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TESTING THE RELATIONSHIP BETWEEN ECONOMIC FREEDOM AND INCOME INEQUALITY IN THE USA

Allen L. Webster

Foster College of Business, Bradley University, Peoria, Illinois, USA

alw@fsmail.bradley.edu

Abstract

The Gini coefficient is used to examine the impact of economic freedom on income inequality among the 50 US states. The degree of economic freedom is provided by the Fraser Institute in Vancouver, Canada. A fixed-effect model based on panel data from 2000-2013 is estimated to determine if differences exist among the four census regions identified by the US Census Bureau. The findings clearly suggest that those states characterized by higher levels of economic freedom exhibit greater income equality. A Dickey-Fuller test for stationary revealed the need for first-differencing and a Granger-causality test concluded that uni-directional causality existed between income distribution and economic liberty.

Key words

Gini Coefficient; Economic Freedom; Income Distribution; Fixed-Effects Models; Granger Causality.

INTRODUCTION

Considerable attention has been devoted to the trend in income inequality in the United States over the past several years. This concern has become a particularly critical issue among economists, business leaders and even the concerned citizenry. Political figures attempting to draw attention to their social conscience have also attached themselves to this pressing socio-economic dilemma.

In his December 4, 2013 speech before the Center for American Progress, President Obama praised the New Deal and the War on Poverty for building “the largest middle class the world has ever known,” but regretfully alluded to the “dangerous and growing inequality and lack of upward mobility that has jeopardized middle-class America’s basic bargain – that if you work hard, you have a chance to get ahead”.

Although somewhat positive in nature, his address carried the message that progress has stalled and income disparities have widened perceptively. Much of the available data seem to support the President's contention. It is generally agreed that income inequality has been on the rise in the United States since the late 1970s. Recent studies have shown that income disparities began to increase in the U.S. economy in the early 1970s and continue today (Webster, 2014; Apergis, et al, 2011; Ram, 2012). Sommeiller and Price (2014), for example, find that between 1979 and 2007, the top 1% income earners took home over one-half (53.9%) of the total increase in US income. Over that same period, the average income of the bottom 99% grew by only 18.9%.

This gap between the income classes has often been cited as a precursor for many socio-economic ills ranging from poverty, impediment to economic growth, elevated crime rates and general social disorder. Over the past several decades, after-tax incomes for the top 1% of households grew 275%. This compares to an 18% rise of the incomes of those in the bottom quintile. According to Piketty and Saez (2003), the top "1%ers" took home a 95% gain in the first three years of the recovery following the Great Recession. Furthermore, Piketty (2014) argues that the central contradiction of capitalism is that it leads to the concentration of wealth in the hands of those who are already rich. He denounces the evils of free markets and the inequities they produce. Piketty (2014) is particularly loathsome of inherited wealth. He contends that it generates slower economic growth and increases the ratio of capital to income which further exacerbates the disparities in incomes.

A common measure of income inequality is the Gini coefficient developed by the Italian economist and statistician Corrado Gini (1997). As an index of inequality, the coefficient (or ratio) ranges between zero and 1.00. The lower the coefficient is, the more equally incomes are dispersed throughout the economy. From all accounts, the US does not compare favorably to many other nations. Figure 1 provides Gini coefficients for some of the 34 nations that constitute the Organization for Economic Cooperation and Development (OECD) (Webster, 2014). Russia, which is not a member nation, is also included for comparison. The upper values represent the coefficients before taxes and transfers while the lower values are the ratios after taxes and transfers have reduced the degree of inequality. It can be seen that only Chile, Turkey and Mexico report after-tax Ginis greater than the US. The after- tax Gini for Russia is unavailable. The OECD averages of 0.316 and 0.463 are also included.

It is interesting to note the net changes in the degree of inequality after taxes and transfers. For example, Canada's Gini coefficient was reduced by over 0.10 through public efforts to combat income inequality.

Clearly, France, Germany and Italy reduced their coefficients the most while Chile, Ireland, Korea and Mexico had very little effect on the coefficients as a result of transfers from the wealthy to the poor. The mean reduction was 0.1106 and the median was 0.117. The less well of in Germany benefited the most as that nation's Gini coefficient dropped by 0.209 while that of Mexico fell the least by 0.018. The US reduction was 0.108. The mean OECD decrease was 0.147.

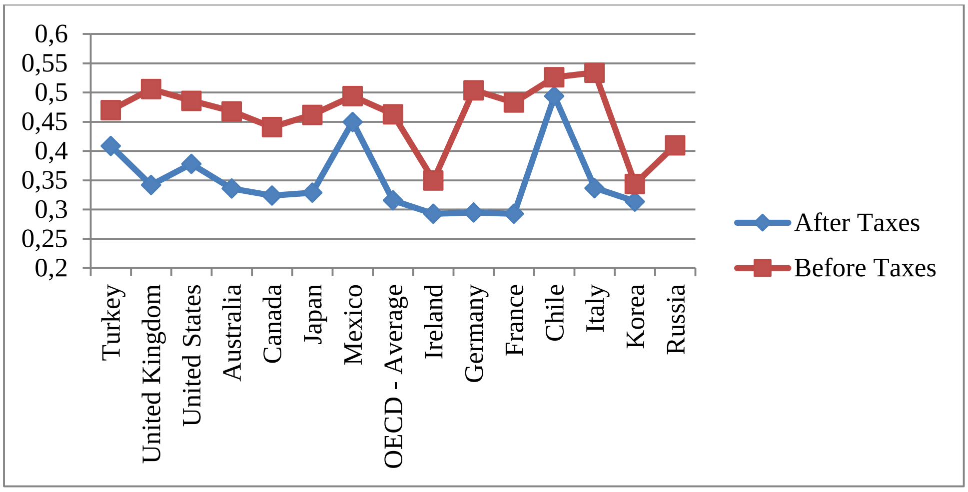

FIG 1. GINI COEFFICIENTS FOR OECD NATIONS, 2011
Source: Webster (2014)

The remainder of the paper is organized as follows: the next section briefly examines commonly cited causes for increases in income inequality in the USA. A definition of economic freedom and the manner in which it is measured is then provided. The research methodology and results of OLS models designed to examine the relationship between economic freedom and income distribution are presented. A fixed-effects model is used to identify any differences in the patterns between economic freedom and income distribution across the four census regions identified by the US Bureau of Census. Granger causality tests are performed to detect any feedback between income distribution and economic freedom.

CAUSES OF INEQUALITY IN THE USA

Many factors are thought to contribute to this growing wage gap. Perhaps foremost is the exportation of manufacturing jobs to foreign sources in poorer nations. This trend in "globalization" has actually created a greater disparity in incomes in both the US as well as those nations to which the jobs are sent. The jobs that are exported are those held by lower-income workers. Upper managerial positions held by higher-paid executives are retained in the US. This has the effect of creating a larger gap between the socio-economic classes here in the US. The jobs sent over-seas are given the more skilled and highly trained workers who are already enjoying incomes in excess of those earned by the lower classes. This results in a larger differential in those nations.

The “diploma gap” has also registered an effect. In 1980 an American with a college degree earned about 30% more than someone who had terminated their education at the high school level. In more recent years, someone with a college education earns roughly 70% more than a high school graduate. The premium for a graduate degree has increased from roughly 50% in 1982 to well over 100% today in some instances (Webster, 2014). Given the explosive rise in the cost of a college education, only those in the more affluent classes can afford more extensive schooling. This has the effect of widening still further any prevailing income differences.

The decline in the level of unionization across the United States has further magnified the separation of the income classes. Evidence has shown (Western & Rosenfeld, 2011; Card & Lemieux, 2004) that incomes are more evenly distributed in areas with higher rates of unionization. In 1983, 20.1% of the labor force or 17.7 million workers were union members. Today those figures stand at 11.3% and 14.5 million members. The impact of such dynamics on the wage gap is evident when it is considered that in 2013 the median weekly earnings for union members was $950, while those who were not union members had median weekly earnings of only $750 (US Department of Labor, January 24, 2014). Further, as fewer union members occupy the work force the gap must widen since the incomes of high-earning management do not depend on union membership.

It is generally recognized that tax laws tend to favor the wealthy. Provisions that affect stock options and capital gains permit the wealthy to shelter their incomes from the tax man. Many argue this lessens the tax burden on the wealthy and has accelerated the separation between the rich and the poor.

ECONOMIC FREEDOM

Although several other forces can be identified that affect income deviation, strong indication persists that income dispersion is also influenced by the degree of economic freedom prevailing in a political or geographical unit. Although no universally accepted definition of economic freedom has been established, it generally refers to the ability of economic participants to make decisions and take actions without restraint from central forces. It is philosophically based on principles ranging from pure laissez-faireism to that advocated by the classical libertarian. Emphasis is placed on reliance on free markets, private property and individual choice.

Most studies rely on the index of economic freedom (EFI) developed by the Fraser Institute in Vancouver, Canada. The Institute provides an annual cardinal measure of the extent of economic freedom prevailing in all 50 U.S. states as well as all Canadian providences. These data are provided in annual reports entitled the *Economic Freedom of North America*. The Institute defines economic freedom as a condition in which

> *Individuals have economic freedom when (a) property they acquire… is protected from physical invasion by others and (b) they are free to use, exchange or give their property as long as their actions do not violate the identical rights of others. Thus, an index of economic freedom should measure the extent to*

which rightly acquired property is protected and individuals are engaged in voluntary transactions.

The indices of economic freedom used by the *Fraser* Institute focus on six areas of concern. Each area contains subcategories as shown in Table 1.

TABLE 1.THE AREAS AND COMPONENTS OF ECONOMIC FREEDOM OF NORTH AMERICA INDEX

Area 1-Size of Government
1A-General Consumption Expenditures by Government as a Percentage of GDP
1B -Transfers and Subsidies as a Percentage of GDP[1]
1C-Social Security Payments as a Percentage of GDP
Area 2-Takings and Discriminatory Taxation
2A-Total Tax Revenue as a Percentage of GDP
2B-Top Marginal Income Tax Rate and the Income Threshold at Which It Applies
2C-Indirect Tax Revenues as a Percentage of GDP
2D-Sales Taxes Collected as a Percentage of GDP
Area 3-Regulation
3A-Labor Market Freedom
3B-Credit Market Regulation
3C-Business Regulations
Area 4-Legal System and Property rights
4A-Judicial Independence
4B-Impartial Courts
4C-Protection of Property Rights
4D-Military Interference in Rule of Law and Politics
Area 5-Sound Money
5A-Money Growth
5B-Standard Deviation of Inflation
5C-Inflation: Most Recent Year
5D-Freedom to Own Foreign-Currency Bank Accounts
Area 6-Freedom to Trade Internationally

Note: [1]Gross state product (GSP) is used in each of these cases when comparing the 50 US states

Each of the areas and their subcategories are largely self-explanatory. However, certain select entries may require further explanation. For example, "Takings and Discretionary Taxation" simply refers to the revenue governments acquire through direct taxation. Discretionary taxation applies only to those individuals engaging in a particular activity. Sales taxes indicated in subcategory 2D refer only to transactions involving taxable retail purchases.

The Institute notes that Areas Four, Five and Six pertain primarily to international comparisons. Since this paper is designed to compare states within the US, only Areas One, Two and Three are used in the analysis.

The index for each component and sub-component is based on a scale from 0 to 10, with 10 indicating the highest degree of economic liberty. The overall index is then compiled as an unweighted average of the three primary areas. A more complete description of the items used to generate the indices can be obtained from any of the annual reports provided by the Fraser Institute.

The indices published by the Institute measure economic freedom at two levels: the sub-national and the all-government. The sub-national level refers to the provincial and municipal governments in Canada and the state and local governments in the United States. At the all-government level the impact of federal governments is measured. All 50 states in the US and the 10 provinces in Canada are included in the Institute's reports. This paper relies only on data from the 50 US states.

A simple mathematical formula is used to mitigate subjective judgments and ordinal rankings that do not permit mathematical manipulation or calculation. Instead, the EFI is a relative valuation in which a cardinal measure comparing each individual geographical region to a set standard is computed. It was constructed by the Institute to represent the underlying distribution of all 10 of the sub-components in Areas 1, 2 and 3. Thus, this index is a relative ranking.

The index assigns a higher score when, for example, component 1A, General Consumption Expenditures by Government as a Percentage of GDP, is smaller in one state or province relative to another. The rating formula is consistent across time to allow an examination of the evolution of economic freedom. In order to construct the overall index without imposing subjective judgments about the relative importance of the components, each area is equally weighted and each component within each area is equally weighted. Thus, Areas 1, 2, and 3 are equally weighted, and each of the components within each area is equally weighted. For example, the weight for Area 1 is 33.3%. Area 1 has three components, each of which received equal weight, or 11.1%, in calculating the overall index.

Objective methods are used to calculate and weigh the components. For all components, each observation is transformed into a number from zero to 10 using the formula

$$\left[\frac{V_{max} - V_i}{V_{max} - V_{min}}\right] * 10 \qquad (1)$$

where V_{max} is the largest value found within a component, V_{min} is the smallest, and V_i is the observation to be transformed. For each component, the calculation includes all data for all years to allow comparisons over time.

Over time, the US has displayed distinct trends in its measures of economic freedom. As Figure 2 displays, near the turn of the century the US reached an index of 8.65 out of 10. This represents the vertex of economic freedom in America. Since then the extent of the nation's commitment to free enterprise has been decreasing steadily.

The data also show that in all three areas seen in Table 1, the US has recorded pronounced declines relative to other nations. This too is reflected in Figure 3. A higher ranking indicates a lower degree of economic freedom relative to other nations. The value for the year 2013 indicates that the U.S. ranked 19th in the world in terms of the measure of economic freedom enjoyed by its residents. Inarguably, the US position relative to other nations has shown a steady decline over the past 30 years. Much of this decline is due not only to deterioration in US policies and practices, but stems also from a relaxation in constraints placed on economic participants in other nations.

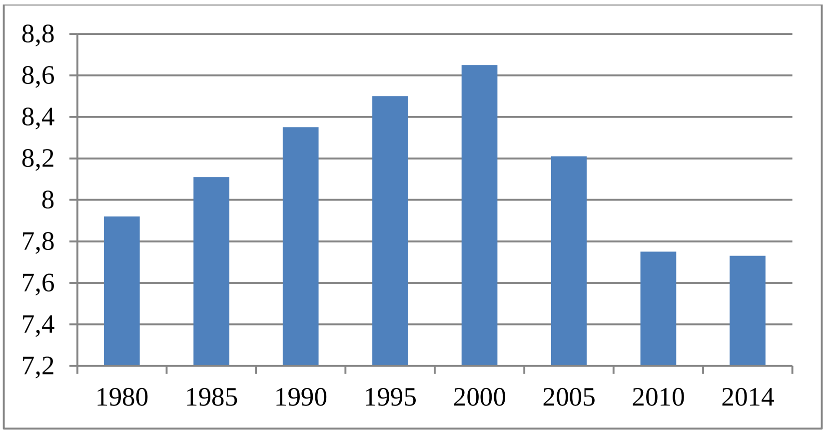

FIG 2. TREND IN ECONOMIC FREEDOM IN THE USA
Source: Fraser Institute (http://www.freetheworld.com/release.html, extracted August 10, 2014)

Traditionally, Hong Kong and Singapore have dominated the top two world-wide positions in terms of promoting free enterprise. Australia, Switzerland, New Zealand and Canada have gained significant prominence across the globe. Sweden and Denmark have also reported impressive gains in economic freedom. Countries experimenting with milder forms of socialism than they did in the past have also gained ground relative to other nations. Estonia, Lithuania and the Czech Republic are notable in that regard. These dynamics have relegated the US to a lesser position world-wide in terms of its index vis-à-vis other nations.

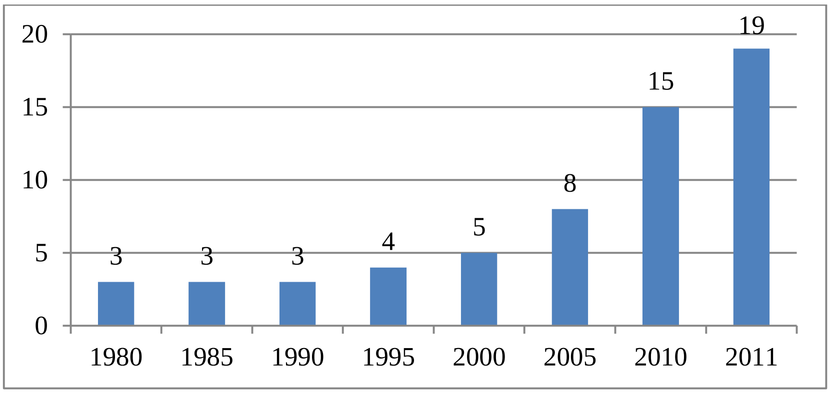

FIG 3. US EFI RANKINGS COMPARED TO OTHER NATIONS
Source: Fraser Institute (http://www.freetheworld.com/release.html, extracted August 10, 2014)

Considerable work has been done in the past that compares nations around the globe in regard to their relationships between income distribution and economic freedom (Gwartney et al, 1996; Carter, 2006; Berggren, 1999; Scully, 2002; Cebula et al, 2013). Similar work comparing U.S. states, however, is much less prevalent. Ashby and Sobel (2008) offer an insightful discourse regarding the impact of economic freedom among the 50 US states. They conclude that "…*changes* (emphasis added) in economic freedom are associated with higher income and higher rates of income growth … and with reductions in relative income inequality". They further contend, however, that the relationship between the prevailing level of economic freedom and income inequality is statistically insignificant.

For the purpose of this paper, panel data were collected for all 50 states for the 14 years from 2000 through 2013. They included the Gini coefficients maintained by the US Bureau of Census and the EFI provided by the Fraser Institute. The control variables included population figures, median income levels, gross state products measured in millions of dollars, percentages of high school graduates and percentage of minorities in each state. These factors were considered here because in many of the studies noted above, they were shown to be statistically significant as explanatory variables of income distribution. With the exception of measures for economic freedom, all were taken from US Census Bureau data and were extracted in the summer of 2014.

Some of the relevant descriptive statistics are displayed in Table 2. The maximum Gini coefficient of 0.499 in both 2000 and 2013 was held by the state of New York thereby indicating the greatest degree of income inequality in both years. New York's Gini coefficient changed over the course of those 14 years, but in 2013 settled at the same ranking it was in the year 2000.

TABLE 2. BASIC DESCRIPTIVE STATISTICS

Variable	Mean	Median	Standard Deviation	Minimum	Maximum
Gini 2013	0.452	0.453	0.0178	0.419	0.499
Gini 2000	0.446	0.445	0.0213	0.402	0.499
EFI 2013	6.560	6.600	0.547	5.400	7.800
EFI 2000	8.236	8.300	0.230	7.600	8.800
Change in Gini	0.006	0.008	0.008	-0.017	0.021
Change in EFI	-1.670	-1.700	0.365	-2.500	-0.900

The minimum Gini coefficients in the years 2000 and 2013 of 0.402 and 0.419 occurred in Alaska and Utah, respectively. Alaska had the greatest degree of income equality in 2000. By 2013, Utah claimed that spot.

Delaware reported the highest degree of economic freedom in the year 2000 with an EFI of 8.8 while West Virginia recorded the minimum Index of 7.6. West Virginia retained that position in 2013 with the lowest measure of economic freedom at 5.4. In 2013 the highest degree of economic freedom prevailed in Mississippi with an index of 7.8.

Changes in both the Gini ratio and the Economic Freedom Index over the 14 year period are also recorded. The maximum change in economic freedom occurred in New Mexico with a reduction of -2.5 while Wyoming reported the smallest change of -0.90. All 50 states, without exception, recorded a reduction in the EFI. Recall, as the index decreases, the measured degree of economic freedom decreases. Thus, in an absolute sense, the extent of economic freedom as define and calculated by the Fraser Institute has fallen.

The minimum and maximum changes in the Gini ratio were -0.017 reported by West Virginia and a 0.021 attributed to Vermont. Keeping in mind that an increase in the ratio indicates greater income inequality, Vermont is guilty of the largest rise in income differences between the haves and the have-nots during that period.

Various studies (Gwartney, et al, 1996; Barro, 2000; Spindler, et al, 2008; Scully 2002) based on a comparison of national economies around the globe have clearly concluded that a positive relationship exists between economic freedom and income equality. As the economic freedom index (EFI) as defined above increases, so does income equality as measured by the Gini coefficient. Thus, the testable hypothesis that increases in equality (decreases in the Gini) are associated with increases in economic freedom is stated as

$$\frac{\partial Gini}{\partial EFI} < 0 \qquad (2)$$

While the studies just cited above affirms this assertion within entire nations, the question remains as to whether that relationship holds internally among the 50 US states. Ashby and Sobel (2008) concluded that the relationship between the prevailing level of economic freedom and income inequality is statistically insignificant. However, Bennett and Vedder (2013) content that an inverted U-shape can best be used to describe the relationship between economic freedom and income inequality. Increases in economic freedom initially result in a rising Gini coefficient. But once some “tipping point” in the level if economic freedom is reached, the level of income inequality begins to wane.

The remainder of the paper examines this relationship between economic freedom and income distribution, both at specific points in time as well as the dynamics of the relationship over time. Equation (2) serves as the testable hypothesis.

MODEL SPECIFICATIONS AND REGRESSION RESULTS

Initial model specifications regressed the Gini ratios from the year 2000 on potential explanatory variables for the same year. In addition to the Economic Freedom Index, control variables for states' population, gross state product measured in millions of dollars, median income and the percentage of the population with a high school degree were included. Only gross state product, median income and the percentage of high school graduates proved statistically significant. The adjusted coefficient of determination reported in at 57%.

In the interest of parsimony, the model was re-specified to include those three variables along with the EFI. The results are displayed in Table 3. All four explanatory variables proved statistically significant at acceptable alpha-values. The coefficients are, of course, quite small since the response variable never exceeds 1.00. Of considerable interest is the fact that the coefficient for the EFI reported to be negative. This reveals that an increase in economic freedom is associated with a reduction in the Gini ratio indicating a movement toward more income equality. Although EFI_{2000} was only marginally significant at the 8.1% level, greater income equality is associated with an elevated degree of economic freedom. These findings are in contrast to those reported by Ashby and Sobel (2008).

Subsequently, a similar model was estimated using the more recent data from 2013 (Table 3). As in Ashby and Sobel (2008), the EFI reported as statistically insignificant. These models offer 'spot checks' on the relationship between economic freedom and income distribution at a specific point in time. A truer measure of the manner in which changes in economic freedom might affect income distribution requires an analysis over some time span. An accurate measure how the levels of economic autonomy might restructure income dispersion is best reflected by an examination of the movements in each factor over time.

In this effort, a model was specified in which the changes in the Gini coefficients over the time period 2000 to 2013 are set as the response variable. The primary explanatory variables include the initial Gini ratio in 2000, the initial EFI in 2000 and the change in the Index over the time period in question. The control variables used in the models above are retained and changes in those variables are added.

Table 3 reveals that the change in the EFI as well as its initial measure at the outset of the time period in 2000 both proved highly significant. Moreover, both carry a negative sign. This suggests that higher levels in the initial measure of economic freedom as defined by the Fraser Institute and increases in that measure over time are associated with a lower Gini coefficient. This reduction in the ratio evidences greater income equality.

The negative correlation between the change in the Gini ratio and the initial EFI_{2000} indicates that states with greater degrees of economic freedom experienced less change in the distribution of income. Higher levels of economic freedom tend to stabilize the current distributional pattern of income. This is perhaps because states that already enjoy a high degree of economic freedom find it more difficult to raise the level of freedom even further.

The change in the Gini ratio was also negatively related to the change in the measure of economic freedom. This is not to say that that an increase in the Index is associated with decrease in the Gini ratio, but is instead related in a negative fashion to changes in that measure. This might suggest that there prevails an inelastic association between these two socio-economic measures. As the change in the Index becomes greater, changes in the Gini ratio diminish.

TABLE 3. REGRESSION RESULTS WITH GINI AS THE RESPONSE VARIABLE

OLS Results With $Gini_{2000}$ as Response Variable				
Variable	Coefficient	Standard Error	t-value	p-value
Constant	0.70819	0.08426	8.40	0.000
EFI_{2000}	-0.01832	0.01027	-1.78	0.081
GSP_{2000}	0.5E-7	0.000	5.24	0.000
Median $Income_{2000}$	-0.121E-5	0.000	-3.41	0.001
%HSED	-0.00096	0.0002	3.89	0.000
Adjusted R^2	56.7%			
Standard Error	0.014			
OLS Results With $Gini_{2013}$ as Response Variable				
Variable	Coefficient	Standard Error	t-value	p-value
Constant	0.55075	0.03322	16.58	0.000
EFI_{2013}	-0.0048	0.00387	-1.25	0.217
GSP_{2013}	0.000002	0.0000004	4.83	0.000
Median $Income_{2013}$	-0.0000004	0.0000003	-1.33	0.189
$\%HSED_{2013}$	-0.00071	0.000323	-2.18	0.034
Adjusted R^2	39.0%			
Standard Error	0.0139			
OLS Results With Change in $Gini_{2013\text{-}00}$ as Response Variable				
Variable	Coefficient	Standard Error	t-value	p-value
Constant	-0.063	0.0627	-1.00	0.3233
EFI_{2000}	-0.0140	0.0058	-2.41	0.0206
$\Delta EFI_{2013\text{-}00}$	-0.0034	0.0010	-2.97	0.0050
GSP_{2000}	0.0000	0.01E-7	0.720	0.0476
$\Delta GSP_{2013\text{-}00}$	-0.0000	0.02E-7	-1.130	0.2670
Median $Income_{2000}$	0.028E-6	0.015E-5	1.860	0.0710
ΔMedian $Income_{2013\text{-}00}$	0.04E-6	0.02E-5	0.210	0.8333
%HSED	0.07E-5	0.014E-2	0.470	0.6410
$\Delta\%HSED_{2013\text{-}00}$	0.03E-4	0.018E-2	1.48	0.1470
Adjusted R^2	55.6%			
Standard Error	0.005			

Perhaps prevailing institutional, political, economic and other social structures that are already in place within a state promote income redistribution. These established qualities already present merely continue the trend toward a redistribution that favors the less fortunate but do so with a diminishing effect. Regardless of any cause-and-effect, the empirical results clearly endorse the contention that elevated levels of economic freedom are associated with greater income equality and that changes in economic freedom correspond positively with changes in income equality.

MEASUREMENTS OF REGIONAL DIFFERENCES: A FIXED EFFECTS MODEL

In a nation with over 317 million residents that covers nearly four million square miles it would surprise no one to learn that different regions will vary noticeably in terms of their socio-economic order. Median incomes, education levels, the industrial mix, reliance on agriculture and a host of other idiosyncratic attributes all vary greatly between and among geographical regions of the nation. It is therefore reasonable to hypothesize that forces will interact differently across the nation producing alternative results in social and economic outcomes. For that reason, it seems prudent to test for regional differences in terms of the relationship between economic freedom and its impact on income distribution.

In that effort, a fixed-effects model is estimated across the four geographical regions identified by the US Census Bureau. These regions are the Northeast, Midwest, South and West. The states that are included in each region are shown in Table 4. The fixed-effects model allows for distinctions among different cross-sectional categorical units such as, in our present case, geographical regions.

Fixed-effects models are well adapted to control for omitted variables that might be correlated with regressors that are categorically-specific and time invariant. It is therefore possible to capture the unadulterated impact of the EFI on income inequality by incorporating as regressors only those variables that measure economic freedom. The model used here identifies the Gini ratio of 2013 as the regressand and includes only the $Gini_{2000}$, the EFI from both 2000 and 2013, the change in the EFI over the time period in question and, of course, the dummy variables for all four regions.

Fixed-effects models rely on within-categorical variation across time. Therefore, they require measurable within-categorical variation of the explanatory variables. In addition, accurate estimation also demands less within-categorical variations in the measurement error of the regressand.

These conditions should cause no problems in the current analysis. Variations in the levels of economic freedom have already been noted in that all 50 states reported drops in their EFIs over the time span under survey. Furthermore, since the same reliable source for the Gini ratio is used throughout the study, measurement error among the states is likely held to a minimum.

The methodology used in fixed-effects models will detect and quantify regional differences in terms of the interplay between economic freedom and income inequality. This is done by allowing the intercept to differ among the cross-sectional categories, but each intercept for each category remains constant over time. The difference between cross-sectional categories is assigned to the intercept and results in constant slope coefficients.

TABLE 4. US CENSUS BUREAU REGIONS

Region 1: Northeast	
Connecticut	New York
Maine	Pennsylvania
Massachusetts	Rhode Island
New Hampshire	Vermont
New Jersey	
Region 2: Midwest	
Indiana	Missouri
Illinois	Nebraska
Iowa	North Dakota
Kansas	Ohio
Michigan	South Dakota
Minnesota	Wisconsin
Region 3: South	
Alabama	Mississippi
Arkansas	North Carolina
Delaware	Oklahoma
Florida	South Carolina
Georgia	Tennessee
Kentucky	Texas
Louisiana	Virginia
Maryland	West Virginia
Region 4: West	
Alaska	Montana
Arizona	Nevada
California	New Mexico
Colorado	Oregon
Hawaii	Utah
Idaho	Washington
Wyoming	

Dummy variables are established for each cross-sectional category. The model is estimated by including data for all four categories but omitting the intercept commonly cited as β_0 in most models. This avoids the "dummy trap" which leads to perfect multicollinearity. The model is specified as

$$y_{it} = \Sigma\beta_{it}X_{it} + \Sigma\beta_i + \varepsilon_{it} \quad (3)$$

where X_{it} are the explanatory variables and β_i refers to the cross-sectional category under examination. As Equation (3) shows, in the absence of β_0, the estimated coefficient for each dummy variable provides a different intercept for each category.

The resulting intercepts allow the model to reflect differences among the different categories.

The fixed-effects model permits the distinct advantage of allowing all data to be used in the regression rather than just those just pertaining to a specific category as is the case with seemingly unrelated regression models (SUR). This permits a larger number of degrees of freedom and is thus likely to be more accurate.

Further, the dummy coefficients for the SUR models can avoid multicollinearity by excluding one of the dummy variables. The ensuing coefficients for the remaining dummy variables represent the change in the intercept when the category is compared to the omitted dummy variable. Since the purpose of this experiment is to capture differences among four census bureau regions, it seems a fixed-effects model is more appropriate. The intent of the fixed effects model is to determine if the intercepts are the same for all four regions. If they are, a fixed-effects model is unnecessary. The determination is based on a hypothesis test framed as

$$Ho: \beta_{North} = \beta_{Midwest} = \beta_{South} = \beta_{West}$$

$$Ha: \text{Not all } \beta_i \text{ are Equal}$$

The appropriate methodology requires an F-test containing values derived from the results of two regression models: a restricted model and unrestricted model. The restricted model does not include a dummy variable for region and contains the constant term, β_0. It is expressed as

$$Y = \beta_0 + \Sigma\beta_{it}X_{it} + \mathcal{E}_{it} \tag{4}$$

By excluding any reference to the different regions and including a single value for β_0, it restricts the four intercepts to equality. The unrestricted model is the fixed-effects model seen as Equation (3). The F-test is calculated as

$$F = \frac{\frac{RSS_R - RSS_U}{q}}{\frac{RSS_U}{n-k-1}} \tag{5}$$

where RSS_R is the residual sum of squares for the restricted model and RSS_U is the residual sum of squares for the unrestricted model. q is the number of restrictions contained in the null hypothesis and equals the number of parametric coefficients set equal to each other. n is the number of observations and k is the number of right-hand side variables in the unrestricted model. Expressed in this manner, the F-statistic measures any improvement in the fit offered by the unrestricted model over that reported by the restricted form.

RSS_R will always be more than RSS_U because some of the variables in the restricted model are constrained and cannot fit the data as well as the unconstrained model. Furthermore, the unrestrained model contains more explanatory variables and will offer a better fit. Consequently, the F-statistic is always positive.

In this present case, q is 4, n is 50 and k = 7. Computations produce an F-value of 3.41 and a p-value of 0.0167. The null hypothesis that the intercepts for all four regions are equal is rejected at the 1.67% level of significance. Clearly, the nature of the relationship between economic freedom and income distribution varies across state boundaries in the US.

Given the cross-sectional nature of the data set, White's test for heteroscedasticity (1980) as modified in Webster (2013) was conducted. Unlike other tests for heteroscedasticity, White's test as modified does not require that the variables proportionally associated with the heteroscedastic variances be identified. Instead, all right-hand side variables used in the fixed-effects specification, and several in their modified forms, are used in the detection method. For example, if the initial regression model carried k regressors, White's test requires the application of Equation (6). The squared residuals from the restricted model are used as the regressand. They are regressed on a battery of potential explanatory variables.

$$\begin{aligned} e^2 = \beta_o + \beta_i X_{it} + \beta_i Y_{it} + \cdots + \beta_i k_{it} + \beta_i X_{it}^2 + \beta_i Y_{it}^2 \\ + \cdots + \beta_i k_{it}^2 + \beta_i X_{it} Y_{it} + \beta_i X_{it} k_{it} + \beta_i Y_{it} k_{it} + \varpi \end{aligned} \tag{6}$$

Four explanatory variables are used in the restricted model: $Gini_{2000}$, EFI_{2000}, EFI_{2013} and the change in the Gini ratio from 2000 to 2013. There are therefore 14 regressors in Equation (6). Testing for heteroscedasticity via this method requires the computation of nR^2 where R^2 is taken from Equation (6). This statistic fits a χ^2 distribution that is used to test the hypothesis that all β_i in Equation (6) are zero. If the null hypothesis is not rejected, it may be concluded that constant error variance prevails and the model is not plagued by heteroscadasticity. The R^2 from Equation (6) is 0.206. With n = 50, $nR^2 = 10.32$. Testing the hypothesis at the 5% level of significance, the critical $\chi_{.05,14} = 23.68$. Since 10.32 < 23.68, the hypothesis of constant error variance is not rejected. It would appear that the model does not suffer from heteroscadasticity. The Durbin-Watson reported as 1.89 indicating no obstruction due to autocorrelation.

TABLE 5. FIXED-EFFECTS REGRESSION RESULTS; CHANGE IN GINI AS REGRESSAND

Results of the Fixed-Effects Model With $Gini_{2013}$ as Response Variable				
Variable	Coefficient	Standard Error	t-value	p-value
$Gini_{2000}$	0.2079	0.0407	5.11	0.000
EFI_{2013}	-0.0048	0.0026	-1.78	0.082
EFI_{2000}	-0.0230	0.0062	-3.73	0.001
Change in EFI	-0.0058	0.0012	-4.83	0.000
NORTHEAST	-0.0557			
MIDWEST	-0.5599			
SOUTH	-0.6190			
WEST	-0.06190			

The results shown in Table 5 are clear and unequivocal. The coefficient for EFI for 2013 is negative and significant. This reveals that states with greater economic freedom in 2013 are going to exhibit greater income equality. This negative correlation attests to the impact economic freedom has on the distribution of income in any geographical unit.

The same interpretation can be applied to the EFI for the year 2000. Here again we find a significant, negative relationship. States displaying pronounced economic freedom at the outset of the time period under examination will report less income inequality later on. It may be concluded that economic freedom as provided by the Fraser Institute promotes greater income equality as measured by the Gini coefficient. The initial Gini ratio is highly significant and carries a positive sign. This indicates that states exhibiting a high degree of income inequality in the year 2000 will continue to experience this condition over time.

TESTS FOR CAUSALITY

A reasonable question focuses on any causal relationship that might exist between income distribution and economic freedom. Therefore, a Granger causality test is offered to test for directional causality between these two socio-economic variables. This test is performed to determine if any past values of one variable may affect present values of a second variable. That is, to address the question as to whether past values of X serve as explanatory variables of Y. Y is then regressed on past (lagged) values of itself as well as lagged values of X. The question as to how many lagged values of each variable should be included in the model is paramount.

It is imperative that the data used in the analysis prove to be stationary and do not exhibit a unit root. Initially, the test for a unit root is based on the Equation (7) such that

$$Y_t = \rho Y_{t-1} + \mu_t \qquad (7)$$

where it is assumed $-1 \leq \rho \leq +1$ and μ_t is merely white noise. If $\rho = 1$, (7) is a nonstationary random walk or stochastic process without drift and cannot be effectively estimated.

It would seem reasonable to then simply test the null hypothesis that $\rho =1$ against the alternative hypothesis that $\rho < 1$. This is written as a one-sided test because if $\rho > 1$, the series is said to be explosive and thus difficult to model. Besides, explosive series of this nature are quite uncommon using economic data due to the cyclical nature of economic activity. If the null is not rejected the series carries a unit root and is nonstationary. That is, if $\rho = 1$, it may be concluded that a unit root prevails and a random walk model without drift results thereby evidencing a nonstationary condition.

However, such a test is not possible since if a unit root exists the t-tests produce biased results. Specifically, the t-values for the coefficient of Y_{t-1} do not follow an asymptotic normal distribution even in the presence of large samples.

It is therefore necessary to find an alternative approach to test for unit root. Perhaps the most common approach is that provided by Dickey-Fuller (1979). The Dickey-Fuller test (DF) is applied by subtracting Y_{t-1} from both sides yielding Equation (8).

$$Y_t - Y_{t-1} = \rho Y_t - Y_{t-1} + \mu_t$$

$$Y_t - Y_{t-1} = (\rho - 1)Y_{t-1} + \mu_t \qquad (8)$$

If $Y_t - Y_{t-1}$ is shown as ΔY_t and δ is set equal to $(\varrho - 1)$, we have

$$\Delta Y_{t-1} = \delta Y_{t-1} + \mu_t \qquad (9)$$

The equivalent test then carries a null hypothesis of $\delta = 0$ with the alternative stated as $\delta < 0$. Since $\delta = (\varrho - 1)$, for $\varrho < 1$ and avoid the unit root, δ must be less than zero. If the null that $\delta = 0$ is not rejected, $\varrho = 1$ and nonstationarity in the series exists making it difficult to conduct any causality test. The t-value used in this hypothesis test is then compared not to the standard t-distribution, but that provided by Fuller (1976) used exclusively for this particular test. Appling Equation (9) to both income inequality and EFI failed to lead to a rejection of the null hypotheses. Thus, it is concluded both suffer unit roots requiring corrective action.

The most common practice is to first-difference the series. Generally, the first differences of time-series are stationary. To determine this, the first differences are taken for both variables and then Equation (9) is applied to those first differences. Both tests for income inequality and economic freedom suggest stationarity.

The Granger causality test can then be applied to these first differences. The complete Granger test for Y and X, for example, involves the comparison of two regressions models, the restricted model and the unrestricted model. To test whether X Granger-causes Y, the restricted model is expressed as

$$\Delta Y_t = \alpha_t + \sum_{i=1}^{p} \alpha_i Y_{t-i} + \mu_t \qquad (10)$$

This is referred to as the restricted model because it contains no reference to the X variable. The coefficients for X are assumed restricted to zero and are therefore held out of the model.

In the absence of more sophisticated computer software that will identify the optimum number of lags, it is advisable to begin with a large number of lags and test for significance in succession starting with the oldest values (largest lags).

The unrestricted model is

$$\Delta Y = \phi_t + \sum_{i=1}^{p} \phi_i Y_{t-1} + \sum_{i=0}^{q} \beta_i X_{t-i} + \nu_t \qquad (11)$$

Here the coefficients for X are not assumed to be zero. If that is so, then past values of X Granger-cause changes in Y.

The null hypothesis is that past values of X do not Granger-cause changes in Y. The null is expressed as Ho: $\beta_i = 0$. If the null is not rejected, it may be concluded that X does not Granger-cause changes in Y.

To test for causality from Y to X, Equations (10) and (11) are reversed setting X as the dependent variable. The same hypothesis test is then performed on this second set of equations.

Final determination is based on the standard Wald F-test that offers a comparison between the restricted and unrestricted estimates as shown by Equation (12).

$$F = \frac{\frac{SSE_R - SSE_U}{q}}{\frac{SSE_U}{n-p-q}} \qquad (12)$$

where SSE_R and SSE_U are the sums of the squared residuals for the restricted and the unrestricted model, n is the number of observations, p and q are in the number of lags in the dependent and independent variables, respectively. The result is compared to a critical F-value with q and n-p-q degrees of freedom.

Four possible outcomes may result. It may be determined that Y Granger-causes X, X Granger-causes Y, there is bilateral causation in that each Granger-causes the other or the two variables may be independent in which neither Granger-causes the other.

After considerable tests to detect significance in lagged values based on the current data set, it was determined that a lag order of three was appropriate for both variables in both tests. A test was first conducted to determine if income inequality might Granger-cause measured levels of economic liberty to vary. That is, the EFI was used as the regressand for both the restricted and unrestricted models. An F-value of 1.14 was reported. This value is below any acceptable critical F-value for this test and carried a p-value of 0.389. It may be concluded that the null hypothesis that income distribution does not Granger-cause economic freedom cannot be rejected. It would appear that there is no directional causality running from income inequality to the freedom index. Past measures of income distribution do not offer any explanatory value for changes in the levels of economic freedom.

Treating income inequality as the dependent variable and using the EFI as the regressor yielded an F-value of 7.25. The associated p-value is 0.0114 allowing a rejection of the null hypothesis that economic freedom does not Granger-cause changes in income distribution. It may be concluded that past levels of economic freedom Granger-cause changes in the degree of income distribution as measured by the Gini coefficient. The significant coefficients for lagged values of the EFI all report negative signs and thereby contend that prevailing measures of economic freedom will in the near future lead to greater economic liberty within the populace.

CONCLUSION

The results presented here suggest that elevated measures of economic freedom are associated with more equal distributions of income. This pattern holds true at specific points in time as well as dynamics measured over the past decade. After controlling for other factors, the Gini ratio for the year 2000 exhibits a negative relationship with the EFI for the same year. Given that increases in the Gini coefficient evidence lower levels of income equality, it may be concluded that in regions with greater measures of economic freedom, a drop in the Gini suggests greater income equality. Incomes become more evenly distributed in areas where economic participants are given greater discretion in their activities in the prevailing economic markets.

These findings seem to run contrary to those reported by Ashby and Sobel (2008). Their paper did not find that measures of economic freedom were negatively related to increased levels of inequality. Perhaps this discord results from the different time periods used in the two studies.

Virtually identical results were found when more recent measures were used in the OLS models. Regressing the Gini coefficient from 2013 on the EFI for the same year suggested again that economic freedom was associated with a more equitable distribution of incomes across the states. These results also conflict with those presented by Ashby and Sobel (2008) but tend to support the findings based on studies comparing different nations around the globe as noted above.

With respect to an examination of the effect of changes in income dispersion over time, we find much a greater degree of accord. This paper demonstrates that changes in the levels of income equality are significantly associated with changes in the degree of economic freedom. Specifically, the change in the Gini over the time span of 2000 to 2013 reported a negative relationship with both the EFI in 2000 as well as EFI_{2013}. Further, the change in that index over the same time period reported a negative relationship with the change in the EFI. This reveals that greater equality can be nurtured by efforts to foster enhanced economic freedom within the economic climate. This conclusion does tend to correspond with that provided by Ashby and Sobel (2008), as well as those presented by Berggren (1999) and Barro (2000).

The fixed-effects model clearly indicated the nature of the relationship between economic freedom and income distribution varies across state boundaries in the US.

Differences among the four US Census Bureau regions in terms of the industrial mix, general economic climate, public policies and the host of other socio-economic factors might explain these differences. Further research should be conducted to determine what attributes and qualities characteristic of the regions are most conducive to promoting this relationship between economic freedom and a more equitable distribution of income.

Certainly, before any firm theoretical foundations can be formulated, additional research into the pressing issue is necessary. Despite the recognized trend in income distribution in the US over the past several decades, it is only recently that concern has been raised as to the consequences such patterned changes might have on our social, economic and political fabric.

It seems self-evident that the current drift in the pattern of income distribution cannot and should not be sustained. The consequences could lead to a regrettable outcome. If effective public policy is to be enacted, a greater understanding of the interaction and dynamics among the forces discussed in this paper is essential.

After correction for non-stationarity using the Dickey-Fuller test and applying first-differencing, the Granger-causality procedure was applied to the data to detect the whether there was any feedback between income measures and the existing degrees of economic liberty as measured by the Economic Freedom Index. It was clearly determined that while earlier measures of income distribution could did not appear to Granger-cause and changes in economic freedom. However, such was not the case for the opposing test. Past measures of the Economic Freedom Index proved to be high significant as explanatory factors for changes in income distribution.

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