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ONE DOLLAR, ONE VOTE*

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This article revisits the relationship between inequality and redistribution in a panel of advanced OECD countries. Using panel data methods that hold constant a variety of determinants of redistributive spending, I find a non-monotonic relationship between pre-tax-and-transfer distribution of income and redistribution. Relative to mean income, a more affluent rich and middle class are associated with less redistribution and a richer poor class is associated with more redistribution. These results are consistent with a one dollar, one vote politico-economic equilibrium: when the income of a group of citizens increases, aggregate redistributive policies tilt towards this group's most preferred policies.

What determines the amount of resources that societies redistribute between their members? The current consensus in the literature, summarised in Persson and Tabellini (2003) and Alesina and Glaeser (2004), is that the pre-tax-and-transfer distribution of income is not a significant determinant of redistribution. The striking contrast between the US and Europe illustrates this 'paradox of redistribution'. The two workhorse models of distribution and redistribution, the normative model of Mirrlees (1971) and the positive theory of Meltzer and Richard (1981), in general predict that higher income inequality leads to more redistribution. However, in reality, the more pre-tax-and-transfer unequal US redistributes less than the more equal Europe.

Contrary to the conventional view in the literature, I argue that there is no paradox if we introduce multiple income inequality statistics in the same empirical framework and interpret the findings as the outcome of a one dollar, one vote politico-economic equilibrium. The motivation for interpreting cross-country differences in redistribution through the lens of this framework is simple. Since different income groups have conflicting goals regarding the redistribution of resources and since, as I show, income is strongly correlated with various measures of political participation, in principle a single inequality statistic (e.g. the Gini coefficient or the distance of median income from mean income) is unlikely to account for all conflicting preferences regarding the size of the welfare state. In other words, when political influence is increasing in income and redistribution responds to the political demands of various groups of voters, one would expect the effect of inequality on redistribution to vary depending on what part of the income distribution is changing.

I measure income inequality with three variables. Inequality at the bottom of the income distribution is given by the ratio of the gross earnings of the worker in the 10th percentile of the distribution to mean gross earnings ($y^{10}/\bar{y}$). The ratio of the gross

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earnings of the worker in the 90th percentile to mean gross earnings ($y^{90}/\bar{y}$) captures the relative affluence of the rich. Finally, the median to mean ratio of gross earnings ($y^{50}/\bar{y}$) measures how median income changes relative to any other income. Figure 1 explains why I summarise inequality with these three variables. The vertical axis measures the unconditional correlation between redistribution (Total Public Social Expenditure as a percentage of gross domestic product (GDP)) and the earnings of individuals located in each percentile of the gross earnings distribution relative to the mean. The horizontal axis measures the percentile of the gross earnings distribution from a sample of advanced OECD countries. The Figure shows, for instance, that the correlation between redistribution and the $y^{10}/\bar{y}$ ratio is around 0.60, when the latter is measured with a three-year lag relative to redistribution.

Clearly, there are three areas of interest. The correlation between redistribution and relative gross earnings is positive for individuals poorer than the median, it falls to around zero for the median and it turns negative for individuals richer than the median. This result arises at the pooled sample but a similar pattern emerges at the cross-country dimension of the sample for separate time periods. As Figure 1 shows, when the inequality ratios are measured with a 15-year lag relative to redistribution, the correlations for the poor and the rich do not decrease much in magnitude. Based on this fact, I hypothesise that the omission of relevant variables is a more serious concern than reverse causation in attempting to estimate a causal link from inequality to redistribution. However, the econometric methodology is meant to address both concerns.

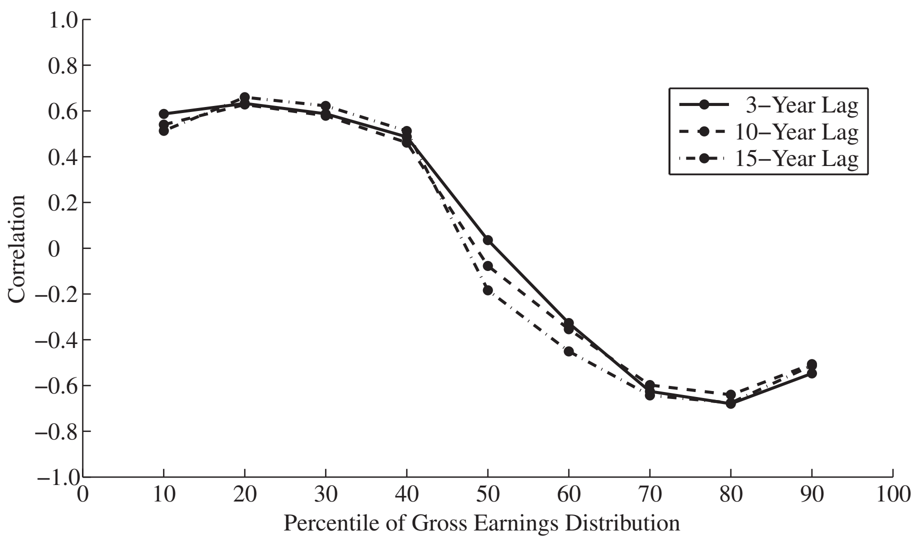

Fig. 1. *Correlation Between Redistribution and Percentile to Mean Ratio of Gross Earnings*

Notes. The horizontal axis measures the percentile of the income distribution. The vertical axis shows the unconditional correlation between redistribution and the gross earnings of every percentile relative to mean gross earnings. For instance, the correlation between redistribution and the 10/mean ratio of gross earnings ($y^{10}/\bar{y}$) is around 0.60, when the latter is measured with a three-year lag relative to redistribution. The correlation is calculated for the pooled sample in levels (similar results are obtained when variables are logged).

To test the hypothesis that inequality is relevant for redistribution, in Section 1, I construct a panel of advanced OECD countries over 1975–2001. My measure of redistribution is standard in the literature and includes expenditures on pensions, survivors and incapacity related benefits, programmes for the poor families, housing, unemployment, active labour market programmes and health. The empirical methodology, however, differs from previous papers in some or all of the following dimensions.[1] First, and most importantly, I introduce three different indices of income inequality in the same empirical framework to capture the idea that redistribution responds to the demands of various groups of voters. Second, I measure inequality using gross earnings data which, as explained next, helps to address a number of econometric issues. Third, I focus on the within-country variation of the data to make my results consistent with a prominent strand of literature which emphasises the role of persistent determinants of redistribution (e.g. institutions, culture and ethnicity; see below for references). In other words, the estimated effect of inequality on redistribution holds constant all 'long-run' factors that cause some countries to redistribute more than others irrespective of differences in their income distributions.

Section 2 presents the results. Overall, I find a robust positive relationship between the relative earnings of the poor and redistributive spending, a solid negative association for the median over mean (the Meltzer and Richard (1981)) ratio and a broadly consistent negative association between relative earnings of the rich and redistributive spending. Therefore, inequality between the tails of the pre-tax-and-transfer distribution, in the sense of a richer rich class and a poorer poor class, is associated with less redistribution. That the median to mean ratio of gross earnings is associated with less redistribution shows how different indices of income inequality are important.

One contribution of the article is to explain why previous empirical tests in the literature have found scant evidence in favour of the basic median voter equilibrium mechanism proposed by Meltzer and Richard (1981). To isolate which of the above three features of the empirical methodology is responsible for my finding that the Meltzer–Richard coefficient has the expected negative sign, I first note that previous studies using close proxies for the median to mean ratio of gross earnings (Persson and Tabellini, 1994; Perotti, 1996; Rodriguez, 1999) have not found a robust negative association between this measure of inequality and redistribution. Second, in my sample, the coefficient of the median to mean ratio of gross earnings is insignificant and often has the wrong sign in various regressions with and without fixed effects (FE), when I omit the relative earnings of the poor and the rich from the specification. As a result, what reverses the sign and the significance of the Meltzer–Richard coefficient is the inclusion of proxies for the poor's and the rich's demand for redistribution. The intuition for this novel empirical result is rather simple and accords with previous theoretical work (Bénabou, 1996; Epple and Romano, 1996; Saint-Paul and Verdier, 1996). When we omit the earnings of the rich and the poor from the regression, we are

[1] A non-exhaustive list of work on the empirical relationship between inequality and redistribution includes: Persson and Tabellini (1994), Lindert (1996), Perotti (1996), Gouveia and Masia (1998), Rodriguez (1999), Milanovic (2000), Moene and Wallerstein (2001), Bassett *et al.* (2003), Persson and Tabellini (2003), Iversen and Soskice (2006), Bellettini and Berti Ceroni (2007), Georgiadis and Manning (2007), Lind (2007), Shelton (2007) and Ramcharan (2010).

implicitly assuming that the median voter is the decisive voter. In reality, however, the political system is more complicated than what the 'one person, one vote' model assumes and various groups of voters may have a 'say' for the equilibrium outcome. In other words, the basic prediction that the median to mean ratio of gross earnings is negatively associated with redistribution should be conditioned on the earnings of other groups of voters to control for the probability that the median is not necessarily the decisive voter.

I provide a battery of robustness checks. The results remain unchanged when many controls and interaction terms with time-invariant determinants of redistribution are included in the regression. I verify my analysis using pensions and statutory income and payroll taxes as proxies for redistribution to address measurement issues in the dependent variable. Dropping influential observations, excluding one country at a time from the sample, redesigning the structure of the panel errors, accounting for country-specific trends and modelling the persistent nature of redistribution do not affect the majority of my findings. Finally, most of the results are verified when the inequality ratios are instrumented and therefore, the empirical estimates are in general robust to the strict exogeneity assumption necessary for the FE estimator.

Section 3 discusses the results. A well-known theoretical proposition is that income and most preferred size of redistribution are negatively related across groups of voters. As I show, this theoretical result finds considerable empirical support in microeconomic data. As a result, the rich prefer less redistribution than the median and the median prefer less redistribution than the poor. Starting from this result, the effects of the inequality variables on redistribution are consistent with what I define as a one dollar, one vote politico-economic equilibrium. In the one dollar, one vote equilibrium, when a group of citizens becomes richer (relative to mean income), redistribution tilts towards its bliss point. For the rich and the median group, this means that redistribution decreases with their income, whereas for poor group, this means that redistribution increases with its income. In other words, the one dollar, one vote definition captures the positive association between group income and distance of group most preferred outcome from equilibrium redistribution. A natural explanation for the one dollar, one vote result is that political influence is not uniform across groups of voters. Using individual-level data from the World Values Survey (WVS), I show that political participation increases in income. As a result, an increase in the income of any group of voters implies more political influence in the equilibrium outcome, in line with the one dollar, one vote result.

The one dollar, one vote result can explain the 'paradox of redistribution'. Between 1980–2001, the growth of redistribution was lower in the US than in a European aggregate of countries. The poor and the middle class became poorer in the US relative to Europe. These two opposing effects on the cross-Atlantic difference in the growth of redistribution tend to cancel out. As the rich became richer in the US (relative to the mean), political influence shifted towards the interests of the American rich class. As the rich became poorer in Europe, political influence shifted away from the interests of the European rich class. Since on both sides of the Atlantic, the rich prefer less redistribution, Europe's growth of redistribution exceeded the US growth.

1. Empirical Methodology

1.1. *Sample*

I use various sources to construct an unbalanced panel of 14 advanced OECD countries over 1975–2001.[2] The selection of the sample is based on the existence of data for income distribution and redistribution. Table 1 presents the summary statistics. The Appendix discusses in more detail the definition and the sources of the variables.

1.2. *Data*

1.2.1. *Redistribution*

The dependent variable is taken from the OECD SOCX Dataset.[3] The measure of redistribution is Total Public Social Expenditure as a percentage of GDP which includes expenditures on old age (pensions), survivors, incapacity related benefits, health, family, active labour market programmes, unemployment, housing and other social policies. I examine the robustness of the results to alternative measures of redistribution by dividing Total Public Social Expenditures into its three main

Table 1
Summary Statistics of Country-level Data

Variable	Mean	SD	Min	Max	Observations
Total public expenditure (RED)	21.21	5.80	10.20	35.10	110
RED net of health and pensions	8.43	3.62	2.40	17.90	110
Pensions	6.96	2.46	2.70	11.50	110
Health	5.83	0.98	3.70	8.50	110
Tax wedge	32.69	9.97	14.10	48.20	112
$y^{90}/\bar{y}$	1.57	0.11	1.32	1.89	90
$y^{50}/\bar{y}$	0.88	0.04	0.71	0.93	90
$y^{10}/\bar{y}$	0.54	0.09	0.30	0.70	90
GDP	19,619	4,051	9,276	33,905	126
Growth	2.11	1.82	−5.40	8.90	126
Share of elderly population	13.16	2.36	8.10	18.00	126
Deadweight loss (DWL)	2.30	0.58	1.01	3.34	122
Voter turnout	78.18	11.23	49.10	95.80	125
Openness	57.77	28.97	16.30	174.00	126
Unemployment rate	7.10	3.27	0.30	16.40	118
Employment ratio	65.78	6.65	51.87	80.53	126
Left vote	35.39	14.54	0.00	56.00	126
Coordination index	3.02	1.40	1.00	5.00	126
Share of labour taxes	62.39	12.54	34.00	86.06	92
Mean income/wage gap	1.18	0.16	0.79	1.47	119
Minimum wage	0.35	0.25	0.00	0.78	107

Notes. These are summary statistics for the pooled sample in which all variables have been averaged using non-overlapping three-year intervals. See the Appendix for the definition and the source of each variable.

[2] The countries are: Australia, Austria, Canada, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, New Zealand, Sweden, UK and USA. Canada is observed mostly in the 1980s. Italy is observed until 1995, that is, before the change in its political institution. For Austria and Ireland, there is only one inequality observation and thus these countries are dropped from the sample in some specifications that require two or more observations per country.

[3] The SOCX Dataset starts in 1980 but my explanatory variables extend back to 1975 because, as explained next, I define a time period to be three years and I lag the regressors by one period.

components, namely health, pensions and expenditures net of health and pensions. As an alternative dependent variable, I also use OECD's Tax Wedge variable that measures the sum of personal income taxes and Social Security contributions.[4]

1.2.2. *Inequality*

An important issue in the empirical investigation is the definition and quality of the inequality variables. The quality of the data has improved in recent years but the data are still far from perfect (Atkinson and Brandolini, 2001). My source is the OECD LFS Dataset which presents gross earnings of full-time workers by deciles of the income distribution. This Dataset has a number of desirable features. First, it offers the possibility to construct a variety of inequality measures and examines different parts of the gross earnings distribution simultaneously. Moreover, the basic political economy model predicts a relationship between pre-tax-and-transfer income inequality and redistribution (Persson and Tabellini, 2000). Using the OECD LFS Dataset, the inequality variables are constructed with gross earnings data and not with data on expenditures, consumption, wages, net earnings or disposable income. As explained below in more detail, using gross instead of net earnings has the important additional advantage that redistribution does not overlap mechanically with the inequality variables.[5]

I map the theory close to the data and proxy for the Meltzer and Richard (1981) ratio with the ratio of the gross earnings of the worker located in the 50th percentile of the income distribution to mean gross earnings, $y^{50}/\bar{y}$.[6] In addition, as discussed in Section 3, the demand of the rich and the poor can also affect redistribution and, in reality, the median voter may not be the decisive voter. I proxy for the rich's and the poor's demand for redistribution with the $y^{90}/\bar{y}$ and the $y^{10}/\bar{y}$ ratios, respectively.[7]

Some of the literature uses the Gini coefficient to proxy for inequality.[8] The intuition developed in the next Section is that the Gini coefficient cannot capture the

[4] In addition, the results below are in general robust when redistribution is measured with Milanovic's (2000) reduction in inequality variables. I prefer to use the Total Public Social Expenditure variable for two reasons. First, the number of observations drops to 32 or 34 when Milanovic's (2000) sample is overlapped with my sample, which increases significantly the variability of the estimates. Second, Lind's (2005) argument that there is a mechanical positive correlation between the Gini coefficient and Milanovic's measure of redistribution may also apply when I regress Milanovic's measure of redistribution on my inequality variables (but it does not apply when I use Total Public Social Expenditure as a measure of redistribution).

[5] The results are not sensitive to the few observations for which gross earnings data are not available.

[6] Perotti (1996) uses the size of the middle class which is close to the $y^{50}/\bar{y}$ variable. He finds that more inequality is associated with less growth, but he rejects that the link between these two variables occurs through redistribution as hypothesised by workhorse models. See also Bassett *et al.* (2003) on this point. Rodriguez (1999) tests the Meltzer and Richard (1981) model carefully using pre-tax median to mean earnings. He presents time series and cross-sectional evidence from US states which reject the median voter model. See also Gouveia and Masia (1998) for a similar result. I explain the difference of my results from these articles in Section 2.2.

[7] Other papers use the y^{90}/y^{10} to proxy for the median's demand and interpret the negative estimated coefficient as failure of the Meltzer and Richard (1981) model. See, for instance, Moene and Wallerstein (2001) who offer an insurance interpretation to explain the negative coefficient and Iversen and Soskice (2006) who also report a negative or close to zero correlation for the 90–50 ratio. In light of the correlations in Figure 1, the negative correlation between the 90–10 and redistribution is not surprising, since both the 90th percentile in the numerator and the 10th percentile in the denominator drive the strong negative correlation. The same intuition applies for the 90–50 ratio. Because the 90–mean ratio is strongly negatively correlated with redistribution and the 50–mean is uncorrelated or weakly negatively correlated with redistribution, the 90–50 ratio is (weakly) negatively correlated with redistribution.

[8] See, for instance, Milanovic (2000), Persson and Tabellini (2003) and Shelton (2007).

non-monotonic relationship between inequality and redistribution. An example that contrasts the distribution of factor income in Sweden and the US in the early 1980s illustrates this point. While Sweden has a Gini coefficient of 0.463 and the US has a Gini coefficient of 0.464, the US is much less equal than Sweden when we examine other dimensions of inequality. According to the OECD LFS Dataset, the y^{90}/y^{10} ratio of gross earnings is 3.79 in the US and only 2.05 in Sweden. If we use the Gini coefficient to summarise the income distribution, then we overlook the fact that inequality between the tails of the distribution is much more pronounced in the US than in Sweden.

1.3. *Empirical Specification*

The estimating equation is:

$$\mathrm{RED}_{i,t} = c_i + \tau_t + \beta_1 \left(\frac{y^{90}}{\bar{y}}\right)_{i,t-1} + \beta_2 \left(\frac{y^{50}}{\bar{y}}\right)_{i,t-1} + \beta_3 \left(\frac{y^{10}}{\bar{y}}\right)_{i,t-1} + \sum_{k=1}^{K} \gamma_k X_{k,i,t-1} + \varepsilon_{i,t}, \quad (1)$$

where $\mathrm{RED}_{i,t}$ is Total Public Social Expenditure as a share of GDP for country $i = 1, \ldots, N$ in period $t = 1, \ldots, T$; $(y^j/\bar{y})_{i,t-1}$ is the gross earnings of the worker in the jth percentile of the gross earnings distribution divided by mean gross earnings for $j = 90,50,10$; $X_{k,i,t-1}$ is other explanatory variables; c_i is a time-invariant country-specific unobservable effect; τ_t is a common unobservable year-specific effect and $\varepsilon_{i,t}$ is the time-varying country-specific idiosyncratic error. Variables are in logs and therefore the regression coefficients denote the elasticities of redistributive spending.[9]

Assessing the causal effect of inequality on redistribution is certainly a challenging task; in general, we cannot expect inequality to be exogenous. The first source of endogeneity is contemporaneous reverse causation and in fact the often reported insignificant and sometimes negative relationship between pre-tax-and-transfer inequality and redistribution may simply reflect the case that societies redistribute to decrease their post-tax-and-transfer economic inequality.[10] Using gross earnings instead of net earnings or consumption to construct the inequality ratios relaxes somewhat this constraint because net earnings vary both 'mechanically' and 'economically' with the fiscal system, whereas gross earnings vary only through the endogenous response of labour supply or the general equilibrium effect on factor prices. To mitigate the concern that redistribution may affect inequality through labour supply or general equilibrium effects, I average all variables using non-overlapping three-year intervals and lag the regressors by one period. For example, in period $t = 2001$, RED_t is an average of redistributive spending over 1999, 2000 and 2001. This is regressed on right-hand-side variables which are averaged over 1996, 1997 and 1998, and therefore, on average, inequality leads redistribution by three years. The averaging procedure is desirable because it removes transitory fluctuations because of economic

9 The coordination index, the government surplus, the growth rate and the minimum wage are not logged because they take negative or zero values. The results are similar when the equation is specified in levels.

10 For an insignificant and sometimes negative relationship between inequality and redistribution, see, for instance, Persson and Tabellini (1994, p. 617), Perotti (1996, p. 170), Persson and Tabellini (2003, p. 43) and Alesina and Glaeser (2004, p. 58).

business cycles, it captures meaningful variations along the political cycle and it reduces serial correlation and measurement error.[11]

Probably, the most serious concern for my empirical investigation is the potential inconsistency that arises from omitting relevant determinants of redistribution. The panel that I construct readily controls for time-invariant observable or unobservable determinants of redistribution. This is important because the literature emphasises the role of long-run, persistent factors that cause some countries to redistribute more than others irrespective of differences in their income distributions. For instance, legal origins, initial technological capabilities and geography, political institutions, such as electoral rules, form of the government, judicial review and federalism (Persson and Tabellini, 2003; Alesina and Glaeser, 2004), persistent cultural characteristics, such as beliefs in a just world (Bénabou and Tirole, 2006) and trust (Tabellini, 2010), ethnic fragmentation (Easterly and Levine, 1997; Alesina and Glaeser, 2004), prospects of upward mobility (Piketty, 1995; Bénabou and Ok, 2001), social beliefs about fairness (Alesina and Angeletos, 2005) and initial wealth inequality (Galor and Zeira, 1993; Deininger and Squire, 1998) vary across countries and explain why some countries adopt more pro-redistributive policies than others. However, these determinants are absorbed by the time-invariant country-specific effect (c_i) since they are fixed within country for the period considered. Because the path of inequality in each country may be the result of certain political institutions, labour market regulations and social perceptions about work ethics, luck and the redistribution of resources, I allow for any arbitrary correlation between the country-specific effects and the regressors.

The inclusion of year-specific FE (τ_t) controls for common shocks in redistribution, such as the technological slowdown caused by the oil shocks in the beginning of the sample and the world rise of rightist movements in the 1980s. The time effects also help to reduce the effects of spurious trends and contemporaneous panel error correlations. Adding $N+T-2$ dummy variables into the model certainly creates a demanding environment to test my hypothesis since it removes a large fraction of the cross-country and within-country variation of the data. However, I adopt this technique because it mitigates the endogeneity problem and helps to identify the effects of inequality on redistribution.

The consistency of the FE estimator is subject to the strict exogeneity assumption. Specifically, factors affecting redistribution in period t but omitted in the error term are assumed to be orthogonal to past, present and future explanatory variables conditional on the FE (Wooldridge, 2002). As discussed in the following, this assumption may be too restrictive and is relaxed in Section 2.3.

2. Income Inequality and Redistribution

2.1. *Main Results*

Table 2, column 1, presents the baseline specification. There is a positive association between the $y^{10}/\bar{y}$ ratio and redistribution. This coefficient is significant at the 1% level. The $y^{50}/\bar{y}$ ratio is negatively associated with redistribution and its coefficient is

[11] The results are, in general, similar when I use non-overlapping two or four-year intervals.

Table 2
Baseline Specification

Dependent variable	RED (1)	RED net of health and pensions (2)	Pensions (3)	Health expenditures (4)	Tax wedge (5)
$y^{90}/\bar{y}$	−1.06	−1.71	−1.13	0.29	−1.26
	(0.23)***	(0.52)***	(0.45)**	(0.42)	(0.34)***
$y^{50}/\bar{y}$	−1.44	−1.22	−0.94	−1.62	−0.44
	(0.63)**	(0.64)*	(1.09)	(1.20)	(0.63)
$y^{10}/\bar{y}$	1.01	1.59	1.47	−0.07	1.02
	(0.26)***	(0.43)***	(0.45)***	(0.41)	(0.37)**
GDP	−0.07	−1.37	0.78	0.81	0.29
	(0.40)	(0.73)*	(0.55)	(0.49)	(0.37)
Growth	−0.025	−0.046	−0.013	−0.007	0.003
	(0.006)***	(0.010)***	(0.007)*	(0.011)	(0.007)
Share of elderly	0.31	0.28	0.53	−0.07	0.42
	(0.13)**	(0.25)	(0.21)**	(0.12)	(0.18)**
DWL	−0.15	−0.03	−0.18	−0.21	−0.17
	(0.09)	(0.18)	(0.06)***	(0.10)	(0.04)***
R^2 (within)	0.70	0.54	0.66	0.64	0.55
Observations	82	82	82	82	82

Notes. The independent variables are defined in the text and in the Appendix. All specifications include year-specific and country-specific fixed effects. Standard errors are clustered by country and are displayed in parentheses. *** Significance at 1%, ** significance at 5%, * significance at 10%. All significance levels are based on t-statistics.

significant at the 5% level. The $y^{90}/\bar{y}$ ratio is negatively associated with redistribution. This coefficient is significant at the 1% level.

In the following columns, I slice Total Public Social Expenditures (RED) into its three major components, namely RED net of Health and Pensions, Pensions and Health. I find a similar pattern in terms of signs and significance for the inequality coefficients when Pensions and RED net of Health and Pensions are used as the dependent variable. However, the inequality variables do not seem relevant in explaining Health expenditures. This is a reasonable result in light of the fact that the one dollar, one vote hypothesis should be stronger for more progressive social programs and health is understood as one of the least progressive social expenditures. The fact that the results do not change significantly when I use the Pensions variable instead of RED shows clearly how the association between inequality and redistribution is not artificial. The reason is that the inequality ratios concern the gross earnings of full-time workers, while by definition pensions are given to retirees.[12] As a result, the inequality measures do not overlap mechanically with this measure of redistribution.

The baseline specification includes four control variables. GDP controls for the size of the tax base and also for Wagner's Law. There is no evidence that GDP affects

[12] Pensions are redistributive within-cohorts because of the progressive benefit formula (Persson and Tabellini, 2000; Feldstein and Liebman, 2002). See, for instance, the report from the Congressional Budget Office in 2006 which shows how for people in the bottom fifth of the earnings distribution, the ratio of benefits to taxes is almost three times as high as it is for those in the top fifth. Tabellini (2000) uses the insight that pensions redistribute within generation to argue that a small degree of altruism of the grandchild towards the grandfathers can sustain equilibrium pensions even in the absence of commitment.

redistribution which is reasonable since the sample consists of developed economies. Economic growth accounts for the countercyclicality of fiscal policy driven by automatic stabilisers (e.g., unemployment compensations).[13] Consistently with previous studies in OECD countries (see Galí and Perotti (2003) – for a discussion) I find strong evidence of countercyclical redistribution. The share of the elderly population holds constant demographic characteristics and the bargaining power of the elderly when demanding more directed transfers, that is, pensions. I find that a larger share of elderly population is associated with more redistribution.[14] Finally, the deadweight loss, a measure of the opportunity cost of taxation, is negatively associated with redistribution. When I omit all these controls, the estimated inequality coefficients are very similar in magnitude but their variability increases somewhat.

Column 5 of Table 2 uses the sum of personal income taxes and payroll taxes (Tax Wedge), a measure of the net tax burden for the average earner, as the dependent variable. I repeat the baseline regression with this measure of redistribution to verify that the results are not sensitive to the cyclical nature of redistribution. Spending on automatic stabilisers decreases during economic booms. If in booms the income benefit for the rich exceeds the benefit for the poor, then the $y^{90}/\bar{y}$ increases and the $y^{10}/\bar{y}$ decreases. This could explain the negative coefficient of the former and the positive coefficient of the latter. The Tax Wedge variable addresses this concern since personal income taxes and Social Security contributions are statutory. Column 5 of Table 2 shows how the estimated coefficients of the rich and the poor remain robust, while the coefficient of the median is still negative but loses significance. In addition, growth enters insignificantly which verifies the initial hypothesis that the Tax Wedge is immune to cyclical fluctuations.

The pattern of the FE emerging from the estimation mimics strongly the pattern of the averaged value of redistribution by country (with a correlation of 0.59). This shows that omitted factors such as political and labour market institutions, ethnicity and social perceptions about redistribution are important determinants of redistribution over the long-run. More formally, the null hypothesis that all country FE are not significant is rejected at the 1% level. The time dummies also enter jointly significantly. Estimation with random effects does not change significantly the estimated inequality coefficients and therefore the results are robust to a potential sensitivity of the FE estimator to measurement error.

The economic significance of the three inequality ratios is also large. According to the first column in Table 2, the conditional elasticity of redistribution with respect to the demand of the rich, the Meltzer–Richard ratio and the demand of the poor is -1.06, -1.44 and 1.01 respectively. To understand the magnitude of these coefficients, suppose that a country redistributes the mean value of RED in the sample, 21.21%. A 10% increase in the three inequality ratios would change redistribution as a percentage of GDP by -2.25, -3.05 and 2.14 percentage points relative to its mean sample value. The magnitude of these variations is reasonable both in the within-country and in the

[13] Growth is not lagged in the regressions because the automatic stabilisation take places contemporaneously.

[14] This result (which also holds for pensions) contrasts with the findings of Perotti (1996) and Bassett *et al.* (2003) who argue that the link between inequality and redistribution weakens once the share of elderly population is controlled for.

cross-country dimension of the sample.[15] Factors that contribute to cross-country differences in the time series variation of inequality include technological change and globalisation (Georgiadis and Manning (2007) for the UK), market-based factors, such as the age or educational premium that affect the relative supply of skills (Gottschalk and Joyce, 1998), but also pre-existing labour market and political institutions, ethnic divisions and cultural beliefs (i.e. the 'FE'). In Section 3.3, I analyse the significance of the results in more detail by showing how changes in inequality can account for the growth of the US–Europe difference in redistribution over 1980–2001.

Table 3 adds other time-varying controls to the baseline specification. Since I analyse the relationship between redistributive expenditures and gross labour earnings inequality but some expenditures are not necessarily financed by labour income taxes, it might be necessary to control for the share of government spending financed by labour income taxes and a measure of the gap between average labour earnings and average income. To calculate the share of government spending financed by labour income taxes, I follow the methodology of Mendoza *et al.* (1994), as updated recently by McDaniel (2007), and obtain measures of the average or 'effective' labour income tax rate. Multiplying the average labour income tax rate by wages and salaries and dividing by government spending, I obtain the variable 'Share of Labour Taxes'.[16] The 'Mean Income/Wage Gap' variable denotes the ratio of average income to average earnings. As columns 1–3 of Table 3 show, the results remain robust to the inclusion of these two variables in isolation or in combination.

In the next columns of Table 3, I include the rate of unemployment, the degree of openness, the share of left vote, the voter turnout, the wage coordination variable and the ratio of the minimum wage to the average wage (in isolation in columns 4–9 and in combination in column 10). The coefficients of these controls present signs in accordance with theory but are not significant at conventional levels. Unemployment is positively associated with redistribution which is expected since unemployment benefits are part of RED. Openness also enters with a positive coefficient which is consistent with Rodrik (1998), given that the size of the country is absorbed by the FE. Left vote enters positively as left-wing parties are associated with more redistributive policies, however, its sign changes when more controls are included. Voter turnout enters positively in the regression as supposedly the poor voters are the ones more elastic to voting. The wage coordination variable enters with a negative coefficient which shows a substitutability between labour market rigidities and cash redistribution. Finally, the coefficient of the minimum wage is positive as more redistributive governments try to raise the minimum relative to the mean wage.

I have controlled for other potentially relevant variables (to save space these results are not reported). Controlling for the percentage of right vote in the last elections, the unionisation of the economy and the budget surplus of the government does not

[15] A 10% increase in the mean to median ratio is an increment of more than two standard deviations. In 1978, France's $y^{50}/\bar{y}$ ratio was 10% smaller than US's ratio. A 10% increase in the rich to mean ratio is somewhat more than one standard deviation increment, whereas a similar change in the poor to mean ratio corresponds to a somewhat less than one standard deviation increment. For instance, in 1999, the $y^{90}/\bar{y}$ ratio is 20% larger in the US than in Sweden and the $y^{10}/\bar{y}$ ratio is 40% smaller in the US than in Sweden.

[16] The variable Share of Labour Taxes is not lagged, since I am interested in controlling for the period t share of expenditures financed by labour income taxes. Lagging this variable makes no difference to the results.

Table 3

Other Time-Varying Controls

Dependent variable: RED	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
$y^{90}/\bar{y}$	−1.61	−1.07	−1.66	−1.11	−1.04	−0.98	−1.15	−1.12	−0.97	−3.18
	(0.39)***	(0.23)***	(0.40)***	(0.26)***	(0.27)***	(0.27)***	(0.21)***	(0.22)***	(0.43)**	(0.94)***
$y^{50}/\bar{y}$	−2.08	−1.41	−2.35	−1.36	−1.19	−1.22	−1.40	−1.42	−1.46	−2.58
	(0.94)**	(0.68)**	(0.98)**	(0.56)**	(0.67)*	(0.63)*	(0.65)**	(0.58)**	(0.69)*	(1.05)**
$y^{10}/\bar{y}$	1.69	1.00	1.84	0.83	0.96	1.03	1.12	1.02	0.95	2.43
	(0.45)***	(0.32)***	(0.42)***	(0.19)***	(0.30)**	(0.34)***	(0.29)	(0.27)***	(0.30)***	(0.64)***
GDP	1.04	−0.03	1.15	0.21	0.11	−0.14	−0.08	−0.04	0.02	1.62
	(0.43)**	(0.40)	(0.42)**	(0.61)	(0.27)	(0.38)	(0.39)	(0.40)	(0.60)	(0.58)**
Growth	−0.009	−0.025	−0.009	−0.026	−0.024	−0.027	−0.027	−0.026	−0.026	−0.006
	(0.004)**	(0.007)***	(0.004)*	(0.007)***	(0.005)***	(0.007)***	(0.007)***	(0.007)***	(0.009)**	(0.008)
Share of elderly	0.52	0.31	0.73	0.26	0.40	0.35	0.31	0.32	0.26	−0.35
	(0.32)	(0.13)**	(0.34)*	(0.13)*	(0.23)	(0.15)**	(0.15)*	(0.14)**	(0.12)*	(0.43)
DWL	−0.31	−0.15	−0.35	−0.10	−0.17	−0.14	−0.14	−0.13	−0.14	−0.29
	(0.09)***	(0.09)	(0.09)***	(0.12)	(0.09)*	(0.10)	(0.10)	(0.10)	(0.13)	(0.21)
Share of labour taxes	0.29		0.28							−0.03
	(0.26)		(0.23)							(0.27)
Mean income/wage gap		−0.05	0.28							0.26
		(0.19)	(0.15)*							(0.23)
Unemployment rate				0.09						0.06
				(0.11)						(0.09)
Openness					0.14					−0.05
					(0.19)					(0.24)
Left vote						0.02				−0.01
						(0.09)				(0.08)
Voter turnout							0.31			1.02
							(0.41)			(0.63)
Coordination								−0.01		−0.03
								(0.01)		(0.02)
Minimum wage									0.16	0.17
									(0.25)	(0.20)
R^2 (within)	0.75	0.70	0.77	0.72	0.71	0.70	0.70	0.70	0.71	0.89
Observations	64	81	64	82	82	74	81	82	77	50

Notes. In all columns, the dependent variable is Total Public Social Expenditures. The independent variables are defined in the text and in the Appendix. All specifications include year-specific and country-specific fixed effects. Standard errors are clustered by country and are displayed in parentheses. *** Significance at 1%, ** significance at 5%, * significance at 10%. All significance levels are based on t-statistics.

change the results shown in Tables 2 and 3 significantly. Adding the employment ratio to the regressions also does not overturn the magnitude or significance of the three inequality coefficients. The employment ratio addresses a potential sensitivity of the poor's coefficient because of the normality of leisure. If a more generous welfare transfer causes low ability workers to exit the labour market or participate only part-time, then the earnings of the worker located in the 10th percentile of the distribution of full-time gross earnings increase artificially with redistribution (for the median and the rich this effect works against finding the negative coefficients).

The inequality variables have been interacted with other determinants of redistribution. First, in the simplest formulation of a voting model, it is assumed that everyone votes but in reality inequality matters for redistribution to the extent that the poor vote. To test this hypothesis, I interact the $y^{10}/\bar{y}$ ratio with the turnout of the voters in the last elections. This coefficient has the expected positive sign.[17] In addition, the coefficients of the other inequality ratios do not change significantly. Second, the interaction of the Meltzer–Richard ratio with the ethnolinguistic fractionalisation (ELF) index of Roeder (2001) provides a test of Lind's (2007) conjecture that only in less fragmented societies does inequality increase redistribution. In accordance to his intuition, I find that less heterogeneous societies redistribute more after an increase in inequality. Finally, I have used interaction terms of political institutions (electoral rules and form of government) with the inequality ratios with no significant difference in the results.

2.2. *Comparison to Other Studies*

The empirical methodology rests on three key features. First, I introduce various indices of income inequality in the same empirical framework. Second, the inequality ratios are constructed with gross earnings data. Third, I focus on the within-country variation of the sample.

The finding that the Meltzer–Richard (1981) ratio enters with the expected negative sign is novel. To isolate which of the three features above is responsible for this result, I first note that previous studies employing close proxies for the Meltzer–Richard ratio have not found the result. Specifically, Persson and Tabellini (1994), Perotti (1996) and Bassett *et al.* (2003) use the (pre-tax) size of the middle class in a cross-section of countries and Gouveia and Masia (1998) and Rodriguez (1999) use pre-tax median over mean earnings in a sample of US states. None of these papers, however, finds robust evidence that a higher Meltzer–Richard ratio is associated with less redistribution.[18] Second, to investigate whether the focus on the within-country variation of the sample explains my result, Table 4 repeats the baseline specification of Table 2, with and without FE, with and without the additional controls and with and without controlling for the $y^{90}/\bar{y}$ and the $y^{10}/\bar{y}$ ratios. As the Table shows, when I drop the variables

[17] The regression also includes the voter turnout independently, as in column 7 of Table 3. A 10% increase of voter turnout relative to its average value in the sample (79%) increases the elasticity of redistribution with respect to the relative earnings of the poor from 1.12 to 1.21.

[18] As in this article, Bassett *et al.* (2003) point out that the median voter model may be too simple to capture the complex relationship between income, political power and redistribution. However, in a cross-section of countries, they find no consistent pattern between social transfers and income share when they substitute the income share of the middle class with the income shares of other quintiles. The difference is that I consider all measures of inequality simultaneously in the same framework.

Table 4

Redistribution and the Meltzer–Richard (1981) Ratio

Dependent variable: RED	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$y^{90}/\bar{y}$	−1.06	−1.20			−0.67	−0.95		
	(0.23)***	(0.45)**			(0.73)	(0.86)		
$y^{50}/\bar{y}$	−1.44	−1.45	0.13	0.46	−1.00	−3.19	−0.53	0.50
	(0.63)**	(0.98)	(0.65)	(0.68)	(1.45)	(1.81)*	(0.63)	(0.95)
$y^{10}/\bar{y}$	1.01	1.09			0.33	1.50		
	(0.26)***	(0.48)**			(0.42)	(0.44)***		
GDP	−0.07		−0.15		0.03		−0.22	
	(0.40)		(0.28)		(0.39)		(0.37)	
Growth	−0.025		−0.028		−0.027		−0.038	
	(0.006)***		(0.007)***		(0.011)***		(0.012)***	
Share of elderly	0.31		0.45		0.84		0.95	
	(0.13)**		(0.19)**		(0.29)**		(0.27)***	
DWL	−0.15		−0.16		−0.43		−0.54	
	(0.09)		(0.11)		(0.24)		(0.24)**	
Country fixed effects	Yes	Yes	Yes	Yes	No	No	No	No
R^2 ('within' for FE)	0.70	0.56	0.63	0.47	0.76	0.51	0.73	0.08
Observations	82	84	82	84	82	84	82	84

Notes. In all columns, the dependent variable is Total Public Social Expenditures. The independent variables are defined in the text and in the Appendix. All specifications include year-specific effects. Standard errors are clustered by country and are displayed in parentheses. *** Significance at 1%, ** significance at 5%, * significance at 10%. All significance levels are based on t-statistics.

for the rich and the poor, the coefficient of the Meltzer–Richard ratio is insignificant and in many cases has the wrong sign, with or without the additional controls and with or without the FE. Therefore, what changes the sign of the Meltzer–Richard coefficient is that I control for the probability that the median voter may not be the decisive voter by introducing other dimensions of income inequality in the same empirical framework. As the Table shows, the use of country-specific FE makes the Meltzer–Richard coefficient statistically significant but the largest fraction of the decreased magnitude of the coefficient comes from introducing the other inequality ratios into the specification.

Most of the literature does not find evidence in favour of a positive association between the Gini coefficient and redistribution. For example, see Persson and Tabellini (2003), the discussion in Alesina and Glaeser (2004) and the references listed therein. Recently, Shelton (2007) tests carefully several prominent theories on the determinants of government expenditures and their composition, including the median voter theory. Using a random effects estimator, he estimates a close to nil impact of the Gini coefficient on social transfers, when the effect is evaluated at the mean value of political rights. He also finds no significant pattern in any other measure of government spending. A notable exception in the literature is Milanovic (2000) who argues that a careful construction of the relevant variables leads to a positive association between the Gini coefficient and redistribution. To test the Gini coefficient in my sample, I use the factor Gini coefficient from Milanovic's accurate calculations. More often than not, I find no significant relationship between the Gini coefficient and redistribution using various specifications with FE. This result, however, does not compare directly with Milanovic because my sample has fewer observations and, perhaps most importantly, because my measure of redistribution differs from Milanovic's measure (see footnote 4 for this point).

My results are also related to Ramcharan (2010), even though they concern income and not wealth inequality. The author looks at the within-county variation in the US over 1890–1930 and shows how a higher land Gini coefficient is associated with less redistribution. Differently from my article, Ramcharan uses a single statistic to summarise wealth inequality. Similarly to my article, he concludes that the negative association between the Gini coefficient and redistribution may reflect the fact that elites had greater political influence.

2.3. *Sensitivity Analysis*

This Section presents a number of additional robustness exercises. None of the robustness checks overturns the results but in some cases one of the three inequality coefficients loses significance.

2.3.1. *Influential observations*

Given the small number of observations and the correlation between the regressors, the results may be sensitive to the inclusion of some outliers. Figure 2 plots the residuals from the regression of redistribution on the three inequality ratios as reported in the first column of Table 2. These and similar graphs for the other regressors show no obvious outliers. A formal procedure is to examine the standardised residuals from the regression. There are six observations that have standardised residuals above 2. Dropping in isolation or all these observations simultaneously from the sample does not change the results significantly. As a further robustness exercise, I calculate the Cook distance which gives a measure of the total influence of each observation on the predicted values of redistribution. Loosely speaking, observations with a distance greater than 1 merit special consideration. The highest value turns out to be around 0.20. Finally, countries are dropped one at a time from the sample. I find that the results are not sensitive to the inclusion of any particular country.

2.3.2. *Structure of panel errors and country trends*

Thus far I have followed an agnostic strategy and considered errors that are robust to any type of correlation within country. Table 5 considers alternative correlation structures. Panel heteroscedasticity allows the variance of the error term to differ across panels. The contemporaneous correlation of the panel errors is addressed by using time effects that capture common trends. I also consider common cross-country autocorrelation of the error term and different autocorrelations across panels. Another concern is that country-specific time trends to inequality and redistribution could drive the results. To address this concern, the specification is augmented by N country-specific time trends.

I consider various estimation strategies to examine the sensitivity of the results to the assumed behaviour of the error term. In the first three columns of Table 5, I perform feasible generalised least squares (FGLS) estimation. It is well known that FGLS estimation has two disadvantages. First, it requires relatively strong assumptions

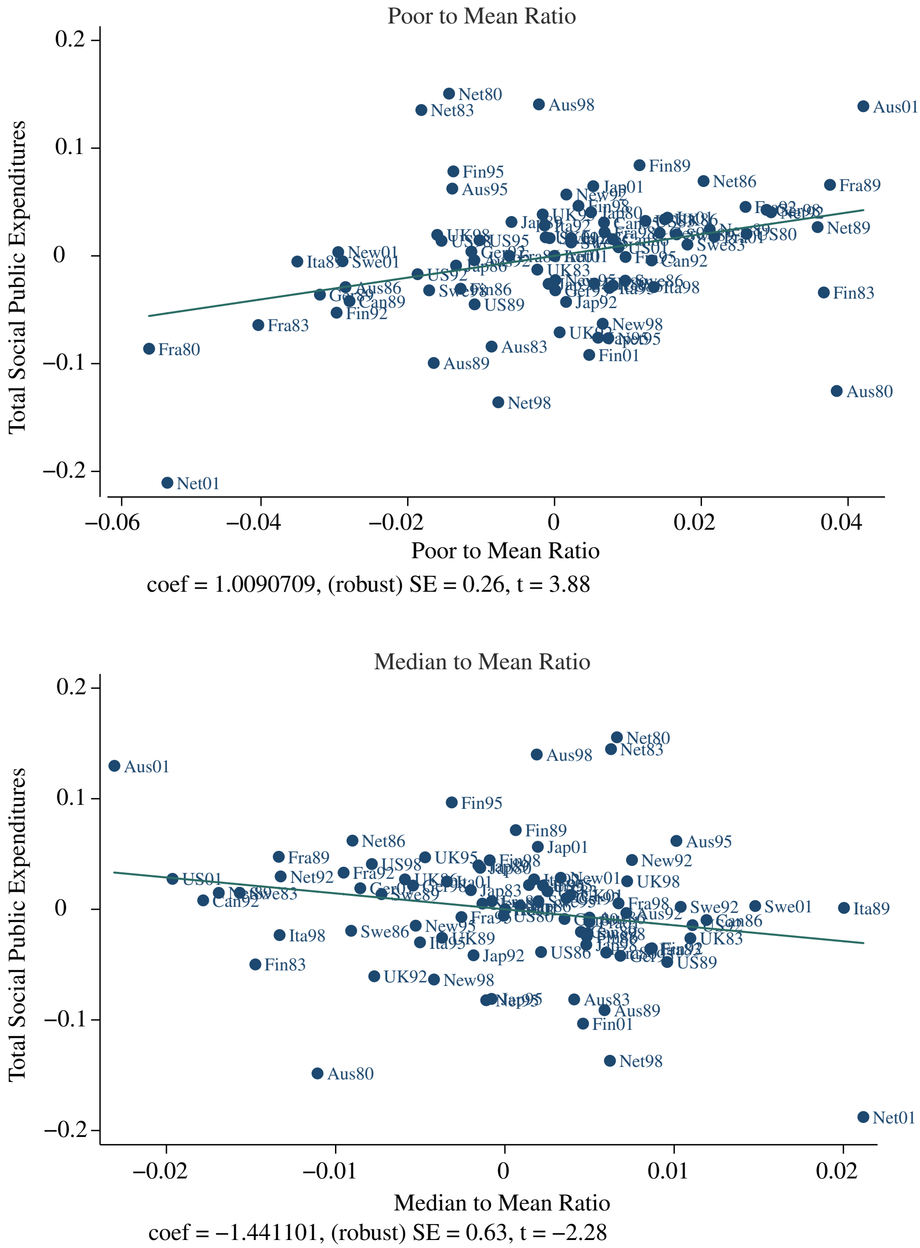

Fig. 2. *Redistribution and Poor, Middle, Rich Income*

Notes. The Figures show the residuals from the conditional relationship between redistribution and the inequality ratios $y^{10}/\bar{y}$, $y^{50}/\bar{y}$ and $y^{90}/\bar{y}$ as estimated in the first column of Table 2.

on the structure of the standard errors. Second, as Beck and Katz (1995) show, it may severely underestimate the variability of the estimates because each round of iteration compounds the uncertainty over the coefficients. Therefore, in the next

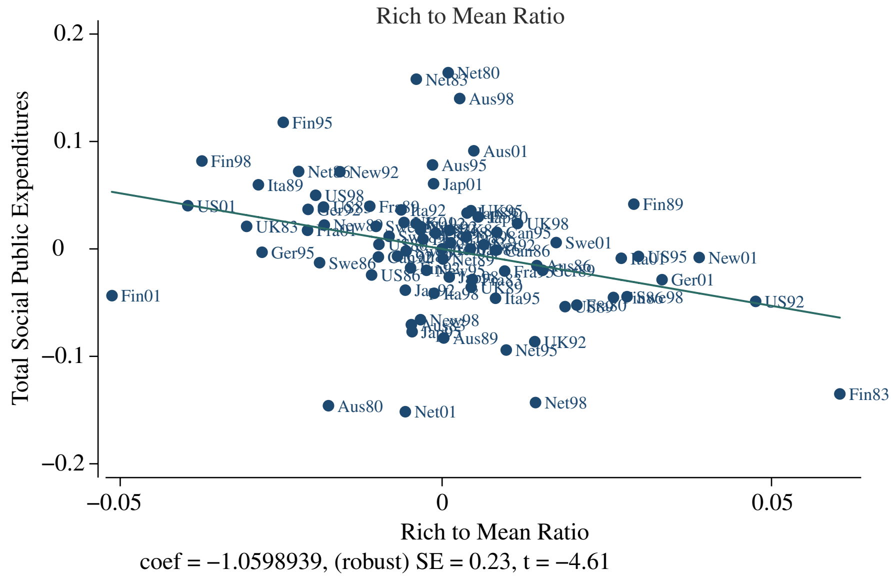

Fig. 2. (*Continued*)

three columns, I use Panel-Corrected Standard Errors which is a strategy more agnostic than FGLS. Finally, in the last two columns, I consider Newey/West standard errors which correct for arbitrary heteroscedasticity and first and second order serial correlation. As Table 5 shows, the results for the coefficients of the rich and the poor remain strong but the coefficient of the median to mean ratio loses significance.

2.3.3. *Persistence of redistribution and endogeneity of inequality*

The effect of inequality on redistribution has not taken into account, thus far, the persistent nature of redistribution. In addition, the consistency of the FE estimator assumes that factors which affect redistribution in period t but are omitted in the error term $\varepsilon_{i,t}$, remain conditionally uncorrelated with past, present and future inequalities. For instance, this assumption is violated if current redistribution affects the future distribution of gross income through labour supply or factor price effects as a result of a sluggish adjustment of the labour market to fiscal policy shocks.

To address these concerns, I model the persistence of redistribution in a dynamic generalised method of moments (GMM) framework. It is well known that when the lagged dependent variable is included into the model, the panel ordinary least squares (POLS) estimator is upward biased (when the error term is positively autocorrelated) and the FE estimator is downward biased by construction (with the bias decreasing as the number of periods increases). In the first two columns of Table 6, the coefficient of the lagged dependent variable is 0.88 for the POLS technique and 0.80 for the FE technique. As a result, the true value of the coefficient lies between these two estimates or at least close to these estimates when taking into account the uncertainty for the point estimates (Bond, 2002).

Table 5

The Structure of Panel Errors and Country-specific Trends

Dependent variable: RED	FGLS (1)	FGLS (2)	FGLS (3)	PCSE (4)	PCSE (5)	PCSE (6)	HAC (7)	HAC (8)
$y^{90}/\bar{y}$	−0.71	−0.77	−0.76	−1.17	−1.27	−1.23	−1.17	−1.17
	(0.32)**	(0.33)**	(0.29)***	(0.39)***	(0.40)***	(0.37)***	(0.64)*	(0.68)*
$y^{50}/\bar{y}$	−0.39	−0.33	−0.25	0.08	0.12	0.36	0.08	0.08
	(0.46)	(0.46)	(0.44)	(0.57)	(0.57)	(0.57)	(0.69)	(0.70)
$y^{10}/\bar{y}$	0.69	0.70	0.70	0.56	0.59	0.49	0.56	0.56
	(0.20)***	(0.20)***	(0.20)***	(0.26)**	(0.25)**	(0.26)*	(0.30)*	(0.30)**
Panel heteroscedasticity	Yes	Yes	Yes	Yes	Yes	Yes		
Autocorrelation	No	Common AR (1)	Panel AR (1)	No	Common AR (1)	Panel AR (1)		
Standard errors							HAC (1)	HAC (2)
Observations	82	80	80	82	82	82	82	82

Notes. In all columns, the dependent variable is Total Public Social Expenditures. All regressions include the controls of the first column of Table 2. The independent variables are defined in the text and in the Appendix. All specifications include year-specific effects, country-specific effects and country-specific time trends. Standard errors are displayed in parenthesis. *** Significance at 1%, ** significance at 5%, * significance at 10%. All significance levels are based on t-statistics. FGLS, feasible generalised least squares estimation; PCSE, panel-corrected standard errors; HAC, heteroscedasticity and autocorrelation robust standard errors. The estimated common AR (1) coefficient is 0.12 in column 2 and 0.11 in column 5.

Table 6

Dynamic Regressions

Dependent variable: RED	POLS (1)	FE (2)	SGMM (3)	SGMM (4)	DGMM (5)	DGMM (6)
RED_{t-1}	0.89	0.80	0.88	0.89	0.80	0.71
	(0.03)***	(0.11)***	(0.04)***	(0.05)***	(0.11)***	(0.14)***
$y^{90}/\bar{y}$	0.09	−0.01	0.22	0.18	−0.01	−0.33
	(0.08)	(0.31)	(0.11)**	(0.14)	(0.28)	(0.36)
$y^{50}/\bar{y}$	−0.52	−0.11	−0.58	−0.64	−0.10	−0.69
	(0.17)***	(0.40)	(0.17)***	(0.16)***	(0.35)	(0.41)*
$y^{10}/\bar{y}$	0.15	0.20	0.13	0.16	0.19	0.38
	(0.07)**	(0.20)	(0.06)**	(0.06)***	(0.18)	(0.27)
p-values			0.99/0.38	0.94/0.38	0.81/0.42	0.56/0.39
Regressors			Pre	Endo	Pre	Endo
Observations	77	77	77	77	63	63

Notes. In all columns, the dependent variable is Total Public Social Expenditures. All specifications include the controls of the first column of Table 4. The independent variables are defined in the text and in the Appendix. All specifications include year-specific effects. The first column does not include country-specific fixed effects, whereas the second column includes country-specific fixed effects. In all other columns, country fixed effects are differenced out. Robust standard errors are displayed in parentheses. *** Significance at 1%, ** significance at 5%, * significance at 10%. POLS, panel ordinary least squares estimation; FE, fixed effects estimation; DGMM, difference generalised method of moments estimation as in Arellano and Bond (1991); SGMM, system generalised method of moments estimation as in Blundell and Bond (1998). The row p-values refer to the null of the hypothesis 'H_0: No Second/Third Order Autocorrelation in the Original Residuals'. The row Regressors indicates 'Pre' when the regressors are treated as predetermined or sequentially exogenous variables and 'Endo' when the regressors are treated as endogenous variables. The matrix of instruments is 'collapsed' (see Roodman, 2006).

Table 6 examines the performance of the model under two estimation procedures, namely the system GMM method of Blundell and Bond (1998) in columns 3 and 4, and the difference GMM method of Arellano and Bond (1991) in columns 5 and 6.[19] The system estimator differs from the difference estimator because it stacks together the first differenced equation and the level equation in a system of equations. In columns 3 and 5, all regressors are treated as 'predetermined' or 'sequentially exogenous' variables, that is the model is estimated under the assumption that $\varepsilon_{i,t}$ can be correlated with future regressors but $\varepsilon_{i,t}$ remains orthogonal to contemporaneous regressors. Since I have already lagged the inequality ratios by one period, this implies that $\varepsilon_{i,t}$ can be correlated with inequality in period t and in other future periods but $\varepsilon_{i,t}$ remains orthogonal to inequality in period $t-1$ and in earlier periods. In this case, valid instruments are first and deeper lags of the instrumenting variable for the differenced equation and, for the system GMM, the zero lag of the instrumenting variable in differences for the levels equation. In columns 4 and 6, all regressors are treated as 'endogenous' variables, that is the model is estimated under the assumption that $\varepsilon_{i,t}$ can be correlated with future and contemporaneous regressors but $\varepsilon_{i,t}$ remains orthogonal

[19] Although relaxing the strict exogeneity assumption is promising, the problem is that too many instruments may severely overfit the model and dynamic panel data estimators are not expected to behave perfectly for small N panels (Roodman, 2006). For instance, one common problem is that J-tests lack power because the number of instruments exceeds the total number of groups. In this case, the Hansen/Sargan test for overidentification tends to over-accept the null hypothesis, with p-values of 1.00 being common. Notwithstanding these caveats, a sound robustness practice is to consider different modelling strategies and examine the sensitivity of the results to the assumed specification.

to past regressors. Since I have already lagged the inequality ratios by one period, this implies that $\varepsilon_{i,t}$ can be correlated with inequality in period $t-1$ and in other future periods but $\varepsilon_{i,t}$ remains orthogonal to inequality in period $t-2$ and in earlier periods. In this case, valid instruments are second and deeper lags of the instrumenting variable for the differenced equation and, for the System GMM, the first lag of the instrumenting variable in differences for the levels equation.

Table 6 shows that the estimated coefficient of the lagged dependent variable in the System GMM method is close to the interval of true values (as implied by the FE and the POLS techniques). The difference GMM method, however, appears to deliver downward biased estimates.[20] As the Table shows, the results remain robust for the poor and the median but not for the rich. Specifically, the coefficients of the $y^{10}/\bar{y}$ and the $y^{50}/\bar{y}$ variables have the correct signs under both methods and they are always significant under the system GMM method. The coefficient of the $y^{90}/\bar{y}$ variable has the correct sign but it loses significance under the difference GMM method and it has the wrong sign under the system GMM method.

3. Discussion

This Section discusses and interprets the results. First, I define the one dollar, one vote equilibrium and explain why the empirical results are consistent with this definition. Then, I propose some mechanisms that can explain the one dollar, one vote result, with a focus on the channel of political participation. To conclude, I present the implications of my results for the US–Europe ‘paradox of redistribution’.

3.1. *One Dollar, One Vote*

I define the one dollar, one vote equilibrium as a politico-economic equilibrium in which aggregate outcomes tilt towards the preferences of a specific group of citizens as this group’s income increases (relative to the mean income).

The empirical results in Section 2 are consistent with this definition. To see this, start with the well-known theoretical proposition that an individual’s most preferred size of redistribution is inversely related to individual income (Romer, 1975; Roberts, 1977). This proposition finds strong empirical support. Alesina and La Ferrara (2005) show that richer individuals prefer less redistributive policies for the case of the US. I extend their exercise for my panel of OECD countries using individual-level data from the WVS Dataset and specifically from its Four-Wave Integrated Data File (2006) over 1981–2004. Table 7 presents summary statistics and the Appendix contains more details for the variables.

Support for redistribution is measured with two variables. The first codes answers that range from 1 if respondents think that ‘Incomes should be made more equal’ to 10 if they believe that ‘We need larger income differences as incentives’. The second ranges from 1 when the respondent believes that it is ‘Very important to eliminate big

[20] This result is consistent with Blundell and Bond (1998) because inequality and redistribution are fairly persistent processes and the number of periods is small in which case the bias of the FE estimator can be important. Bobba and Coviello (2007) apply this argument in the democracy and education literature and show the downward bias of the difference GMM estimator.

Table 7

Summary Statistics of Individual-Level Data

Variable	Mean	SD	Min	Max	Observations
MOREEQ	0.46	0.49	0	1	41,037
ELIMINEQ	0.60	0.49	0	1	11,993
Discuss with friends	0.69	0.45	0	1	57,965
Politics important	0.44	0.49	0	1	43,472
Sign petitions	0.61	0.49	0	1	56,216
Income	5.38	2.68	1	10	48,203
Female	0.53	0.49	0	1	58,365
Married	0.58	0.49	0	1	58,117
Age	43.92	17.31	15	97	58,200
Children	0.76	0.42	0	1	58,445
No high school	0.24	0.42	0	1	26,164
College	0.14	0.34	0	1	26,164

Notes. These are summary statistics for the variables taken from the WVS. The sample is restricted to the 14 OECD countries in the macro-sample. See the Appendix for the definition of each variable.

income differences' to 5 if it is 'Not at all important'. To ease the interpretation, I recode the variables in binary form with 1 denoting more support for redistribution and 0 denoting less support for redistribution. The results are not sensitive to the specific recoding procedure. The two variables are called MOREEQ and ELIMINEQ.

The survey codes pre-tax income in 10 scales which are standardised at the country level. A higher value for the income variable means that the respondent belongs to a household of a higher income scale. The regressions control for gender, age, the presence of children, marital status and education. I assume that support for redistribution of individual i living in country c at time t is characterised by a latent variable:

$$R^*_{ict} = d_c + \tau_t + \beta Y_{ict} + \gamma \boldsymbol{X}_{ict} + \varepsilon_{ict}, \tag{2}$$

where d_c and τ_t are country and year dummies, Y_{ict} is the income scale of the respondent, $\boldsymbol{X}_{ict}$ is the vector of controls and ε_{ict} is the error term. The variable R^*_{ict} is not observed but the variable R_{ict}, taking a value of 0 for low support and a value of 1 for high support for redistribution, is observed. Assuming that the distribution of the error term is logistic, I estimate a logit model.

The first two columns of Table 8 show that support for redistribution decreases with income. The point estimates suggest that the likelihood of finding someone who supports redistribution decreases by around 2.1% each time we move up one income scale. I have repeated this regression for all countries separately. In all cases, the point estimate for income is negative and most of the times the coefficient is significant at the 1% level.[21] Restricting the sample for specific time periods or estimating an ordered logit model makes no difference for the results.

Since support for redistribution decreases as income increases, the poor prefer more redistribution than the median and the median prefers more redistribution than the rich. When the initial equilibrium redistribution lies between the bliss points of the poor and the rich, the result in Section 2 that redistribution increases in the income of

[21] The two exceptions (across 28 specifications) are the coefficient in the MOREEQ regression for Germany and the coefficient in the ELIMINEQ regression for France.

Table 8

Support for Redistribution and Political Participation

Dependent variable	MOREEQ (1)	ELIMINEQ (2)	Politics important (3)	Discuss politics (4)	Petitions (5)	Politics important (6)	Discuss politics (7)	Petitions (8)
Income	−0.021	−0.022	0.010	0.016	0.015	0.011	0.019	0.024
	(0.002)***	(0.003)***	(0.002)***	(0.002)***	(0.002)***	(0.007)	(0.006)***	(0.006)***
Income squared						−0.0001	−0.0003	−0.0008
						(0.0006)	(0.0005)	(0.0005)
Female	0.053	0.029	−0.028	−0.085	0.031	−0.028	−0.085	0.031
	(0.008)***	(0.011)***	(0.010)***	(0.07)***	(0.008)***	(0.010)***	(0.07)***	(0.008)***
Married	−0.025	−0.021	0.027	0.032	0.012	0.027	0.032	0.011
	(0.010)**	(0.015)	(0.008)***	(0.009)***	(0.009)	(0.008)***	(0.009)***	(0.009)
Age	−0.000	0.002	0.005	0.003	0.000	0.004	0.003	0.000
	(0.0003)	(0.0004)***	(0.001)***	(0.001)***	(0.001)	(0.001)***	(0.001)***	(0.001)
Children	0.000	0.016	−0.047	−0.020	0.024	−0.047	−0.020	0.024
	(0.011)	(0.017)	(0.010)***	(0.008)***	(0.008)***	(0.010)***	(0.008)***	(0.008)***
No high school	0.069	0.027	−0.120	−0.146	−0.134	−0.120	−0.146	−0.133
	(0.012)***	(0.014)*	(0.011)***	(0.010)***	(0.012)***	(0.011)***	(0.010)***	(0.012)***
College	−0.036	−0.024	0.091	0.115	(0.110)	0.092	0.116	(0.111)
	(0.013)***	(0.020)	(0.012)***	(0.010)**	(0.012)***	(0.012)***	(0.010)**	(0.012)***
Observations	17,425	8764	19,334	20,136	19,899	19,334	20,136	19,899

Notes. All coefficients are marginal effects (e.g. the coefficient −0.021 in the first column for income denotes that the probability that a respondent answers *MOREEQ* = 1 decreases by 2.1% when the respondent's income increases by one scale). The independent variables are defined in the text and the Appendix. All specifications include year- and country-specific fixed effects. Standard errors are clustered at the town size level and are displayed in parentheses. *** Significance at 1%, ** significance at 5%, * significance at 10%.

the poor implies that as the poor become richer, aggregate spending tilts closer to their most preferred outcome. When median income increases relative to mean income, the median's net benefit from redistribution decreases. The one dollar, one vote definition holds in the Meltzer and Richard (1981) model since the median is always the decisive voter. This system also holds in an extended framework in which the median is not exclusively the decisive voter but in which the median's decisiveness increases with its relative income. In either framework, the median prefers less redistribution and also has the political power to impose less redistribution as an equilibrium outcome. As a result, redistribution decreases which is consistent with the results of Section 2.[22] Finally, it is easy to see that the one dollar, one vote definition is also satisfied for the high-income group since the rich in general prefer less redistribution.

3.2. *Explanations*

When the rich become richer, their most preferred size of redistribution decreases. As a result, the one dollar, one vote definition is satisfied if the rich's political influence increases with income. In this case, the rich group will impose less redistribution as an equilibrium outcome shifting the politico-economic outcome closer to its preferences. On theoretical grounds, the positive association between income and pivotal role of the rich can be explained by the ideas developed by Grossman and Helpman (1996). As a richer rich class has more to lose from redistribution, the incentive for the rich to contribute to political parties and tilt policy closer to their preferences is expected to increase with income. Bénabou (1996) uses this general insight to introduce a non-monotonic relationship between inequality and redistribution. Campante (2008) lays out a model that describes this behaviour, with specific reference to campaign contributions. Bartels (2009) analyses the responsiveness of US senators to the preferences of their constituents and finds that the voting behaviour of senators is considerably more responsive to the opinions of the wealthy.

When we contrast the individual-level evidence with the panel outcomes of Section 2, we see a one-to-one correspondence between preferences for redistribution and aggregate policies for the middle and the rich class: both groups prefer less redistribution as they become richer and aggregate spending decreases in the earnings of these two groups. In contrast, using individual-level data, I find no evidence that the poor prefer more redistribution when they get richer but in the country-level data a more affluent poor class is associated with a broader welfare state. To understand this, consider the implications of the inverse relationship between income and preference

[22] Epple and Romano (1996) argue that private alternatives to public services may introduce an equilibrium in which the median income voter is not decisive. Instead, the decisive voter has an income that lies below the median. They describe this finding as an 'ends against the middle' result in which high-income voters join the poor to oppose public spending. In the Epple and Romano (1996) world, my finding that the median to mean ratio is negatively associated with redistribution can be interpreted as an increase in the decisiveness of the 'middle' relative to the political influence of the 'ends'. The caveat is that the model applies mostly to expenditures with private alternatives, such as health and education, and not so much to expenditures like welfare benefits for the poor. See, however, Casamatta *et al.* (2000) for an 'ends against the middle' result in the provision of pensions. See also Bellettini and Berti Ceroni (2007) for an 'ends against the middle' result that does not rely on the existence of private alternatives, but instead it is based on borrowing constraints. Interestingly, the authors also find a positive coefficient of the poor and a negative coefficient of the median but in their regressions the dependent variable is educational spending.

for redistribution for the poor. The first implication is that in general the poor prefer less redistribution when they get richer. The second implication is that the poor's bliss point always involves more redistribution than that of the other groups.. The evidence using country-level data that redistribution increases in poor's gross earnings implies that the second effect dominates the first. In other words, in response to a richer low-income group policy moves closer to the poor's most preferred outcome which involves more redistribution than the redistribution mostly preferred by the other groups. This is the essence of the one dollar, one vote definition.

The question then becomes why policy moves closer to the poor's bliss point as their income increases. One natural explanation is that political participation increases with income.[23] Rosenstone and Hansen (1993) and McCarty *et al.* (2006) find a strong positive association between income and political participation in the US. To verify that these results are not special to the US, I present evidence from the WVS Dataset.

Political participation is proxied by 14 different indicators (see Appendix, for more details).[24] Variables are in binary form such that 1 means higher political participation. I assume that political participation of individual i, living in country c at time t is characterised by a latent variable:

$$P^*_{ict} = d_c + \tau_t + \beta Y_{ict} + \boldsymbol{\gamma} \mathbf{X}_{ict} + \varepsilon_{ict} \tag{3}$$

where d_c and τ_t are country and time dummies, Y_{ict} is the income scale of the respondent, $\mathbf{X}_{ict}$ is the vector of controls used in equation (2) and ε_{ict} is the error term. I do not observe P^*_{ict}, but instead I observe a variable P_{ict}, which equals 0 for low and 1 for high political participation.

Columns 3–5 in Table 8 present strong evidence that political participation increases with income. The Table presents results for the three variables with the highest number of observations. The point estimates for income's coefficient in all 14 specifications are positive and with one exception all coefficients are significant at the 1% level. The estimates presented in the Table indicate that the likelihood of participating politically each time we move up one income scale increases by 1.0–1.6%. The absolute cost barrier to political participation implied by these estimates is important, especially since I control for other proxies of income including education.[25] Finally, in columns 6–8, I introduce a quadratic specification for income to investigate how the effect of income on political participation depends on income. As the negative coefficient of

[23] The political participation explanation is expected to become even more powerful when examining the relationship between wealth inequality and redistribution. See Perotti (1996) for some discussion of wealth inequality and how it relates to redistribution and growth, and Deininger and Squire (1998) for empirical evidence on the relationship between land inequality and growth. See Ramcharan (2010) for the relationship between land inequality and redistribution in the US counties over 1890–1930.

[24] These are: Interest in Politics; Participation in Local Political Acts; Belong to Political Party; Join Boycotts; Sign Petitions; Participation in Lawful Demonstration; Adherence to Unofficial Strikes; Politics Important in Life; Discussion of Political Matters with Friends; Unpaid Work for Political Parties; Unpaid Work for Local Political Acts; Unpaid Work for Labour Unions; Belong to Labour Union; Active in Labour Union.

[25] Consider, as an example, the case of US in 1999. A 1% increase in the probability to sign a petition is associated with approximately $4,700. A respondent belonging to a household with yearly pre-tax income of $30,000 has approximately 10% lower probability to sign a petition than a respondent with similar demographic and educational characteristics who belongs to a household that earns $75,000. The former is approximately 7% less likely to be affiliated with a political party than the latter.

squared income shows, the effect of income on political participation is stronger for the poor than for the rich. The results presented in columns 3–8 of Table 8 are robust when restricting the sample to specific countries or time periods.

To summarise, these findings indicate that income is positively associated with political participation across the income spectrum. The poor do not contribute cash in campaigns, however, they do engage in other activities that influence policy and engage in these activities more as their income increases. A richer poor class demonstrates more and participates in political parties, voluntary political clubs and labour unions. If the stronger political voice of the poor dominates their decreased preference for redistribution, then redistribution increases in poor's income.[26]

The political participation explanation assumes that policy is sufficiently elastic to rich's and poor's political demands, that is, that the position of the decisive voter is not rigid. I also consider some alternative explanations that can rationalise why redistribution tilts towards the preferences of groups of citizens with higher income without assuming the reallocation of political power.

The theories of Becker (1983) and Becker and Mulligan (2003) offer an explanation for the rich's coefficient. With a progressive system of taxes, the distortionary effects of labour income taxation are higher at the right tail of the income distribution. Therefore, a richer high-income class could be associated with less redistribution if the median voter internalises these increasing distortions when demanding redistribution.

The 'social affinity theory' of Kristov *et al.* (1992) can potentially rationalise the effects at the left tail of the distribution. The authors develop a model in which political agents organise into pressure groups and devote resources to affect redistribution, as in Peltzman (1980). The equilibrium redistribution is determined in the following manner: one feels closer to people that have a similar level of income. The theory predicts a positive sign for the gap between the poor's and the middle's earnings coefficients in a regression of redistribution on this income gap.[27] I note that the authors allow for the interpretation that social affinity operates through social mobility as in Bénabou and Ok (2001). Therefore, when the median income gets closer to the income of the poor, redistribution increases.[28]

An alternative plausible explanation comes from relaxing the assumption of unidimensionality in the Meltzer and Richard (1981) model. On both sides of the Atlantic, modern welfare states include means-tested redistributive programmes. Examples of such programmes are the Medicaid and Food-Stamps in the US, family and solidarity

[26] Models which are reduced to the maximisation of a weighted average of citizens' utility, such as the utilitarian model or its positive version, the probabilistic voting model, typically do not explain the poor's positive coefficient. The reason is that as poor's income increases, the 'Pareto weight' attached to the poor class declines and hence the planner or the politician tilts redistribution closer to the median's or the rich's most preferred policy. This result holds under general conditions and is the essence of utilitarian redistribution (see Hellwig, 1986, for these conditions). The channel of political participation offers a microfoundation that supports a positive relationship between the 'Pareto weight' and income.

[27] The estimated coefficient of the y^{10}/y^{50} variable is of similar magnitude to the estimated coefficient of the $y^{10}/\bar{y}$ variable. As a result, the macro data are not sufficiently informative to distinguish between the political participation explanation and the social affinity theory. See also Lindert (1996) for careful tests of the social affinity theory over 1960–1981.

[28] See also Grüner and Schils (2007) who establish an unconventional voting behaviour of the politically decisive middle class in a model of wealth redistribution. Specifically, the authors show how middle class voters form a coalition with the poor when redistribution gives access to entrepreneurial rents.

allowances in many European countries and parts of the medical insurance in France. The statutory nature of redistribution provides an explanation for the results because when transfers are declining with income, a richer poor class becomes less eligible for targeted transfers and a higher fraction of the tax revenues can be appropriated by the middle class. If political power remains in the hands of the median class, then the median's increased benefit from receiving transfers will change the politico-economic trade-off in favour of more redistribution.

3.3. *The US–Europe Difference in Redistribution*

In the Introduction, I referred to the striking divergence of the data from the theory to motivate the widely held belief that inequality does not matter for redistribution. In light of my results it is worthwhile to revisit this issue. Consider the US *versus* a European aggregate of Finland, France, Netherlands and Sweden (two Continental and two Nordic welfare states).[29] The European aggregate redistributed approximately 26.4% of its GDP in 2001, compared to 23.8% in 1980. US redistribution increased from 13.3% in 1980 to 14.4% in 2001. As a result, the European growth of redistribution exceeded the US growth by approximately 2.7% (not percentage points) over 1980–2001.

Consider, as an illustrative example, the estimates of column 9 in Table 3 (which hold constant the minimum wage). Over 1977–98, the growth of the $y^{10}/\bar{y}$ ratio is approximately −2% in Europe but almost −19% in the US. Multiplying the difference (16.8%) with the estimated coefficient of the poor leads to an approximately 15.9% change in redistribution. The growth of the $y^{50}/\bar{y}$ ratio is more than −1% in Europe but less than −11% in the US. Multiplying the difference (10.6%) with the estimated coefficient of the median leads to an approximately −15.4% change in redistribution. According to these estimates, Europe increased redistribution relative to the US in 2001 relative to 1980 because the European poor did not become as poor as the American poor (both relative to their mean) and US increased redistribution relative to Europe because the American median voter became poorer. These two opposing effects tend to cancel out. Finally, the growth of the $y^{90}/\bar{y}$ ratio is −2.9% in Europe and 0.6% in the US. Multiplying the difference (−3.5%) by the estimated coefficient of the rich leads to an approximately 3.4% change in redistribution.

Assuming that the position of the pivotal voter is not rigid over this period of time, the conclusion from these calculations is that the faster growth of the rich's income in the US allowed the rich class to devote resources and tilt policy closer to their bliss point which involves less progressive policies. As a result, the growth of redistribution in the US lagged the growth of redistribution in Europe.[30]

4. Conclusion

The results of this article challenge the conventional view that income inequality does not matter for redistribution. Instead of focusing on a single summary inequality

[29] Germany and Italy have missing observations and cannot be included in this calculation.
[30] Actually, these changes in inequality over-predict the difference in the growth of redistribution across the Atlantic (but I have not factored in the other determinants of redistribution).

statistic, I show how looking at different parts of the income distribution offers an explanation for the observed cross-country differences in redistribution. In a panel of OECD countries over 1975–2001, I find that when a group of voters becomes richer (relative to the mean), redistribution tilts closer to its bliss point. This result is consistent with the argument that more money is associated with more political power. The idea that political influence is not uniform across voters explains why the literature has found scant evidence in favour of the basic politico-economic model of inequality and redistribution.

Appendix

Country-level Data

Redistribution

I use the Total Gross Public Expenditure variable from the SOCX Dataset of OECD. It includes expenditures on old age (pensions), survivors, incapacity related benefits, health, family, active labour market programs, unemployment, housing and other social policies. The data for the construction of redistribution net of health and pensions in Table 2 is also available at the SOCX Dataset. Redistribution is expressed as a percentage of GDP. The variable Tax Wedge is taken from the Allard and Lindert (2006) Dataset. The variable is defined as the sum of personal income tax and employee plus employer social security contributions together with any payroll tax less cash transfers for the average worker. This is expressed as a percentage of labour costs.

Inequality

The inequality variables are from OECD's Labour Force Statistics Dataset. I quote the description of the variables from OECD: 'This [...] contains gross earnings of full-time workers by earnings percentiles and mean earnings, in national currency units. The series are a mixture of hourly, daily, weekly, monthly, and annual earnings and are specified in the country notes. These data were first collected and used in the tables, charts, and analysis on earnings dispersion presented in various editions of the OECD Employment Outlook: 1993 (Chapter 5), 1996 (Chapter 3), 1997 (Chapter 2), 1998 (Chapter 2)'.

Wage coordination index

Taken from Kenworthy (2001). The coordination index takes values from 1 which denotes fragmented wage bargaining, a bargaining process confined mostly to large enterprises and plants, to 5 which denotes centralised bargaining by peak confederations or government imposition of wage schedules.

Share of elderly, surplus, voter turnout, left and right vote, presidentialism and single member district electoral system, mean income/wage gap, minimum wage

All these variables are taken from the LIS-CWS Dataset (Huber *et al.* 2004). The variable, share of elderly, measures the percentage of the population older than 65 years. The variable government surplus is expressed as a fraction of GDP. Voter turnout refers to the percentage of the population that voted in the last elections. Left and right vote are expressed as fractions of total votes directed towards leftist and rightist parties respectively in the last elections. For the interaction of the inequality variables with institutions, I use two variables. The variable presidentialism is a binary dummy and the single member district variable takes the value 0 under proportional representation, 1 under modified proportional representation and 2 under plurality systems. See the LIS-CWS Dataset (2004) for the classification of the parties into left, central and right and the definition of the constitutional variables. The mean income/wage gap is defined as the ratio of the mean income of wage and salaried employees in national currency units at current prices to the average earnings of production workers in national currency units at current prices. Finally, the variable minimum wage is defined as the ratio of the minimum wage to the average production wage.

Openness, growth and GDP
These variables are taken from the Penn Tables, Version 6.2 of Heston *et al.* (2006). Openness is defined as exports plus imports as a percent of GDP. Growth is the growth rate of real GDP per capita in constant 2,000 prices. GDP is real GDP per capita in 1996 International Dollars (Laspeyers).

Unemployment rate and employment ratio
The unemployment rate variable is taken from OECD's Dataset on Labour Market Statistics and is expressed in percentage points. The employment ratio is defined as the proportion of an economy's working-age population that is employed and it is taken from the OECD's LFS Dataset. It is also expressed in percentage points.

Ethnolinguistic fractionalisation
The ELF index is taken from Roeder (2001). The ELF index is defined as one minus the probability that two randomly chosen persons from a population belong to the same ethnic, linguistic or racial group. A higher ELF index denotes a more heterogeneous population.

Deadweight loss (DWL)
The DWL variable is constructed by dividing the top corporate marginal tax rate by government's share in GDP. The former is taken from the World Tax Dataset and the latter from the Penn Tables, Version 6.2. A higher value for DWL denotes higher distortions per unit of government activity.

Share of labour taxes
Share of labour taxes is defined as the product of the 'effective' tax rate with wage and salaries divided by final government consumption. The effective tax rate is taken from Mendoza *et al.* (1994) and McDaniel (2007). The variable, wages and salaries, is taken from OECD. The variable, final government consumption, is also taken from OECD.

Individual-level Data

The individual-level data are taken from the WVS.

Support for redistribution
The first question (*E*035) asks respondents to choose a number from 1 if they believe that 'Incomes should be made more equal' to 10 if 'We need larger income differences as incentives'. The second question (*E*146) asks to rank from 'Very Important' (1) to 'Not at all Important' (5) the statement 'How important it is to eliminate big income inequalities between citizens?' In both cases, the variables are recoded to take higher values for more support for redistribution. MOREEQ takes the value 1 if $11 - E035 > 5$ and 0 otherwise. ELIMINEQ takes the value 1 if $6 - E036 > 3$ and 0 otherwise. The results are not sensitive to the construction of the cutoff.

Political participation
All variables are recoded such that higher values denote more political participation. Some variables in the WVS are binary. Non-binary variables are recoded into a binary form to ease the

comparison. The 14 variables that are used are: politics important in life (*A*004; 1 if *A*004 = 1,2), discussion of political matters with friends (*A*062; 1 if *A*062 = 1,2), belong to labour union (*A*067; binary), belong to political party (*A*068; binary), participation in local political acts (*A*069; binary), unpaid work for labour unions (*A*084; binary), unpaid work for political parties (*A*085; binary), unpaid work for local political acts (*A*086; binary), active in labour union (*A*101; 1 if *A*101 = 2). Interest in politics (*E*023; 1 if *E*023 = 1,2), sign petitions (*E*025; 1 if *E*025 = 1), join boycotts (*E*026; 1 if *E*026 = 1), participation in lawful demonstration (*E*027; 1 if *E*027 = 1) and adherence to unofficial strike (*E*028; 1 if *E*028 = 1).

Income and other socioeconomic controls

Each respondent is placed in a country-specific income scale (*X*047). Female (*X*001) takes the value 1 if the respondent is female and 0 otherwise, Age (*X*003) measures the age of the respondent, children (*X*011) takes the value 1 if the respondent has children and 0 otherwise, married (*X*007) takes the value 1 if the individual is married and 0 otherwise, no high school (*X*025) takes the value 1 if the respondent has not attended the equivalent of high school in the US and 0 otherwise. College (*X*025) takes the value 1 if the individual has completed at least college education and 0 otherwise.

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