Revisiting Macroeconomic Activity and Income Distribution in the United States

John A. Bishop, Lester A. Zeager, and Yizhen Zhao

Department of Economics
East Carolina University
Greenville, NC 27858

Abstract

February 26, 2016

Revisiting Macroeconomic Activity and Income Distribution in the United States

I. Introduction

Interest in the distributional impacts of modern macroeconomic policies can be traced back at least as far as an exchange between Arthur Burns and James Tobin. Burns contended that inflation harmed the poor the most[1] and the claim became widely known as "inflation is the cruelest tax." Tobin (1972) countered that there was no evidence to support this claim and noted that unemployment has both "distributional effects as well as deadweight losses."

Following this exchange, Blinder and Esaki (1978, 604) sought to "address this issue directly and quantitatively" by exploring the influence of the unemployment and inflation rates on the size distribution of income among U.S. families using data for 1947-1974. Later, Jantti (1994) and Bishop, *et al.* (1994) reexamined these relationships using data through 1989. They also expanded the analysis to include budget deficits and trade policies, both of which moved to the forefront of macroeconomic policy in the late 1980's. Today we have access to 21 years of data that were not available to those researchers, covering the deepest recession of the postwar period. This paper revisits the influence of macroeconomic forces on U.S. income inequality using these data.

The studies by Blinder and Esaki (1978) and Jantti (1994) measured inequality using *income shares* by quintile, but this formulation conflates two dimensions of the distribution of incomes that we want to distinguish, their level and dispersion. As the latter is our main interest, we measure inequality using Lorenz ordinates by quintile, from which we can infer the effects of macroeconomic variables on the U.S. Lorenz curve using the powerful dominance methods of

[1] From the *Tax Review*, as quoted in Palmer (1973) and Blinder and Esaki (1978).

Atkinson (1970). To clarify some results, we also explore the effects on income levels using quintile conditional means, which allow inferences about first-order dominance.

All of the earlier studies consider other influences – beyond the unemployment and inflation rates – on the distribution of family incomes, if only a demographic or major policy change. We follow this precedent by including other macro policy variables: the budget deficit, public transfers, and an index of openness to world markets. The last two variables reach their highest levels in the additional years covered in this study.

The paper is organized as follows. Section II explains the model and the estimation methods and gives a brief description of our data sample. Section III presents our estimation results and compares them to the findings from earlier studies. In Section IV we discuss of the most striking changes in the findings when we use the larger sample. Section V summarizes our main conclusions.

II. The Model, Estimation Methods, and Data

In this section we discuss our estimation procedure. Blinder and Esaki (1978) select ordinary least squares (OLS). While aware of its potential problems, the evidence available to them suggested that OLS was adequate for their purpose. In his follow-up analysis, Jäntti (1994) opted for a feasible generalized least squares method, primarily for improved efficiency. For the sample period we consider, results from variation ratio test (Appendix B, Tables B1d and B2d) show that most data series not only have autocorrelated and heteroskedastic error terms, but also are mean-averting and cointegrated (Appendices C and D). Hence, we adopt the Engle-Granger (1987) two-step procedure and the seemingly unrelated regression (SUR) technique to improve efficiency. As SUR permits error terms to be correlated among different income quintiles, the shocks to household incomes can impact low- and high-income households in smaller or larger magnitudes. The choice of SUR also makes our results comparable with those of Bishop, *et al* (1994).

Our estimation procedure involves two steps. The first stage uses ordinary least squares to estimate the equation,

$$L_{jt} = a_{j0} + \sum_{\mathrm{i}=1}^{\mathrm{N}} a_{ji} X_{it} + e_{jt},$$

where L_{jt} is the jth quintile Lorenz ordinate in year t, X_{it} is the ith explanatory variable in year t, and e_{jt} is an error term. Our explanatory variables include the unemployment rate, Consumer Price Index, a time trend ($t = 1,2, \ldots, m$), public transfers, federal budget deficit, and an index of openness to world markets. The coefficients $a_{jo} \cdots a_{jN}$ in equation (1) provide estimates of the ong-run equilibrium relationship between the income distribution variable and the explanatory variables. The residual term, e_{jt}, measures the divergence from equilibrium of the dependent variable, L_{jt}, and will be used in the second-stage regression.

The second-stage equation involves first-order differences of the dependent and explanatory variables, with the residuals (e_{jt}) from the first-stage regression included as a correction for autocorrelation. It takes the form

$$\Delta L_{jt} = b_{j0} + \sum_{\mathrm{i}=1}^{\mathrm{N}} b_{ji} \Delta X_{it} + \; b_{jN+1} e_{jt-1} + u_{jt},$$

where ΔL_{jt} is the first-order-difference of the jth quintile Lorenz ordinate in year t, ΔX_{it} is the first-order difference of ith explanatory variable in year t, e_{jt-1} is the residual from the jth regression in equation (1) with a one period lag, and u_{jt} is a random disturbance.

The estimated coefficients b_{ji} and b_{jN+1} provide information on the impacts of the explanatory variables on the income distribution. Any deviations from long-run equilibrium triggered by changes in the explanatory variables require adjustments; the estimated b_{ji} reveal how the deviations are eliminated. To achieve a new long-run equilibrium, the $b_{ji} (i = 0 \dots N + 1)$ must have a negative sign. For example, a positive e_{jt} implies that L_{jt} is above the long-run equilibrium and must decrease to reach the new equilibrium (ΔL_{jt} will be negative). Moreover, if the absolute value of b_{ji} is less than unity, then the variable L_j converges monotonically. The absolute value of b_{ji} also indicates the speed of convergence; the closer to unity, the faster is the convergence.

For the estimation, we use time series data from 1950 through 2010, the first and last years for which all of our variables are available. For comparisons with earlier studies, we also estimate the model using a restricted sample (1950-89).[2] Table 1 gives summary statistics for the explanatory variables. The unemployment rate has a slightly higher mean in the full sample than in the restricted sample and reaches its peak during the Great Recession (2009). Inflation, which we measure by annual changes in the Consumer Price Index in the urban areas (CPI-U), also has a higher mean in the full sample, but reaches its peak (1980) in the restricted sample. Expressed as a percentage of GDP, public transfers have a higher mean in the full sample and grow to their highest level in the last year of the full sample (2010). The largest federal budget deficits, given as a percentage of GDP, occur in the full sample, but the means are identical in the two samples. The most striking difference between the samples is in the degree of openness to world markets (defined as exports plus imports, expressed as a percentage of GDP), which is far higher in the full sample in both the maximum and mean values. See Appendix A for the definitions of and the sources for our variables.

[2] The samples in Bishop, *et al.* (1994) and Jäntti (1994) begin in 1947 and 1948, respectively, and end in 1989.

III. Estimation Results

We begin with tests for unit roots (random walks) and cointegration of our variables. Following Hayes *et al.* (1990), who find that the quintile income shares for the U.S. follow a random walk from the late 1940s to the early 1980s, we test for random walks in both dependent variables – Lorenz ordinates and quintile conditional means – and the five explanatory variables: unemployment, inflation, public transfers, budget deficits, and the openness index. The purpose of this test, together with the cointegration test, is to determine whether there is some long-run equilibrium relation tying the dynamics of these variables together (cointegration exists).

Using the full data sample (1950-2010), we apply four different tests: (i) Dicky and Fuller (1979) test, (ii) Phillips-Perron (1988) test, (iii) Kwiatkowski, Phillips, Schmidt, and Shin (1992) test, and (iv) Lo and MacKinlay (1988) variance ratio test to determine if the variable of interest is a unit-root process.[3] For each test, we consider two or three possible data-generating mechanisms. Appendix Tables B1-B3 present the specifications. Appendix B reports and discusses the detailed results of the random walk tests.

The results of the random walk tests can be summarized as follows. First, the estimates of the autoregressive terms for all Lorenz ordinates and quintile mean incomes are very close to unity, which is consistent with random walks in the income distribution across time. For all the data generating mechanisms listed in Table B1, B2 and B3, the test statistics for all the variables are either positive (the series is mean-averting) or the test cannot reject the null at conventional significance levels in most cases. Therefore, the null hypothesis of a random walk cannot be rejected.

[3] Bishop, Formby and Sakano (1994) consider only the first two tests.

Given the finding that both the dependent and independent variables follow a random walk, we test whether they are jointly cointegrated. We use two tests to identify a cointegration vector: (a) trace test (Johansen (1988)) and (b) maximum eigenvalue test (Johansen and Juselius (1990)). Here we restrict our attention to only five of the explanatory variables in each test. The Johansen-Juselius (1990) method only provides critical values up to the five-variable case, as it is difficult and misleading to identify a large number of cointegrating vectors from just 60 data points. Finally, the explanatory variables are highly likely to have a long-run relationship with dependent variable over a 60 year period.

We report the results in Appendix C. In brief, the tests reveal that, in almost every case, we fail to reject the null hypothesis of no cointegration. Therefore, it is highly likely that the explanatory variables and the income distribution variables are cointegrated, which affirms a long-run equilibrium relationship among these variables. The tests also indicate that there are at least two cointegrating vectors in most cases. These findings are consistent with the findings of Bishop, *et al* (1994) for sample period from 1947 to 1989. The existence of cointegration also implies that – in the sense of Granger (1969) – a causality relationship is present in the model. Therefore, we conclude that the explanatory variables interact with one another in a general equilibrium sense, while at the same time systematically causing changes in the income distribution across time.

We turn next to the results from the second-stage estimates of the cointegration equations. The first-stage estimates from equation (1) are shown in Appendix D. Table 2 presents the parameter estimates and p-values for equation (2), where the dependent variables (ΔL_{jt}) are first differences in quintile Lorenz ordinates. In each quintile, the sign of the lagged error correction is negative and significant, so the Lorenz ordinates converge monotonically to a long-run equilibrium. From the absolute values of the coefficients of the lagged error correction (ranging from 0.46 to 0.64), we can see that the speed of convergence is moderate. The Durbin-Watson statistics in Table 2 indicate that the residuals (u_{jt}) in equation (2) are stationary, which indicates that the error-correction term eliminates the autocorrelation.

Table 2 indicates that a positive change in the unemployment rate has an ambiguous effect on income inequality. We obtain negative and significant coefficients in the bottom four quintiles, but a positive and significant coefficient in the top quintile, yielding a Lorenz crossing. Likewise, a positive change in the CPI has an ambiguous effect on inequality. Here we obtain a negative and significant coefficient in the bottom quintile, but positive and significant coefficients in the top four quintiles.

Three of the control variables (the time trend, budget deficits, and the openness index) have unambiguous effects on income inequality. The time trend reduces Lorenz ordinates in all quintiles, implying greater income inequality. Larger budget deficits increase inequality, while a greater openness index has the opposite effect – raising Lorenz ordinates in all quintiles, and thus reducing inequality. In contrast, larger public transfers have an ambiguous effect on inequality, because the coefficient is positive and significant in the bottom quintile, but negative and significant in the top three quintiles.

Table 3 provides a summary of the signs and significant coefficients in the error-correction model, estimated with data for 1950-89 and 1950-2010, with and without control variables in each sample (including a time trend, not reported in Table 3). The former period is similar to those used in the studies by Jäntti (1994) and Bishop, *et al.* (1994). As in Bishop *et al.* (1994, Table 4), a higher unemployment rate reduces the Lorenz ordinates in the bottom quintiles while a higher inflation rate raises the Lorenz ordinates in all quintiles in the shorter sample, with the control variables included. For both variables, the effect on inequality is unambiguous, but in opposite directions. The 1950-89 results without the control variables in Table 3 are closer to the specification of Jäntii (1994), although his dependent variable is the quintile income shares rather than the Lorenz ordinates. Still, in the bottom quintile the income share and Lorenz ordinate are identical, and there he obtains a negative and significant coefficient for the unemployment rate and a positive and significant coefficient for the inflation rate, as we do.

Two control variables, budget deficits (reflecting the means of financing government spending) and public transfers (capturing redistributive policies) have similar effects to those found in Bishop, *et al.* (1994). Larger budget deficits significantly reduce the Lorenz ordinates almost everywhere (except the top quintile, where we find the opposite). Larger public transfers significantly increase the Lorenz ordinates at some or all quintiles and therefore, *increase* income inequality unambiguously. Our remaining control variable, the openness index (not included in any of the previous studies), significantly reduces the Lorenz ordinates in all quintiles, and thus increases income inequality in the 1950-89 sample.

When we use the full sample (1950-2012), our findings change. Whether or not we include the control variables, a higher unemployment rate generates a Lorenz crossing (lower ordinates in the bottom quintiles, but a higher ordinate at the top) instead of Lorenz dominance (as in the earlier period). Likewise, a higher inflation rate generates a Lorenz crossing (a lower ordinate at the bottom quintile, but higher ordinates at the top), instead of Lorenz dominance as before. Thus, the unemployment and inflation rates have ambiguous effects on the distribution of income in the full sample.

In the extension to the full sample, we also find striking changes for two of the control variables. Public transfers generate a Lorenz crossing (a higher ordinate in the bottom quintile, but lower ordinates in the top quintiles), instead of Lorenz dominance as before. Budget deficits reduce Lorenz ordinates as before, but now in all quintiles. Most striking of all, greater openness to world markets raises (instead of lowering, as before) the Lorenz ordinates in each quintile, and therefore *reduces* income inequality.

Table 4 presents further comparisons with the prior literature as well as extensions of those findings to the full sample. The summary results for 1950-89 are comparable to those in Bishop, et al. (1994, Table 4, column 1), where the dependent variables are conditional quintile means in first differences. For the two key macro variables (unemployment and inflation rates), the two studies have similar findings. An increase in the unemployment rate reduces all or some conditional means (and if some, raises no others), which implies a first-order dominance relation. An increase in the CPI, however, raises one or more conditional means and reduces none, which implies the opposite first-order dominance relation.

The control variables differ across the studies (duties versus the openness index and inclusion of demographic controls) and here some differences in results emerge. Bishop *et al.* (1994) find that an increase in public transfers has no significant effect on the conditional means, but we find significant increases in the second and third quintiles for 1950-89. Both studies find that larger budget deficits increase the conditional mean incomes in the lower deciles. However, the studies differ on the impact of the foreign sector. Whereas they report that higher duties on foreign trade (which reduce the volume of trade) raise the conditional mean in the top quintile, we find that greater openness *reduces* the conditional means in the lower quintiles during the same period.

When we extend the analysis to the larger sample (1950-2010) in Table 4, some interesting changes emerge. Most importantly, increases in *either* unemployment or inflation rates reduce the conditional mean income in the bottom quintile, in contrast to the earlier period. That is, both unemployment and inflation are harmful to the poor. Furthermore, budget deficits *reduce* the conditional means (in the three middle quintiles), whereas they raised them (in the three bottom quintiles) in the earlier period.

IV. Discussion

In this section we suggest possible interpretations for two striking differences in the findings when we extend the sample to 2010. The most striking difference in Table 3 is that greater openness to world markets *increases* income inequality in all quintiles in 1950-89, but *reduces* inequality in all quintiles when we extend the sample to 2010, the years of the greatest openness to international trade. The second difference that stands out in Table 3 is the change in the impact of pubic transfers on income inequality. Instead of raising the Lorenz ordinates in all quintiles, and thus reducing income inequality unambiguously (as in 1950-89), public transfers raise the Lorenz ordinate in the bottom quintile, but *lower* the Lorenz ordinates in the middle quintiles. We discuss each of these changes in turn.

To understand the change in the effect of openness to world markets, notice first that greater openness reduces income *levels* Table 4 – in the lower quintiles in 1950-89 and also in the top quintile in the larger sample. Hence, the change is not in the direction of the impact on incomes. Also, falling incomes do not imply that greater openness reduces well-being, because greater competition also reduces prices. They do imply, however, that incomes decline more at the top of the distribution than at the bottom, so that inequality diminishes. We cannot pinpoint the reasons for this pattern, but we know that the composition of foreign trade is clearly changing as its volume increases. The Economic Report of the President (2015, 74, Figure 7-4) shows that real imports and exports of services have risen by more than 400 and 600 percent, respectively, since 1980. Included in the emergence of services trade is "outsourcing," which has affected skilled workers who once seemed immune from the threat of foreign competition.

With respect to public transfers, there have been substantial policy shifts since 1989, notably the welfare reforms during the Clinton Administration (greater reliance on the earned income tax credit) and the expansion of social welfare benefits (unemployment insurance, food stamps) during the Great Recession by the Obama Administration. It appears from our results in the full sample that U.S. public transfers have become more focused on the bottom quintile. This shift reminds us of some perceptive reasoning by Tullock (1971), articulated well before the shift took place. Drawing on the seminal contributions of Downs (1957) and Riker (1962), and on his collaboration with James Buchanan (Buchanan and Tullock, 1962), Tullock (1971, 382) reaches the familiar conclusion about the effects of political redistribution, "The reasoning so far would indicate that the people toward the top of the bottom 51 percent might receive much more than the people at the lower end." In the next sentence, however, he envisions another possibility:

> The only restriction on the delivery of the bulk of the resources transferred from the wealthy to the upper end of the bottom coalition (other than charitable instincts on the part of the members of the upper end) would seem to be the possibility that the wealthy would attempt a coalition with the very poor.

He sees some signs of attempts to form such a coalition, e.g., calls for strict means-testing of transfers, mostly voiced by wealthy persons and entirely consistent with their self-interest, as they could substantially increase transfers to the poor, eliminate transfers to the middle, and realize a "profit" for themselves in the switch. He speculates further that:

> This particular coalition has so far foundered largely because of miscalculations by the poor. The poor realize that the interests of the wealthy are clearly not

> congruent with their interests, but they do not realize that the interests of people between the twentieth and fifty-first percentile of the income distribution are also not identical to theirs. They therefore tend to favor a coalition with the second group rather than the former.

Since that time, a politically potent coalition between the rich – especially the newly rich prospering from the rise of the financial and information technology industries – and the poor has arisen. It is most clearly evident in the elections of Barak Obama, whose campaigns raised impressive amounts of money from the new rich and dramatically increased voter turnout among persons with low incomes – especially African-Americans and the young. Indeed, the recent rise of "outsider" candidates might be fueled by the reaction of the angry middle, which has lost political power to the new coalition.

V. Conclusion

We re-examine the influence of the inflation and unemployment rates on the size distribution of income among U.S. families using 21 years of additional data not available in previous studies, including the deepest recession since World War II. We control for, among other things, changes in openness to the world economy and in public transfers, both of which reached their highest levels in the years added to the sample in this study.

Using the larger sample alters some of the findings of previous studies. A higher unemployment rate creates a Lorenz crossing (lower ordinates in the bottom quintiles, but a higher ordinate at the top), rather than simply increasing income inequality (lower ordinates at the bottom). Likewise, a higher inflation rate creates a Lorenz crossing (a lower ordinate at the bottom, but higher ordinates at the top), rather than unambiguously reducing inequality. Greater openness to the world economy unambiguously reduces income inequality in the larger sample, but has the opposite effect in the smaller sample. Public transfers have an ambiguous effect on inequality in the larger sample, but unambiguously reduce inequality in the smaller one. We suggest possible reasons for the changes in the effects of public transfers and openness to international trade.

REFERENCES

Atkinson, Anthony, "On the Measurement of Inequality," *Journal of Economic Theory,* 2 (1970), 244-263.

Bishop, John A., John P.Formby, and Ryoichi Sakano, "Evaluating Changes in the Distribution of Income in the United States," *Journal of Income Distribution,* 4 (1 1994), 79-105.

Blinder, Alan S. and Howard Y. Esaki, "Macroeconomic Activity and Income Distribution in the Postwar United States," *The Review of Economics and Statistics,* 60 (Nov. 1978), 604-609.

Buchanan, James M. and Gordon Tullock, *The Calculus of Consent,* Ann Arbor: University of Michigan Press, 1962.

Dickey, D. A. and W.A. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root". *Journal of the American Statistical Association* 74 (366 1979): 427–431.

Downs, Anthony, *An Economic Theory of Democracy,* New York: Harper & Row, 1957.

Economic Report of the President, 2015.

Hayes, K. J., S. Slottje, S. Porter-Hudak, and G. Scully,. "Is the Size Distribution a Random Walk", *Journal of Econometrics*, 43 (1990), 213-26.

Jäntti, Markus, "A More Efficient Estimate of the Effects of Macroeconomic Activity on the Distribution of Income," *The Review of Economics and Statistics*, 76 (May 1994), 372-378.

Johansen, S., "Statistical Analysis of Cointegration Vectors", *Journal of Economic Dynamic and Control*, 12 (1988), 231-54.

Johansen, S. and K. Juselius K, "Maximum Likelihood Estimation and Inference on Coingration – with Application to the Demand for Money", *Oxford Bulletin of Economics and Statistics,* 52 (1990): 169-210.

Kwiatkowski, D., P.C.B. Phillips, P. Schmidt, P.; and Y. Shin, "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root". *Journal of Econometrics*, 54 (1–3 1992): 159–178.

Lo, A. W.; and A. C. MacKinlay. "Stock Market Prices Do Not Follow Random Walks: Evidence from a Simple Specification Test." *Review of Financial Studies*. Vol. 1 (1988), 41–66.

Palmer, John L., *Inflation, Unemployment and Poverty,* Lexington, MA: D.C. Heath, 1973.

Phillips, P. C. B. and P. Perron, "Testing for a Unit Root in Time Series Regression," *Biometrika*, 75 (2 1988), 335–346.

Riker, William, *The Theory of Political Coalitions*, New Haven: Yale University Press, 1962.

Tobin, James, "Inflation and Unemployment," *American Economic Review,* 60 (May 1970), 261-269.

Tullock, Gordon, "The Charity of the Uncharitable," *Economic Inquiry,* 9 (December 1971), 379-392.

Table 1
Summary Statistics of the Explanatory Variables for Two Periods

Explanatory Variable	1950–1989			1950–2010		
	Minimum	Maximum	Mean	Minimum	Maximum	Mean
Unemployment Rate	2.80 (1953)	9.50 (1982)	5.58	2.80 (1953)	9.80 (2009)	5.68
Change in CPI	–0.07 (1955)	9.80 (1980)	2.56	–0.69 (2009)	9.80 (1980)	3.23
Public Transfers (% of GDP)	0.03 (1952)	0.13 (1983)	0.08	0.03 (1952)	0.14 (2010)	0.09
Budget Balance (% of GDP)	–0.03 (1950)	0.04 (1975)	0.01	–0.03 (1950)	0.10 (2009)	0.01
Openness Index*	5.92 (1953)	15.54 (1989)	9.67	5.92 (1953)	29.20 (2010)	14.35

Budget deficits are positive; negative numbers indicate a budget surplus
* (Exports + Imports) / GDP

Table 2A
1950–2010 Second-Stage Estimation Results by Quintile:
Error-Correction Model

Explanatory Variable	**Change in Lorenz Ordinates**				
	ΔL_1	ΔL_2	ΔL_3	ΔL_4	ΔL_5
Constant	0.04 (0.00)	0.03 (0.00)	0.05 (0.00)	0.08 (0.00)	0.08 (0.00)
Lagged Error Correction	–0.65 (0.00)	–0.48 (0.00)	–0.46 (0.00)	–0.47 (0.00)	–0.52 (0.00)
Change in Unemployment Rate	–0.10 (0.00)	–0.14 (0.00)	–0.13 (0.00)	–0.06 (0.00)	0.09 (0.00)
Change in CPI	–0.05 (0.00)	0.00 (0.50)	0.07 (0.01)	0.20 (0.00)	0.34 (0.00)
Time Trend	–0.00 (0.00)	–0.00 (0.00)	–0.01 (0.00)	–0.01 (0.00)	–0.01 (0.00)
Change in Public Transfers*	0.29 (0.00)	–0.02 (0.42)	–0.28 (0.05)	–0.46 (0.02)	–0.55 (0.01)
Change in Budget Deficit*	0.47 (0.00)	0.90 (0.00)	1.06 (0.00)	1.15 (0.00)	0.94 (0.00)
Change in Openness Index**	0.06 (0.07)	0.47 (0.00)	0.58 (0.00)	0.42 (0.01)	0.55 (0.00)
Test Statistics					
R-Square	0.52	0.45	0.36	0.28	0.24
Dep. Var. Mean	4.74	15.91	32.78	56.47	82.56
Durbin-Watson	1.58	1.60	1.80	1.88	1.97
F Statistics	8.04	6.00	4.13	2.94	2.38
Joint p-value	0.00	0.00	0.00	0.01	0.04

The numbers in parentheses are p-values for the estimated coefficients

*Expressed as a percentage of GDP
**Openness index = (Exports + Imports) / GDP

Table 2B
1950-2010 Estimation Results by Quintile:
Error Correction Model

Explanatory Variable	**Changes in Conditional Mean Income**					
	ΔR_1	ΔR_2	ΔR_2	ΔR_2	ΔR_2	Top 5 %
Constant	0.11 (0.00)	0.20 (0.00)	0.31 (0.00)	0.42 (0.00)	0.65 (0.00)	0.23 (0.00)
Lagged Error Correction	-0.37 (0.00)	-0.32 (0.00)	-0.32 (0.00)	-0.28 (0.00)	-0.11 (0.00)	-0.12 (0.00)
Change in Unemployment Rate	-0.03 (0.00)	0.03 (0.00)	0.08 (0.00)	0.12 (0.00)	0.11 (0.00)	-0.03 (0.03)
Change in CPI	-0.05 (0.00)	0.01 (0.27)	0.03 (0.07)	0.09 (0.00)	0.31 (0.00)	0.11 (0.01)
Time Trend	-0.00 (0.00)	-0.01 (0.00)	-0.01 (0.00)	-0.01 (0.00)	-0.01 (0.00)	-0.00 (0.00)
Change in Public Transfers*	-0.07 (0.07)	-0.56 (0.00)	-0.64 (0.00)	-0.68 (0.00)	0.99 (0.00)	1.23 (0.00)
Change in Budget Deficit*	-0.12 (0.00)	-0.39 (0.00)	-0.88 (0.00)	-1.28 (0.00)	-2.41 (0.00)	-1.28 (0.00)
Change in Openness Index**	-0.08 (0.01)	0.06 (0.18)	-0.35 (0.00)	-0.78 (0.00)	-1.26 (0.00)	-1.11 (0.00)
Test Statistics						
R-Square	0.44	0.39	0.35	0.29	0.16	0.11
Dep. Var. Mean	3.22	7.54	11.49	16.27	30.29	12.23
Durbin-Watson	1.83	1.85	1.89	1.68	1.22	1.59
F Statistics	5.79	4.85	3.96	3.07	1.40	0.90
Joint p-value	0.00	0.00	0.00	0.01	0.23	0.50

The numbers in parentheses are p-values for the estimated coefficients

*Expressed as a percentage of GDP
**Openness index = (Exports + Imports) / GDP

Table 3
Summary of Signs and Significant Coefficients in the Error-Correction Model for Two Periods
(Dependent variables: first differences of the quintile Lorenz ordinates)

Independent Variables (in first differences)	1950-89		1950-2010	
	No Control Variables	Control Variables	No Control Variables	Control Variables
Unemployment Rate	[1–4] , 5	[1–4]	[1–3] , 4–5	[1–4] , 5
Consumer Price Index	1–5	1–5	[1–2] , 4–5	[1] , 3–5
Public Transfers*		1–5		1 , [3–5]
Budget Deficit*		[1–4] , 5		[1–5]
Openness Index**		[1–5]		1–5

* Expressed as a percentage of GDP **Openness Index = (Exports + Imports)/GDP Additional variable is a time trend.
The significance level is 10 percent

Table 4
Summary of Signs and Significant Coefficients in the Error-Correction Model for Two Periods
(Dependent variables: first differences of the conditional quintile means)

Independent Variables (in first differences)	1950–1989 (with controls)	1950–2010 (with controls)
Unemployment Rate	[1–3] , 5	[1] , 2–5
Change in CPI	2–3 , [5]	[1] , 3–5
Budget Deficit*	[2–5]	[1–5]
Public Transfers*	1–3 , [5]	[1–4] , 5
Openness Index**	[1–5]	[1, 3–5]

*Expressed as a percentage of GDP **Openness Index = (Exports + Imports) / GDP Square brackets denote negative signs.
Quintiles included only if the p-value < 0.1 for the variable and the F-test for the equation has a p-value < 0.1. Additional control is a time trend. The significance level is 10 percent.

APPENDIX A
Definition and Measurement of Variables

Dependent Variables:

Rx: *x*th conditional mean income, where Rx is equal to the quintile income share (that is, share of total income times mean family income). We consider quintiles, *R1, R2, R3, R4*, and *R5*. The data are from *Current Population Reports*, Series P-60.

Lx: Lorenz ordinate, where *Lx* is equal to the cumulative share of family income. Incomes are ranked from lowest to highest and the ordinates are measured in percentage terms. We consider five Lorenz ordinates, *L1=.2, L2=.4, L3=.6, L4=.8*, and *L5=.95*, where *L5* is the combined share of family income of the bottom 95 per cent of families. Computed from *Current Population Reports*, Series P-60.

Independent Variables

Unemployment: Unemployment rate of all workers, measured in percentage terms.

Inflation: Inflation rate computed as a first difference of logarithm of GNP deflator, which is a ratio of nominal GNP to real GNP, measured in percentage terms.

Public Transfers: Ratio of government (federal and local) transfer payments (to persons) to nominal GNP, measured in percentage terms.

Budget Deficit: Ratio of government deficit, including both federal and local governments to nominal GNP, measured in percentage terms.

Openness: The Ratio of the country's total trade (the sum of exports plus imports) to the country's gross domestic product. (Source: Federal Reserve Economic Data).

APPENDIX B
Unit Root Tests and Random Walks

Table B1 shows the unit root test results for quintile conditional means. Table B2 shows the results for Lorenz ordinates. Table B3 shows the results for explanatory variables: unemployment, inflation, transfers, budget deficit, and openness index.

The first row in Table B1a-b, Table B2a-b, and Table B3a-b shows estimated autoregressive coefficients, where each variable is regressed on its own one-period lag (for example, $L1_t = \rho L1_{t-1} + \mu_t$). The rest rows report test statistics and the p-values from Dicky-Fuller Test (Table B1a, B2a and B3a) and Philips-Perron Test (Table B1a, B2a and B3a). In each test, we consider three different specifications of alternatives: (1) an AR(1) process with no drift and no trend, (2) an AR(1) process with a drift and (3) an AR(1) process with a drift and a time trend. Failure to reject the null indicates a unit root.

In addition, Table B1c and B2c report the results from Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) Test (1992). The test assesses the null hypothesis that a univariate time series is trend stationary against the alternative that it is a nonstationary unit root process. Table B1d and B2d report the results from Variation Ratio Test (1988). The null assumes the variable of interest is a random walk. When error terms e_t are not $i.i.d.$, the alternative is that e_t are correlated. When error terms are $i.i.d.$, the alternative is that e_t are either dependent or not identically distributed (for example, heteroscedastic). Furthermore, the first row in Table B1d and B2d report the estimated variation ratio. When ratios are asymptotically equal to one, the series to be tested is a random walk. When ratios are less than one, the series is mean-reverting. When ratios are greater than one, the series is mean-averting.

We also tested the quintile conditional means and Lorenz ordinates for the existence of possible $I(2)$ processes. We reject the unit root of $I(2)$for all variables, providing still more evidence that the conditional means and Lorenz ordinate measures of the income distribution follow random walks. Test results are available from the authors upon request.

TABLE B1
Unit Root Tests for Quintile Conditional Means, 1950-2010

	B1a Dicky-Fuller Test (1979)							
	No Drift, No Trend						**Critical Values**	
	R1	**R2**	**R3**	**R4**	**R5**	**Top 5%**	**5%**	**10%**
ρ	1.00	1.00	1.01	1.01	1.01	1.01	--	--
test stat	0.73	1.27	1.76	2.25	2.73	1.80	-1.95	-1.61
p-value	(0.87)	(0.95)	(0.98)	(0.99)	(1.00)	(0.98)	--	--
	With Drift						**Critical Values**	
ρ	0.95	0.94	0.95	0.96	0.98	0.98	--	--
test stat	-2.30	-3.12	-3.08	-2.66	-1.66	-1.07	-2.91	-2.59
p-value	(0.17)	(0.03)**	(0.03**)	(0.09*)	(0.45)	(0.70)	--	--
	With Time Trend						**Critical Values**	
ρ	0.99	0.99	0.99	1.00	1.01	0.95	--	--
test stat	-0.17	-0.35	-0.18	-0.06	0.17	-0.88	-3.49	-3.17
p-value	(0.99)	(0.99)	(0.99)	(0.99)	(1.00)	(0.95)	--	--
	B1b Philips-Perron Test (1988)							
	No Drift, No Trend						**Critical Values**	
ρ	1.00	1.00	1.01	1.01	1.01	1.01	--	--
test stat	0.73	1.27	1.76	2.25	2.73	1.80	-1.95	-1.61
p-value	(0.87)	(0.95)	(0.98)	(0.99)	(1.00)	(0.98)	--	--
	With Drift						**Critical Values**	
ρ	0.95	0.94	0.95	0.96	0.98	0.98	--	--
test stat	-2.30	-3.12	-3.08	-2.66	-1.66	-1.07	-2.91	-2.59
p-value	(0.17)	(0.03**)	(0.03**)	(0.09*)	(0.45)	(0.70)	--	--
	With Time Trend						**Critical Values**	
ρ	0.99	0.99	0.99	1.00	1.01	0.95	--	--
test stat	-0.17	-0.35	-0.18	-0.06	0.17	-0.88	-3.49	-3.17
p-value	0.99	0.99	0.99	0.99	1.00	0.95	--	--
	B1c Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) Test (1992)							
	No Deterministic Trend						**Critical Values**	
test stat	0.56	0.63	0.71	0.76	0.78	0.76	0.46	0.35
p-value	(0.03**)	(0.02**)	(0.01**)	(0.01**)	(0.01**)	(0.01**)	--	--
	With Deterministic Trend						**Critical Values**	
Test stat	0.20	0.20	0.19	0.18	0.08	0.09	0.15	0.12
p-value	(0.02**)	(0.02**)	(0.02**)	(0.02**)	(0.10**)	(0.10**)		
	B1d Variation Ratio Test (1988)							
	e_t is not $i.i.d.$						**Critical Values**	
ratio	1.16	1.23	1.18	1.26	1.46	1.24	--	--
test stat	1.52	1.81	1.34	2.38	3.23	1.55	1.96	1.64
p-value	0.13	(0.07*)	0.18	(0.02**)	(0.00**)	(0.12)	--	--
	e_t is $i.i.d.$						**Critical Values**	
ratio	1.16	1.23	1.18	1.26	1.46	1.24	--	--
test stat	1.26	1.77	1.43	2.05	3.63	1.85	1.96	1.64
p-value	(0.21)	(0.08*)	(0.15)	(0.04**)	(0.00**)	(0.06*)	--	--

TABLE B2
Unit Root Tests for Lorenz Ordinates, 1950-2010

	B2a Dicky-Fuller Test (1979)						
	No Drift, No Trend					**Critical Values**	
	L1	**L2**	**L3**	**L4**	**L5**	**5%**	**10%**
ρ	1.00	1.00	1.00	1.00	1.00	--	--
test stat	-1.04	-1.56	-1.51	-1.16	-0.56	-1.95	-1.61
p-value	(0.27)	(0.11)	(0.12)	(0.22)	(0.44)	--	--
	With Drift					**Critical Values**	
ρ	0.98	1.01	1.01	1.00	0.97	--	--
test stat	-0.53	0.20	0.33	0.09	-0.73	-2.91	-2.59
p-value	(0.88)	(0.97)	(0.98)	(0.96)	(0.83)	--	--
	With Time Trend					**Critical Values**	
ρ	0.92	0.89	0.89	0.89	0.88	--	--
test stat	-1.56	-2.10	-2.46	-2.53	-2.33	-3.49	-3.17
p-value	(0.80)	(0.54)	(0.37)	(0.33)	(0.43)	--	--
	B2b Philips-Perron Test (1988)						
	No Drift, No Trend					**Critical Values**	
ρ	1.00	1.00	1.00	1.00	1.00	--	--
test stat	-1.04	-1.56	-1.51	-1.61	-0.56	-1.95	-1.61
p-value	(0.27)	(0.11)	(0.12)	(0.22)	(0.44)	--	--
	With Drift					**Critical Values**	
ρ	0.98	1.01	1.01	1.00	0.97		
test stat	-0.53	0.20	0.33	0.09	-0.73	-2.91	-2.59
p-value	(0.88)	(0.97)	(0.98)	(0.96)	(0.83)		
	With Time Trend					**Critical Values**	
ρ	0.92	0.89	0.89	0.89	0.88		
test stat	-1.56	-2.10	-2.46	-2.53	-2.33	-3.49	-3.17
p-value	(0.80)	(0.54)	(0.37)	(0.33)	(0.43)		
	B2c Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) Test (1992)						
	No Deterministic Trend					**Critical Values**	
test stat	0.51	0.70	0.70	0.66	0.56	0.46	0.35
p-value	(0.04**)	(0.01**)	(0.01**)	(0.02**)	(0.03**)	--	--
	With Deterministic Trend					**Critical Values**	
test stat	0.18	0.18	0.19	0.20	0.19	0.15	0.12
p-value	(0.03**)	(0.02**)	(0.02**)	(0.02**)	(0.02**)	--	--
	B2d Variation Ratio Test (1988)						
	$\boldsymbol{e_t}$ **is not** $\boldsymbol{i.i.d.}$					**Critical Values**	
ratio	1.08	1.06	0.95	0.92	0.82	--	--
test stat	0.77	0.43	-0.36	-0.67	-1.60	1.96	1.64
p-value	(0.44)	(0.67)	(0.72)	(0.50)	(0.11)	--	--
	$\boldsymbol{e_t}$ **is** $\boldsymbol{i.i.d.}$					**Critical Values**	
ratio	1.08	1.06	0.95	0.92	0.82	--	--
test stat	0.64	0.44	-0.38	-0.66	-1.43	1.96	1.64
p-value	0.52	0.66	0.71	0.51	0.15	--	--

TABLE B3:
Unit Root Tests for Explanatory Variables, 1950-2010

	B2a Dicky-Fuller Test (1979)						
	No Drift, No Trend					**Critical Values**	
	unem	**cpi**	**trans**	**budget**	**openness**	**5%**	**10%**
ρ	1.00	0.91	1.01	0.80	1.03	--	--
test stat	-0.01	12.39	1.93	-2.31	5.13	-1.95	-1.61
p-value	(0.64)	(1.00)	(0.99)	(0.02**)	(1.00)	--	--
	With Drift					**Critical Values**	
ρ	0.78	1.02	0.99	0.70	1.02	--	--
test stat	-2.46	5.31	-0.28	-3.29	1.90	-2.91	-2.59
p-value	(0.13)	(1.00)	(0.92)	(0.02**)	(1.00)	--	--
	With Time Trend					**Critical Values**	
ρ	0.74	0.97	0.92	0.60	0.95	--	--
test stat	-2.81	-2.51	-1.63	-3.88	-1.33	-3.49	-3.17
p-value	(0.20)	(0.34)	(0.77)	(0.02**)	(0.87)	--	--
	B2b Philips-Perron Test (1988)						
	No Drift, No Trend					**Critical Values**	
ρ	1.00	1.03	1.01	0.80	1.03	--	--
test stat	-0.01	12.39	1.93	-2.31	5.13	-1.95	-1.61
p-value	(0.64)	(1.00)	(0.99)	(0.02**)	(1.00)	--	--
	With Drift					**Critical Values**	
ρ	0.78	1.02	0.99	0.70	1.02	--	--
test stat	-2.46	5.31	-0.28	-3.29	1.90	-2.91	-2.59
p-value	(0.13)	(1.00)	(0.92)	(0.02**)	(1.00)	--	--
	With Time Trend					**Critical Values**	
ρ	0.74	0.97	0.92	0.60	0.95	--	--
test stat	-2.81	-2.51	-1.63	-3.88	-1.33	-3.49	-3.17
p-value	(0.20)	(0.34)	(0.77)	(0.02**)	(0.87)	--	--
	B2c Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) Test (1992)						
	No Deterministic Trend					**Critical Values**	
test stat	0.32	0.78	0.72	0.57	0.75	0.46	0.35
p-value	(0.10*)	(0.01**)	(0.01**)	(0.03**)	(0.01**)	--	--
	With Deterministic Trend					**Critical Values**	
test stat	0.10	0.20	0.15	0.06	0.20	0.15	0.12
p-value	(0.10*)	(0.02**)	(0.05**)	(0.02**)	(0.02**)	--	--
	B2d Variation Ratio Test (1988)						
	e_t is not $i.i.d.$					**Critical Values**	
ratio	1.04	1.80	1.06	0.98	0.89	--	--
test stat	0.41	4.52	0.63	-0.17	-0.51	1.96	1.64
p-value	(0.68)	(0.00**)	(0.53)	(0.87)	(0.61)	--	--
	e_t is $i.i.d.$					**Critical Values**	
ratio	1.04	1.80	1.06	0.98	0.89	--	--
test stat	0.35	6.27	0.48	-0.19	-0.87	1.96	1.64
p-value	0.73	(0.00**)	0.63	0.85	0.39	--	--

Notes: p-values are reported in parenthesis. * rejects the null at 5% significance level. ** rejects the null at 10% significance level.

APPENDIX C
Cointegration Tests

Table C1 shows the results from the Johansen (1988) Cointegration Test for the quintile conditional mean incomes. Table C2 shows the corresponding results for Lorenz ordinates. The interpretation of the test results are as follows. For trace tests, we can reject the null hypothesis that the number of cointegration vectors is less than or equal to one against the alternative that the number of cointegrating vectors is more than one at both 5% and 1% significance levels. Therefore, there are at least two cointegrating vectors among the income distribution and the explanatory variables. The maximum eigenvalue test results indicate a number of cointegrating vectors in each set of variables. In all cases we fail to reject the null hypothesis $H_2: r < 2$ against $r = 3$. In most cases, however, we can reject the null hypothesis $H_2: r < 1$ against $r = 2$. Thus, we conclude that there are two or more cointegrating vectors in most cases.

Table C1
Cointegration Test for Quintile Conditional Mean Incomes, 1950-2010

	C1a. Trace Test							
		Conditional Mean Incomes					Critical Values	
H_2	R1	R2	R3	R4	R5	Top 5%	5%	1%
$r \leq 5$	4.82	5.68	5.82	5.80	5.45	4.15	9.16	12.76
$r \leq 4$	14.28	14.06	13.74	13.13	12.53	11.53	20.26	25.08
$r \leq 3$	28.77	29.87	29.63	28.80	28.09	27.04	35.19	41.19
$r \leq 2$	48.93	50.24	50.68	51.36	57.85	52.64	54.08	61.27
$r \leq 1$	86.22	88.78	89.20	90.04	96.57	91.17	76.97	85.33
$r = 0$	211.22	214.58	214.82	215.78	225.10	223.33	103.85	113.42
	C1b. Maximum Eigenvalue Test							
		Conditional Mean Incomes					Critical Values	
H_2	R1	R2	R3	R4	R5	Top 5%	5%	1%
$r \leq 5$	4.82	5.68	5.82	5.80	5.45	4.15	9.16	12.76
$r \leq 4$	9.46	8.38	7.92	7.33	7.08	7.39	15.89	20.16
$r \leq 3$	14.49	15.81	15.89	15.67	15.56	15.50	22.30	27.06
$r \leq 2$	20.16	20.37	21.06	22.55	29.76	25.61	28.59	33.73
$r \leq 1$	37.29	38.54	38.52	38.68	38.72	38.52	34.81	40.30
$r = 0$	125.00	125.80	125.61	125.74	128.53	132.17	40.96	46.76

Table C2
Cointegration Test for Lorenz Ordinates, 1950-2010

C2a. Trace Test							
	Conditional Mean Incomes					Critical Values	
H_2	L1	L2	L3	L4	L5	5%	1%
$r \leq 5$	1.86	1.92	2.05	2.17	2.19	3.84	6.63
$r \leq 4$	13.80	14.54	13.24	11.02	9.14	15.49	19.94
$r \leq 3$	27.92	28.47	26.37	23.88	22.19	29.80	35.47
$r \leq 2$	54.96	54.43	51.06	48.37	46.97	47.86	54.58
$r \leq 1$	90.43	89.11	86.40	85.80	82.50	69.82	77.82
$r = 0$	155.66	162.28	163.29	166.42	162.29	95.75	104.96
C2b. Maximum Eigenvalue Test							
	Conditional Mean Incomes					Critical Values	
H_2	L1	L2	L3	L4	L5	5%	1%
$r \leq 5$	1.86	1.92	2.05	2.17	2.19	3.84	6.63
$r \leq 4$	11.93	12.61	11.19	8.85	6.95	14.26	18.52
$r \leq 3$	14.13	13.93	13.13	12.86	13.05	21.13	25.86
$r \leq 2$	27.04	25.96	24.68	24.49	24.78	27.59	32.72
$r \leq 1$	35.46	34.69	35.34	37.43	35.53	33.88	39.37
$r = 0$	65.23	73.17	76.89	80.62	79.79	40.08	45.87

Notes:

1. For the trace tests, the alternative hypothesis is $H_a: r \geq r^*$, where $r^* = 5,4, \dots, 1$ in H_1 space of $r = 6$. For the maximum eigenvalue tests, the alternative hypothesis is $r = r^* + 1$, where $r^* = 5, \dots, 0$ in H_1 space of $r = r^* + 1$.

2. The five variables evaluated with each quintile conditional mean income or Lorenz ordinates are unemployment, inflation, transfers, budget deficit, and openness index.

3. For quintile conditional means, we specify that there are intercepts in the cointegrating relations but no trends in the data. For Lorenz ordinates, we specify that there are intercepts in the cointegrating relations and there are linear trends in the data, in which case, we use a model of deterministic cointegration, where the relations eliminate both stochastic and deterministic trends in the data.

APPENDIX D
Stage 1 Cointegration Estimates

Table D1
Cointegration Equation of Quintile Conditional Means, 1950-2010
(61 Observations)

	Quintile Conditional Mean Incomes					
	R1	**R2**	**R3**	**R4**	**R5**	**Top 5%**
Intercept	2.81	7.09	10.07	13.38	22.52	8.65
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Unemployment	-0.02	0.07	0.15	0.23	0.43	0.14
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CPI	-0.17	-0.21	-0.19	-0.10	0.27	0.20
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Public Transfers	0.13	0.28	0.41	0.58	0.99	0.37
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Budget Deficit	-0.44	-1.10	-1.96	-2.93	-6.87	-3.33
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Openness Index	-0.60	-2.66	-4.02	-5.85	-13.39	-6.49
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Durbin-Watson	0.83	0.84	0.90	0.81	0.64	0.70

Table D3
Cointegration Equation of Lorenz Ordinates, 1950-2010
(61 Observations)

	Quintile Conditional Mean Incomes				
	L1	L2	L3	L4	L5
Intercept	5.03	17.08	35.86	59.70	84.40
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Unemployment	-0.09	-0.15	-0.14	-0.08	0.04
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
CPI	-0.23	-0.42	-0.57	-0.60	-0.40
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Public Transfers	0.03	0.02	0.02	0.05	0.07
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Budget Deficit	0.31	0.85	1.17	1.36	1.14
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Openness Index	0.76	1.06	1.41	1.72	1.68
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Durbin-Watson	1.18	1.00	1.00	1.09	1.21

Notes:

1. p-values are reported in parenthesis. Due to the autocorrelation in error terms, standard deviations of parameter estimates are not efficient. Significance is exaggerated.
2. Table1 and Table 2 provide the first stage results. These results represent a long-run stationary linear relationship between explanatory variables and dependent variables. However, the existence of such a relationship does not necessarily imply a causation. The linear stationary relationship may not be unique and there can be more than one cointegration vector. For this reason, we enclose the first-stage regression results only in the appendix.
3. The critical values for Durbin-Watson test at 1% significance level with sample size of 60 are $d_L = 1.25$, $d_U = 1.60$. All reported test statistics are smaller than d_L. This indicates that the error terms in ordinary least square regressions are positively autocorrelated.