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Income Inequality and Economic Freedom Revisited: Are Freedom and Equality Conflicting Values? Evidence from the twenty-first Century

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ABSTRACT
This empirical research addressed the short – and long-run relationship between economic freedom (and its subcomponents) and income inequality using a panel of 102 countries between 2000 and 2018. The results of employing an autoregressive distributed lag model showed that economic freedom has a detrimental impact on income inequality measured by any of the main inequality indicators. However, the results point to a relatively inelastic relationship. Additionally, the study explored the interactions between the subcomponents of economic freedom and income inequality, again pointing to a rigid relationship. While the size of the government and legal property rights increase income inequality, deregulation exerts the opposite effect. This paper closes with future guidelines for research.

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1. Introduction

The concept of economic freedom emphasizes the idea that individuals should have favorable conditions to freely pursue their interests, where the state should only interfere in the fundamental societal issues that are out of the scope of an individual agent: protection, law, justice, and provision of essential public goods. In addition, it emphasizes that the economy should rely on individual agents' interactions (supply vis-à-vis demand), supported by the classic liberal idea that individuals promote social welfare by promoting their self-interest. Hence, economic freedom stresses the importance of private property, free domestic and international markets, the rule of law, and the government's limited role.

This work revisits the empirical relationship between economic freedom and income inequality, a debate that started soon after the forthcoming release of the first measures (indexes) of Economic Freedom – the level of freedom for an individual to engage in an economic decision. Although this is a contradictive discussion, economic evidence has supported the idea that economic freedom fosters economic growth, whereas economic growth reduces (or, at least, does not increase) income inequality.

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This empirical work measures the relative income inequality based on cross-sectional time-series data. In that sense, it is important first to note some initial points. So far, no standard empirical model established in the economic literature allows controlling and estimating as precisely as needed income inequality. Also, one should be aware of the data advantages and disadvantages based on the measurement theory of income inequality.[1] Furthermore, the authors would like to stress that all the data is collected from the most recent dataset versions available for research and is independently modeled to allow international (between countries) comparison. Finally, the research found that economic freedom is detrimental to income inequality due to lower government support (perhaps in income acquisition opportunities) and increased property value. However, that relationship is inelastic, suggesting that economic freedom is not a fundamental factor behind income inequality.

This study contributes to the literature in three ways. First, this paper employs the most recent database versions available for income inequality (SWIID version 9.0) and economic freedom (EFW 2020 Panel Dataset and IEF 2021 version), which allows constructing a balanced panel while most of the previous studies made use of unbalanced frameworks. Second, and to the best of one's knowledge, this is the first study on this topic to document a model specification test for omitted variables and the overall model specification. Third, using the autoregressive distributed lag model (ARDL) methodology introduced by Pesaran and Smith (1995), the possible endogeneity problems that arise if income inequality impacts economic freedom and the possible spurious regression problems (caused by non-stationary variables) are explicitly accounted for (e.g. Asteriou et al., 2021; Cho et al., 2023).

The empirical argument is constructed as follows. Section 2 presents a detailed literature review regarding the freedom-inequality nexus of previous research. Section 3 introduces the data and methodology used in this analysis. This paper employed an error correction model through a linear transformation of an ARDL for a panel of 102 countries in the twenty-first century to study economic freedom's short – and long-run impacts (and its subcomponents) on income inequality. Section 4 presents the estimation results, and Section 5 discusses them. Additionally, Section 6 addresses a sensitivity analysis made to changes in the main variable of interest and the dependent variable to check the robustness of the empirical results. Finally, Section 7 closes with concluding remarks.

2. Literature Review

Scholars have assessed the impact of economic freedom (EF) on a range of economic outcomes[2] since the first indicators of EF were created. As Hall and Lawson (2014) reported, of the 198 articles (at the time) where economic freedom was imputed as an independent variable in an empirical model, 134 studies found a 'positive' economic outcome, while only eight papers documented a 'negative' outcome associated with it. For example, economic literature consistently supports the idea that economic freedom fosters economic growth (Barro, 1996; Carlsson & Lundström, 2002; De Haan & Sturm, 2000). Nevertheless,

[1] See, amongst others, Jenkins (1991) and Silber (2012).

[2] See, amongst others, Berggren (2003); Krieger and Meierrieks (2016); Gwartney et al. (1996); Feldmann (2017); Bengoa and Sanchez-Robles (2003); Graeff and Mehlkop (2003); Gehring (2013); Dreher et al. (2012).

most studies that found a 'negative' outcome related to EF assess the relationship between economic freedom and income inequality.

The theoretical foundation of economic freedom and pre-taxes and transfers income inequality is complex and dependent on various factors. For example, lower government support to individuals and households through social welfare policies – such as social benefits, healthcare, and education access – may lead those already disadvantaged to face economic hardship, widening the income gap. Other mechanisms may also impact income inequality. For example, Alvaredo et al. (2013) concluded that there is a strong correlation between decreases in top tax rates and increased pre-tax top income shares.

Economic freedom increases labor/business and financial market flexibility. However, this flexibility may impact income inequality in opposite directions. Tridico (2018) concludes that labor market flexibility and trade unions lost power contributed to increased income inequality. Other theses in the literature assert that lower regulation in the labor market improves wage earnings and job mobility. Financial and business liberalization also have contradictory views. Trade openness is, again, also not without controversy. While some scholars theorize the positive benefits of creating opportunities for greater specialization and productivity gains (higher wages), others claim that globalization and trade liberalization are one of the causes of rising inequalities. In short, theoretical conceptions are diverse and contradictory.

The empirical basis of the freedom-inequality nexus is ambiguous as well. To the best of one's knowledge, this debate first started with the work of Berggren (1999), that found that while increases (changes) in economic freedom reduce income inequality, its levels were related to lower equality. Carter (2007) showed that Berggren (1999) errs in interpreting his results since his model is mathematically equivalent to a distributed lag model and found a U-shaped relationship between economic freedom and income inequality. The marginal effect of economic freedom in inequality turns (from positive to negative) at an Economic Freedom of the World Index (EFW) of 4.028. The author concluded that economic freedom increases inequality in the long run, mainly because it reduces government redistribution, thus documenting the existence of a trade-off.

Clark and Lawson (2008) employed a 2SLS model to estimate the relationship between economic growth, tax policy, and economic freedom on income inequality. Their results suggested that economic freedom increases are associated with income inequality decreases. Bergh and Nilsson (2010) decomposed the EFW index into five major subcomponents and found that freedom to trade internationally and deregulation positively affect income inequality.

To reduce the disparities that arise from the different political structures and countries' institutions, Apergis et al. (2014) investigated the causal relationship between income inequality and economic freedom in the United States between 1981 and 2004, employing a Granger-causal analysis within a panel error correction model. The authors' results showed bidirectional causality both in short and the long run and suggested (i) that high-income inequality may lead the states to increase redistribution policies, therefore reducing economic freedom levels; and (ii) similar to Ashby and Sobel (2008), income inequality is reduced with increases in economic freedom, both in short and in the long-run. Bennett and Vedder (2013) examined the dynamic relationship between 1979 and 2004, employing a fixed-effects regression. The authors found evidence that economic freedom reduces income inequality, dependent on the initial level of EF, suggesting an inverted U relationship. Additionally, the authors evidenced an existing lag between economic freedom changes and income inequality decreases as the former takes time to consider its effects. Pérez-Moreno and Angulo-Guerrero (2016) constructed an unbalanced panel for 28 EU members between 2000 and 2010 and examined the relationship between the decomposed EFW index and income inequality. The results suggested that smaller government sizes and deregulation increase income inequality while access to sound money, legal systems, and property rights have no statistical significance. Also, the link freedom to trade internationally is only negative and significant in the old EU-15 countries.

Turning to more recent international studies, Sturm and De Haan (2015) used data on inequality from the Standardized World Income Inequality Database (v2) to construct an unbalanced panel of 108 countries between 1971 and 2010, split into eight five-year intervals. The authors' results suggested no robust relationship between economic freedom and income inequality. Apergis and Cooray (2017) employed linear and non-linear co-integration analyses with an unbalanced panel of 138 countries. The authors' results on the linear baseline highlighted a negative relationship between the overall EFW index and the five EFW subindexes and income inequality. In the non-linear approach, the authors employed a Panel Smooth Transition Regression. They concluded that above the threshold overall EFW index value of 5.428, the effect of economic freedom on income inequality is negative. Similarly, above the threshold values of 5.236, 4.435, 3.873, 4.908, 5.801 (size of the government, legal system, and property rights, sound money, freedom to trade internationally, regulation, respectively), the impact of economic freedom on income inequality is negative, supporting the idea of an inverted U-shaped relationship.

The extensive work of Bennett and Nikolaev (2017) showed how previous works addressing this topic are sensitive to the country and time sample as well to the inequality measure used by reproducing the previous studies of Bergh and Nilsson (2010) and Carter (2007) using six different inequality measures for 112 countries between 1970 and 2010. Moreover, the authors employed a dynamic system GMM to an unbalanced panel of 91 countries and inequality data extracted from the Standardized World Income Inequality Database (v5). Their results again supported the idea that the relationship between economic freedom and income inequality is extremely sensitive to the inequality measure used. In some specifications, the decomposed subcomponents of the IEF index are significant, while in others, they are not.

De Haan and Sturm (2017) used a dynamic fixed-effects model to examine 121 countries from 1975 to 2005. They found that financial liberalization can increase income inequality, which may be affected by financial development and the quality of political institutions. Specifically, the increase in the Gini coefficient was found to be higher in countries with higher-quality political institutions. Kwon (2018) used a two-stage least-squares regression to analyze 20 advanced industrial countries from 1988 to 2009 and found that liberalization can indirectly increase income inequality by increasing financial activity in these advanced industrial economies. Graafland and Lous (2018) used a panel of 21 OECD countries between 1990 and 2014. They concluded that the subcomponents of economic freedom – fiscal freedom, freedom to trade internationally, and deregulation – decrease income equality, whereas access to sound money decreases income inequality. Finally, Kwon (2019) employed an unbalanced panel and fixed effects Prais-Winsten regressions

Table 1. List of countries.

Albania	Cote d'Ivoire	Honduras	Luxembourg	Peru	Thailand
Angola	Croatia	Hong Kong	Malaysia	Philippines	Tunisia
Argentina	Cyprus	Hungary	Malta	Poland	Turkey
Armenia	Czech Republic	Iceland	Mauritius	Portugal	Ukraine
Australia	Denmark	India	Mexico	Romania	United Kingdom
Austria	Dominican Republic	Indonesia	Moldova	Russia	United States
Bangladesh	Ecuador	Iran	Mongolia	Rwanda	Uruguay
Belgium	Egypt	Ireland	Morocco	Serbia	Venezuela
Bolivia	El Salvador	Israel	Mozambique	Sierra Leone	Vietnam
Botswana	Estonia	Italy	Namibia	Singapore	Zambia
Brazil	Fiji	Jamaica	Netherlands	Slovakia	Zimbabwe
Bulgaria	Finland	Japan	New Zealand	Slovenia	
Cameroon	France	Jordan	Nicaragua	South Africa	
Canada	Gabon	Kazakhstan	Niger	South Korea	
Chile	Germany	Kenya	Norway	Spain	
China	Ghana	Latvia	Pakistan	Sweden	
Colombia	Greece	Lesotho	Panama	Switzerland	
Costa Rica	Guatemala	Lithuania	Paraguay	Tanzania	

to study 14 advanced economies from 1995 to 2010 and found that financialization positively affects income inequality, which becomes more significant at higher levels of financial liberalization.

More recently, De Soysa and Vadlamannati (2021) used a panel database of 128 countries from 1990 to 2017 and two instrumental variable estimations to show that economic freedom increases the Gini Index, although the impact is relatively small. Melki (2022) used panel country data and robust standard errors to find that investment is negatively associated with inequality for shallow levels of property rights, but this relationship disappears as property rights improve. Karakotsios et al. (2020) used a pooled mean-group estimation method on a panel of 58 countries between 1995 and 2016 to study the causal relationship between income inequality and economic freedom and found a positive trade-off. Finally, Saccone (2021) analyzed an unbalanced panel of 76 developed and developing countries between 1980 and 2014 and found that higher economic freedom levels are associated with lower income shares of the bottom 80% while increasing the income shares of the top 10% and 5%. Lawson and Dean (2021) revisited Saccone's work, studied income decile levels (instead of shares), and employed a panel of 75 countries for the same period. The authors then refuted Saccone's results by concluding that economic freedom corresponds with higher incomes for all income decile levels (in absolute).

3. Data and Methodology

The data is available for 102 countries from 2000 to 2018. The chosen countries follow the principle of maximum information. The selected period was chosen based on complete data availability and variable transformations' absence.[3] Table 1 summarizes the country sample.

Since the number of cross series is superior to the analysis period, this study computed a micro panel. Therefore, *Stata 17* assumes a balanced panel at first. The dependent variable is the Household Equivalized Market Gini Index (pre-taxes and pre-transfers) from

[3] Once, until 2000, the EFW index was available only in a 5-year time period.

the latest Standardized World Income Inequality Database (SWIID) version 9.1, released in May 2021 following Solt (2020). The Gini Index equals the Gini Coefficient times 100 and varies from 0 (perfect income equality) to 100 (perfect income inequality), measuring relative inequality. Solt (2020) defined the Gini Index as 'the average difference in income between all pairs in a population, divided by twice the average income in the population' (p. 2). Following this definition, a higher Gini Index indicates that high-income individuals receive a significant portion of the income distribution. Therefore, an increase in the Gini Index is equivalent to an increase in income inequality. Also, note that the dependent variable is on a pre-tax and transfers basis once this study tries to address the economic system before any government intervention (i.e. market creation) and economic freedom subcomponents already incorporate measures of financial assistance such as tax and transfers and subsidies.

The data from SWIID was chosen due to its high coverage, homogeneous comparability, and time availability. Nevertheless, there is often a trade-off between data coverage and quality. SWIID's latest versions recognize this and incorporate the underlying uncertainty into the estimated Gini index parameters.

The primary independent variable for Model I will be the Economic Freedom of the World (EFW) Index 2020 version, first developed by Gwartney et al. (1996) and updated subsequently. The EFW index was constructed to measure a country's economic freedom level, i.e. the degree to which countries' institutions and policies protect individuals and their properties from others' hostility. EFW index is placed on a zero-to-ten scale, having higher scores in countries whose institutions provide infrastructures for private ownership and voluntary exchange (higher economic freedom). The overall EFW index is a composite indicator that averages five subcomponents also placed on a zero-to-ten scale (calculated from around 42 distinct variables collected and harmonized from different sources).[4] Data on the EFW index was collected from Fraser Institute.

Regarding Model II, the main study variables will be the decomposed EFW index into its five major areas from the same database. These areas are, as mentioned above, also placed on a zero-to-ten scale and follow the same reasoning. To comprehensively understand the potential impacts of income inequality, it is important to examine the subcomponents of the major areas in detail. Therefore, the five major areas and their subcomponents are summarized in Table 2.

This study includes the variables commonly used in the freedom-inequality literature as control variables and other controls based on economic rationality.

The unemployment rate (the share of the labor force without work but available and actively seeking employment) was used as a control variable for short-term variations in income distribution. Data were obtained from the World Bank Database and modeled by the International Labour Organization (ILO), which provides comparable international labor statistics estimates. For example, Mocan (1999) found 'that an increase in structural unemployment increases the income share of the highest quintile and decreases the share of the bottom sixty percent of the population' (p.132). Likewise, Jäntti (1994) found evidence that unemployment is associated with regressive effects on income inequality.

The Real Gross Domestic Product per capita at chained Purchase Parity Power (2017 US dollars) from Penn World Table was added as a proxy for real economic growth. Famously,

[4] Full methodology, data collection and indicators review can be accessed in Gwartney et al. (2020, pp. 213–225).

Table 2. Decomposition of the five major subcomponents of economic freedom.

Area 1: Size of the Government
A. Government Consumption as a percentage of total consumption
B. Transfers and subsidies as a percentage of GDP
C. Government enterprise and investment as a percentage of GDP
D. Top Marginal Tax Rate
E. State ownership of assets
Area 2: Legal System and Property Rights
A. Judicial independence: no interference by the government or parties in disputes.
B. Impartial courts: a trusted legal framework for private businesses
C. Protection of Property Rights
D. Military interference in the rule of law and politics
E. Legal system's integrity
F. Legal Enforcement of contracts
G. Regulatory restrictions on the sale of real property
H. Reliability of Police
I. Gender Legal Rights Adjustment
Area 3: Sound Money
A. Money Growth (Average annual growth of the money supply in the last five years minus average annual growth of real GDP in the last ten years)
B. Standard Deviation of Inflation in the last five years
C. Annual Inflation in the most recent year
D. Freedom to own foreign currency bank
Area 4. Freedom to Trade Internationally
A. Tariffs
B. Regulatory trade barriers
C. Difference between the official exchange rate and the black-market rate
D. Controls of the movement of capital and people
Area 5: Regulation
A. Credit Market Regulation
B. Labour Market Regulation
C. Business Regulations

Source: simplified from Gwartney et al. (2020).

Kuznets (1955) theorizes that inequality increases in the first stages of economic development and then declines as this development proceeds. This framework assumes an inverted-U relationship between economic growth and income inequality that was not followed in this study based (i) on the argument of Roine and Waldenström (2015) regarding the mismatch between the long-run trends in inequality and the Kuznets relationship and (ii) the argument of Piketty (2015) who refuted Kuznets with data for more than one century. Deininger and Squire (1997) found a systematic relation between economic growth and increases in income of the poorest quintile. The Share of the Working Population in the Industry Sector was used to control the population's demographic structure. The data was extracted from World Bank, also modeled by ILO. Gustafsson and Johansson (1999) concluded that the economy's composition impacts income distribution, and increases in the industrial sector promote equality. More industrialized countries show less income inequality measured by any indicator.

Last, to control for other economic and non-economic factors that may influence income inequality, one included the Human Development Index (HDI) from United Nations Development Program. The HDI is a multidimensional construction based on life expectancy, education, and living standards, varying between 0 and 1. It is expected to reduce inequality by increasing societies' development. Table 3 summarizes our variables.

We can opt for the ARDL estimator or the Arellano and Bond (1991) estimator to perform the empirical analysis. The Arellano and Bond (1991) estimator was designed for

Table 3. Variables, acronym, and source.

Variables	Code	Database
Household Market Gini Index	*gini_mkt*	Standardized World Income Inequality Database
Economic Freedom of the World Index	*efw*	Fraser Institute Database
– Size of the Government	*govsize*	Fraser Institute Database
– Legal System and Property Rights	*lpr*	Fraser Institute Database
– Sound Money	*money*	Fraser Institute Database
– Freedom to Trade Internationally	*trade*	Fraser Institute Database
– Regulation	*reg*	Fraser Institute Database
Unemployment Rate	*unrate*	World Bank Database
Share of the Population in Industry	*shareindustry*	World Bank Database
Real GDP per Capita at chained PPPs (in mil. 2017 US $)	*rgdp_pc*	Penn World Table
Human Development Index	*hdi*	United Nations – Human Development Data Center

datasets with many panels and few periods, requiring no autocorrelation in the idiosyncratic errors. Although what can be considered a long period is not apparent, we can infer that it will be around ten periods of time. More often than not, it is the moment from which it is recommended to carry out unit root tests on longitudinal data. We think 19 years put our research out of range for the Arellano and Bond estimator.

Moreover, the Autoregressive Distributed Lag (ARDL) model, introduced by Pesaran and Smith (1995), is a time series econometric model that can be used to estimate the long-run and short-run dynamics of a dependent variable based on its own lagged values and the lagged values of one or more independent variables. On the other hand, the Arellano and Bond (1991) approach is a dynamic panel data estimator that can be used to control for unobserved individual-level heterogeneity in longitudinal data. Furthermore, the advantages of using an ARDL model relative to the Arellano and Bond approach are: (i) ARDL models are more flexible than Arellano and Bond's (1991) models regarding the types of variables that can be included. ARDL models can include both stationary and non-stationary variables, while the Arellano and Bond model requires all variables to be stationary; (ii) ARDL models are simpler to estimate and interpret than Arellano and Bond models, which require a more complex estimation procedure and assumptions about the distribution of individual-level effects; (iii) ARDL models are robust to serial correlation in the error term, while Arellano and Bond's models are not (this means that ARDL models can provide more accurate estimates of the long-run and short-run dynamics of the dependent variable when there is serial correlation in the error term); ARDL models are more appropriate for non-stationary variables, which are common in economic time series data (Arellano and Bond's models assume that all variables are stationary, which may not be the case in many empirical applications). Another point that inclined us toward the ARDL model is that it is easy to control for the presence of cross-sectional dependence using the estimator of Driscoll and Kraay (1998) with an ARDL specification.

Methodologically, this study modeled a dynamic error correction through a linear transformation of an ARDL once it allows the decomposition of both short and long-run impacts on income inequality. Additionally, as argued in Nkoro and Uko (2016), this specification (i) allows series to be I(0), I(1), or on the borderline between them; (ii) avoids spurious regression problems due to non-stationary variables; (iii) is robust to endogeneity, that is, a correlation between explanatory variables and the error term, once all variables are assumed to be endogenous (Pesaran & Shin, 1999). More specifically, this model is a transformation of autoregressive distributed lag, known as UECM-ARDL (unrestricted error correction mechanism of an autoregressive distributed model), proposed by Pesaran et al. (2001).

This model estimates in a single equation the short-run parameters (variables in first differences), the parameter of the cointegrating vector (variables in levels lagged once), and the ECM (error correction mechanism, also called speed of adjustment) coefficient (dependent variable in levels lagged once). However, the long-run parameters cannot be directly interpreted from the estimated parameters and must be computed. It was done by computing the ratio between the variables, lagged once, by the parameter of ECM, and multiplying it by – 1.

Variables are transformed into natural logarithms and denoted with 'l.' The first differences are denoted with 'Δ.' There were two reasons for using index logarithms in this investigation. The first one was computational. This transformation has several benefits: (i) taking logarithms can help smooth out the volatility of the index, making it easier to identify trends and patterns in the data; (ii) changes in the logarithms index can be interpreted as percentage changes in the original index (this can make it easier to compare the magnitude of changes across different periods or indices); and (iii) taking logarithms can linearize relationships between variables, making them easier to estimate and interpret in regression analysis. The second one was theoretical and in line with the literature. Taking the natural logarithm of an index is a common transformation in economics and finance research, especially when the index is subject to high volatility or contains large values. It is also important to note that taking the logarithm of an index does not change the underlying data or the interpretation of the indices. Instead, it transforms the values into a different scale that can be easier to work with in some contexts. For example, it is well known that economic agents do not behave linearly regarding increases (e.g. saturation, risk aversion, asymmetries, thresholds, etc.).

Equations (1) and (2) show ARDL (1,1) of models (1) and (2), and equations (3) and (4) show the re-parametrized relationships[5], respectively:

$$\begin{aligned} lgini\_mkt_{it} = {} & \alpha_i + \beta_1 lgini_{mkt_{it}-1} + \beta_2 lefw_{it} + \beta_3 lefw_{it-1} + \beta_4 lrgdp\_pc_{it} \\ & + \beta_5 lrgdp\_pc_{it-1} + \beta_6 lunrate_{it} + \beta_7 lunrate_{it-1} + \beta_8 lshareindustry_{it} \\ & + \beta_9 lshareindustry_{it-1} + \beta_{10} lhdi_{it} + \beta_{11} lhdi_{it-1} + \epsilon_{it} \end{aligned} \tag{1}$$

$$\begin{aligned} lgini\_mkt_{it} = {} & \alpha_i + \beta_1 lgini_{mkt_{it}-1} + \beta_2 lgovsize_{it} + \beta_3 lgovsize_{it-1} + \beta_4 ltrade_{it} \\ & + \beta_5 ltrade_{it-1} + \beta_6 lmoney_{it} + \beta_7 lmoney_{it-1} + \beta_8 lreg_{it} + \beta_9 lreg_{it-1} \\ & + \beta_{10} llpr_{it} + \beta_{11} llpr_{it-1} + \beta_{12} lrgdp\_pc_{it} + \beta_{13} lrgdp\_pc_{it-1} \\ & + \beta_{14} lunrate_{it} + \beta_{15} lunrate_{it-1} + \beta_{16} lshareindustry_{it} \\ & + \beta_{17} lshareindustry_{it-1} + \beta_{18} lhdi_{it} + \beta_{19} lhdi_{it-1} + \epsilon_{it} \end{aligned} \tag{2}$$

$$\begin{aligned} \Delta lgini\_mkt_{it} = {} & \alpha_i + \beta_1 \Delta lefw_{it} + \beta_2 \Delta lrgdp_{pc_{it}} + \beta_3 \Delta lunrate_{it} + \beta_4 \Delta lshareindustry_{it} \\ & + \beta_5 \Delta lhdi_{it} + \varphi_1 lgini_{mkt_{it-1}} + \gamma_1 lefw_{it-1} + \gamma_2 lrgdp_{pc_{it-1}} \\ & + \gamma_3 lunrate_{it-1} + \gamma_4 lshareindustry_{it-1} + \gamma_5 lhdi_{it} + \epsilon_{it} \end{aligned} \tag{3}$$

[5] See Best (2008) for the mathematical error correction re-parametrization.

Table 4. Descriptive statistics (time-period and last sample year) and cross-sectional dependence.

	Descriptive Statistics					Cross-Sectional Dependence		
Variable	Obs	Mean	SD	Min	Max	CD test	Corr	Abs (corr)
gini mkt	1,858	46.92783	6.67551	21.9	72.4	3.91***	0.014	0.738
Efw	1,900	7.063873	0.9495357	2.66883	9.094872	48.78***	0.167	0.461
Govsize	1,900	6.821252	1.112137	3.984469	9.443358	20.8***	0.07	0.347
Lpr	1,900	5.782959	1.622162	2.227683	8.998176	48.25***	0.164	0.426
Money	1,899	8.325259	1.524233	0	9.922187	30.87***	0.105	0.364
Trade	1,900	7.340604	1.275999	1.832772	9.761062	12.12***	0.04	0.441
Reg	1,900	7.050736	1.04348	2.497275	9.429423	86.28***	0.282	0.429
rgdp pc	1,919	20887.21	18559.68	378.086	111703.6	242.89***	0.798	0.815
Unrate	1,919	7.881042	5.454019	0.21	35.27	29.7***	0.097	0.417
Shareindustry	1,919	21.316	7.143942	2.55	40.53	20.55***	0.065	0.61
Hdi	1,938	0.7371785	0.1483754	0.262	0.956	286.67***	0.942	0.944
Last sample year								
gini mkt	71	45.55775	5.802651	21.9	57.4			
Efw	102	7.207367	0.9154938	3.3086	9.030783			
Govsize	102	6.775397	1.098307	4.378478	9.423184			
Lpr	102	5.861082	1.534578	2.336457	8.681434			
Money	102	8.635634	1.417271	0.6921011	9.869045			
Trade	102	7.469451	1.211039	2.919964	9.491209			
Reg	102	7.295271	1.055549	2.551541	9.429423			
rgdp pc	101	25942.04	21982.45	378.086	111703.6			
Unrate	101	6.824257	5.052569	0.47	26.91			
Shareindustry	101	20.78624	6.595335	6.19	37.5			
Hdi	102	0.7795588	0.1341658	0.391	0.956			

Notes: *** denotes statistical significance at the 1% level; Stata command *sum* and *xtcd* were used.

$$\begin{aligned}\Delta lgini\_mkt_{it} = {} & \alpha_i + \beta_1 \Delta lgovsize_{it} + \beta_2 \Delta ltrade_{it} + \beta_3 \Delta lmoney_{it} + \beta_4 \Delta lreg_{it} \\ & + \beta_5 \Delta llpr_{it} + \beta_6 \Delta lrgdp\_pc_{it} + \beta_7 \Delta lunrate_{it} + \beta_8 \Delta lshareindustry_{it} \\ & + \beta_9 \Delta lhdi_{it} + \varphi_2 lgini_{mkt_{it-1}} + \gamma_1 lgovsize_{it-1} + \gamma_2 ltrade_{it-1} \\ & + \gamma_3 lmoney_{it-1} + \gamma_4 lreg_{it-1} + \gamma_{4i5} llpr_{it-1} + \gamma_6 lrgdp\_pc_{it-1} \\ & + \gamma_7 lunrate_{it-1} + \gamma_8 lshareindustry_{it-1} + \gamma_9 lhdi_{it-1} + \epsilon_{it} \end{aligned} \tag{4}$$

From this point forward, equations (3) and (4) will refer to models (1) and model (2), respectively. The α_i express the constant (intercept), β_j and γ_i with $j = 1, \ldots, 19$ and $i = 1, \ldots, 9$ represent the estimates and φ_i with $i = 1, 2$ denotes the speed of adjustment of both models.

To first inspect our variables, we computed the summary statistics and performed Pesaran's CD test following Pesaran (2004), under the null of cross-sectional independence, to test for the presence of cross-sectional dependence (CSD) amongst our series. CSD analysis is crucial when working with panel data, especially when there are large cross-series and a short time. Ignoring CSD has consequences on the first-order properties of panel estimators (Sarafidis & Wansbeek, 2012). In short, it occurs when units in the same cross-section are correlated or, in other words, when the error term is not independent and identically distributed across time and cross-series. The descriptive statistics and the CD-test are presented in Table 4. The results show evidence of CSD for all variables. Therefore, further tests and estimation techniques need to account for it.

To avoid spurious regressions due to multicollinearity, one used the Variance Inflation Factor (VIF) following Belsley et al. (1980) and examined the correlation matrix of both models. Results are shown in Tables 5 and 6, respectively. Although *lrgdp_pc* and *lhdi* (in levels) have relatively high values, which is expectable since human development is associated with economic development, none is superior to 10. Therefore, we assume no multicollinearity problems as the rule of thumb is fulfilled (VIF less than 10).[6]

Regarding our correlation matrix for Models (1) and (2), the only high correlation observed is, again and expected, between *lrgdp_pc* and *lhdi*. Note that both variables were employed because they were used as different proxies to control different aspects. As the goal is to analyze the variable of interest, the authors assume the potentially biased estimates of these variables.

Since that in the presence of cross-sectional dependence, first-generation unit root tests are no longer reliable (Westerlund et al., 2016), the cross-sectionally augmented Im, Pesaran, and Shin (CIPS) test was used as argued in Pesaran (2007), under the null of nonstationary. Regarding both models, in levels, only the aggregate and decomposed indicators of economic freedom (without the time trend) seem to be stationary since the null is rejected, ruling out an ECM based on cointegration. In their first differences, all our variables are stationary. Results are shown in Table 7. As we are in the presence of the I(0) and I(1) series, one is in condition to employ the ARDL methodology. Note that nothing in the CIPS test suggests using a time trend better suits our model.

After carrying out the analysis of our variables, one now turns to panel data estimation techniques. First, following Breusch and Pagan (1980), one employed the Breusch and Pagan Lagrange-Multiplier test – under the null that the variance across entities is zero $[var(u_i) = 0]$ – to test if random effects are preferable to pooled OLS. By rejecting the null, one concludes that the Pooled OLS is not appropriate as results indicate that, in fact, panel effects exist. Therefore, Hausman's specification Test (Hausman, 1978) between fixed and random effects estimations was employed to choose the appropriate estimator. Hausman's specification test tests two different estimators (consistent *vis-à-vis* efficient) under the null that there are no systematic differences. In short, rejecting the null favors fixed – rather than random-effects estimation. Results can be seen for both tests in Table 8. As the null is rejected, one concludes that the within-estimator better suits our model.

The modified Wald test was employed to test for group-wise heteroscedasticity with the fixed-effects model, following Greene (2000) and under the null of homoscedasticity. One computed the Pesaran test under the null of no contemporaneous correlation to confirm the presence of contemporaneous correlation. Finally, the Wooldridge test was used to check the serial correlation[7] presented in Drukker (2003) and derived from Wooldridge (2010, p. 176) under the null of no first-order serial correlation. Results can be seen in Table 9.

The previous tests corroborate the presence of heteroscedasticity, cross-sectional dependence, and serial correlation. Hence, the author chose the Driscoll and Kraay estimator following Hoechle (2007) and derived from Driscoll and Kraay (1998) to produce robust standard errors for coefficients estimated by fixed-effects where the error structure is

[6] This rule of thumb is explicitly analyzed in O'brien (2007). The author concludes that even with VIF values exceeding the rule of thumb of 10 (and mean VIF of 4), one can confidently derive conclusions, since the model does not suffer from multicollinearity. Although Model (1) presents a mean VIF of 4.25, one still follows the conclusions of O'brien (2007) since VIF values are inferior to 10.

[7] Serial correlation (or autocorrelation) occurs when observations of the error term are correlated with each other. Note that $\varepsilon_{it} = \rho * \varepsilon_{it-1} + \mu_{it} where - 1 < \rho < 1$. If $\rho \neq 0$ there is presence of first-order serial correlation.

Table 5. Correlation matrix for model (1) and model (2).

	lgini_mkt	lefw	lrgdp_pc	lunrate	lshareindustry	lhdi
lgini_mkt	1.0000					
lefw	0.1561	1.0000				
lrgdp_pc	0.0114	0.6321	1.0000			
lunrate	0.2949	0.0433	0.1884	1.0000		
lshareindustry	−0.1454	0.3053	0.6077	0.2921	1.0000	
lhdi	−0.0801	0.6257	0.9256	0.2303	0.6943	1.0000

	Δgini_mkt	Δlefw	Δlrgdp_pc	Δlunrate	Δlshareindustry	Δlhdi
Δgini_mkt	1.0000					
Δlefw	0.0050	1.0000				
Δlrgdp_pc	−0.0661	0.0447	1.0000			
Δlunrate	0.1771	−0.0305	−0.1847	1.0000		
Δlshareindustry	−0.0871	0.0189	0.2155	−0.1845	1.0000	
Δlhdi	0.0006	0.0664	−0.3007	−0.0748	0.2414	1.0000

	lgini_mkt	lgovsize	lreg	llpr	lmoney	ltrade	lrgdp_pc	lunrate	lshareindustry	lhdi
lgini_mkt	1.0000									
lgovsize	−0.1053	1.0000								
Lreg	0.1097	0.0301	1.0000							
Llpr	0.1723	−0.2557	0.6942	1.0000						
lmoney	0.1762	0.0248	0.5856	0.5306	1.0000					
ltrade	0.1596	0.0277	0.6368	0.6863	0.6761	1.0000				
lrgdp_pc	0.0114	−0.3004	0.5427	0.7671	0.4485	0.6040	1.0000			
lunrate	0.2949	−0.1574	0.0207	0.1017	0.0366	0.0366	0.1884	1.0000		
lshareindustry	−0.1454	−0.0984	0.1869	0.3399	0.1962	0.1962	0.6077	0.2921	1.0000	
Lhdi	−0.0801	−0.2130	0.5079	0.7189	0.4119	0.4119	0.9256	0.2303	0.6943	1.0000

	Δl gini_mkt	Δl govsize	Δl reg	Δl lpr	Δl money	Δl ltrade	Δl rgdp_pc	Δl unrate	Δlshareindustry	Δl hdi
Δlgini_mk	1.0000									
Δlgovsize	−0.0107	1.0000								
Δlreg	−0.0146	0.6727	1.0000							
Δllpr	−0.0010	0.6356	0.6656	1.0000						
Δlmoney	0.0198	0.1201	0.2007	0.1658	1.0000					
Δltrade	0.0131	0.5035	0.4945	0.4980	0.1078	1.0000				
Δlrgdp_pc	−0.0661	0.0241	0.0455	0.0390	0.1104	−0.006	1.0000			
Δlunrate	0.1771	−0.0761	−0.0224	−0.019	0.0071	−0.006	−0.185	1.000		
Δlshare industry	−0.0871	0.0250	0.0186	0.0371	0.0000	0.0142	0.2155	−0.185	1.0000	
Δlhdi	0.0006	0.0229	0.0348	0.0716	0.1007	0.0261	0.3007	−0.075	0.2414	1.0000

Notes: The post-estimation Stata command *pwcorr* was used (after *reg*).

Table 6. Variance inflation factor.

Model (1)			Model (2)		
Variables	VIF	1/VIF	Variables	VIF	1/VIF
lhdi	8.961	0.112	lhdi	9.528	0.105
lrgdp pc	7.408	0.135	lrgdp_pc	9.383	0.107
lshareindustry	2.031	0.492	llpr	4.137	0.242
lefw	1.767	0.566	ltrade	3.126	0.32
lunrate	1.103	0.907	lreg	2.452	0.408
Mean VIF	4.254		lshareindustry	2.143	0.467
Δrgdp pc	1.157	0.865	lmoney	2.082	0.48
Δlhdi	1.146	0.872	lgovsize	1.362	0.734
Δlshareindustry	1.115	0.897	lunrate	1.122	0.891
Δlunrate	1.066	0.938	**Mean VIF**	3.926	
Δlefw	1.005	0.995	Δlreg	2.439	0.41
Mean VIF	1.098		Δlgovsize	2.421	0.413
			Δllpr	2.286	0.437
			Δltrade	1.557	0.642
			Δlrgdp pc	1.168	0.856
			Δlhdi	1.153	0.867
			Δlshareindustry	1.117	0.895
			Δlunrate	1.078	0.928
			Δlmoney	1.06	0.943
			Mean VIF	1.587	

Notes: Postestimation Stata command *vif* was used (after *reg*).

assumed to be cross-sectionally dependent, autocorrelated and heteroscedastic. This estimator is also used when handling the above conditions by, amongst others, Fuinhas et al. (2017) and Marques et al. (2018).

Finally, after verifying that the error structure is autocorrelated, specification tests (ResetL and ResetS) were employed for both Model (1) and (2) for omitted variables and model specification, following DeBenedictis and Giles (1998) and DeBenedictis and Giles (1999). The null hypothesis of both tests implies that the model is correctly specified (from omitted variables and overall specification). Both tests suggest that the model presented in this paper is correctly specified, and the estimation does not suffer from omitted variable problems. The results can be seen in Table 10. These previous results give confidence to the estimations presented below. To the best of one's knowledge, no empirical study regarding this topic presented any specification test applied to their econometric models.

4. Results

The first estimation with Driscoll-Kray standard errors can be seen in Table 11. As explained by Asteriou and Hall (2011), 'incorporating additional coefficients will necessarily increase the fit of the regression equation (that is, the value of the R^2 will increase), but the cost will be a reduction of the degrees of freedom.' (p. 276). Accordingly, and following Hendry (1995) general-to-specific modeling approach, we excluded from the specifications – of both model (1) and (2) – controls that present no statistical significance at a 10% level. This criterion to achieve a parsimonious model was confirmed recurring to information criterion decision (AIC and SBIC) and, as a robust check, a Wald test of linear hypotheses about the parameters to exclude. Therefore, the variable $\Delta lhdi_{it}$ was excluded in both models.

Table 7. CIPS unit root test for both models.

	Model (1)	
	Without Trend	With Trend
lgini_mkt	10.121	8.156
lefw	−5.726***	−1.938**
lrgdp_pc	8.845	8.096
lunrate	4.493	6.773
lshareindustry	6.589	3.540
lhdi	7.001	3.128
Δlgini_mkt	−5.974***	−4.957***
Δlefw	−21.820***	−16.342***
Δlrgdp_pc	−10.326***	−7.343***
Δlunrate	−12.308***	−9.100***
Δlshareindustry	−14.396***	−11.460***
ΔlHDI	−16.487***	−12.832***
	Additional variables of Model (2)	
	Without Trend	With Trend
lgovsize	−4.406***	−2.857***
lreg	−5.757***	−3.010***
llpr	−1.972**	0.257
lmoney	−6.332***	−1.954**
ltrade	−1.313***	−0.942
Δlgovsize	−22.657***	−18.975***
Δlreg	−25.278***	−20.469***
Δllpr	−18.441***	−14.288***
Δlmoney	−22.204***	−17.857***
Δltrade	−11.142***	−20.328***

Notes: ***, ** denotes statistical significance at 1% and 5% levels, respectively; Stata command *multipurt* was used.

Table 8. Breusch and Pagan Lagrangian multiplier test for random effects.

Model (1)	Model (2)
1404.20***	1262.90***

Notes: *** denotes statistical significance at the 1% level; post estimation, Stata command *xttest0* was used (after xtreg).

Table 9. Specification tests.

	Model (1)	Model (2)
Modified Wald test for groupwise heteroskedasticity	79614.68***	67184.63***
Pesaran's test of cross-sectional independence	6.094***	6.493***
Wooldridge test for autocorrelation	119.992***	119.556***

Notes: *** denotes statistical significance at the 1% level; post estimation, Stata command *xttest2,* and *xtcsd, pesaran* were used (after *xtreg*); Stata command *xtserial* was employed.

Nevertheless, note that one preserves the main variables of interest (both short and long-run) for deriving conclusions, even with no statistical significance. Additionally, under this idea, the authors removed any variable of interest that does not present statistical significance (at least at a 10% significance level) in both the short and the long-run: $\Delta ltrade_{it}$ and $ltrade_{it-1}$ were excluded, concluding that freedom of trade has not impacted income inequality in the years of study. One reason for this result is the chosen sample (country and

Table 10. DeBenedictis-Giles specification reset test.

DeBenedictis-Giles Specification Reset Test	Model (1)
DeBenedictis-Giles ResetL1 Test = 2.304	*P*-Value > F(2, 1642) (0.1001)
DeBenedictis-Giles ResetL2 Test = 2.150	*P*-Value > F(4, 1640) (0.0724)
DeBenedictis-Giles ResetL3 Test = 1.641	*P*-Value > F(6, 1638) (0.1321)
DeBenedictis-Giles ResetS1 Test = 2.304	*P*-Value > F(2, 1642) (0.1414)
DeBenedictis-Giles ResetS2 Test = 2.150	*P*-Value > F(3, 1641) (0.0775)
DeBenedictis-Giles ResetS3 Test = 1.641	*P*-Value > F(4, 1640) (0.1444)
	Model (2)
Debenedictis-Giles ResetL1 Test = 1.740	*P*-Value > F(2, 1631) (0.1758)
Debenedictis-Giles ResetL2 Test = 1.468	*P*-Value > F(4, 1629) (0.2094)
Debenedictis-Giles ResetL3 Test = 1.270	*P*-Value > F(6, 1627) (0.2682)
Debenedictis-Giles ResetS1 Test = 1.526	*P*-Value > F(2, 1631) (0.2178)
Debenedictis-Giles ResetS2 Test = 1.120	*P*-Value > F(3, 1630) (0.3397)
Debenedictis-Giles ResetS3 Test = 0.845	*P*-Value > F(4, 1629) (0.4967)

Notes: Stata command *resetxt* was used.

Table 11. The first estimation of model (1) and model (2).

Dependent variable: $\Delta lgini_{mkt\,it}$	Model (1)	Model (2)
Constant	0.1559*** (0.0425)	0.1524*** (0.0406)
Δlefw	0.0007 (0.0040)	–
Δlgovsize	–	0.0040 (0.0026)
Δltrade	–	0.0025 (0.0018)
Δlreg	–	−0.0085** (0.0040)
Δllpr	–	0.0020 (0.0027)
Δlmoney	–	0.0011 (0.0020)
ΔlHDI	−0.0101 (0.0149)	−0.0171 (0.0166)
Δlshareindustry	−0.0090** (0.0038)	−0.0091** (0.0038)
Δlrgdp_pc	−0.0050** (0.0020)	−0.0049** (0.0022)
Δlunrate	0.0073*** (0.0012)	0.0071*** (0.0012)
ECM ($lgini\_mkt_{t-1}$)	−0.0397*** (0.0130)	−0.0387*** (0.0124)
$lefw_{t-1}$	0.0077* (0.0038)	–
$lgovsize_{t-1}$	–	0.0061*** (0.0015)
$lreg_{t-1}$	–	−0.0057 (0.0035)
$ltrade_{t-1}$	–	−0.0008 (0.0021)
$llpr_{t-1}$	–	0.0048* (0.0023)
$lmoney_{t-1}$	–	0.0034** (0.0016)
$lrgdp\_pc_{t-1}$	−0.0039*** (0.0009)	−0.0039*** (0.0012)
$lHDI_{t-1}$	−0.0115*** (0.0037)	−0.0112*** (0.0036)
$lunrate_{t-1}$	0.0029*** (0.0007)	0.0030*** (0.0007)
$lshareindustry_{t-1}$	0.0031** (0.0013)	0.0031** (0.0011)
Statistics		
Observations	1704	1699
Within R-squared	0.1319	0.1417
F	F(11,17) = 431.03***	F(19,17) = 9473.59***

Notes: ***, **, * denotes statistical significance at 1%, 5%, and 10% levels, respectively; Driscoll Kray standard errors in parentheses; Stata command xtscd, fe was used.

time) and the unlikely chance of capturing long-term general relationships/equilibriums amongst countries. Still, this result aligns with Savvides (1998), which found no statistical significance between trade liberalization and income inequality.

Therefore, Models (1) and (2) are now specified in Equations (5) and (6):

$$\begin{aligned}\Delta lgini\_mkt_{it} &= \alpha_i + \beta_1 \Delta lefw_{it} + \beta_2 \Delta lrgdp_{pc_{it}} + \beta_3 \Delta lunrate_{it} + \beta_4 \Delta lshareindustry_{it} \\ &\quad + \varphi_1 lgini_{mkt_{it-1}} + \gamma_1 lefw_{it-1} + \gamma_2 lrgdp_{pc_{it-1}} + \gamma_3 lunrate_{it-1} \\ &\quad + \gamma_4 lshareindustry_{it-1} + \gamma_5 lhdi_{it-1} + \epsilon_{it} \end{aligned} \tag{5}$$

Table 12. Estimation of the adjusted model.

Dependent variable: $\Delta lgini_{mkt\,it}$	Model (1)	Model (2)
Constant	0.1556*** (0.0424)	0.1549*** (0.0414)
Δlefw	0.0006 (0.0040)	–
Δlgovsize	–	0.0045 (0.0028)
Δlreg	–	−0.0080** (0.0040)
Δllpr	–	0.0027 (0.0028)
Δlmoney	–	0.0007 (0.0020)
Δlshareindustry	−0.0092** (0.0038)	−0.0095** (0.0039)
Δlrgdp_pc	−0.0052** (0.0019)	−0.0052** (0.0020)
Δlunrate	0.0073*** (0.0012)	0.0072*** (0.0012)
ECM ($lgini\_mkt_{t-1}$)	−0.0395*** (0.0129)	−0.0388*** (0.0123)
$lefw_{t-1}$	0.0076* (0.0037)	–
$lgovsize_{t-1}$	–	0.0057*** (0.0014)
$lreg_{t-1}$	–	−0.0059* (0.0033)
$llpr_{t-1}$	–	0.0046* (0.0024)
$lmoney_{t-1}$	–	0.0031* (0.0018)
$lrgdp\_pc_{t-1}$	−0.00395*** (0.0009)	−0.0040*** (0.0010)
$lhdi_{t-1}$	−0.0111*** (0.0036)	−0.0099*** (0.0034)
$lunrate_{t-1}$	0.0029*** (0.0007)	0.0029*** (0.0007)
$lshareindustry_{t-1}$	0.0031** (0.0013)	0.0029** (0.0011)
Statistics		
Observations	1704	1699
Within R-squared	0.1318	0.1405
F	F(10,17) = 428.74***	F(16,17) = 1441.92***

Notes: ***, **, * denotes statistical significance at 1%, 5%, and 10% levels, respectively; Driscoll Kray standard errors are in parentheses; Stata command *xtscc, fe* was used.

$$\begin{aligned}\Delta lgini_{mkt\,it} = {} & \alpha_i + \beta_1 \Delta lgovsize_{it} + \beta_2 \Delta lmoney_{it} + \beta_3 \Delta lreg_{it} + \beta_4 \Delta llpr_{it} \\ & + \beta_5 \Delta lrgdp_{pc_{it}} + \beta_6 \Delta lunrate_{it} + \beta_7 \Delta lshareindustry_{it} + \varphi_2 lgini_{mkt_{it-1}} \\ & + \gamma_1 lgovsize_{it-1} + \gamma_2 lmoney_{it-1} + \gamma_3 lreg_{it-1} + \gamma_4 llpr_{it-1} + \gamma_5 lrgdp_{pc_{it-1}} \\ & + \gamma_6 lunrate_{it-1} + \gamma_7 lshareindustry_{it-1} + \gamma_8 lhdi_{it-1} + \epsilon_{it} \qquad (6)\end{aligned}$$

Although the former specifications are simpler than (3) and (4), the author should stress that these models could be more parsimonious. Due to omitted variable bias problems empirical formalization becomes difficult as income inequality cannot be modulated with a parsimonious specification. Many factors can influence the dependent variable. This task would be easier if an empirical model were established in economic literature. That, however, is not the case. Hence, one must cope with some loss of the degrees of freedom, especially in Model (2), to perform the current estimations. The estimation of the adjusted model can be seen in Table 12.

In order to check structural breaks and control for outliers, one employed a box plot analysis of the residuals and added dummy variables (0 and 1) following Fuinhas and Marques (2012). Pesaran et al. (2001) supported this procedure once the authors argued that the asymptotic theory of bounds test approach (ECM) is not affected by the inclusion of zero-one dummy variables. Figure 1 shows the box plot analysis:

Figure 1 shows that the model presents outliers problems that disturb the estimation.

Therefore, the dummy approach explained earlier was used after checking which year/country was a cause of concern following the rule of a standard deviation greater than 0.01. Table 13 shows the included dummies and a probable explanation for the events that caused these outliers.

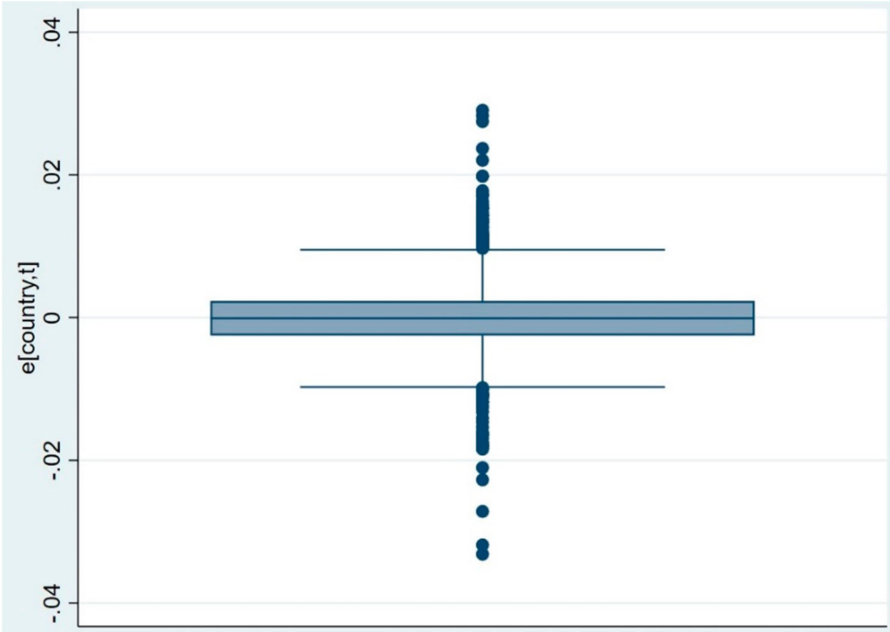

Figure 1. Box plot of the residuals for shock control.

The previous dummies are, thus, added to the adjusted model to control for shocks. Table 14 shows the estimation of both models controlled for shocks and structural breaks:

Although the short-run elasticities are directly derived from the estimations, the long-run elasticities need to be calculated, as mentioned above, by the following equation (7):

$$Long\ run\ computed\ elasticities = -\frac{\gamma_{ji}}{\varphi_i} \tag{7}$$

Subscript denomination follows the above explanation. Table 15 presents the short-run impacts and long-run elasticities for adjusted and controlled shocks in Model (1) and Model (2).

5. Discussion

This section discusses the results shown in Table 15 for the short – and long-run impacts of both the main variables of interest and the control variables on income inequality. Note that the short-run impacts can be interpreted as percentage point increases[8] in the growth rate of the dependent variable, whereas the long-run impacts are elasticities. Additionally, note that given the nature of inequality measures, a positive impact is undesirable once it increases income inequality.

[8] As they are expressed in logarithmic differences.

Table 13. Plausible events that might explain outliers and/or structural breaks, description, and dummies code.

Country	code	Description
Poland	*po2003*	**2003:** June – Referendum voted in favor of joining the European Union.
Poland	*po2004*	**2004:** Poland adheres to the European Union. Prime Minister Miller resigns.
	po2017	**2017:** Anti-Muslim Elk riots; political and military instability also involving the US.
Uruguay	*ur2004*	**2004:** First time a party other than the Colorado Party or National Party had held power since the formation of both (1830).
	ur2011	**2011:** In January 2011, the government issued and sold new bonds in national currency for more than $1.25 billion, placing the local currency share of the public debt at about 40%.
	ur2012	**2011:** In January 2011, the government issued and sold new bonds in national currency for more than $1.25 billion, placing the local currency share of the public debt at about 40%.
Slovenia	*slo2006*	**2006:** The employers revoked the General Collective Agreement for the Private Sector (GCAPS). The Slovenian parliament adopted the Law on Collective Agreements (LCA).
Iceland	*ice2009*	**2009:** Strong public pressure from the 2009 financial crisis anticipated the parliamentary elections.
Spain	*sp2010*	**2010:** The highest unemployment rate in 13 years (over 20%).
	sp2011	**2011:** Parliamentary elections were won by the Conservative Popular Party, which announced new austerity policies.
Paraguay	*pa2012*	**2012:** President Fernando Luga impeached.
Brazil	*br2016*	**2016:** President Dilma Roussef impeached.
Chile	*ch2013*	**2013:** Michelle Bachelet (Socialist candidate) is elected.
All	*id2009*	**2009:** Lehman Brothers triggered an international financial crisis that spread around the world and froze the international monetary system
Other included dummies:		The United Kingdom in 2011; Canada in 2018; Slovakia in 2016; Chile in 2001; Spain in 2001; Paraguay in 2018; Norway in 2002 and 2006.

Adjusted and Controlled Model (1)

From a first look at the overall estimation of Model (1), one concludes that the composite EFW index has no impact in the short run, given the lack of statistical significance. Although the coefficient of $\Delta lefw$ is revealed to be inconclusive (documented as a positive signal on the adjusted model and a negative signal on the controlled model), it seems that the effect of Economic Freedom exerts some time before impacting the economic mechanisms behind income inequality. This outcome underscores the logical relationship by which economic reforms depend on political decisions and institutional systems (therefore, are embedded in the political system). There is (i) a time gap between political decision-making and the associated economic adjustments and (ii) a strict dependence between economic policies and the nature of the political system (Acemoglu et al., 2015).

The control variables present robust results when facing the adjusted model against the controlled model and are consistent with the predicted signs. The real gross domestic product per capita is statistically significant at 5% and 1% (respectively), and all else equal, an increase of 1 percentage point of Δlrgdp_pc decreases income inequality by 0.0052 percentage points. The unemployment rate is statistically significant at 1%, implying that unemployment impacts income distribution with the immediate effect of increasing income inequality. An increase of 1 percentage point in Δlunrate is associated with an increase of approximately 0.007 percentage points (in both specifications) in income

Table 14. Estimation controlled for shocks.

Dependent variable: $\Delta lgini_{mkt\,it}$	Model (1)	Model (2)
Constant	0.1350*** (0.0397)	0.1310*** (0.0393)
Δlefw	−0.0017 (0.0029)	–
Δlgovsize	–	0.0045** (0.0017)
Δlreg	–	−0.0076** (0.0035)
Δllpr	–	0.0016 (0.0030)
Δlmoney	–	−0.0003 (0.0022)
Δlshareindustry	−0.0110*** (0.0031)	−0.0116*** (0.0031)
Δlrgdp_pc	−0.0052*** (0.0016)	−0.0051*** (0.0018)
Δlunrate	0.0070*** (0.0014)	0.0070*** (0.0013)
ECM ($lgini\_mkt_{t-1}$)	−0.0357*** (0.0122)	−0.0353*** (0.0118)
$lefw_{t-1}$	0.0076** (0.0031)	–
$lgovsize_{t-1}$	–	0.0063*** (0.0015)
$lreg_{t-1}$	–	−0.0057** (0.0027)
$llpr_{t-1}$	–	0.0064** (0.0027)
$lmoney_{t-1}$	–	0.0026 (0.0016)
$lrgdp\_pc_{t-1}$	−0.0033*** (0.0009)	−0.0032*** (0.0009)
$lhdi_{t-1}$	−0.0113*** (0.0029)	−0.0112*** (0.0037)
$lunrate_{t-1}$	0.0022*** (0.0005)	0.0022*** (0.0005)
$lshareindustry_{t-1}$	0.0035** (0.0013)	0.0032*** (0.0011)
Dummys	***	***
Statistics		
Observations	1704	1699
Within R-squared	0.3470	0.3570
F	F(33,17) = 3285.68***	F(39,17) = 5274.57***

Notes: ***, ** denotes statistical significance at 1% and 5% levels, respectively; Driscoll Kray standard errors are in parentheses; Stata command *xtscc, fe* was used.

inequality, all else equal. Finally, the share of the population working in the industrial sector, a proxy for countries' industrialization level, is statistically significant at a 1% level and is associated with decreased income inequality in the short run. *Ceteris paribus*, an increase of 1 percentage point in the share of the population working in the industrial sector implies a decrease of 0.009 and 0.011 percentage points (respectively) in income inequality.

When turning to the long-run elasticities of Model (1), results suggest that the overall EFW index is statistically significant at 10% and 5% (adjusted and controlled model, respectively) and, all else equal, an increase of 1% in economic freedom implies an increase of 0.1979% and 0.2132% in income inequality. Although there is evidence of a positive trade-off, these results suggest a relatively inelastic relationship between economic freedom and income inequality, possibly because the economy re-structures after internalizing economic freedom effects, reducing its importance as a primary influencer on income inequality. Within Model (2), the general result is broken down and examined in the context of the impacts that lower government size, [de] regulation, and property rights exert.

The unemployment rate is statistically significant at 5%, and an increase of 1% reveals a positive impact of 0.0742% and 0.0618% (respectively) in the Gini index, increasing income inequality. However, the variable *unrate* is used to control short-run impacts on the income distribution and does not capture structural unemployment but rather the unemployment rate at a given time. Therefore, in line with Martínez et al. (2001), this result suggests that the contribution of the unemployment rate to overall inequality is limited (despite clearly affecting the preponderance of unemployed individuals in the bottom income quintiles, ibid).

Table 15. Short-run impacts and long-run computed elasticities.

Dependent variable: $\Delta lgini_{mkt\,it}$	Model (1) adjusted	Model (1) controlled for shocks	Model (2) adjusted	Model (2) controlled for shocks
Short-Run Impacts				
Δlefw	0.0006	−0.0017	–	–
Δlreg	–	–	−0.0080**	−0.0076**
Δlgovsize	–	–	0.0045	0.0045**
Δllpr	–	–	0.0027	0.0016
Δlmoney	–	–	0.0007	−0.0003
Δlshareindustry	−0.0092**	−0.0110***	−0.0095**	−0.0116***
Δlrgdp_pc	−0.0052**	−0.0052***	−0.0052**	−0.0051***
Δlunrate	0.0073***	0.0070***	0.0072***	0.0070***
Speed of Adjustment				
ECM	−0.0395***	−0.0357***	−0.0388***	−0.0353***
Long-Run Computed Elasticities				
$lefw_{t-1}$	0.1919* (0.1072)	0.2132** (0.0970)	–	–
$lreg_{t-1}$	–	–	−0.1515* (0.0888)	−0.1622* (0.0958)
$lgovsize_{t-1}$	–	–	0.1471** (0.0736)	0.1784** (0.0919)
$llpr_{t-1}$	–	–	0.1175*** (0.0458)	0.1805*** (0.0529)
$lmoney_{t-1}$	–	–	0.0801 (0.0523)	0.0726 (0.0531)
$lunrate_{t-1}$	0.0742** (0.0356)	0.0618** (0.0306)	0.0759** (0.0370)	0.0638** (0.0319)
$lshareindustry_{t-1}$	0.0787*** (0.0727)	0.0974*** (0.0276)	0.0753*** (0.0236)	0.0903*** (0.0307)
$lrgdp\_pc_{t-1}$	−0.0998** (0.0507)	−0.0924* (0.0494)	−0.1024** (0.0515)	−0.0913* (0.0499)
$lhdi_{t-1}$	−0.2800** (0.1287)	−0.3171** (0.1330)	−0.2557** (0.1073)	−0.3180***(0.1252)

Notes: ***,**,* denotes statistical significance at 1%, 5%, and 10% levels, respectively; standard errors are in parentheses; Stata command *nlcom* was used to calculate the long-run elasticities.

Economic growth negatively impacts (decreases) income inequality, which follows the economic literature, although the results also suggest an inelastic relationship. All else equal, an increase of 1% in economic growth decreases income inequality by approximately 0.1%. This result is in line with Gallup et al. (1998). Although they had found a robust positive relationship between economic growth and the income of people experiencing poverty (both in the short – and the long run), they also concluded that, as the overall income grows, the income distribution can remain unchanged (depending on the initial level), only meaning that people experiencing poverty 'do not fall behind.' Nevertheless, this average result also depends on the period chosen and should not be taken for granted.

The Human Development Index is statistically significant at 5%. An increase of 1% in *hdi* reduces the Gini index by 0.28% and 0.3171%, meaning that developing life conditions, education, and living standards is the most important (as policy targets) to reduce inequality. This finding aligns with Cingano's (2014) report that one of the most important policy strategies to reduce inequality relies on increasing human capital and allowing access to income acquisition opportunities.

Perhaps the most contradictive result is the positive impact of *shareindustry* on income inequality (at a 1% significance level), which means that an increase of 1% in the number of people working in the industry increases the Gini by 0.079% and 0.097% (respectively). This result may occur because of the country sample. However, it is a debate that future research should address regarding the causes of this phenomenon. One can advance a possible explanation: technological enhancements have allowed greater added value for the goods produced in the industrial sector (as capital substituted labor), whereas an increase in labor would ultimately lead to a decrease in productivity in industry and conduct to a stagnation (or even a decrease) of wages. Therefore, the income distribution of the industry gains would favor the capital owners over the labor suppliers.

On the other hand, the skill-biased technological change argument may apply here, as expressed in Katz and Autor (1999). Demand shift for higher-skilled sectors (such as services and high-tech) increases the skill and pay gap in industrialized countries. We leave future research lines for this matter.

Adjusted and Controlled Model (2)

Turning one's attention to Model (2), adjusted and controlled results suggest relevant implications. In the short run, only $\Delta lgovsize$ and $\Delta lreg$ (the main variables of interest) seem to impact income inequality, as both are statistically significant at a 5% level. This outcome shows that the effect of a one percentage point increase in *govsize* (smaller government size) will have an immediate impact of increasing income inequality by 0.0045 percentage points, perhaps mainly through lower public enterprise and public investment needed to boost employment and economic development capable of influencing income structure in the short run. Conversely, an increase of 1 percentage point in *reg* decreases income inequality by 0.008 and 0.0076 percentage points (respectively). This outcome may be due to the impact of lower red-tape costs for businesses and an incentive to (or, at least, fewer entrance barriers costs too) entrepreneurship, which impact income distribution at a higher rate than the loss of labor protection, which is associated with decreases in income inequality (Checchi & García-Peñalosa, 2008). Also, more accessible access to the financial system or greater job mobility can be the reason for decreases in income inequality. The control variables in Model (2) approximately follow the same discussion applied to Model (1) above and remain robust to the different specifications.

Focusing on the long-run elasticities of the main variables of interest, one concludes that access to sound money does not impact the Gini index once it presents no statistical significance in both adjusted and controlled Model (2). On a contrary note, $lreg_{t-1}$, $lgovsize_{t-1}$, and $llpr_{t-1}$ are statistically significant at 10%, 5%, and 1% levels, respectively, for both adjusted and controlled Model (2). Moreover, an increase of 1% in *reg* is associated with a decrease in income inequality of 0.1515% and 0.1622% (respectively). This result suggests that the same short-run impacts of *reg* also apply in the long run after the economy accommodates its effects, supporting labor, business, and financial deregulation.

Regarding the variable govsize, all else equal, an increase of 1% in *govsize* increases income inequality by 0.1471% and 0.1784% (respectively). These results suggest the detrimental impact on income inequality exerted by lower government spending and lower government support to further equality of opportunity. The latter is partially another advocate for the skill premium, ultimately affecting income structure through inequality of opportunities.

Another interesting result is the positive impact of *lpr* on the Gini index, where an increase of 1% in *lpr* increases the Gini index by 0.1175% and 0.1805%, respectively. One argues, in line with Carter (2007), that the protection of property rights mostly favors the ones who have more property '(…) as this protection increases tenure security for the owner, which in turn is expected to increase the value of the property itself' (p. 489), as well as protecting the *status quo.* Even though the results align with the part of Katz and Autor's (1999) argument, the authors suggest another line for future research. The above

Table 16. Additional variables for robustness analysis.

Variables	Code	Database
Palma Ratio	*palma*	Global Consumption and Income Program
Theil's Index	*theil*	Global Consumption and Income Program
Atkinson Measure	*atkinson*	Global Consumption and Income Program
Index of Economic Freedom	*ief*	Heritage Foundation

Table 17. Descriptive statistics (robustness analysis).

	Descriptive Statistics				
Variable	Obs	Mean	SD	Min	Max
Ef	1924	63.23763	10.22033	21.4	90.2
palma	1569	3.213457	2.235576	0.8015007	14.43498
theil	1569	0.3694447	0.1975702	0.0869404	1.040622
atkinson	1569	0.4818797	0.1828306	0.1569068	0.887352

Notes: Stata command *sum* and *xtcd* were used, respectively.

results suggest a relatively inelastic relationship between the independent and dependent variables.

Again, the discussion regarding the control variables is the same as Model (1) since they present robust results and approximately the same coefficient value without changing their signs. Finally, the speed of adjustment of both adjusted and controlled Model (1) and Model (2) align with the theoretical requirements for the ECM specification, as they are negative, between − 1 and 0, and statistically significant at a 1% level. The low value of the speed of adjustment supports the idea presented above that the economic adjustments of changes in Economic Freedom take time to exert their effects.

6. Robustness Analysis

This section will present a robustness analysis of Model (1). Note that the previous tests employed in Section 2 are not presented to preserve space. However, all the tests and statistics used before were employed for these variables. The same conclusions hold to employing the ARDL methodology and apply here too. Variables are transformed into natural logarithms and first differences.

The Palma Ratio, Theil's Index, and Atkinson Measure are available from 2000 to 2015 from the Global Consumption and Income Program, producing income and consumption datasets equivalized to more than 160 countries to allow international comparisons. In addition, the Heritage Foundation's Index of Economic Freedom is available from 2000 to 2018. Table 16 summarizes.

The summary statistics of the variables used in this analysis are shown in Table 17. However, it should be noted that the variables employed may not impact all countries by the same measure, which may limit the robustness of the analysis. Attempts were made to overcome these limitations by applying dummies to the country sample to study development levels (low-, middle-, and high-income countries). However, this was impossible due to multicollinearity problems that this study could not surpass. Given the former, the usual caveats apply to the following framework.

Table 18. Estimation controlled for shocks with IEF as the independent variable.

Dependent variable: $\Delta lgini_{mkt\,it}$	Model (4)
Constant	0. 1280*** (0.0401)
Δlief	0.0091** (0.0041)
Δlshareindustry	−0.0096*** (0.0032)
Δlrgdp_pc	−0.0038*** (0.0012)
Δlunrate	0.0072*** (0.0014)
ECM ($lgini\_mkt_{t-1}$)	−0.0382*** (0.0120)
lef_{t-1}	0.0082** (0.0033)
$lrgdp\_pc_{t-1}$	−0.0035*** (0.0008)
$lhdi_{t-1}$	−0.0085*** (0.0022)
$lunrate_{t-1}$	0.0024*** (0.0005)
$lshareindustry_{t-1}$	0.0034*** (0.0011)
Dummys	***
Statistics	
Observations	1727
Within R-squared	0.3474
F	F(33,17) = 51891.53***

Notes: ***, ** denotes statistical significance at 1% and 5% levels, respectively; Driscoll Kray standard errors are in parentheses; Stata command *xtscc, fe* was used.

Table 19. Long-run Elasticities with IEF as the primary independent variable.

Long-Run Computed Elasticities	
$lief_{t-1}$	0.2156** (0.0950)
$lunrate_{t-1}$	0.0628** (0.0286)
$lshareindustry_{t-1}$	0.0886*** (0.0287)
$lrgdp\_pc_{t-1}$	−0.0917** (0.0403)
$lhdi_{t-1}$	−0.2224** (0.0949)

Notes: ***, ** denotes statistical significance at 1% and 5% levels, respectively; standard errors are in parentheses; Stata command *nlcom* was used.

The first robustness analysis uses the IEF index instead of the EFW index. Note that this new model (let it be Model (4)) maintains the exact specification as Model (1) and dummy control as explained above in Section 3 [Equation (5)].

Table 18 presents the estimation with the IEF index as a primary independent variable:

Based again on equation (7), and once the short-run elasticities can be directly derived from the estimation above, Table 19 shows the long-run elasticities with the EIF index as an independent variable:

As noted in Table 19 above, this study's estimation results are strongly robust to changes in the independent variable of interest. The only difference is the positive and statistically significant at a 5% coefficient of *Δlief,* suggesting that economic freedom might increase income inequality in the short run. Nonetheless, these robust results confirm that economic freedom increases income inequality in the long run (measured by both most important indicators of EF available) while suggesting a relatively inelastic relationship. Moreover, the long-run economic freedom elasticities present the same value with both EF indexes, even with different methodological specifications. The other variables are also robust to EF measures changes, meaning that the model is well-constructed and specified.

One now turns to estimate three different regressions using different dependent variables: Palma ratio, Theil's index, and Atkinson measure. Table 20 shows the regression results.

Table 20. Estimation controlled for shocks with different dependent variables.

	Dependent Variables		
	Palma	Theil	Atkinson
Constant	0.3614*** (0.1047)	−0.2414* (0.1332)	−0.2252** (0.0807)
Δlefw	0.0029 (0.0171)	0.0083 (0.0143)	0.0216 (0.0154)
Δlshareindustry	0.0703 (0.0755)	0.0230 (0.0594)	−0.0180 (0.0418)
Δlrgdp_pc	−0.00563** (0.0755)	−0.00512** (0.0239)	−0.0295 (0.0207)
Δlunrate	0.0297** (0.0129)	0.0282** (0.0102)	0.0375*** (0.0097)
ECM (ly_{t-1})	−0.3026*** (0.0505)	−0.3569*** (0.0681)	−0.3429*** (0.0625)
$lefw_{t-1}$	0.1124*** (0.0345)	0.1055*** (0.012)	0.0835*** (0.0218)
$llrgdp\_pc_{t-1}$	−0.0506*** (0.0136)	−0.0541*** (0.0135)	−0.0306*** (0.0066)
$lhdi_{t-1}$	−0.0923** (0.0422)	−0.0233 (0.0275)	−0.0253 (0.0165)
$lunrate_{t-1}$	0.0239*** (0.0058)	0.0234*** (0.0055)	0.0182*** (0.0046)
$lshareindustry_{t-1}$	0.0334** (0.0117)	0.0259** (0.0120)	0.0094 (0.0110)
Dummies	***	***	***
Statistics			
Observations	1421	1421	1421
Within R-squared	0.2034	0.2361	0.2252
F	F(20,14) = 433.54***	F(20,14) = 1160.39***	F(19,14) = 674.68***

Notes: ***, **, * denotes statistical significance at 1%, 5%, and 10% levels, respectively; Driscoll Kray standard errors are in parentheses; Stata command *xtscc, fe* was used.

Table 21. Long-run elasticities for different dependent variables.

Long-Run Elasticities	Palma	Theil	Atkinson
$lefw_{t-1}$	0.3715*** (0.0873)	0.2956*** (0.0709)	0.2436*** (0.0580)
$lunrate_{t-1}$	0.0791*** (0.0196)	0.0656*** (0.0172)	0.0532*** (0.0098)
$lshareindustry_{t-1}$	0.1104*** (0.0408)	0.0726** (0.0360)	0.0274 (0.0335)
$lrgdp\_pc_{t-1}$	−0.1672*** (0.0500)	−0.1515*** (0.0375)	−0.0892*** (0.0335)
$lhdi_{t-1}$	−0.3051*** (0.1008)	−0.0652 (0.0733)	−0.0739* (0.0436)

Notes: ***, **, * denotes statistical significance at 1%, 5%, and 10% levels, respectively; Stata command *nlcom* was used.

Again, the short-run impacts of Economic Freedom on income inequality are directly seen in Table 20 above. However, these models have no statistical relationship in the short run. Hence, although the results could be more conclusive, one still concludes that changes in Economic Freedom take their time to affect the economy. The rest of our variables keep their signs and approximated values, which reveals that the former model is robust to changes in the primary variable of interest and changes in the dependent variable.

Turning to the long-run impacts of economic freedom on income inequality, Table 21 summarizes.

This sensitivity analysis again shows that the long-run relationship between Economic Freedom and income inequality is positive, implying a positive trade-off, measured by any inequality measures employed above but also pointing to an inelastic relationship. Furthermore, the controls used in this paper reveal themselves robustly in the long run, leaving possible lines of research.

7. Conclusion and Policy Implications

This work revisited the relationship between economic freedom and relative income inequality using the most recent databases available for research. This contradictive debate had flaws regarding non-stationarity and endogeneity problems. This study explicitly addressed those problems by employing an error correction model (derived from an autoregressive distributed lag model) robust to these data characteristics.

The results illustrate a positive but inelastic relationship between economic freedom and income inequality in the long run. On the one hand, in line with Cingano (2014), increases in economic freedom seem to drive the concentration of wealth. This result comes mainly (i) because of reduced government support, directly on the income structure and indirectly through fewer income acquisition opportunities and, to a lower extent, (ii) through increasing property value (both tangible and intangible), which directly benefits the ones who own more property and does not exert any redistribution effects. On the other hand, despite the unclear channels through which deregulation impacts income inequality, the evidence shows that lower regulation helps improve income equality. As asserted in the seminal work of Katz and Autor (1999), institutional changes, such as market deregulations, can impact wage structures and consequently influence income inequality. The impact of sound money on income inequality follows the results of Jäntti and Jenkins (2010), which found no statistical significance. In closing, this empirical research dismantles the routes through which the subcomponents of economic freedom aggregate into the general conclusion that economic freedom positively impacts income inequality.

This paper, however, does not advocate that economic freedom should or should not be pursued. The main finding of this paper asserts that the relationships between changes in economic freedom (and its subcomponents) and income inequality (measured by four different indicators) are positive but relatively inelastic. If this empirical assessment adds anything, this economic discussion – in the terms it has been discussed – adds few solutions for dealing with income inequality once the relationships presented above seem somewhat rigid. This macroeconomic discussion could not extrapolate practical policy guidelines when employing these measures if this were not reason enough. Given the inelastic relationship, the authors suggest that future research addresses the impact on the income distribution of lower regulation on labor/business/financial markets by employing a quantile approach to income distribution to address income shares specifically. Also, future research should reimagine the possible interactions between economic freedom (measured by macroeconomic data), income inequality, and (amongst others) economic growth within a multi-equation framework to test if the benefits of economic freedom on the overall economic system surpass the costs of increased income inequality.

As mentioned, despite its detrimental impacts on income inequality, the relationship between economic freedom and income inequality is inelastic, suggesting that it is not the fundamental factor impacting and explaining income inequalities. If otherwise, economic freedom triggers other economic mechanisms that may promote income equality (such as human capital increases, economic development, or investment capacity), and policy guidelines are not as straightforward as entirely rejecting freedom benefits. Nevertheless, this study concludes that policies that aim to increase economic freedom must at least account for the harmful effects on income inequality outputs and acquisition opportunities.

Ethical Approval

The research did not involve human participants and/or animals.

Consent to Participate

All the authors have approved and consented to the manuscript submission.

Consent to Publication

All authors have given their consent for the publication of the manuscript.

Author Contributions

DM: Conceptualisation, Methodology, Writing and original draft preparation, Data curation, Investigation, Formal analysis, Visualisation. JAF: Writing – review and editing, Supervision, Validation, and Investigation.

Disclosure Statement

No potential conflict of interest was reported by the authors.

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Availability of Data and Materials

Data is available on request from the corresponding author.

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