Economic freedom and income inequality: further evidence from 58 countries in the long-run

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Abstract

This study employs panel data for 58 countries from 1980-2010, to investigate the dynamic relationship between economic freedom and income inequality. Both linear and non-linear (Panel Smooth Threshold Regression) cointegration estimation methods are used to identify a long-run equilibrium relationship between the overall economic freedom index and its components, and income inequality. The linear long-run parameter estimates for the entire panel of countries show that the association is negative, while the non-linear long-run parameter estimates indicate that above a threshold point the association between economic freedom and income inequality is negative, while below this threshold point the association is positive.

Keywords: economic freedom, Economic Freedom Index, income distribution

1 INTRODUCTION

Economic freedom is a multifaceted concept which can have differing relations with income distribution at different stages of economic freedom. Kuznets (1955) hypothesized that as economic growth occurs, inequality in the distribution of income may initially increase with structural change, and then decrease in the long run, beyond a certain point. This proposition known as the inverted U-hypothesis, has relevance for the relationship between economic freedom and income distribution. At the early stages of economic freedom, income inequality can increase due to market allocation benefiting high income groups, however, in later stages of economic freedom, as income rises and the standard of living overall improves, income inequality can fall. If the income of low-income groups grows at a faster rate than that of other income groups, then greater economic freedom will lead to greater equality in income distribution. The question this paper attempts to answer is whether economic freedom leads to greater equality in income distribution.

Studies undertaken on freedom and income distribution are sparse. Additionally, there is no consensus in the literature as to the net impact of economic freedom on income distribution. Therefore, the contribution of this study to the literature is threefold. First, we employ both linear and non-linear panel (Panel Smooth Threshold Regression – PSTR) cointegration methods to investigate the relationship between economic freedom and income inequality. The non-linear estimation methodologies provide clear information on the relationship between economic freedom and income inequality changes over different stages of economic freedom. The only studies that account for the presence of thresholds in the economic freedom-income inequality relationship are those by Carter (2006) and Bennett and Vedder (2013).

We advance upon these studies by using the PSTR estimation method which allows the income inequality-economic freedom coefficient to vary not only across countries, but also over time. We also employ the Economic Freedom Index (EFI) of Gwartney et al. (2006) to measure economic freedom, and the Gini coefficient to measure income inequality. Another consideration is that some EFI components could lead to growth and reduce inequality more than other components. However, the current literature investigates the effects of different EFI components on economic growth. For this reason, this paper also investigates the impact of the different components of the EFI (i.e., size, taxation and labour structure) on income inequality. Third, there could be regional heterogenenity in the relationship between the EFI and income inequality. Accordingly, we account for this regional heterogeneity by conducting the analysis not only at the overall level, but also separately for each region.

The paper is structured as follows. Section 2 discusses the literature. Section 3 presents the data including unit root tests. Section 4 presents empirical results for the linear model, and robustness tests across components of the EFI and across regions. Section 5 presents results for the non-linear model and section 6 concludes.

2 THE LITERATURE

Studies which investigate the relationship between economic freedom and income inequality include those by Berggren (1999), Scully (2002), Ashby and Sobel (2008), Carter (2006) and Apergis et al. (2013). Berggren (1999) argues that an increase in economic freedom leads to lower taxes, a relaxation of regulations, and higher economic growth. Scully (2002) examines the role of economic freedom in income distribution for a pooled sample of 26 countries using the EFI of Gwartney et al. (2006). The author concludes that economic freedom leads to greater equality with the existence of a marginal trade-off between growth and income inequality. Employing an unbalanced panel of 126 countries and the EFI, Carter (2006) argues that higher levels of economic freedom can increase income equality by extending income-earning prospects, but also reduce equality by diminishing income redistribution opportunities. He finds that the latter effect outweighs the former except at very low levels of freedom; therefore, implying a trade-off between economic freedom and income equality, with the impact of economic freedom increasing at higher levels of freedom. Clark and Lawson (2008) investigate the role of tax policy in income distribution and they document evidence in favour of increased income equality due to progressive taxation with high top marginal tax rates. Bennett and Vedder (2013) investigate the dynamic relationship between economic freedom and income inequality across the U.S. states. They provide robust evidence showing that increases in economic freedom are associated with lower income inequality. However, they also provide supportive evidence that the relationship depends on the initial level of economic freedom, implying that there may be an inverted U-shaped relationship with the inflection point being explicitly determined.

Ashby and Sobel (2008) investigate the link between economic freedom and income distribution across U.S. states using the Economic Freedom of North America Index (EFNA), as introduced by Karabegovic and McMahon (2005). Their results highlight that increases in economic freedom correspond with lower inequality. Apergis et al. (2013) examine the relationship between income inequality and

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economic freedom across the U.S. Their findings document bi-directional causality between economic freedom and income inequality in both the short- and long- run.

Given the inconclusive results on the relationship between economic freedom and inequality, in the empirical analysis that follows we extend upon the literature by accounting for non-linear (i.e., threshold) effects, and investigate the relationship between economic freedom and inequality across components of the EFI and across regions. Our results demonstrate that the relationship between economic freedom and income inequality is non-linear.

3 DATA

3.1 DATA

The data for this study include annual observations over 1980-2010 for an unbalanced panel of 58 countries which constitutes a representative cross-section of the regions covering Europe, Asia, the Americas (North and South) and the Pacific. The list of countries is provided in the appendix. The unbalanced panel is associated with a number of countries in which data availability prompted for using data over the 1991-2010 time span. These countries included in this particular sample were: Bulgaria, Croatia, Poland, Romania, Slovakia and Slovenia.

The variables include: the **Gini coefficient** (GI), proxying for inequality is the dependent variable. The Gini coefficient can vary anywhere from 0 (perfect income equality) to 1 (perfect income inequality). The primary independent variable of interest is the **Economic Freedom Index** (EFI). The EFI measures the degree of economic freedom in five main areas: (1) Size of Government; (2) Legal System and Security of Property Rights; (3) Sound Money; (4) Freedom to Trade Internationally; (5) Regulation. Within these five main areas there are 24 components. Each component is measured from 0 ("no economic freedom") to 10 ("full economic freedom"). Gwartney and Lawson (2002) argue that economic freedom declines when taxes, government expenditures, and regulations are substituted for personal preference, voluntary exchange, and market coordination. Accordingly, we examine the effects not only of the overall index, but also individual components of the index that are important for creating greater equality in income, on income distribution. These components include: the size of government (SIZE), proxied by the sum of general government consumption expenditures as percentage of GDP + total transfers and subsidies as percentage of GDP + social security payments as percentage of GDP, total tax revenues as percentage of GDP (TAX), and labour market freedom (LAB), proxied by the sum of the minimum wage legislation index + union density.

Other control variables used are based upon the previous literature. We include **Per capita income** (CAPY) to capture the level of development of a country (Carter 2006), the **Unemployment rate** (U), and the **share of population over the age 65** (POP65) to capture welfare expenditures of the government (Carter, 2006), and the **college attainment rates** (COL) to measure literacy (Ashby and Sobel, 2008).

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While some data for variables go back to 1970, a complete panel with no missing values is currently available for all countries for only the years 1980-2010. Therefore our study covers this time period. The economic freedom data are from the Fraser Institute's Economic Freedom of the World (*Free the World.com* site) index compiled by Gwartney et al. (2006). The rest of the data series are from *Datastream*.

3.2 UNIT ROOT TESTS

We begin the analysis by examining the presence of cross-sectional dependence. Panel unit root tests of the first-generation can lead to spurious results (because of size distortions) if significant degrees of positive residual cross-section dependence exist and are yet ignored. Consequently, the implementation of second-generation panel unit root tests is desirable only when it has been established that the panel is subject to a significant degree of residual cross-section dependence. In cases in which cross-section dependence is not sufficiently high, a loss of power might result if second-generation panel unit root tests that allow for cross-section dependence are employed. Therefore, before selecting the appropriate panel unit root test, it is crucial to provide some evidence on the degree of residual cross-section dependence.

The cross-sectional dependence (CD) statistic by Pesaran (2004) is based on a simple average of all pair-wise correlation coefficients of the OLS residuals obtained from standard augmented Dickey-Fuller (1979) regressions for each variable in the panel. Under the null hypothesis of cross-sectional independence, the CD test statistic follows asymptotically a two-tailed standard normal distribution. The results reported in table 1 uniformly reject the null hypothesis of cross-section independence, providing evidence of cross-sectional dependence in the data. The statistical significance of the CD statistics signifies the presence of cross dependence, irrespectively of the number of lags (i.e., from 1 to 4) included in the ADF regressions.

TABLE 1

Cross-section dependence (CD) test: cross-section correlations of the residuals in ADF(p) regressions

Variables	Lags			
	1	2	3	4
EFI	[0.00][a]	[0.00][a]	[0.00][a]	[0.04][b]
GI	[0.00][a]	[0.00][a]	[0.00][a]	[0.00][a]
SIZE	[0.00][a]	[0.00][a]	[0.00][a]	[0.00][a]
YTAX	[0.00][a]	[0.00][a]	[0.02][b]	[0.03][b]
LAB	[0.00][a]	[0.02][b]	[0.00][a]	[0.02][b]
CAPY	[0.03][b]	[0.00][a]	[0.02][b]	[0.03][b]
U	[0.00][a]	[0.04][b]	[0.00][a]	[0.00][a]
POP65	[0.00][a]	[0.00][a]	[0.02][b]	[0.01][a]
COL	[0.00][a]	[0.00][a]	[0.00][a]	[0.01][a]

Notes: under the null hypothesis of cross-sectional independence the CD statistic is distributed as a two tailed standard normal. Results are based on the test of Pesaran (2004). Figures in parentheses denote p-values. Significance levels: [a](l%) and [b](5%).

Two second-generation panel unit root tests are employed to determine the degree of integration in the respective variables. The Pesaran (2007) panel unit root test does not require the estimation of factor loading to eliminate cross-sectional dependence. Specifically, the usual ADF regression is augmented to include the lagged cross-sectional mean and its first difference to capture the cross-sectional dependence that arises through a single-factor model. The null hypothesis of the Pesaran (2007) test implies the presence of a unit root across all variables under investigation. The bootstrap panel unit root tests by Smith et al. (2004) utilize a sieve sampling scheme to account for both the time series and cross-sectional dependence in the data through bootstrap blocks. All four tests by Smith et al. (2004) are constructed with a unit root under the null hypothesis and heterogeneous autoregressive roots under the alternative hypothesis. The results of these panel unit root tests are reported in table 2 and support of the presence of a unit root in all variables under consideration.

Table 2

Panel unit root tests

Variables	Pesaran CIPS	Pesaran CIPS*	Smith et al. t-test	Smith et al. LM-test	Smith et al. max-test	Smith et al. min-test
EFI	-1.35	-1.41	-1.52	3.02	-1.25	1.45
ΔEFI	-5.62[a]	-5.33[a]	-5.42[a]	18.93[a]	-6.72[a]	6.53[a]
GI	-1.28	-1.26	-1.25	3.15	-1.39	1.28
ΔGI	-5.69[a]	-5.31[a]	-6.34[a]	17.51[a]	-7.85[a]	7.31[a]
SIZE	-1.14	-1.22	-1.28	2.36	-1.43	1.25
ΔSIZE	-6.44[a]	-6.58[a]	-5.73[a]	16.74[a]	-8.74[a]	6.56[a]
YTAX	-1.52	-1.52	-1.34	1.23	-1.29	1.18
ΔYTAX	-7.49[a]	-6.42[a]	-5.63[a]	15.98[a]	-7.81[a]	6.75[a]
LAB	-1.35	-1.34	-1.36	1.22	-1.25	1.26
ΔLAB	-7.54[a]	-6.16[a]	-6.65[a]	18.85[a]	-8.53[a]	7.64[a]
CAPY	-1.22	-1.33	-1.34	1.51	-1.34	1.48
ΔCAPY	-5.54[a]	-5.51[a]	-5.82[a]	18.95[a]	-6.84[a]	6.81[a]
U	-1.29	-1.30	-1.44	1.12	-1.37	1.26
ΔU	-7.23[a]	-6.63[a]	-5.81[a]	18.98[a]	-7.66[a]	7.22[a]
POP65	-1.36	-1.33	-1.35	1.19	-1.34	1.28
ΔPOP65	-5.52[a]	-5.44[a]	-7.83[a]	17.92[a]	-5.78[a]	5.79[a]
COL	-1.26	-1.31	-1.34	1.46	-1.22	1.35
ΔCOL	-5.71[a]	-5.52[a]	-6.59[a]	17.85[a]	-5.74[a]	6.53[a]

Notes: Δ denotes first differences. A constant is included in the Pesaran (2007) tests. Rejection of the null hypothesis indicates stationarity in at least one country. CIPS = truncated CIPS test. Critical values for the Pesaran (2007) test are -2.40 at 1%, -2.22 at 5%, and -2.14 at 10%, respectively, [a] denotes rejection of the null hypothesis. Both a constant and a time trend are included in the Smith et al. (2004) tests. Rejection of the null hypothesis indicates stationarity in at least one country. For both tests the results are reported at lag = 4. The null hypothesis is that of a unit root.*

Given that the respective variables are integrated of order one, we perform the Pedroni (1999, 2004) heterogeneous panel cointegration test to determine whether a long-run equilibrium relationship exists as follows:

$$GI_{it} = \alpha_i + \delta_i t + \beta_{1i} EFI_{it} + \beta_{2i} CAPY_{it} + \beta_{3i} U_{it} + \beta_{4i} POP65_{it} + \beta_{5i} COL_{it} + \varepsilon_{it} \tag{1}$$

where $i = 1, ..., N$ for each state in the panel and $t = 1, ..., T$ refers to the time period. The parameters α_i and δ_i allow for the possibility of country-specific fixed effects and deterministic trends, respectively. In light of the specification of equation (1), we proceed with testing the null hypothesis of no cointegration, $\rho_i = 1$, based on a unit root test of the residuals:

$$\varepsilon_{it} = \rho_i \varepsilon_{it-1} + w_{it} \tag{2}$$

where the estimated residuals, ε_{it}, represent deviations from the long-run equilibrium relationship. Following Pedroni (1999, 2004) both within-dimension and between-dimension approaches to panel cointegration tests are performed. The panel tests based on the within-dimension approach (panel v, panel ρ, panel *PP*, and panel *ADF-statistics*) essentially pool the autoregressive coefficients across different states for the unit root tests on the estimated residuals, taking into account common time factors and heterogeneity across states. The group mean panel tests based on the between-dimension approach (group ρ, group *PP*, and group *ADF-statistics*) are founded on averages of the individual autoregressive coefficients associated with the unit root tests of the residuals for each state in the panel. All seven test statistics, as shown in Panel A of table 3, reject the null hypothesis of no cointegration at the 1% significance level.

TABLE 3

Panel cointegration tests, FMOLS and DOLS estimates

Panel A: Panel cointegration tests

Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	39.89566*	Group ρ-statistic	-39.78508*
Panel ρ-statistic	-38.80943*	Group PP-statistic	-38.24369*
Panel PP-statistic	-38.78062*	Group ADF-statistic	-7.69045*
Panel ADF-statistic	-7.83287*		

*Notes: both the panel and group mean panel tests are distributed asymptotically as standard normal. Of the seven tests, the panel v-statistic is a one-sided test in which large positive values reject the null hypothesis of no cointegration. For the remaining test statistics, large negative values reject the null hypothesis of no cointegration. Statistical significance at the 1% significance level is denoted by *.*

Panel B: FMOLS and DOLS long-run parameter estimates

FMOLS

$$GI = \underset{(14.6)^*}{2.613} - \underset{(-12.3)^*}{0.128}\, EFI + \underset{(10.8)^*}{0.224}\, CAPY - \underset{(-6.19)^*}{0.129}\, U + \underset{(5.86)^*}{0.068} POP65 - \underset{(-7.14)^*}{0.115} COL$$

Adj. $R^2 = 0.57$ LM = 1.12 [0.42] RESET = 1.37 [0.25]

DOLS

$$GI = 1.984 - 0.109\,EFI + 0.207\,CAPY - 0.114\,U + 0.059POP65 - 0.103COL$$
$$(11.4)^{*}\ (-8.53)^{*}\quad (8.75)^{*}\quad (-5.86)^{*}\quad (5.25)^{*}\quad (-5.92)^{*}$$

Adj. $R^2 = 0.63$	LM = 1.38	RESET = 1.54
	[0.33]	[0.18]

*Notes: t-statistics and probability values are reported in parentheses and brackets, respectively. LM is the Lagrange multiplier test for serial correlation. RESET is Ramsey's regression equation specification error test. Statistical significance at the 1% level is denoted by *.*

4 EMPIRICAL RESULTS

4.1 LONG-RUN PANEL PARAMETER ESTIMATES

With the presence of cointegration we employ the fully modified OLS (FMOLS) technique for heterogeneous cointegrated panels (Pedroni, 2000) to arrive at the long-run parameter estimates.[1] Panel B of table 3 displays the long-run parameter estimates of equation (1) based on FMOLS. The Economic Freedom Index yields a statistically significant negative coefficient with respect to the Inequality Index (i.e., Gini). Both per capita income and the share of population over 65 render a statistically significant positive impact on income inequality, while both the unemployment rate and the college attainment rate exert a statistically significant negative impact on income inequality.

The results show that a one unit increase in the Economic Freedom Index decreases income inequality by about 0.128. These findings are consistent with other studies in the relevant literature (Berggren, 1999; Scully, 2002; Ashby and Sobel, 2008). By contrast, our findings are not consistent with those provided by Carter (2006), which could be significantly attributed to the differentiation in methodologies used across the two studies. Furthermore, the positive estimates, associated with per capita income and the share of population over 65, support the notion that both variables contribute to higher income inequality. The positive association between income inequality and income per capita suggest a trade-off between higher income per capita and greater income equality. A positive relationship between income and inequality are supported in the studies of Voitchovsky (2005), Forbes (2000), Li and Zhou (1998), Barro (1998), while Deininger and Squire (1996) do not find a strong relationship between growth and changes in aggregate inequality. Voitchovsky (2005) suggests that redistributive policies in industrial nations such as progressive taxation and social welfare payments could foster growth through its effect on lower segments of the distribution, and yet also dampen growth through their effects on the top end of the distribution. Forbes (2000) similarly, suggests that a larger share of government spending on basic needs such as health and education, which are positively associated with growth, tend to be negatively correlated with inequality, leading to a positive relation between income inequality and growth. A similar argument is put forward by Li and

[1] The estimates from either FMOLS or Dynamic OLS (DOLS) are asymptotically equivalent for more than 60 observations (Banerjee, 1999).

Zhou (1998) who find that greater income inequality could lead to higher economic growth if government consumption enters the utility function.

The results indicate negative coefficients on both the unemployment rate and college attainment which are consistent with the notion that a better educated population leads to lower income inequality, while higher unemployment rates also yield higher probabilities of a more income unequal economy. In terms of the income inequality-unemployment nexus, the results indicate a trade-off between income inequality and unemployment, which is consistent with the presence of a trade-off between higher per capita income and greater income equality. Browning and Johnson (1984) argue that redistributive policies are not costless. The disposable incomes of lower income groups are increased, by reducing the disposable income of higher income groups. As taxes are levied on labour income, this has negative implications for labour supply. There is also a view in the literature that increases in unemployment and wage inequality are "alternative" results of changes in the structure of the demand for labour which could imply a trade-off between income inequality and increasing unemployment. Panel B in table 3 also provides robust evidence in terms of Dynamic OLS (DOLS) estimates recommended by Saikkonen (1992) and Kao and Chiang (2000). These additional findings confirm those reached earlier.

4.2 ROBUSTNESS TESTS: ACROSS THE COMPONENTS OF THE EFI INDEX

While table 3 presents the long-run parameter estimates for the entire panel of states, we present the long-run parameter estimates across the components of the EFI index defined in the data section, i.e. size, taxation and labour structure. The new results are reported in table 4.

Table 4

Panel cointegration tests, FMOLS and DOLS estimates (across components of the EFI index)

Panel A

Panel cointegration tests-SIZE			
Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	42.18239*	Group ρ-statistic	-42.37562*
Panel ρ-statistic	-40.93472*	Group PP-statistic	-40.67329*
Panel PP-statistic	-40.29804*	Group ADF-statistic	-8.89451*
Panel ADF-statistic	-8.72373*		

Panel cointegration tests-TAX			
Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	41.88934*	Group ρ-statistic	-40.56052*
Panel ρ-statistic	-39.72385*	Group PP-statistic	-39.95413*
Panel PP-statistic	-39.45478*	Group ADF-statistic	-7.14677*
Panel ADF-statistic	-7.78523*		

Panel cointegration tests-LAB

Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	36.37494*	Group ρ-statistic	-35.60842*
Panel ρ-statistic	-34.47081*	Group PP-statistic	-34.89521*
Panel PP-statistic	-34.27423*	Group ADF-statistic	-6.16094*
Panel ADF-statistic	-6.36586*		

Notes: similar to table 2.

Panel B: FMOLS and DOLS long-run parameter estimates

SIZE (FMOLS)

$$GI = 1.784 - 0.174\,EFI + 0.247\,CAPY - 0.148\,U + 0.042POP65 - 0.149COL$$
$$(10.4)^{*}\ (-9.28)^{*}\quad (8.71)^{*}\quad (-7.28)^{*}\quad (4.31)^{*}\quad (-6.58)^{*}$$

Adj. $R^2 = 0.61$ LM = 1.04 [0.48] RESET = 1.31 [0.29]

SIZE (DOLS)

$$GI = 1.427 - 0.158\,EFI + 0.219\,CAPY - 0.126\,U + 0.033POP65 - 0.125COL$$
$$(8.51)^{*}\ (-7.13)^{*}\quad (6.68)^{*}\quad (-5.84)^{*}\quad (4.17)^{*}\quad (-5.92)^{*}$$

Adj. $R^2 = 0.65$ LM = 1.38 [0.41] RESET = 1.62 [0.21]

TAX (FMOLS)

$$GI = 0.652 - 0.116\,EFI + 0.219\,CAPY - 0.138\,U + 0.036POP65 - 0.119COL$$
$$(1.51)^{***}\ (-4.74)^{*}\quad (6.36)^{*}\quad (-4.64)^{*}\quad (4.15)^{*}\quad (-5.11)^{*}$$

Adj. $R^2 = 0.49$ LM = 1.25 [0.39] RESET = 1.50 [0.17]

TAX (DOLS)

$$GI = 0.538 - 0.102\,EFI + 0.194\,CAPY - 0.116\,U + 0.028POP65 - 0.098COL$$
$$(1.94)^{**}\ (-4.58)^{*}\quad (5.82)^{*}\quad (-4.29)^{*}\quad (4.04)^{*}\quad (-4.84)^{*}$$

Adj. $R^2 = 0.54$ LM = 1.59 [0.28] RESET = 1.64 [0.15]

LAB (FMOLS)

$$GI = 0.349 - 0.086\,EFI + 0.184\,CAPY - 0.125\,U + 0.032POP65 - 0.138COL$$
$$(1.04)\ (-4.27)^{*}\quad (5.19)^{*}\quad (-4.28)^{*}\quad (4.03)^{*}\quad (-5.62)^{*}$$

Adj. $R^2 = 0.45$ LM = 1.21 [0.43] RESET = 1.57 [0.15]

LAB (DOLS)

$$GI = 0.327 - 0.065\,EFI + 0.152\,CAPY - 0.112\,U + 0.025POP65 - 0.125COL$$
$$(0.85)\ (-4.14)^{*}\quad (4.85)^{*}\quad (-4.09)^{*}\quad (4.36)^{*}\quad (-5.81)^{*}$$

Adj. $R^2 = 0.49$ LM = 1.49 [0.35] RESET = 1.76 [0.11]

*Notes: ** denotes statistical significance at 5%. The remaining are similar to table 3.*

Panel A documents the presence of a cointegrating relationship between income inequality and each component of the EFI index. Once such a long-run equilibrium relationship is acknowledged across the EFI components, Panel B reports the long-run estimations. The empirical findings highlight the negative association between income inequality measures and the three metrics of the EFI index, while the remaining model coefficients retain their expected sign. Comparing the estimates in Panel B, we find substantial variation in the magnitude of the estimated coefficients. In particular, the highest (negative) impact on income inequality comes from the SIZE component, e.g. 0.174 vs 0.116 and 0.086 for the TAX and the LABOR freedom components, respectively. In other words, reducing the intervention of the public-government sector in the economy will act as an incentive for economic forces to boost economic activity and, thus to reduce stronger income inequality, relatively to lower taxation and deregulation of the labour market. Finally, all three estimated equations satisfy a number of diagnostic criteria, giving our estimates a higher validity. These results raise questions regarding the effectiveness of activities supporting higher economic freedom to reduce income inequality.

4.3 ROBUSTNESS TESTS: ACROSS REGIONS

We repeat the above analysis across regions and the new results are reported in table 5. These regional results indicate (once again) the presence of a long-run association between income inequality and both the overall and the disaggregated components of the EFI index (Panel A).

Table 5

Panel cointegration tests, FMOLS and DOLS estimates (across regions)

Panel A

Panel cointegration tests-Europe (23 countries, 647 obs.)			
Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	45.83471*	Group ρ-statistic	-44.54372*
Panel ρ-statistic	-44.32449*	Group PP-statistic	-44.79045*
Panel PP-statistic	-44.87053*	Group ADF-statistic	-9.54522*
Panel ADF-statistic	-9.72654*		

Panel cointegration tests-America (17 countries, 527 obs.)			
Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	40.92286*	Group ρ-statistic	-38.62267*
Panel ρ-statistic	-38.55437*	Group PP-statistic	-38.33426*
Panel PP-statistic	-38.63792*	Group ADF-statistic	-7.10773*
Panel ADF-statistic	-7.16739*		

Panel cointegration tests-Asia (17 countries, 527 obs.)			
Panel Test Statistics:		**Group Mean Panel Test Statistics:**	
Panel v-statistic	48.91256*	Group ρ-statistic	-46.21378*
Panel ρ-statistic	-46.19875*	Group PP-statistic	-46.54384*
Panel PP-statistic	-46.31126*	Group ADF-statistic	-10.44378*
Panel ADF-statistic	-10.85403*		

Panel cointegration tests-Pacific (2 countries, 62 obs.)

Panel Test Statistics:		Group Mean Panel Test Statistics:	
Panel v-statistic	42.65420*	Group ρ-statistic	-40.70633*
Panel ρ-statistic	-40.75298*	Group PP-statistic	-40.41245*
Panel PP-statistic	-40.13529*	Group ADF-statistic	-7.29745*
Panel ADF-statistic	-7.31277*		

Notes: similar to table 3.

Panel B: FMOLS long-run parameter estimates

Europe (FMOLS)

$$GI = 1.235 - 0.283\, EFI + 0.319\, CAPY - 0.239\, U + 0.138 POP65 - 0.257 COL$$
$$(14.6)^{*}\ (-8.29)^{*}\quad (8.25)^{*}\quad (-9.55)^{*}\quad (6.72)^{*}\quad (-5.16)^{*}$$

Adj. $R^2 = 0.67$ LM = 1.18 [0.42] RESET = 1.45 [0.21]

Europe (DOLS)

$$GI = 1.069 - 0.247\, EFI + 0.296\, CAPY - 0.206\, U + 0.119 POP65 - 0.236 COL$$
$$(9.84)^{*}\ (-7.73)^{*}\quad (7.38)^{*}\quad (-7.91)^{*}\quad (5.90)^{*}\quad (-4.82)^{*}$$

Adj. $R^2 = 0.69$ LM = 1.47 [0.34] RESET = 1.72 [0.15]

Americas (FMOLS)

$$GI = 1.458 - 0.225\, EFI + 0.276\, CAPY - 0.163\, U + 0.079 POP65 - 0.271 COL$$
$$(8.17)^{*}\ (-6.38)^{*}\quad (6.94)^{*}\quad (-5.23)^{*}\quad (4.82)^{*}\quad (-7.63)^{*}$$

Adj. $R^2 = 0.65$ LM = 1.29 [0.37] RESET = 1.31 [0.28]

Americas (DOLS)

$$GI = 1.174 - 0.203\, EFI + 0.238\, CAPY - 0.139\, U + 0.061 POP65 - 0.248 COL$$
$$(6.72)^{*}\ (-5.62)^{*}\quad (6.37)^{*}\quad (-5.49)^{*}\quad (4.45)^{*}\quad (-6.41)^{*}$$

Adj. $R^2 = 0.67$ LM = 1.50 [0.29] RESET = 1.64 [0.22]

Asia (FMOLS)

$$GI = 1.018 - 0.109\, EFI + 0.138\, CAPY - 0.072\, U + 0.058 POP65 - 0.170 COL$$
$$(4.29)^{*}\ (-5.32)^{*}\quad (4.57)^{*}\quad (-4.61)^{*}\quad (3.84)^{*}\quad (-5.97)^{*}$$

Adj. $R^2 = 0.60$ LM = 1.39 [0.35] RESET = 1.84 [0.10]

Asia (DOLS)

$$GI = 0.862 - 0.095\, EFI + 0.125\, CAPY - 0.063\, U + 0.052 POP65 - 0.156 COL$$
$$(4.46)^{*}\ (-5.14)^{*}\quad (4.72)^{*}\quad (-4.38)^{*}\quad (4.05)^{*}\quad (-5.51)^{*}$$

Adj. $R^2 = 0.66$ LM = 1.52 [0.31] RESET = 1.67 [0.16]

Pacific (FMOLS)

$$GI = 1.338 - 0.169\ EFI + 0.122\ CAPY - 0.035\ U + 0.039 POP65 - 0.153 COL$$
$$(4.57)^{*}\ \ (-5.88)^{*}\ \ (4.19)^{*}\ \ (-3.84)^{*}\ \ (3.65)^{*}\ \ (-5.27)^{*}$$

Adj. R^2 = 0.52	LM = 1.62	RESET = 1.95
	[0.22]	[0.13]

Pacific (DOLS)

$$GI = 1.187 - 0.144\ EFI + 0.109\ CAPY - 0.032\ U + 0.034 POP65 - 0.138 COL$$
$$(4.38)^{*}\ \ (-5.36)^{*}\ \ (4.47)^{*}\ \ (-4.01)^{*}\ \ (3.91)^{*}\ \ (-4.97)^{*}$$

Adj. R^2 = 0.58	LM = 1.77	RESET = 1.62
	[0.18]	[0.19]

Notes: similar to table 3.

Panel B reports the long-run regional estimations. The new empirical findings highlight the negative association between income inequality measures and the three metrics of the EFI index, while the remaining model coefficients still retain their expected sign. Comparing the estimates in Panel B, we find substantial variation in the magnitude of the estimated coefficients across regions. In particular, the highest (negative) impact on income inequality comes from the European area, e.g. 0.273 vs 0.225, 0.109 and 0.169 for America, Asia and Pacific, respectively. This is possibly due to the very high tax rates in Europe, and then in the American region, which are used to finance government consumption. Therefore reducing the size of the public-government sector in the European economies will provide more solid support to economic forces to boost economic activity and, thus, to reduce income inequality in a stronger manner. Once again, all four regional estimated equations satisfy a number of diagnostics.

5 A NON-LINEAR APPROACH

There is a particular strand in the literature that supports the way an economic freedom index changes over time, i.e. it takes time for economic freedom reforms to affect the course of income inequality (Berggren, 2003; De Haan et al., 2006; Gwartney et al., 2006; Hall et al., 2010). According to this literature, the employment of a linear methodological framework like the one used above tends not to capture the exact form of the relationship under scrutiny. At the same time, the economies included in our sample are at different stages of development, while they run different economic institutions and policies through which they affect not only the course of the real economy, but also distribution policies and, therefore, the income inequality pattern. According to Kuznets (1955), as economies grow, inequality rises until a critical level of income is reached. In this case, inequality begins to decline. In other words, initially the benefits coming from the growth process are creamed off by the upper part of the income distribution, while beyond a threshold growth point these benefits reach the lower part of the same distribution.

In terms of our analysis, initial levels of economic freedom benefit the upper level participants of the income distribution. The primary explanation lies in the fact that in the early stages of expansion, investments are undertaken extensively by those who possess the physical and human capital resources that allow the economy to grow in terms of entrepreneurship and higher trade transactions, thus leading to higher income inequality. As economic freedom keeps rising and growth continues, new economic opportunities are disseminated among all who could not be part of the growth process. Eventually, the low-level participants in economic distribution are also capable of reaping the benefits of economic growth and, thus, income inequality starts to decline. According to Barro (2000), the link between economic growth and income inequality is based on the assumption that the "distribution of political power is more egalitarian than the distribution of economic power", indicating the inability of policies fighting income inequality to induce improvements in equality, for reasons probably related to rent-seeking and corruption. Moreover, growth can lead to increased inequality due to the presence of mechanisms that affect the revenues used to finance redistribution policies. In particular, such revenues are raised through distortionary taxation that provides a disincentive to work. If the power of such disincentives is high, then particular groups in the population, especially those who are near the eligibility level for transfer programs, may become dependent on the government for transfers, which leads to stagnation in incomes. By contrast, those who remain in the labour market continue to acquire higher levels of human capital and thus they can experience income gains; as a result, an increase in income inequality is observed (Cox and Alm, 1995). Finally, Vedder et al. (1988) argue that growth could not reduce inequality due to the presence of the crowding out of private sector charity and the capitalization of public transfer payments (Gruber and Hungerman, 2007; Tullock, 1986).

We make use of the Panel Smooth Threshold Regression (PSTR) model, proposed by Gonzàlez et al. (2005) and Fok et al. (2005), which authorizes a smooth transition, for a number of thresholds, as well as for a continuum of regimes. This approach presents two main advantages: first of all, a PSTR specification allows the income inequality-economic freedom coefficient to vary not only across countries, but also over time. Secondly, this approach allows for a smooth change in country-specific correlation depending upon the threshold variables. This methodological approach allows our study to analyze the role of income in determining income inequality-economic freedom non-linearity by dividing the data depending upon their per capita GDP.

Let us assume the simplest case of a PSTR with two extreme regimes and a single transition function to illustrate the income inequality-economic freedom relationship;

$$GI_{it} = \alpha_i + \beta_0{'}EFI_{i,t} + \beta_1 EFI_{it}\ x\ \Gamma(q_{it}; \gamma, c) + \delta{'}z_{it} + \varepsilon_{it} \qquad (3)$$

where z_{it} is a k-dimensional vector of control variables defined above, α_i represents the individual fixed effects, and ε_{it} is the error term. The transition function is continuous and depends on the threshold variable q_{it}; c is a vector of location parameters. Finally, the parameter γ determines the slope of the transition function.

Gonzàlez et al. (2005) propose a testing procedure in the following order: (1) test linearity against the PSTR model, and (2) determine the number r of transition functions. The test of linearity in the PSTR model can be done by testing H_0: $\gamma = 0$ or H_0: $\beta_1 = 0$. But under the null hypothesis, the test will be non-standard in both cases, and the PSTR model contains unidentified nuisance parameters. A possible solution is to replace the transition function $\Gamma(q_{it}; \gamma, c)$ by its first-order Taylor expression around $\gamma = 0$ and to test an equivalent hypothesis in an auxiliary regression. We then obtain:

$$GI_{it} = \alpha_i + \theta_0{'}EFI_{it} + \theta_1{'}EFI_{it}q_{it} + \delta{'}z_{it} + \varepsilon^*_{it} \quad (4)$$

Since θ_i parameters are proportional to the slope parameter of transition function γ, testing the linearity of income inequality-economic freedom model against PSTR consists of testing H_0: $\theta_1 = 0$ against H_1: $\theta_1 \neq 0$. If we denote SSR_0 the panel sum of squared residuals under H_0 and SSR_1 the PSTR model with m regimes, then the corresponding F-statistic is then defined by: $LM_F = [(SSR_0 - SSR_1) / mK] / [SSR_0 / (TN\text{-}N\text{-}mK)]$, which follows an F test with mK and TN-N-mK degrees of freedom. Finally, T, N and K stand for the number of time, number of countries and number of exogenous variables, respectively.

Before proceeding with the non-linear estimates, we check the linearity vs the non-linearity case. The results of the specification tests are presented in table 6. The table shows the p-value of both the Lagrange multiplier and the Likelihood-ratio test for the null hypothesis of linearity against the alternative of the PSTR specification. The findings highlight that the null hypothesis of linearity is rejected at the 1% significance level. The results imply that there exists a non-linear relationship between the Economic Freedom Index and income inequality in our country sample.

Table 6
LM and LR test of linearity

	Statistic	p-value
Lagrange Multiplier test (LM)	5.81	0.00
Likelihood Ratio test (LR)	23.48	0.00

*Notes: * denotes significance at 1%.*

Table 7 presents the empirical findings of the income inequality-economic freedom relationship, actualized by the PSTR model. As PSTR starts with defining the degree of non-linearity and the number of thresholds (no remaining heterogeneity), our preliminary findings guide us in our selection of the number of transition functions. In our case, the residual sum of squares and the criteria of information lead us to choose one threshold level and one transition function. Our specification is based on the linear model estimation reported above, showing the overall effect of the Economic Freedom Index on income inequality. This linear model is used as benchmark for any other specifications and allows us to show the effect of economic freedom, in addition to its effects that appear after thresholds captured by the non-linear model.

Table 7
Non-linear estimates

Aggregate EFI

$$
GI = \begin{cases} \underset{(4.52)^*}{0.143}\,EFI^b + \underset{(5.11)^*}{0.253}\,CAPY - \underset{(-4.85)^*}{0.148}\,U + \underset{(5.29)^*}{0.053}\,POP65 - \underset{(-5.61)^*}{0.249}\,COL \\[2ex] \underset{(-4.09)^*}{-0.184}\,EFI^a + \underset{(5.67)^*}{0.286}\,CAPY - \underset{(-5.93)^*}{0.175}\,U + \underset{(5.29)^*}{0.091}\,POP65 - \underset{(-6.39)^*}{0.218}\,COL \end{cases}
$$

$q = 6.673$ LM_F [0.01]

EFI-SIZE

$$
GI = \begin{cases} \underset{(4.78)^*}{0.161}\,EFI^b + \underset{(5.39)^*}{0.264}\,CAPY - \underset{(-4.83)^*}{0.151}\,U + \underset{(5.72)^*}{0.060}\,POP65 - \underset{(-4.88)^*}{0.237}\,COL \\[2ex] \underset{(-5.14)^*}{-0.238}\,EFI^a + \underset{(5.94)^*}{0.316}\,CAPY - \underset{(-6.37)^*}{0.163}\,U + \underset{(6.11)^*}{0.083}\,POP65 - \underset{(-6.83)^*}{0.260}\,COL \end{cases}
$$

$q = 6.218$ LM_F [0.00]

EFI-TAX

$$
GI = \begin{cases} \underset{(5.31)^*}{0.133}\,EFI^b + \underset{(5.24)^*}{0.258}\,CAPY - \underset{(-4.81)^*}{0.109}\,U + \underset{(5.34)^*}{0.039}\,POP65 - \underset{(-5.86)^*}{0.198}\,COL \\[2ex] \underset{(-5.97)^*}{-0.174}\,EFI^a + \underset{(5.13)^*}{0.285}\,CAPY - \underset{(-4.93)^*}{0.126}\,U + \underset{(4.27)^*}{0.053}\,POP65 - \underset{(-5.34)^*}{0.219}\,COL \end{cases}
$$

$q = 4.856$ LM_F [0.00]

EFI-LAB

$$
GI = \begin{cases} \underset{(4.99)^*}{0.124}\,EFI^b + \underset{(5.42)^*}{0.168}\,CAPY - \underset{(-4.82)^*}{0.086}\,U + \underset{(4.81)^*}{0.028}\,POP65 - \underset{(-5.24)^*}{0.178}\,COL \\[2ex] \underset{(-4.28)^*}{-0.153}\,EFI^a + \underset{(5.48)^*}{0.219}\,CAPY - \underset{(-4.11)^*}{0.107}\,U + \underset{(4.02)^*}{0.036}\,POP65 - \underset{(-4.72)^*}{0.212}\,COL \end{cases}
$$

$q = 4.137$ LM_F [0.00]

EFI-EUROPE

$$
GI = \begin{cases} \underset{(5.32)^*}{0.158}\,EFI^b + \underset{(5.26)^*}{0.285}\,CAPY - \underset{(-4.31)^*}{0.125}\,U + \underset{(4.58)^*}{0.077}\,POP65 - \underset{(-5.61)^*}{0.214}\,COL \\[2ex] \underset{(-5.86)^*}{-0.165}\,EFI^a + \underset{(6.32)^*}{0.316}\,CAPY - \underset{(-5.64)^*}{0.193}\,U + \underset{(4.82)^*}{0.105}\,POP65 - \underset{(-6.44)^*}{0.237}\,COL \end{cases}
$$

$q = 5.319$ LM_F [0.00]

EFI-AMERICA

$$
GI = \begin{cases} \underset{(5.24)^*}{0.144}\,EFI^b + \underset{(5.37)^*}{0.256}\,CAPY - \underset{(-4.76)^*}{0.104}\,U + \underset{(4.91)^*}{0.082}\,POP65 - \underset{(-5.17)^*}{0.203}\,COL \\[2ex] \underset{(-5.57)^*}{-0.162}\,EFI^a + \underset{(6.25)^*}{0.309}\,CAPY - \underset{(-5.48)^*}{0.152}\,U + \underset{(4.90)^*}{0.102}\,POP65 - \underset{(-5.83)^*}{0.231}\,COL \end{cases}
$$

$q = 7.742$ LM_F [0.00]

EFI-ASIA

$$GI = \begin{cases} \begin{array}{lllll} 0.131\ EFI^{b} + & 0.256\ CAPY - & 0.147\ U + & 0.086\ POP65 - & 0.236\ COL \\ (5.47)^{*} & (5.39)^{*} & (-4.95)^{*} & (4.82)^{*} & (-5.48)^{*} \end{array} \\ \\ \begin{array}{lllll} -0.172\ EFI^{a} + & 0.304\ CAPY - & 0.186\ U + & 0.114\ POP65 - & 0.249\ COL \\ (-5.59)^{*} & (6.15)^{*} & (-5.37)^{*} & (4.98)^{*} & (-6.13)^{*} \end{array} \end{cases}$$

$q = 4.360$ $\quad LM_F$ [0.00]

EFI-PACIFIC

$$GI = \begin{cases} \begin{array}{lllll} 0.153\ EFI^{b} + & 0.214\ CAPY - & 0.107\ U + & 0.048\ POP65 - & 0.226\ COL \\ (5.61)^{*} & (4.68)^{*} & (-4.58)^{*} & (4.49)^{*} & (-5.48)^{*} \end{array} \\ \\ \begin{array}{lllll} -0.169\ EFI^{a} + & 0.265\ CAPY - & 0.147\ U + & 0.055\ POP65 - & 0.205\ COL \\ (-5.26)^{*} & (6.10)^{*} & (-5.42)^{*} & (4.96)^{*} & (-6.02)^{*} \end{array} \end{cases}$$

$q = 5.082$ $\quad LM_F$ [0.00]

Notes: ^{b} denotes "below" and ^{a} denotes "above", q *is the threshold parameter. Figures in brackets denote p-values, while those in parentheses denote t-statistics. The LMF statistic measures whether the regime switching is significant or not, i.e. the test of linearity versus PSTR. The remaining notes are similar to those in table 3.*

The economic freedom threshold appears at 6.67. Nevertheless, the EFI effect below this level is positive and statistically significant and above this level it is negative and statistically significant. Indeed, under the regime of higher economic freedom (>6.67), other things being equal, an increase of 1 unit in economic freedom reduces income inequality by 0.184, whereas in the first regime (<6.67) the effect of the index is positive and statistically significant. These empirical findings suggest that countries with an economic freedom index below this level tend to experience higher inequality when the index indicates stronger economic freedom, whereas countries with an economic freedom index above this level tend to experience reductions in inequality when the index displays higher levels of economic freedom. Finally, the statistical significance of the LM_F statistic rejects the null hypothesis that the coefficient of the slope coefficient is zero. Similar results are also obtained in terms of the components of the EFI index, with the stronger results coming from the SIZE component.

In terms of the components of the EFI index, all three dimensions of that index support the presence of homogeneous results. With respect to the SIZE, TAX and LAB component, the thresholds appear at 6.22, 4.86 and 4.14, respectively. The EFI-SIZE, EFI-TAX and EFI-LAB effects below these levels are positive and statistically significant and above this level they are negative and statistically significant. In both regimes the strongest effect appears in the EFI-SIZE case, indicating that reducing the degree of government intervention below the corresponding threshold level tends to display higher inequality when the index is on the rise, whereas with the component of the index above this level tends to display reductions in inequality for extra degrees in government intervention declines. The statistical significance of the LM_F statistic in all three cases leads to the rejection of the linear "income inequality-economic freedom" relationship. Finally, the results remain robust across the geographical regions, with the testing procedure to substantially recommend the non-linear approach. The thresholds turn out to be 5.32, 7.74, 4.36 and 5.08, for the case of Europe, America, Asia and Pacific country samples, respectively.

6 POLICY IMPLICATIONS AND CONCLUSIONS

This study used panel-country data to study the dynamic relationship between economic freedom and income inequality. While the literature has devoted considerable attention to studying the above nexus, the results provided having been inconclusive, the dynamic aspects have been studied primarily through a linear framework.

Unlike previous researchers, we employed both a linear and a non-linear panel cointegration model to identify the long-run equilibrium relationship between economic freedom and income inequality for the entire panel of countries and at the components of economic freedom index level. The linear long-run parameter estimates for the entire panel of countries showed that the association under study was negative, while the non-linear long-run parameter estimates indicated that above a threshold point the association between economic freedom and income inequality retained its negative sign, while below this threshold point the association turned out to be positive. In other words, the empirical findings, in terms of the non-linear long-run model, displayed that beginning from a low level of economic freedom, a higher level of this index generated more inequality as the participants in the upper part of the income distribution benefit relatively more than the lower-level participants. As increases of the index continue, then the lower-level participants tend to experience larger relative income gains, while these findings were robust to the components of the Economic Freedom Index as well.

With regard to the policy implications of our findings, the results suggest that once the threshold is overtaken, economic freedom should promote greater equality in income distribution. The movement from the low phase of economic freedom to the high phase will not occur automatically. Reducing the size of government and increased human capital accumulation will permit progression to the next phase. The results additionally suggest that countries face a trade-off between higher per capita income and income equality, and unemployment and income inequality. The nature and extent of the perceived "trade-off" between unemployment and inequality are also subject to policy interventions. The results demonstrate the risks of the initial phase of economic freedom where an increase in inequality requires a pro-active stance by policy makers to take freedom to the next phase. The distributional impact of any exogenous shock is not predetermined. Even if a shock primarily or disproportionately affected particular income-earners, some of these costs can be redistributed (through fiscal and other measures) in such a way as to mitigate the net impact on inequality.

APPENDIX

The study includes the following countries:

Europe = Austria, Belgium, Bulgaria, Croatia, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Norway, Poland, Portugal, Romania, Slovakia, Slovenia, Spain, Sweden, Switzerland, U.K.

America = Argentina, Bolivia, Brazil, Chile, Costa Rica, El Salvador, Equator, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Uruguay, U.S., Venezuela.

Asia = Brunei, Hong-Kong, India, Indonesia, Israel, Japan, Jordan, Kuwait, Malaysia, Nepal, Oman, Pakistan, Qatar, Singapore, South Arabia, South Korea, Taiwan.

Pacific = Australia, New Zealand.

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