensity which has been increasing over time (Hundie & Daksa, 2019). Evidences show that demand for energy in Ethiopia has been increasing due to different factors like population growth, economic growth, and urbanization. Ethiopia depends primarily on biomass fuels as the main energy sources which is characterized as inefficient (Ethiopian Environmental Protection Agency (EEPA), 2004; Hundie & Daksa, 2019; Mondal et al., 2018; Taka et al., 2020). The increasing consumption of inefficient biomass energy increases energy intensity, which in turn positively influences CO2 emissions. Due to heavy reliance on the traditional biomass, which accounts for 91% of energy consumed, CO2 emissions in Ethiopia have increased from

 al et al., 2018). They further projected that the GHG emissions in Ethiopia will be doubled between

2012 and 2030 due to growth in energy consumption. The long-run estimates show that the impact of energy intensity on $CO_2$ emissions in Ethiopia during the study period is weak. This finding is consistent with the finding of Taka et al. (2020) for Ethiopia. Taka et al. (2020) argued that though the role of energy intensity in Ethiopia was less pronounced, still it is vital enough to consider it in national level policies. This is important because improving energy intensity through utilizing renewable energy, for instance, is good not only for environmental quality but also for economic growth of developing countries (see Rafindadi & Ozturk, 2017; Ulucak & Bilgili, 2018). Some scholars (M. A. Baloch, Ozturk, et al., 2020) recommend that countries need to undertake energy innovation in order to circumvent the negative impact of energy consumption on environmental quality.

The estimated ARDL model is free from any diagnostic problems. The statistical tests conducted using a 5% significance level showed that the model satisfies the assumption of normality, no autocorrelation, and homoscedasticity$(\text{p-value}=0.5214)$. Moreover, the model is well specified as confirmed by the Ramsey RESET test. This implies that the model can produce unbiased estimations.

Table 5
Error Correction Representation for the Selected ARDL Model

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| $\Delta\ln(CO_2)_{t-1}$ | 1.346797 | 0.099324 | 13.55963 | 0.0009 |
| $\Delta\ln Y_t$ | -0.232192 | 0.059026 | -3.933728 | 0.0293 |
| $\Delta\ln Y_{t-1}$ | -2.097325 | 0.086741 | -24.17915 | 0.0002 |
| $\Delta\ln Y_{t-2}$ | -1.328143 | 0.080872 | -16.42269 | 0.0005 |
| $\Delta(\ln Y)_t^2$ | 0.010083 | 0.005396 | 1.868739 | 0.1585 |
| $\Delta(\ln Y)_{t-1}^2$ | 0.220752 | 0.008405 | 26.26297 | 0.0001 |
| $\Delta(\ln Y)_{t-2}^2$ | 0.136727 | 0.008091 | 16.89799 | 0.0005 |
| $\Delta\ln Gini_t$ | -0.207335 | 0.021238 | -9.762612 | 0.0023 |
| $\Delta\ln Gini_{t-1}$ | -0.396941 | 0.041232 | -9.627055 | 0.0024 |
| $\Delta\ln Gini_{t-2}$ | 0.491378 | 0.023650 | 20.77724 | 0.0002 |
| $\Delta\ln P_t$ | 1.683299 | 0.069344 | 24.27451 | 0.0002 |
| $\Delta\ln P_{t-1}$ | 0.659536 | 0.093819 | 7.029893 | 0.0059 |
| $\Delta\ln P_{t-2}$ | -0.200945 | 0.050025 | -4.016901 | 0.0277 |
| $\Delta\ln UR_t$ | -0.630763 | 0.159233 | -3.961259 | 0.0287 |
| $\Delta\ln UR_{t-1}$ | 5.073643 | 0.214773 | 23.62332 | 0.0002 |
| $\Delta\ln UR_{t-2}$ | 8.315893 | 0.361179 | 23.02430 | 0.0002 |
| $\Delta\ln Ind_t$ | 0.136771 | 0.005163 | 26.49147 | 0.0001 |
| $\Delta\ln Ind_{t-1}$ | 0.101071 | 0.007217 | 14.00420 | 0.0008 |
| $\Delta\ln Ind_{t-2}$ | 0.035779 | 0.006584 | 5.433775 | 0.0122 |
| $\Delta\ln EI_t$ | 0.094727 | 0.005855 | 16.17883 | 0.0005 |
| $\Delta\ln EI_{t-1}$ | -0.015277 | 0.007462 | -2.047399 | 0.1331 |
| $\Delta\ln EI_{t-2}$ | 0.032172 | 0.004134 | 7.782615 | 0.0044 |
| $ECT_{t-1}$ | -2.687607 | 0.076942 | -34.93033 | 0.0001 |

The error-correction representation of the ARDL model is given in Table 5. The result shows that all predictors and their lagged values affect the dynamics of $CO_2$ emissions in Ethiopia during 1979–2014. The coefficient of the lagged error term ($ECT_{t-1}$) is -2.69 and statistically significant at 1% level of significance. This finding reaffirms the long-run relationship established through the ARDL approach. More specifically, the deviation from the long-term $CO_2$ emissions path due to certain shock is adjusted by 268.9% over the next year. The $ECT_{t-1}$value is less than − 1 and indicates that the error-correction process fluctuates around the long-term value in a dampening situation rather than monotonically converging to the equilibrium path directly. However, once

this process is complete, convergence to the equilibrium path is rapid (Hundie & Daksa, 2019; Loayza & Ranciere, 2005; Narayan & Smyth, 2005; Uzar & Eyuboglu, 2019).

## 3.3. Stability of the Estimated Model

Misspecification of model in time series data may produce biased results (Hansen, 1992) which may reduce the explaining power of the empirical findings (Farhani & Ozturk, 2015). This study used the cumulative sum (CUSUM) and cumulative sum squares (CUSUMSQ) of the recursive residuals to test for structural stability of the estimated coefficients. Figures 1 and 2 show that the estimated coefficients in the error-correction model are stable over the study period.

This confirms that the selected $CO_2$ emissions model can be used to understand the decision-making policy because the impact of policy changes regarding the explanatory variables of $CO_2$ emissions will not significantly affect the level of $CO_2$ emissions, since the parameters in this equation seems to follow a stable pattern during the estimation period (Halicioglu, 2009).

## 3.4. Granger Causality Test Result

Before conducting Granger causality test, it requires determining appropriate VAR lag length. This study considers lag length selected by Akaike's information criterion (AIC) because it has superior power properties for small sample data (Liew, 2014; Lütkepohl, 2006) and provides efficient and consistent results (Shahbaz et al., 2015). Accordingly, the selected optimal lag length is 3 (Table 6). The unit root test result confirmed that the maximum order of integration is one and there are variables that are stationary at level. This implies that the variables under consideration are a mixture of I(1) and I(0) and the appropriate Granger causality test method under such circumstance is the TY approach (Chindo et al., 2014; Menyah & Wolde-rufael, 2010; Toda & Yamamoto, 1995).

Table 6
VAR Lag Order Selection Criteria

| Lag | LogL | LR | FPE | AIC | SC | HQ |
|---|---|---|---|---|---|---|
| 0 | 266.1742 | NA | 2.21e-17 | -15.64692 | -15.28413 | -15.52485 |
| 1 | 669.6360 | 586.8536 | 2.88e-26 | -36.22037 | -32.95526 | -35.12176 |
| 2 | 827.6944 | 153.2687 | 2.06e-28 | -41.92087 | -35.75344 | -39.84572 |
| 3 | 1033.915 | 99.98575* | 6.45e-31* | -50.54030* | -41.47056* | -47.48861* |
| * indicates lag order selected by the criterion | | | | | | |

The TY approach estimates augmented VAR (4). Then, coefficients of the first three lags of each independent variable are tested whether they affect the dependent variables in the VAR (4) system.

Table 7
Toda-Yamamoto Granger Causality Test Result

| Dependent variables | Sources of causation | | | | | | |
|---|---|---|---|---|---|---|---|
| | $lnCO_2$ | lnGDP | lnGini | lnEI | lnInd | lnP | lnUR |
| | $\chi^2(3)$ | $\chi^2(3)$ | $\chi^2(3)$ | $\chi^2(3)$ | $\chi^2(3)$ | $\chi^2(3)$ | $\chi^2(3)$ |
| $lnCO_2$ | - | 4.76 | 11.31** | 3.84 | 35.61*** | 16.80*** | 16.81*** |
| lnGDP | 88.31*** | - | 863.54*** | 133.93*** | 204.56*** | 135.04*** | 23.51*** |
| lnGini | 38.72*** | 39.75*** | - | 13.60*** | 4.05 | 17.25*** | 40.92*** |
| lnEI | 22.55** | 11.38*** | 15.06*** | - | 14.56*** | 6.14 | 15.04*** |
| lnInd | 38.66*** | 39.26*** | 107.27*** | 29.25*** | - | 84.57*** | 28.26*** |
| lnP | 15.51*** | 19.24*** | 39.80*** | 33.37*** | 33.35*** | - | 39.80*** |
| lnUR | 264.42*** | 30.34*** | 264.42*** | 13.29*** | 258.76*** | 81.42*** | - |
| Statistically significant at 1% (***), 5% (**) and 10% (*) level | | | | | | | |

Since seemingly unrelated regression (SURE) produces more efficient coefficients (Zellner, 1962), this study estimates the VAR (4) systems using SURE. Further, SURE makes computing modified Wald test statistic very simple (Rambaldi & Doran, 1996).

Table 7 presents TY Granger causality test results. As the main objective of this study to is exploring the relationship between economic growth, income inequality and $CO_2$ emissions, interpretation of the Granger causality test result shall focus on results pertaining to these variables. Table 7 shows that there is unidirectional causality running from $CO_2$ emissions to economic growth without feedback. Granger causality between $CO_2$ emissions and income

inequality is bidirectional. This finding shows that $CO_2$ emissions affects economic growth, mainly the component that comes from the agricultural sector. This is because the Ethiopian agriculture mainly depends on weather condition specially rainfall. This situation may exacerbate income inequality between urban and rural people, which in turn has a feedback effect on economic growth and environmental degradation. Income inequality, energy intensity, industrialization, urbanization and population size Granger cause $CO_2$ emissions with feedback effect.

The result also denoted that all variables except $CO_2$ emissions Granger cause economic growth with feedback from economic growth. The implication is that income inequality, energy intensity, population size, industrialization and urbanization influence economic growth. Income inequality Granger causes economic growth with feedback effect, which corroborated with the finding of Hailemariam et al. (2020). This implies that the status of income inequality dependents on the level of economic growth and the level of income inequality determines economic growth. The finding also revealed that energy intensity, industrialization and urbanization depends on the level of economic growth. The result reveals that income inequality, industrialization, population and urbanization Granger cause $CO_2$ emissions, while carbon dioxide Grange causes all variables. This implies that there is bidirectional Granger causality between the aforementioned variables and $CO_2$ emissions.

## 3.5. Robustness Analysis

## 3.5.1. Dynamic OLS (DOLS)

Robustness of the long-run coefficients estimated from the ARDL approach was checked through applying dynamic OLS (DOLS) method. The DOLS procedure has many advantages. The core benefit of the DOLS method is that it fixes the potential endogeneity problem and small sample bias. It can also be applied regardless of whether the underlying regressors are integrated of I(0), I(1) or mutually integrated. Besides, DOLS estimators provide asymptotically efficient cointegrating vectors (A. Baloch et al., 2017; Begum et al., 2015). The DOLS estimators also remove potential effects of autocorrelation (Hailemariam et al., 2019).

Estimated coefficients through the DOLS method are presented in Table 8. The result shows that estimators from DOLS are consistent with the ARDL result, according to sign and statistical significance of the coefficient except for lnP and lnGini. The coefficient of lnP is changed from positive and statistically significant under ARDL to negative and insignificant under DOLS method. The coefficient of income inequality (lnGini) becomes statistically significant under DOLS method, while the sign remains the same. This may arise from the fact that the ARDL approach to cointegration is better in precision of estimation and reliability of statistical inferences as argued by Panopoulou and Pittis (2004) because it fully corrects for the second-order asymptotic bias effects of cointegration. Therefore, this study relies on results from ARDL regarding interpretation of the long-run coefficients and related inferences. However, cointegration test result is consistent with the result obtained through ADRL approach, which indicates that the variables under consideration are cointegrated.

**Table 8:** Long Run Coefficients of DOLS Estimator and Cointegration Test

| Long-run Coefficients DOLS Estimator | | | Cointegration Test-Hansen Parameter Instability | |
|---|---|---|---|---|
| Variable | Coefficient | P-value | | |
| lnY | 0.364678 | 0.0028 | Lc Statistic | 0.154817 |
| $(lnY)^2$ | -0.031483 | 0.0077 | Stochastic Trends (m) | 7 |
| lnGini | 0.214274 | 0.0003 | Deterministic Trends (k) | 0 |
| lnP | -0.095461 | 0.1712 | Excluded Trends (p2) | 0 |
| lnUR | 0.773050 | 0.0002 | Prob. | 0.2 |
| lnInd | 0.162089 | 0.0000 | | |
| lnEI | 0.257286 | 0.0003 | $H_0$: Series are cointegrated | |
| R-squared | 0.9984 | | Long-run variance: 3.13E-06 | |

*Source: Own Computations Using EViews 10*

## 3.5.2. Sasabuchi–Lind–Mehlum (SLM) U-Test for a Quadratic Relationship

The ARDL and DOLS results revealed that the relationship between $CO_2$ and economic growth in Ethiopia follows an inverted U-shape. Nevertheless, Lind and Mehlum (2010) contend that the traditional test of U-shaped or inverse U-shaped relationship is too weak since it fails to provide a sufficient condition for the presence of U-shape relationship. For instance, it may erroneously detect a quadratic relationship as U-shape when the true relationship is convex but monotone over relevant data values (A. Baloch et al., 2017; Lind & Mehlum, 2010). Therefore, exact U-test that provides both necessary and sufficient conditions for the existence of the inverse U-test needs to be applied. To address these problems this study applied the Sasabuchi–Lind–Mehlum (SML) U-test

The model for SLM U-test as specified in A. Baloch et al.(2017) can be stated as follows:

$$CO_{2t} = \alpha \ln Y_t + \beta (\ln Y_t)^2 + \varepsilon_t \quad (8)$$

The U-shape hypothesis is expressed as:

$H_0: (\alpha + 2\beta \ln Y_{\min} \leqslant 0)$ and/or $(\alpha + 2\beta \ln Y_{\max} \geqslant 0)$ against the inverse U-shaped hypothesis given by $H_1: (\alpha + 2\beta \ln Y_{\min} \geqslant 0)$ and/or $(\alpha + 2\beta \ln Y_{\max} \leqslant 0)$ where $\ln Y_{\min}$ and $\ln Y_{\max}$ denote the minimum and maximum value of natural logarithm of GDP. If the null hypothesis is rejected, inverted U-shape or the EKC hypothesis holds. Results of the U-test are presented in Table 9. Table 9 shows that, the lower bound slope of $\ln Y$ is 0.361 and the upper bound slope of $\ln Y$ is -0.270. Both coefficients are statistically significant at 1% level of significance. The result discloses that the null hypothesis of monotone or U-shaped relationship is rejected in favor of the inverted U-shape/EKC. This implies that the relationship between economic growth and $CO_2$ emissions follows the inverted U-shape/EKC hypothesis which is in line with the result obtained from the ARDL and DOLS approaches.

Table 9
U-test for U-shaped Relationships (Dep't var: $CO_2$ emissions)

| | Lower bound | Upper bound |
|---|---|---|
| Interval | 4.694 | 6.316 |
| Slope | 0.361 | -0.270 |
| t-value | 21.165 | -8.832 |
| $P > \|t\|$ | 0.000 | 0.000 |
| Extreme point | 5.621682 | |
| Overall test of presence of an Inverse U shape: | | |
| t-value | 8.83 | |
| $P > \|t\|$ | 0.000 | |
| 95% Fieller interval for extreme point: [5.4976219; 5.7570405] | | |
| Test: $H_1$: Inverse U shape vs. $H_0$: Monotone or U shape | | |
| *Source: Own Computation using Stata 16* | | |

## 4. Conclusion And Recommendation

Ethiopia aims to achieve three interrelated goals: first, achieving middle-income status by 2025; second, reducing income inequality; and third, limiting GHG emissions to around today's 150 Mt $CO_2$e by 2030. Striving to achieve the first goal affects the latter two goals in an unpredictable manner as revealed in empirical evidence. The main objective of this study was, therefore, to examine the nexus between income inequality, economic growth and $CO_2$ emissions in Ethiopia by incorporating other key factors affecting $CO_2$ emissions between 1979 and 2014.

The present study employed unit root tests that consider structural breaks for stationarity test. ARDL bounds test and DOLS approaches to cointegration were used to test the long-run relationship between variables under consideration. The current study employed U-test to examine whether the relationship between environmental quality and economic growth supports the EKC hypothesis. Moreover, direction of causal relationship among variables of the study was tested using Toda-Yamamoto Granger causality test approach.

Results of the ARDL and the DOLS approaches show that there is a co-integrated, stable and robust long-run relationship among income inequality, energy intensity, economic growth and its square, industrialization, urbanization and population size during the study period. The impact of economic growth on $CO_2$ emissions in Ethiopia is not linear; rather, it follows an inverted U-shape supporting the EKC hypothesis during the study period. This implies that economic growth takes place at the expense of environmental degradation during the early stage of economic development process. Economic growth then improves environmental quality in later stage of development process once a certain level of development is achieved. Because people's demand for clean environmental and efficiency of environmental regulations increase at this stage of economic development. The impact of income inequality on environmental degradation in Ethiopia is sensitive to the choice of the econometric techniques. Results from the ARDL bounds test approach to cointegration shows that income inequality does not have any statistically significant influence on the environmental quality. According to this result, redistributing income from the wealthy class to the poor class increases propensity to emit first and then decreases it. This makes income inequality neutral in determining the environmental degradation. However, estimates of the DOLS reveal that income inequality intensifies environmental degradation which is consistent with the PEA and Veblen's emulation theory. Urbanization, industrialization, and energy intensity consistently increase $CO_2$ emissions in Ethiopia during the study period.

Results from the TY Granger causality show that Granger causality runs from income inequality, industrialization, population and urbanization to $CO_2$ emissions with feedback effect. There is a bidirectional causal relationship between economic growth and income inequality. Moreover, $CO_2$ emissions

Granger causes economic growth with no feedback response. This implies that economic growth, income inequality and $CO_2$ are interconnected in Ethiopia during the study period. Policy intervention which is designed to affect any one of the three variables may have potential direct or indirect impact on the remaining two variables. Therefore, policy intervention intended to influence any of the three variables should consider the interconnection among them.

Based on findings the study, Ethiopia may implement the following policy recommendations are provided in order to improve environmental quality in Ethiopia. First, Ethiopia needs to accelerate its economic growth sustainably under the climate resilient green growth strategy which will reduce the negative environmental impacts of the growing economic activities. This requires sustainable use of natural resources, encouraging green practices, and implementing strong environmental policies and strategies. Promoting economic growth may also take place through shifting the economic structure from service and agriculture sector to high productive and energy efficient industrial sector. Accelerating economic growth may shorten time span at which the decoupling between economic growth and environmental degradation will start. Environmentally friendly economic growth process is also crucial in addressing the growing income inequality and poverty which reduce human pressure on the natural environment. Improvement in economic growth also decreases energy intensity and transforms industrialization and urbanization process in the country. These improvements help contribute to reduction in $CO_2$ emissions. Moreover, it is necessary for Ethiopia to reduce population growth rate and optimize the population structure.

Second, reducing income inequality has twofold advantages: (1) it reduces $CO_2$ emissions, and (2) it ignites economic growth and reduce social tensions on the other hand. Higher income inequality decelerates the rate of economic growth through different ways. It increases socio-political instability; it increases credit constraint to invest in human capital; and it adversely affects capital accumulation. Deterioration of growth rates due to the increasing income inequality pulls people into severe poverty where they are enforced to degrade natural resources like forest and soil. This results in environmental degradation. Moreover, increasing income inequality may lengthens the decoupling time between environmental quality and economic growth.

Third, $CO_2$ emissions reduction requires decreasing energy intensity/improving energy efficiency through designing appropriate energy conservation strategies with due attention to rural areas shifting from traditional biomass energy to modern energy sources like electricity. Additionally, Ethiopia needs to reduce its energy intensity through energy efficient appliances. The adoption of energy efficient appliances both by households and industrial companies also helps in reducing CO2 emissions. The country can also improve its environmental quality by replacing fossil fuel energy with clean energy sources. Accelerating universal electrification and promoting improved cookstove can also contribute to energy efficiency and environmental quality.

## Declarations


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## Figures

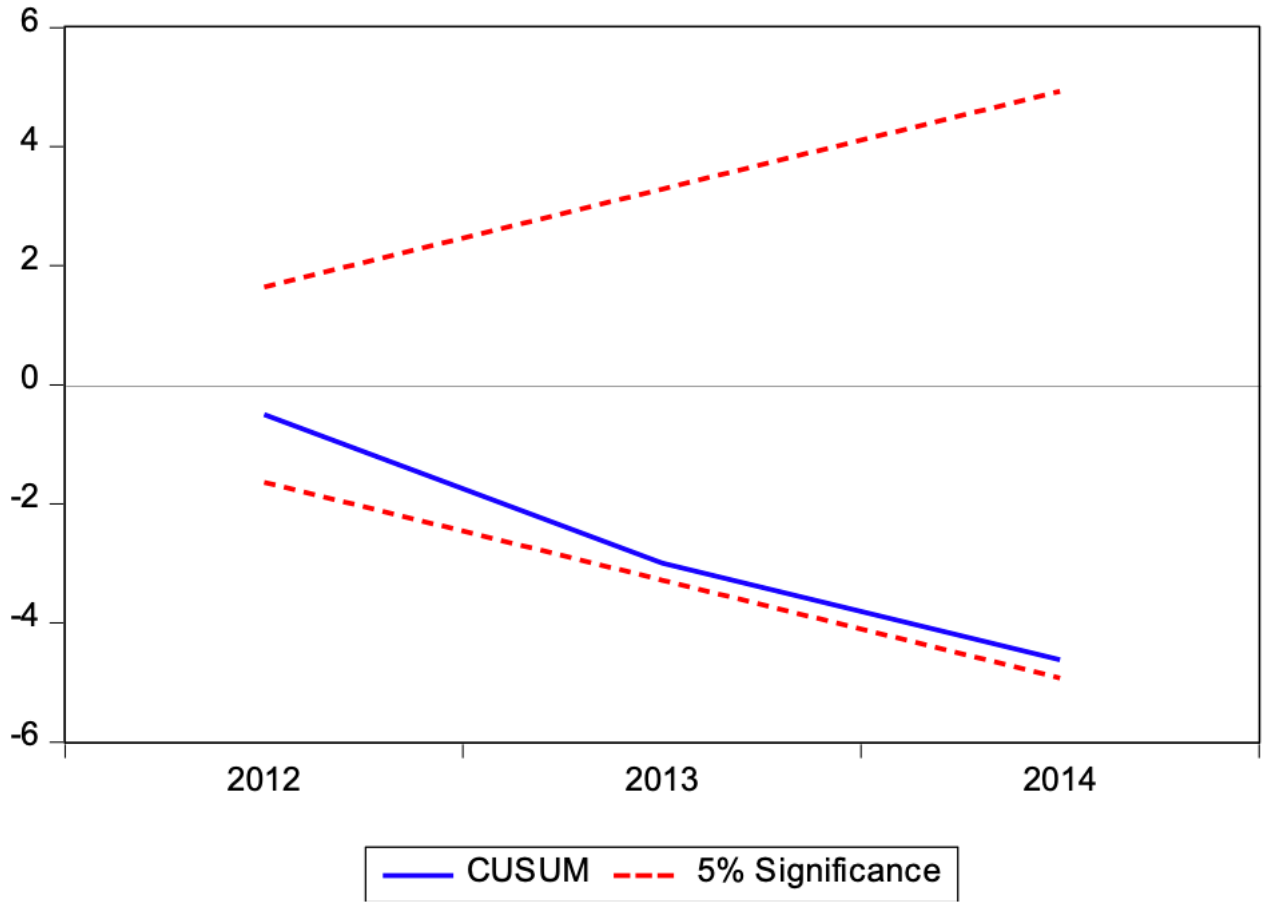


**Figure 1**

Plot of Cumulative Sum of Recursive Residuals

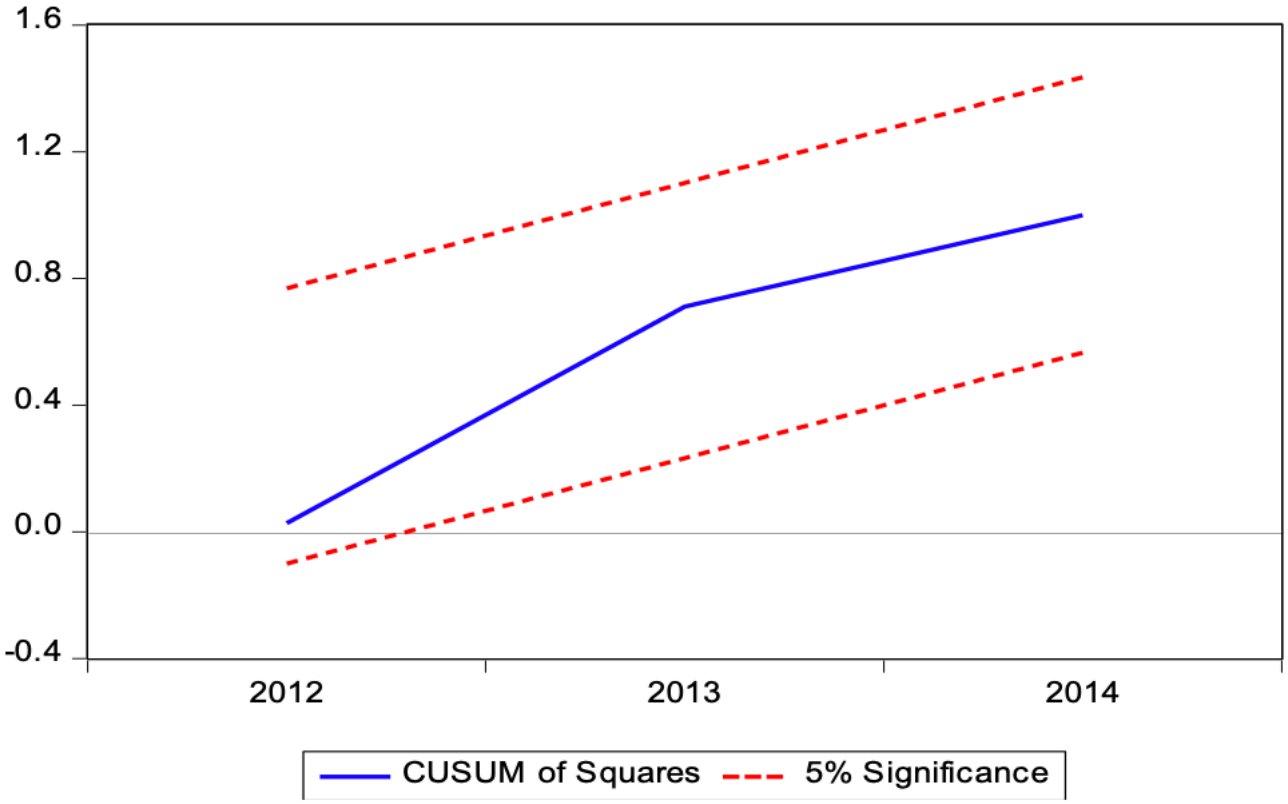


**Figure 2**

Plot of Cumulative Sum of Squares of Recursive Residuals

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