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## Contrasting rational choice and behavioural perspectives

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# INCOME INEQUALITY, REDISTRIBUTION AND POVERTY: Contrasting rational choice and behavioural perspectives

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## Abstract

Based on the standard axiom of individual utility maximization, rational choice has postulated that higher income inequality translates into greater redistribution by shaping the median voter's preferences. While numerous papers have tested this proposition, the literature has remained divided over the appropriate measure for redistribution. Revisiting the original contribution by Meltzer and Richard, the present paper argues that the median voter hypothesis implies that relative redistribution should increase in line with inequality. An empirical test based on 110 observations from the Luxembourg Income Study fails to find any support for the hypothesis. By contrast, voters' actual preferences offer a better guide to understanding redistributive outcomes. The findings challenge the narrow concept of human motivation that underpins rational choice, and point to the importance of fairness orientations that have been emphasized in behavioural economics.

Keywords: income distribution, redistribution, median voter theorem, behavioural economics

JEL classification: D31; D03; H23; H55

## 1. Introduction: redistribution and poverty

At the danger of over-simplification, income poverty is a function of two factors: the level of average incomes and their distribution between households and persons. Holding income levels constant, poverty will generally be more severe when incomes are distributed more unevenly (see e.g. Kanbur 2005). Consequently, countries with comparable income levels can have very different outcomes in terms of poverty incidence and depth. While economic growth increases the level of average incomes, it is generally more effective in alleviating poverty when the initial distribution of incomes is more equitable or when it is accompanied with a reduction in inequality (White 2001; Dagdeviren et al. 2002). Even though growth has helped to reduce poverty in a large number of countries since the mid-1990s, Fosu (2011) concludes in his recent review of poverty trends that 'further progress could have occurred under [a] relatively [more] favourable income distribution'.

It is thus not surprising that redistribution, broadly defined as the use of tax and transfer policies to reduce income inequality, has re-entered the mainstream of the poverty debate—much like income inequality itself has been 'brought in from the cold' by the economics discipline in the mid-1990s (Atkinson 1997; see also Kanbur and Lustig 2000). Whereas redistributive instruments are generally more developed in the advanced countries—where relative poverty has remained a policy concern—developing countries such as Brazil are now using cash-transfer programmes (along with other policy tools, such as minimum wage legislation) to reduce poverty and to put a dent into sky-high inequality. By contrast, tax and transfer systems have only a negligible impact on inequality in other Latin American countries such as Guatemala or Columbia. Even among developed economies, the welfare-state literature has found a wide gulf between the redistributive efforts made in Nordic countries and in the liberal market economies of the Anglo-Saxon world (see e.g. Korpi and Palme 1998).

What explains these differences in the extent of redistribution? Mainstream rational choice theory has postulated an automatism under which higher initial income inequality will lead to higher redistribution. This would be good news for those concerned with poverty eradication, since redistribution would be in greater supply precisely where it is needed most to redress inequities generated by the market and the social context in which it operates. In an influential paper, Meltzer and Richard (1981) have argued that the median voter's interest in redistribution will be greater in more unequal societies. Since self-interested politicians want to maximize their chance of gaining or retaining power, they will strive to translate the median voter's preferences into policy action. In democratic politics, this mechanism should translate higher initial inequality into higher subsequent redistribution.[2]

The Meltzer-Richard hypothesis, as the proposition has become known, draws on the standard assumptions of rational choice—individuals are rational actors who maximize their own, narrowly defined utility—and relies on methodological individualism to extrapolate from the postulated (rather than observed) individual behaviour to predict developments at the macro-level. Behavioural economics has found many of these assumptions wanting, and pointed to the bounds of rationality (see for example Kahneman and Tversky 1984). By drawing on insights from neighbouring disciplines (that range from psychology to political sociology), it has also questioned the narrow definition of utility as material gain. While the simplistic concept of human motivation makes the agents of rational choice theory behave like 'rational fools' (see Sen 1977), the well-established research on social alignments and value orientations offers a more nuanced understanding of individual voting behaviour (see Lipset and Rokkan 1967; Flanagan 1987; Knutsen 1995; Dalton 1996).

[2] Romer (1975) and Roberts (1977) had made similar arguments earlier, and all modern median voter theories of course find their intellectual heritage in Schumpeter's (1942) seminal book 'Capitalism, Socialism and Democracy' and Downs' (1957) volume 'An economic theory of democracy'.

The hypothesis developed by Meltzer and Richard is readily testable: Do more unequal societies redistribute more? Unsurprisingly, many papers have sought to address this question (for example, Milanovic 2000; Kenworthy and Pontusson 2005; de Mello and Tiongson 2006; Lupu and Pontusson 2011).[3] Overall, the literature has arrived at the unsatisfactory conclusion that the answer partly depends on how 'redistribution' is defined. By and large, papers that look into 'absolute redistribution' (the absolute reduction in the Gini coefficient) concluded that more unequal societies, indeed, redistribute more (see e.g. Kenworthy and Pontusson 2005). By contrast, papers that have measured 'relative redistribution' (the reduction of the Gini coefficient relative to its initial level) have not found any correlation between market inequality and subsequent redistribution (see e.g. de Mello and Tiongson 2006; Lupu and Pontusson 2011). Mixed findings have also emerged from a related body of literature on welfare spending, and tests of reduced-form models in the new growth literature of the 1990s (see e.g. Persson and Tabellini 1994; Perotti 1996; and Bassett et al. 1999).

In this context, the present paper aims to the make the following contributions: (1) As a contribution to theory, it revisits the original paper by Meltzer and Richard to deduct a valid test with the appropriate measure for redistribution. (2) As a contribution to econometric analysis, it uses an expanded dataset from the Luxembourg Income Study (LIS) to test the hypothesis. (3) And lastly, the paper explores behavioural approaches to understanding support for redistribution, namely the observed preferences of voters for equity and redistribution, and submits this alternative explanation to an empirical test. The paper concludes by reviewing the utility of the two different approaches and discusses their commonalities and limitations.

There are a number of issues this paper will not address. For instance, one could argue that incomes are flows that accrue from a stock of assets (be it physical or human capital). What would the rational (or indeed the actual) voter have to say about asset redistribution, especially since wealth inequality is even greater than income inequality (Davies 2008)? Likewise, the relative returns to capital and labour—and to labour of different skill levels—are not god-given, so pre-tax pre-transfer inequality itself can be shaped by policies that fall short of outright asset redistribution (such as strong collective bargaining rights or minimum wages that set a wage floor). Further, there is an interesting debate on the trade-offs between different transfer schemes and their effectiveness in terms of poverty reduction. Does targeting imply that political support for them will be weaker, and hence that benefits will be stingier than in universal schemes? And finally, poverty is not only a function of incomes, but a broader phenomenon of social exclusion and the lack of rights and entitlements. All these questions warrant debate, but this paper will set them aside to focus on a much narrower topic: the redistribution of incomes through taxes and transfers.

## 2. Revisiting the Meltzer-Richard hypothesis

While several studies have sought to establish a relationship between pre-government inequality and the extent of redistribution, most of them have suffered from the lack of reliable data for market inequality and/or used proxy variables for redistribution, such as the size of social expenditures or public transfers.[4] Mahler and Jesuit (2006) were among the first to provide reliable cross-country time-series data for both concepts on the basis of the LIS. They report the Gini index for the inequality of private sector incomes,[5] which presents the desired measure of the initial distribution of incomes (i.e. before taxes and transfers), as well as for the distribution of disposable incomes (i.e.

[3] There is a further body of literature that has sought to exploit differences between regions within the same country to test the Meltzer and Richard hypothesis, which are numerous to cite.

[4] See e.g. Perotti (1996), Moene and Wallerstein (2003) and de Mello and Tiongson (2006); notable exceptions are Milanovic (2000), Bradley et al. (2003) and Kenworthy and Pontusson (2005).

[5] This paper uses the terms 'market incomes' and 'private sector incomes' interchangeably to describe all pre-tax, pre-transfer incomes received by private households. See footnote 10 for a definition in terms of LIS variables.

after taxes and transfers). Based on this, researchers have a choice between measuring fiscal redistribution as the *absolute* difference between the two Gini coefficients, or as the change in the Gini coefficient due to taxes and transfers *relative* to its initial level. Both the absolute and relative measures are frequently used in the literature on inequality and redistribution, and the justification for using either concept crucially depends on the research context.[6]

To determine which of the two measures is theoretically more appropriate for the narrow purpose at hand, it is necessary to revisit some of the details of the original model by Meltzer and Richard (1981). They assume that taxes are levied against all private sector incomes using a linear tax rate, and that all tax receipts are spent on distributing equal lump sums among citizens.[7] These simplifying model assumptions make it possible to calculate by how much a given Gini coefficient would be reduced as a result of a given tax rate. A full proof is supplied in Appendix 1, but one can intuitively understand the process as a shift of the Lorenz curve from its original position towards the 45 degree line (that would imply perfect equality). The magnitude of this shift, and hence the extent of redistribution, depends on the share of the lump-sum receipts and of private sector incomes in total disposable income. Fortunately, the distribution of incomes from both sources is known: The Gini for the remaining private sector incomes is equal to the initial distribution of private sector incomes, $G_p$, and the Gini for incomes from lump-sum redistribution, $G_l$, is zero (since all individuals receive equal lump sums).

Further, Meltzer and Richard assume that all proceeds from taxation are redistributed, so the total sum of incomes does not change. The share of the lump sum receipts in total incomes is thus equal to the tax rate $t$, and the share of the remaining private sector incomes in total income is equal to $1 - t$. We therefore know the distribution of both income components and their relative weight in the overall post-tax, post-transfer distribution. Since Gini coefficients cannot be easily decomposed, this information would be insufficient to calculate the Gini coefficient for total disposable incomes in any real-world application. This is due to the fact that the relative position of individuals usually differs between any two income distributions (see e.g. Shorrocks 1982). However, in the model world of Meltzer and Richard, each individual's income grows by the same amount so that their relative position does not change when transfer receipts are added.[8] The Gini coefficient of disposable incomes, $G_d$, can thus be calculated as a weighted average of the two income components $G_p$ and $G_l$, where the weights are given by $1 - t$ and the tax rate $t$:

$$G_d = (1-t)\times G_p + t \times G_l \tag{1}$$

Since all persons receive the same lump-sum payment, the Gini coefficient $G_l$ takes the value of zero and equation (1) can be simplified into:

$$G_d = (1-t)\times G_p \tag{1'}$$

It is easy to see that at a tax rate of zero, post- and pre-government inequality are identical (and hence no redistribution takes place), but that as the tax rate rises, the Gini for disposable income decreases until it eventually reaches zero when all income is taxed and redistributed.

For Meltzer and Richard, political conflict is therefore about determining the tax rate $t$. They start from the premise that '[u]nder majority rule, the voter with median income is decisive' and, in line

[6] See, for example, the arguments in favour of the absolute measure in Kenworthy and Pontusson (2005); examples for studies base on the relative measure of inequality include Bradley et al. (2003) or Mahler (2004).
[7] Hence, taxation itself has no impact on inequality, and redistribution is solely achieved through the transfer system. This is of course a gross oversimplification, but it corresponds to the real world in so far as Mahler and Jesuit (2006) find that about three-quarters of fiscal redistribution can be attributed to the transfer system.
[8] This condition is crucial; unless it is satisfied (i.e. virtually in all real-world applications), it is not possible to average Gini coefficients.

with standard theory, assume that '[t]he decisive voter chooses the tax rate that maximizes his utility' (Meltzer and Richard 1981: 920). The median voter's utility is given by the cost that taxation imposes on her or him (which is given by $t \times y_d$, where $y_d$ is her or his own income) and the benefit from lump-sum redistribution (which is proportionate to $t \times \bar{y}$, where $\bar{y}$ is mean income). Even after taking into account potentially adverse effects of taxation on incentives, Meltzer and Richard show that 'the tax rate rises as mean income rises relative to the income of the decisive voter' (which corresponds to median income; see ibid.: 923). The ratio of mean over median income is a common metric for inequality, and monotonically related to the Gini coefficient under the assumption that distribution of incomes follows a lognormal pattern (see Lopez and Servén 2006). Returning to the two measures for redistribution, absolute redistribution, $\Delta G^{abs}$, can be defined as the absolute difference between the two Gini coefficients,

$$\Delta G^{abs} = G_p - G_d \tag{2}$$

and relative redistribution, $\Delta G^{rel}$, as the absolute difference between the two Gini coefficients divided by the initial level of the Gini coefficient:

$$\Delta G^{rel} = \frac{G_p - G_d}{G_p} \tag{3}$$

Substituting (1') into equations (2) and (3) leads to:

$$\Delta G^{abs} = t \times G_p \tag{2'}$$

$$\Delta G^{rel} = t \tag{3'}$$

Therefore, the relative change of the Gini coefficient is directly proportional to the tax rate, while the absolute change is a function of both the tax rate and the initial Gini coefficient for private sector incomes.

Since Meltzer and Richard postulate that greater market inequality leads to a higher tax rate, the identity in equation (3') implies that *relative* redistribution is the best proxy for the tax rate $t$, which they postulate to rise with greater inequality. It is therefore appropriate to investigate how market inequality influences relative redistribution, and to test the relationship between $G_p$ and $\Delta G^{rel}$. Note that equation (2') shows that *absolute* redistribution will increase with the Gini coefficient for market incomes even if the tax rate remains constant. The finding of a positive association between market inequality and absolute redistribution would therefore not confirm the Meltzer-Richard hypothesis that the tax rate $t$ rises with inequality.

Although this 'model world' might seem removed from reality, the two equations are helpful to think about redistribution in the real world. As can be seen in equation (2'), we would expect absolute redistribution to increase with market inequality even if the characteristics of the tax and transfer system remain largely unchanged. Incidentally, this process of 'automatic stabilization' is what seems to have been at work over the 1980s and 1990s in rich countries where the welfare state partially compensated for the surge in market inequality (see Kenworthy and Pontusson 2005). Equation (3') implies that only the characteristics of the tax and transfer system (for which $t$ is the short-hand) will influence relative redistribution (regardless of the initial level of inequality). Studies concerned that want to assess the effect of different entitlement rules and the progressiveness of taxation on redistribution are thus well-advised to focus on relative redistribution. Note, however, that the tax and transfer system in the Meltzer and Richard model is very crude and that different effects might be observed in the real world.

## 3. Do more unequal societies redistribute more?

The discussion of the Meltzer and Richard model in the previous section leads to a readily testable hypothesis, namely that relative redistribution $\Delta G^{rel}$ is a direct function of the initial level of inequality for private sector incomes, $G_p$:

$$\text{H}_1\text{:}\ \Delta G^{rel} = f(G_p) \tag{4}$$

This relationship should hold true both *within* countries over time and *between* countries, at least as far as electoral democracies are concerned. It is thus appropriate to test the hypothesis on a dataset that contains repeated observations across countries. The LIS provides such a source and is generally recognized as the best compilation of standardized household income datasets that allow for such an analysis (Atkinson 2004a). In their initial publication, Mahler and Jesuit (2006) provided a total 59 data points from 13 countries for inequality of private sector incomes and of disposable incomes, and hence for redistribution (see also Bradley et al. 2003). In February 2008, they released an updated dataset with 68 observations from 14 industrialized countries.[9] Since then, the LIS has significantly expanded its coverage and now includes observations from Latin American countries (Colombia, Brazil, and Guatemala) as well as Asia (Republic of Korea and Taiwan, Province of China). In total, the relevant income concepts can be retrieved for 110 surveys from 26 countries and territories.[10]

The resulting dataset (which is reproduced in Appendix 4) contains the desired cross-sectional and inter-temporal variation. The oldest observation dates back to 1967 (Sweden) and the newest are from 2006 (Brazil, Guatemala, and Republic of Korea). The panel is unbalanced, and the number of observations ranges from ten observations in Canada to only one data-point in seven countries (Austria, Brazil, Colombia, Estonia, Guatemala, Republic of Korea, and the Slovak Republic; see Appendix 3). This still leaves 19 countries that have at least two data points needed to study variation across time. Among these, there is a predominant trend towards higher inequality in the primary distribution of incomes. Some 15 countries showed a rise in the Gini coefficient for private sector incomes (i.e. before taxes and transfers), while only one displayed stability and three a decline (see Table 1).

[9] David K. Jesuit and Vincent A. Mahler, Fiscal Redistribution Dataset, Version 2, February 2008.

[10] The LIS database contains further datasets that record only net income, so that no comparison between incomes before and after taxes and transfers can be made. The definition of private sector income follows Mahler and Jesuit and refers to the sum of LIS variables 'Market income' (MI), 'Private transfers' (PRIVATI) and 'Other cash income' (V36). Disposable income is derived by adding 'Transfer income' (TRANSI) and subtracting 'Mandatory payroll taxes' (PAYROLL) and 'Income taxes' (V11). Standard LIS procedures are used to top- and bottom-code and to obtain equivalized per capita income. The results are consistent with Jesuit and Mahler's 2008 dataset and the LIS key figures as of mid-2011. Note that all data were extracted before the launch of the new LIS template on 31 October 2011 that brought some changes to the definition of income concepts (in particular the inclusion of non-monetary income to disposable household income).

**Table 1: Trends in inequality and redistribution in 19 countries and territories, 1970s to 2000s**

| Country / territory | Observations | | | Trends in inequality (annual change, in Gini points)* | | Trends in redistribution (annual change)* | |
|---|---|---|---|---|---|---|---|
| | First | Last | n = | Private sector incomes | Disposable incomes | Absolute redistribution | Relative redistribution |
| Poland | 1999 | 2004 | 2 | 1.015 | 0.630 | 0.384 | -0.022 |
| Belgium | 1992 | 1997 | 2 | 0.640 | 0.532 | 0.108 | -0.444 |
| United Kingdom | 1969 | 2004 | 9 | 0.505 | 0.291 | 0.214 | 0.175 |
| Czech Republic | 1992 | 2004 | 3 | 0.490 | 0.441 | 0.049 | -0.400 |
| Finland | 1987 | 2004 | 5 | 0.449 | 0.291 | 0.158 | -0.119 |
| Germany | 1973 | 2004 | 9 | 0.402 | 0.038 | 0.364 | 0.524 |
| United States | 1979 | 2004 | 7 | 0.330 | 0.293 | 0.036 | -0.092 |
| Australia | 1981 | 2003 | 6 | 0.329 | 0.142 | 0.186 | 0.193 |
| Israel | 1979 | 2005 | 6 | 0.327 | 0.256 | 0.071 | -0.040 |
| Norway | 1979 | 2004 | 6 | 0.284 | 0.128 | 0.156 | 0.121 |
| France | 1979 | 1994 | 4 | 0.236 | -0.053 | 0.289 | 0.430 |
| Taiwan, POC | 1981 | 2005 | 7 | 0.212 | 0.157 | 0.055 | 0.159 |
| Canada | 1971 | 2004 | 10 | 0.186 | 0.035 | 0.151 | 0.256 |
| Sweden | 1967 | 2005 | 8 | 0.177 | 0.009 | 0.167 | 0.217 |
| Denmark | 1987 | 2004 | 5 | 0.072 | -0.147 | 0.219 | 0.468 |
| Ireland | 1987 | 2004 | 2 | 0.000 | -0.098 | 0.099 | 0.197 |
| Switzerland | 1982 | 2004 | 5 | -0.033 | -0.200 | 0.167 | 0.443 |
| Romania | 1995 | 1997 | 2 | -0.061 | -0.009 | -0.051 | -0.097 |
| Netherlands | 1983 | 2004 | 5 | -0.242 | -0.045 | -0.197 | -0.218 |
| *Average* | *1982* | *2003* | *5.4* | *0.280* | *0.142* | *0.138* | *0.092* |

Notes: * Based on Gini coefficients multiplied by 100. Countries are ordered in descending order by trends in private sector Gini coefficients. Annual change is calculated by fitting a regression line through all observations from a country.

Source: Luxembourg Income Study Database (LIS); analysis of micro-data completed between February and May 2011.

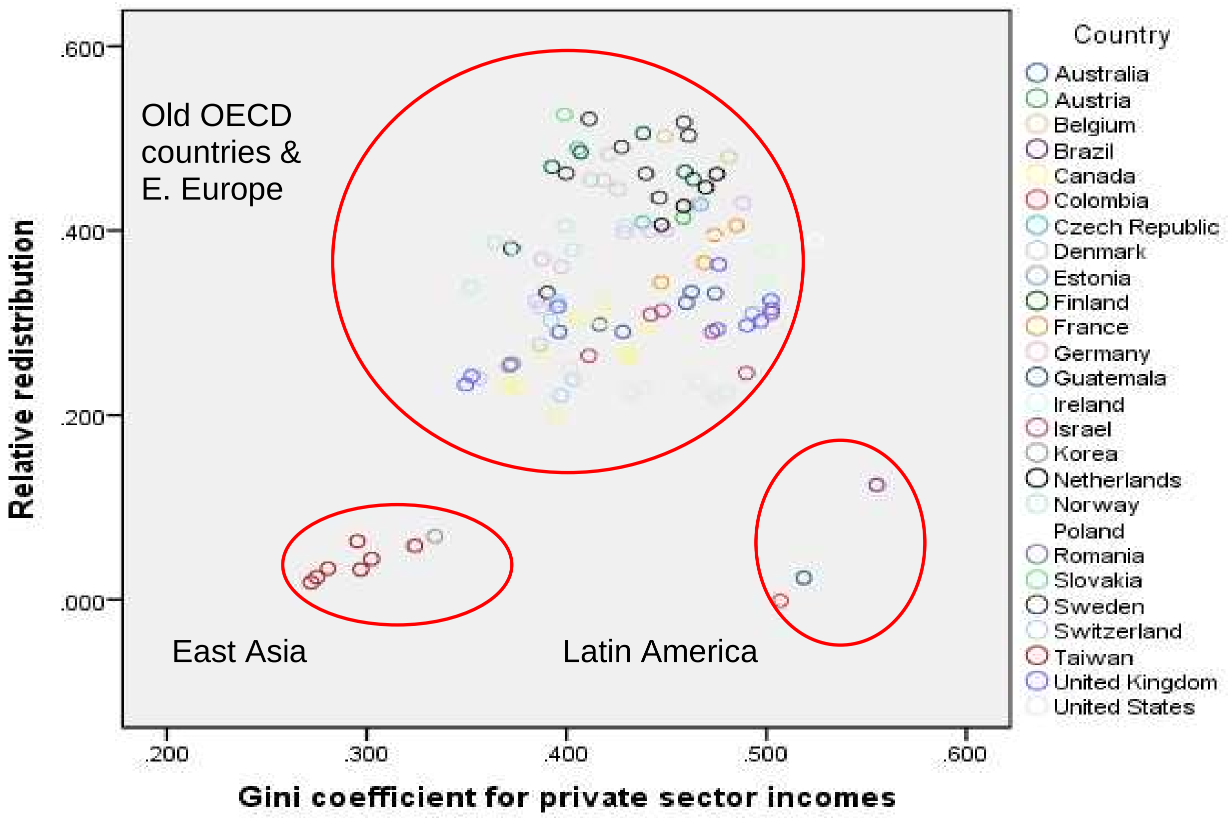
However, the data also show that welfare states have been relatively resilient in rich countries and that their social security and tax systems have dampened the rise in inequality over the past decades (see Pierson 1996). In the typical country, inequality of private sector incomes rose by 0.28 Gini points per year, but inequality of disposable income grew at a rate of 'only' 0.14 Gini points (see Table 1). This implies that an increase in absolute redistribution offset about half the rise in inequality. Nonetheless, there is substantial variation in both inequality trends and countries' responses (see also Atkinson 2004b). Among the countries with long time series,[11] the United Kingdom, Australia and Norway compensated about half the increase in private sector inequality through greater redistribution. By contrast, Germany—and to a lesser extent Sweden and Canada—responded with a sharp increase in

[11] For four countries, only two data points are available, and a further seven dropped out of the sample since only one observation was available.

redistribution so that inequality of disposable incomes only rose marginally.[12] At the other end of the spectrum, in the United States and Israel the tax and transfer system absorbed only a minuscule fraction of the rise in private sector inequality, which translated almost unfettered into greater inequality of disposable incomes. In both countries, relative redistribution actually *declined* as inequality of private sector incomes rose—in sharp contrast to the predictions of the Meltzer and Richard model.

Figure 1 provides a scatter-plot for the relationship between market inequality and relative redistribution. Immediately striking are two clusters of outliers: In the lower left of the graph, observations from East Asia combine low market inequality with low redistribution (see Hwang 2004). By contrast, the three Latin American countries in the lower right of the graph combine high market inequality and low redistribution (see Huber et al. 2006; Goñi et al. 2008). Both findings correspond to the literature on redistribution in these two regions. For the main cluster of observations, where the developed economies can be found, no clear pattern emerges. However, on closer inspection, it appears that repeated observations from the same country—for example, from Canada or France—roughly corresponds to the pattern predicted by Meltzer and Richard.

**Figure 1: Gini for private sector income and relative redistribution in 26 countries and territories**




Source: Luxembourg Income Study Database (LIS); analysis of micro-data completed between February and May 2011.

[12] Note that the reunification in 1990 contributed to the surge in inequality as incomes in the former German Democratic Republic are significantly below those in the West. However, large income transfers—notably pensions and unemployment benefits—reduce the inequality of disposable incomes substantially.

The scatter-plot has two implications for the empirical strategy: (a) The presence of outliers suggests that the applicability of the Meltzer and Richard model might be confined to the developed countries, which can be expected to have stronger and more mature democratic institutions. All models will therefore be estimated on the full sample and on a reduced sample that excludes observations from East Asia and Latin America. (b) While there is no apparent cross-country relationship, the expected relationship might still hold within countries. It is therefore useful to distinguish between-country from within-country effects, and to run separate models for these.

The scatter-plot also calls into question the utility of a pooled cross-section, time-series analysis. Such models imply that the same relationship can be observed between and within countries. Of course, if the underlying assumptions of the median voter theory are valid, the Meltzer and Richard model has universal applicability: the same mechanism should be at work regardless of whether one compares between countries or within countries over time. Model (1a) in Table 2 therefore presents the standard random effects model (with robust-cluster standard errors)[13] for the pooled dataset largely on a priori theoretical grounds, and with caveats about its analytical utility and statistical validity (see also Kenworthy 2007).

The model yields no support for the Meltzer and Richard hypothesis: the regression coefficient on the Gini for private sector incomes remains insignificant (even if one applies a generous 0.10 threshold). Although the East Asian and Latin American observations included in the dataset are from democracies and hence the mechanism proposed by Meltzer and Richard should apply to them, one could of course argue that their democratic institutions are less mature than those in advanced industrialized countries. Institutional weaknesses could therefore explain the failure to confirm the hypothesis. However, the results do not change when the observations from East Asia and Latin America are excluded, as is done in model (1b). On the contrary, the regression coefficient moves further from significance (p-value: 0.189). Note that no control variables are added to the regression; the median voter theorem postulates a universal relationship that is not conditional on the presence of specific conditions (other than majority voting).[14] A Hausman test (see notes to Table 2) confirms the initial caveats about the suitability of a pooled analysis.

The failure to establish a relationship between initial inequality and subsequent redistribution is in line with previous studies based on pooled datasets cited above. But can the median voter hypothesis possibly explain variation in redistribution between countries? Models (2a) and (2b) present a test of the between-country effect, essentially a regression on the mean of all observations from the same country. This removes the within-country variation, while using all available observations—an approach that is preferable to arbitrarily selecting a single observation from each country. As in the pooled model, the regression coefficients on the Gini for private sector incomes are insignificant, regardless of whether the full or the reduced sample is used (p-values: 0.959 and 0.867, respectively).[15]

[13] See Huber et al. (2006: 956f.) for a discussion of the different empirical approaches to deal with pooled cross-section time-series datasets.

[14] See the helpful note by Kenworthy (2007) on regressions in macro-comparative analysis. The author refrains from the common technique of adding control variables in order to achieve a significant regression coefficient. For those who are nonetheless interested: The coefficient on the Gini for private sector incomes remains insignificant when the two most obvious control variables are added, namely the unemployment rate and the share of the population aged 65 years and above. The p-values drop to 0.614 (full sample) and 0.775 (sample excluding observations from East Asia and Latin America).

[15] Again, the coefficient on the Gini for private sector incomes remains insignificant when the unemployment rate and the share of the population aged 65 years and above are added as control variables. The p-values are 0.999 (full sample) and 0.433 (sample excluding observations from East Asia and Latin America); the coefficient also carries the positive sign in the latter case.

**Table 2: Regression results for models with private sector inequality an explanatory variable (dependent variable: relative redistribution)**

| Variable / Model | (1) Random effects, robust cluster SE | | (2) Between-country effects | | (3) Within-country effects | | (4) Within-country effects, with controls | |
|---|---|---|---|---|---|---|---|---|
| | (1a) full sample | (1b) OECD | (2a) full sample | (2b) OECD | (3a) full sample | (3b) OECD | (4a) full sample | (4b) OECD |
| pi_gini (Gini, private sector incomes) | 0.854<br>(0.530) | 0.360<br>(0.264) | 0.028<br>(0.536) | 0.093<br>(0.547) | 0.584***<br>(0.107) | 0.581***<br>(0.111) | 0.192<br>(0.146) | 0.168<br>(0.153) |
| unemp (unemployment rate) | | | | | | | 0.457***<br>(0.160) | 0.468***<br>(0.165) |
| oldage (population 65+ years) | | | | | | | 0.900**<br>(0.360) | 1.056**<br>(0.410) |
| Constant | -0.040<br>(0.233) | 0.198<br>(0.115) | 0.297<br>(0.237) | 0.330<br>(0.241) | 0.074<br>(0.046) | 0.102**<br>(0.048) | 0.094**<br>(0.048) | 0.105**<br>(0.053) |
| n | 110 | 99 | 110 | 99 | 110 | 99 | 110 | 99 |
| Number of clusters / groups | 26 | 21 | 26 | 21 | 26 | 21 | 26 | 21 |
| $R^2$ | 0.146 | 0.029 | 0.000 | 0.002 | 0.266 | 0.261 | 0.371 | 0.376 |

Notes: $R^2$ refers to $R^2$ (overall) for the random effects model (1), to $R^2$ (between) for the between-effects model (2), and to $R^2$ (within) for the within-effects model (3) and (4).
Standard errors are given in parenthesis; those in model (1) refer to robust cluster standard errors.
***, **, and * denote significance at risks levels 0.01, 0.05 and 0.10, respectively.
A Hausman test was performed to confirm the consistency of the regression coefficient obtained from the random effects model (1a) with the coefficient from the fixed effects model (3a). It produced a test statistic of -0.27 and failed to meet the assumptions of the Hausman test.

Source: Based on Luxembourg Income Study (pi_gini and dependent), ILO (unemp), World Bank (oldage) and Statistics Bureau of Taiwan, Province of China (unemp and oldage for Taiwan, POC). For details see Appendix 2.

The results have so far been disappointing for the Meltzer and Richard hypothesis. But maybe unobserved institutional characteristics that vary between countries obscure a relationship, which nonetheless holds within countries? Models (3a) and (3b) therefore provide a fixed effects model that tests the within-country relationship. The results appear to offer overwhelming support to this ‘weak’ hypothesis. Although the explanatory power of the regression is modest (within $R^2 = 0.266$), the coefficients are highly significant and robust to the exclusion of observations from East Asia and Latin America. Two interpretations offer themselves for the contradictory results from the between- and within-country models: Unobserved country characteristics—say, differences in the electoral process (see Iversen and Soskice 2006)—could obscure the median voter’s influence, which only becomes evident once they are controlled for by introducing country dummies. The fixed effects model would then be the only valid test, and the results would offer sufficient support for the median voter theorem.

However, the within-country effect could also be due to a different mechanism. Recall that Meltzer and Richard built a very rudimentary model of redistribution under which all income is taxed at a flat rate and the entire revenue is redistributed in the form of equal lump-sum benefits. In the real world, benefits are means-tested and income taxes are generally progressive (with a few exceptions, such as

Switzerland[16]) (see also Prasad and Deng 2009). The automatic stabilization of inequality that results from such a progressive tax and transfer system would be greater than what we would expect in the 'model world' of Meltzer and Richard. If changes in a country's demographic structure causes greater market inequality, relative redistribution might increase not as a result of changes in welfare generosity (or the hypothetical tax rate *t*), but due to the very same demographic shifts.

Rich countries have of course experienced a large increase in unemployment since the early 1970s, and low fertility and rising life expectancy have led to a steady growth in the share of the elderly population. An alternative explanation for the results of the fixed effects model would be that relative redistribution has grown as a result of these structural shifts, rather than due to increased market inequality and subsequent changes in the tax rate.[17] Models (4a) and (4b) therefore repeat the within-country analysis with two control variables, the unemployment rate and the share of the population aged 65 years and above. From the viewpoint of the Meltzer and Richard model, the results should be robust to the addition of additional control variables: they should not have any impact on relative redistribution, which is sufficiently explained by variations in initial inequality.

In model (4a) with the full sample, the unemployment rate and the share of the population aged 65 years and above both turn out to be highly significant predictors for within-country changes in relative redistribution (at the 0.01- and 0.05-level, respectively). It appears that structural changes in the labour market and demography sufficiently explain within-changes in inequality. Once these factors are controlled for, changes in the initial inequality of private sector incomes no longer carry any explanatory power and the regression coefficient loses its significance. Moreover, the explanatory power of the model improves (within $R^2 = 0.371$), which indicates that the private sector inequality was a poor proxy for the underlying demographic and labour market trends. The results are robust to the exclusion of observations from East Asia and Latin America, as can be seen in model (4b). In sum, within-country changes in redistribution offer no convincing support for the Meltzer and Richard hypothesis; even a rudimentary model of structural changes since the 1970s offers a better explanation that renders the effects of private sector inequality—the key variable of the rational choice model—obsolete.

## 4. Alternative explanations for variance in redistribution and the perspective of behavioural economics

Can behavioural economics explain differences in redistribution where rational choice has failed? As argued in the previous section, rational choice theory seems to be a poor guide to understand either differences in redistribution between countries and changes within countries. Recall that the two central assumptions of the model were that the political system responds to demands of the median voter, and that the median voter seeks to maximize her own, narrowly defined utility. At least one of these assumptions must be faulty, and an extensive literature has indeed discussed their respective shortcomings (for a short review see Kaufman 2009). One part of this literature, with many contributions from political science, has concentrated on the question how the political system translates preferences into policy outcomes. Various contributions have investigated differences between proportional representation and majority voting, the impact of voter turnout or how the structure of inequality will influence coalitions between different income groups (see e.g. Bassett et al. 1999; Tanninen 1999; Austen-Smith 2000; Cukierman and Spiegel 2003; Iversen and Soskice 2006; Borck 2007; Solt 2008; Mahler 2008; Lupu and Pontusson 2011). Others have argued that

[16] According to the LIS data, Switzerland's Gini coefficient *increased* (rather than decreased) as a result of taxation in 2000 and 2004. In both cases, the (limited) redistribution was achieved exclusively though transfers.
[17] See Perotti (1996) and Bassett et al. (1999) for earlier studies that control for the share of the population aged 65 and above. See also Bradley et al. (2003), who show that higher unemployment increased both pre-tax pre-transfer inequality and subsequent redistribution. They restrict their analysis to the working age population, and therefore do not consider the effect of aging societies.

social security systems have unclear *a priori* distributive outcomes and that they primarily serve an insurance purpose (Moene and Wallerstein 2003). Hence, greater risk exposure should increase support for these schemes (see Cusack et al. 2006).

More fundamentally, questions have arisen about the underlying *Menschenbild* (view of the nature of man) of rational choice—do people only consider their own advantage when voting? While the proposition that demand for redistribution should increase with rising inequality is unproblematic within the rational choice framework, it collapses when voters are not only motivated by the maximization of their own, narrowly defined utility. Behavioural economics has challenged this paradigm and explored the role of social norms in explaining actual, observed human behaviour and studied the role of altruism, inequality aversion and fairness (see, for example, Fehr and Fischbacher 2004; Fehr and Schmidt 2005).

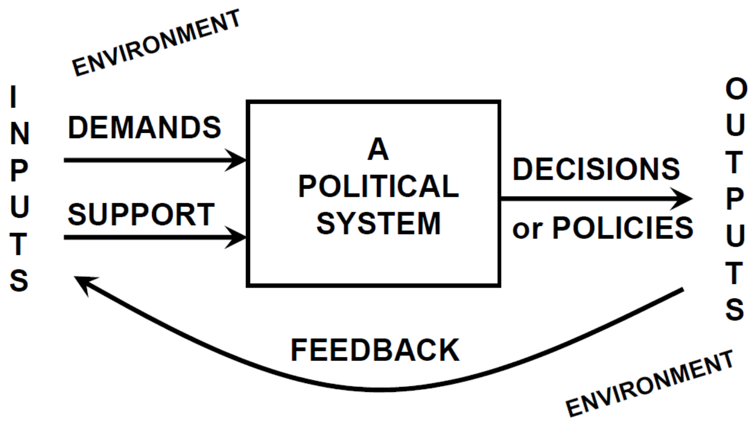
One prominent approach within behavioural economics has been to conduct experiments with groups of individuals who are asked to distribute small amounts of money between themselves and another person (Fehr and Schmidt 2005). Results from the 'Dictator game' (where the transfer is simply set by the subject) and the 'Ultimatum game' (where the recipient can reject the split) have been interpreted as evidence that individuals behave altruistically by passing on part of their endowment. Further, in the 'Ultimatum game' their counterparts are willing to forego a small gain by rejecting splits that they perceive as overly unfair (see Andreoni et al. 2008). Interestingly, while altruism appears to be a universal phenomenon, there is some variation between countries and communities (see Cardenas and Carpenter 2008, for a review). Similarly, Falk et al. (2008) show that fairness intentions matter, and that individuals frequently prefer an option seen as 'fair' over an alternative that maximizes only their own utility—a finding that is not reconcilable with the standard assumptions of rational choice.

However, the sample size and coverage of these experiments are too small to gain reliable information on cross-national (and inter-temporal) variations in inequality aversion, altruism and fairness intentions. Building on large cross-national survey datasets, political sociology has studied the role of value orientations in shaping people's preference for equity and their support for redistribution (Blekesaune and Quadango 2003; Luebker 2004). Unlike the rational choice literature, this political sociology approach leaves room for social norms and individual belief systems as intervening factors to shape support for redistribution (that is no longer a direct function of initial market inequality; see e.g. Kuhn 2009a and 2009b). Simply put, people are thought to evaluate a given level of inequality against their own ideas of what is fair and just to arrive at an assessment of inequality. While the literature has shown that economic inequality has an adverse effect on people's life satisfaction (see Verme 2011), it also points out that Americans are more tolerant of inequality than Europeans (Alesina et al. 2004). These differences should matter: If people see inequality as a problem in need of remedy, they are more likely to support state action in the form of redistribution. What follows is that, if tolerance of inequality varies between societies, voters (median or not) in two countries with identical levels of inequality might differ in their enthusiasm for government redistribution – something that defies the logic of the rational choice approach.

Consistent with the above, the literature has shown that support for redistribution is not simply a function of inequality, but that different societies evaluate income inequality differently and also display differences in their support for redistribution (Alesina and Angeletos 2005; Luebker 2004 and 2007). It is a reasonable hypothesis that these differences will influence the degree to which national tax and transfer systems redistribute income. Research on social welfare responsiveness has in fact shown a close association between citizen's demands for equity and welfare state generosity (at least as far as rich countries are concerned; see Brooks and Manza 2006 and 2007; see also Burstein 1998). The key challenge is that the causality might run in the other way—generous welfare states might well not be a response to citizen's demands, but could have generated their own support through performance over time (Kenworthy 2009). Socialization in a particular welfare regime type undoubtedly shapes social norms by providing a benchmark of what can reasonably be expected, and hence also evaluations of inequality and support for redistribution. The post-war division and subsequent reunification of Germany provides for an insightful natural experiment: East Germans, who were brought up in a nominally socialist state, expect a far greater welfare state than their West German compatriots, even when other individual-level factors are controlled for (Alesina and Fuchs-Schündeln 2007).

The question which way the causality runs has probably no clear answer—it would seem plausible that it in fact runs *both* ways. Easton's (1957, 1965) system analysis of the political life provides a useful theoretical framework for this (see Figure 2). For him, the political system generates outputs (such as welfare payments and redistribution) that are evaluated by the citizenry and, through a feedback loop, influence the inputs that feed into the political system in the form of demands and support. Incidentally, Easton's ideas were at the vanguard of the 'behaviouralist revolution' of political science (Farr 1995) that discovered the individual as a unit of analysis, and aimed at providing predictive and causal explanations of political behaviour (a trend that has arrived in the economics discipline half a century later).

**Figure 2: Easton's system model of political life**




Source: Based on Easton (1957: 384).

From this perspective, support for redistribution could then be shaped by previous performance of the welfare state, and explain why the welfare state is maintained through popular support once it is established. However, the present paper has a more limited concern, namely to contrast median voter theorem that relies on the 'hypothetical' preference for redistribution as deducted from the utility maximization axiom of rational choice with an analysis that draws on actual, observed preferences of real individuals. Are they a better guide to understand why welfare states do not converge onto the level of redistribution predicted by rational choice?

A related controversy has focused on measurement issues, particularly the treatment of pensions. In countries where pensions are provided though public social insurance schemes, people save less in working years but pay compulsory social security contributions (which are usually matched by their employers). When people reach retirement age, their private sector income often falls to zero and they live from transfer payments in the form of old-age pensions. In countries without such public systems, people pay into private, capital-based schemes during their working lives and in retirement receive annuities (which are usually counted as private sector incomes). Thus, both inequality of market incomes and redistribution will be lower in the latter class of countries, while the degree of market inequality and redistribution will be 'inflated' in countries that provide pensions though social insurance systems (or as universal state pensions; see Bradley et al. 2003: 208).

One approach to address this observation has been to exclude the elderly population from the analysis and compute measures for inequality and redistribution for the working-age population (Bradley et al.

2003; Kenworthy and Pontusson 2005; see also Mahler and Jesuit 2006). An alternative is to adjust income concepts by including social insurance and state pensions into a concept of 'primary income' (in line with private pensions), and by treating payroll taxes analogous to savings and including them in 'adjusted disposable income' (Jesuit and Mahler 2010).[18]

While it is a valid argument that the pensions system will have profound impacts on inequality and redistribution, this leads to a more fundamental question: Do we want to control for these differences when analyzing redistribution and welfare states? After all, old-age pensions are not fundamentally different from other types of social insurance, such as unemployment, sickness or invalidity benefits—the design of which will likewise lead to different redistributive outcomes.[19] Employees (and usually employers as well) pay contributions, and receive benefits when certain qualifying conditions are met. Not all who contribute to a scheme will also receive a benefit—workers who never become unemployed will not receive unemployment benefits, and those who do not reach pension age will not receive a pension. While benefits are often linked into previous contributions, they also reflect other, social objectives. For example, time spent in education or caring for children are frequently credited as contribution periods, and spouses who survive a beneficiary typically receive a survivor's benefit (for which no extra contributions have been made). Often, a substantial part of benefits is funded not out of contributions, but out of general tax revenue.[20]

Social insurance institutions are one of the main mechanisms for welfare states to redistribute incomes, and their design is of central importance for redistributive outcomes (see Korpi and Palme 1998; Kraus 2004; Conde-Ruiz and Profeta 2007). By comparison, private pensions have different distributive outcomes (Behrendt 2000). Pension systems are thus subject to intense political debate, and even incremental transitions from one model to another go hand-in-hand with intense conflict.[21] Excluding the pension system from the comparative analysis of welfare regimes and redistribution would mean to miss a large part of the picture. The empirical analysis in the following section will therefore maintain the dependent variable for relative redistribution as introduced in the previous section (i.e. based on the total population).

## 5. Can voters' actual preferences for distribution account for differences in redistribution between countries?

To test the preposition that actual (as opposed to assumed) public opinion matters for policy outcomes, we need to find an appropriate way to measure public support for redistribution. In other words, we need to operationalize the independent variable that is of interest from the perspective of behavioural economics (and, it must be added in all fairness, from the perspective of political sociology and political science that have studied the role of social norms and values long before behavioural economics discovered them for the economics discipline). The International Social Survey Programme (ISSP) is the most reputable and most commonly used source for this type of

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[18] Note that 'disposable' income becomes somewhat of a misnomer since social insurance contributions are mandatory and not disposable to the household.

[19] See Statistical Appendix Part B in ILO (2010) for a comprehensive overview of the different social insurance systems. For redistributive outcomes of in the case of different sickness benefit systems, see Khan and Jansson (2008).

[20] For example, Germany's 2011 federal budget contains an allocation of 115.2 billion Euros for subsidies to social insurance schemes (including 71.4 billion Euros to the pension system). This is the largest single expenditure item, accounting for 37.7 per cent of total federal expenditure (see Bundesfinanzministerium, Übersichten zum Bundeshaushaltsplan 2011, Teil II: Funktionenübersicht; Berlin, not dated).

[21] Refer to Korpi and Palme (1998) for a typology, and see the examples of Germany (introduction of an additional, private pillar to the pension system) or the United States ('Obamacare') for examples of conflict around incremental change to existing social insurance institutions.

analysis (see e.g. Alesina and Angeletos 2005; Osberg and Smeeding 2006).[22] The consortium started in 1984 with four members (Australia, Germany, Great Britain, and the United States) and has since expanded to a total of 48 member countries, including several newly industrialized and developing countries. One of the questions in the module on Social Inequality addresses support for government redistribution directly:

How much do you agree or disagree with each statement about differences in income? It is the responsibility of the government to reduce the differences in income between people with high incomes and those with low incomes.

The Social Inequality module has so far been in the field in 1987, 1992, 1999, and 2009, and an identical question was also included in the Role of Government module in 1985, 1990 and 1996.[23] Respondents were asked to record their answers on a five-point Likkert-scale that ranges from 'Strongly agree' to 'Strongly disagree'. The proportion of those who either agreed or strongly agreed with government action to reduce income differences is a good proxy for the prevalence of support for redistribution.[24]

Although there is a large overlap between the LIS and the ISSP, the match between the two sources is not perfect. For instance, the Latin American countries covered by the ISSP (Argentina and Chile) differ from those included by LIS (Brazil, Colombia, and Guatemala) so that there are no common observations from this region. In Asia, both the ISSP and LIS cover Taiwan, Province of China, and the Republic of Korea; they also share a large, common pool of advanced industrialized economies. This leaves the problem that the years to which observations refer do not always match. In order to use as much of the available information as possible, the support for redistribution in a given year was estimated by fitting a linear trend between two neighbouring observations.[25] In some cases where only one neighbouring observation was available, the closest year was used on the assumption that support for redistribution had not changed.[26] These approximations are of course not ideal, and they compromise the quality of within-country trends. However, they are the best possible solutions in a world of non-perfect data and should have less impact on the robustness of between-country comparisons.

[22] An alternative source for cross-national data on views on inequality is the European Values Survey (EVS). However, as the name suggests, the survey covers only European countries.

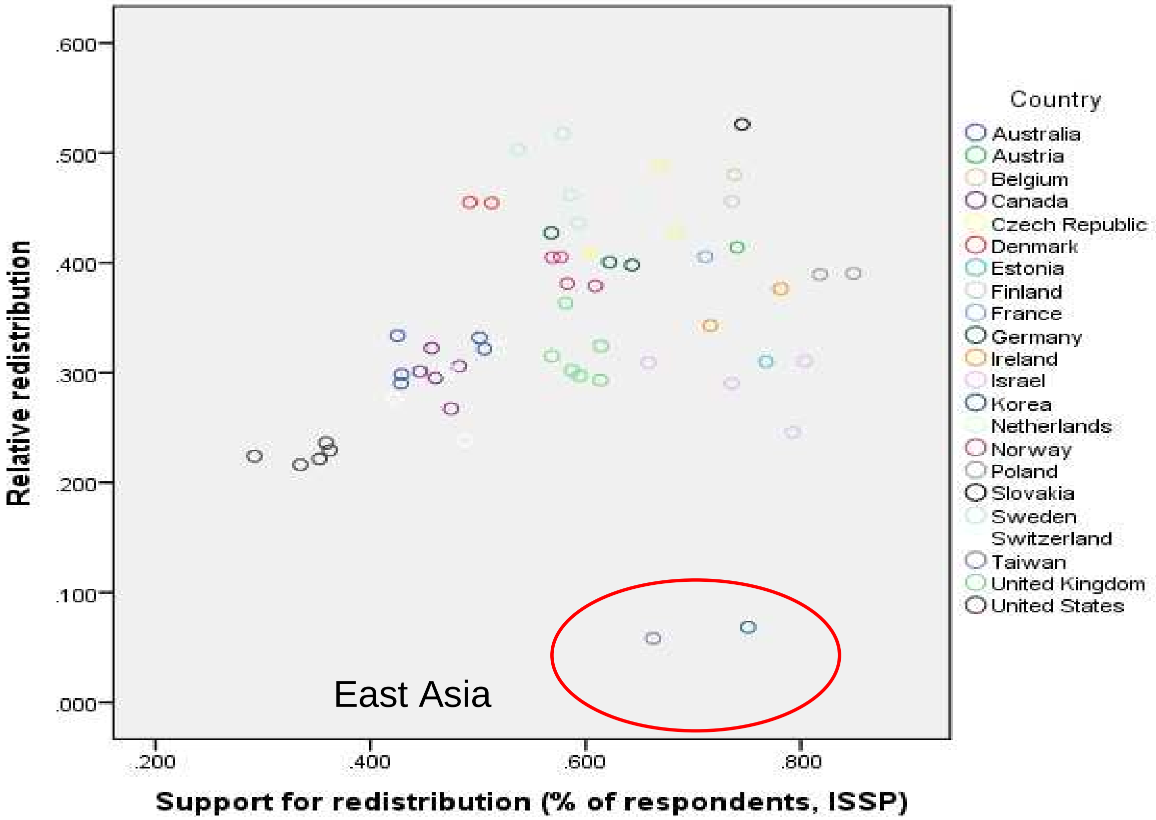
[23] In 2006, the Role of Government module contained a similar question that, however, used a different Likkert-scale to record answers (four categories ranging from 'Definitely should be' to 'Definitely should not be').

[24] Arguably, the mean response is also a valid summary of responses. However, in order to side-step the debate on whether a Likkert-scale is a true interval scale, this paper uses the proportion of respondents who agreed with the statement. Both measures are closely correlated.

[25] For example, the share of the Canadian population that supported redistribution in 1994 was estimated on the basis of data for 1992 and 1996.

[26] For example, for the Republic of Korea a 2009 observation from the ISSP was matched to a 2006 observation from LIS.

**Figure 3: Support for redistribution and relative redistribution in 22 countries and territories**




Source: Based on Luxembourg Income Study (relative redistribution) and ISSP and related sources (support for redistribution). For details see Appendix 2.

Figure 3 displays the 58 observations from 22 countries for which information from the ISSP and LIS is available. The scatter shows a reasonably close, though not perfect association between the two variables: As expected, relative redistribution grows roughly in line with support for redistribution. This is particularly true for between-country variation, but less apparent within countries. However, this is perhaps not surprising, given that the variation within countries is relatively small for either variable (see Kenworthy and McCall 2008, for a more careful analysis of within-country trends). Two outliers can be found on the lower right-hand corner of the scatter; these are the Republic of Korea and Taiwan, Province of China, which combine relatively high support for redistribution with very limited actual redistribution.[27] Latin American countries would most likely be found in the same corner: The three countries for which we have data redistribute incomes on a very limited scale (mean relative redistribution: 0.049; see also Goñi et al. 2008), while public opinion in the region strongly favours redistribution (84.8 per cent of respondents in Argentina and 73 per cent in Chile agreed with the statement introduced above). The paper will later return to a substantive interpretation of these outliers.

In line with the previous design, the first regression will use both the time-series and cross-section component of the pooled sample and estimate a random effects model with cluster-robust standard errors. However, due to the mismatch of years for which observations are available from the two

[27] For a detailed study on the redistributive impact of taxes and transfers in the Republic of Korea, see Sung and Park (2011); for an analysis of inequality trends in Taiwan, Province of China, see Bourguignon et al. (2001).

primary sources (LIS and ISSP), the time-series component of the pooled analysis is not always robust and model (5) is presented with a strong caveat. For the same reasons, no within-effects model is estimated and more trust is placed in the between-effects model (6) that uses country means (and therefore only captures the variation between countries).

In addition to the support for redistribution, the models carry over the share of the population aged 65 years and above and the unemployment rate as control variables. To control for 'automatic stabilization' of inequality through the welfare state that goes beyond the impact of unemployment insurance and old-age pensions, we will also maintain the initial level of private sector inequality as an explanatory variable. As before, the regressions will be estimated for the full sample and for a reduced sample that excludes the East Asian economies. (Recall that the Latin American countries are missing from both samples due to lack of data on public opinion).

The pooled analysis on the full sample in model (5a) produces no significant regression coefficients apart from the highly significant coefficient on the variable 'oldage' (see Table 3). At first sight, the performance of the 'behavioural' model is therefore no better than the previous test of the Meltzer and Richard hypothesis. However, this changes when the two East Asian observations are excluded, as done in model (5b): The regression coefficient on support for redistribution is now highly significant (at the 0.01-level), and the control variable 'share of the population aged 65 years or above' also gains significance (at the 0.05-level). With an $R^2$ of 0.490, the explanatory power of the model is satisfactory. The control variables contribute to this, but the model's performance does not depend on their inclusion. A bivariate random effects model (not reported) produced an $R^2$ of 0.236 and a significant regression coefficient on support for redistribution (b: 0.296, significant at the 0.05-level). These results confirm that, as far as the old industrialized countries are concerned, public support for redistribution is an input into the political system that is strongly associated with actual redistribution at the output side of the system (to use Easton's terminology).

**Table 3: Regression results for models with 'support for redistribution' as an explanatory variable (dependent variable: relative redistribution)**

| Variable / Model | (5) Random effects, robust cluster SE | | (6) Between-country effects | |
|---|---|---|---|---|
| | (5a) full sample | (5b) OECD | (6a) full sample | (6b) OECD |
| pi_gini (Gini, private sector incomes) | -0.003<br>(0.491) | -0.819*<br>(0.404) | 0.180<br>(0.727) | -1.357**<br>(0.598) |
| support (support for redistribution) | 0.169<br>(0.114) | 0.348***<br>(0.086) | -0.008<br>(0.197) | 0.383**<br>(0.161) |
| unemp (unemployment rate) | 0.630<br>(0.436) | 0.335<br>(0.308) | 1.461<br>(1.113) | 0.527<br>(0.794) |
| oldage (population 65+ years) | 2.311***<br>(0.631) | 1.534**<br>(0.676) | 2.600**<br>(1.075) | 1.138<br>(0.800) |
| Constant | -0.115<br>(0.237) | 0.285<br>(0.197) | -0.187<br>(0.258) | 0.558**<br>(0.239) |
| n = | 58 | 56 | 58 | 56 |
| Number of clusters / groups | 22 | 20 | 22 | 20 |
| $R^2$ | 0.395 | 0.490 | 0.468 | 0.458 |

Notes: $R^2$ refers to $R^2$ (overall) for the random effects model (5) and to $R^2$ (between) for the between-effects model (6).

Standard errors are given in parenthesis; those in model (5) refer to robust cluster standard errors.
***, **, and * denote significance at risks levels 0.01, 0.05 and 0.10, respectively.
OECD refers to old OECD member countries and those in Europe; the Republic of Korea is excluded.

Source: Based on Luxembourg Income Study (pi_gini and dependent), ISSP and related sources (support), ILO (unemp), World Bank (oldage) and Statistics Bureau of Taiwan, Province of China (unemp and oldage for Taiwan, POC). For details see Appendix 2.

The between-effects model (6a) of the full sample largely replicates the results of the random effects model and does not produce any significant coefficients with the exception of the variable 'oldage'. Once the East Asian countries are excluded, as done in model (6b), support for redistribution once again becomes a significant explanatory variable (at the 0.05-level). Note that the coefficient on the share of the elderly population loses its significance. This could be due to the same effect as discussed above, namely that a rising share of older people over time automatically leads to greater redistribution through the pension system. However, once the time-series component is removed, the effect becomes weaker. The same happens excluding the East Asian economies that combine a young population with low redistribution. Another observation is that the level of inequality itself becomes a significant predictor (p-value: 0.038). However, this does not lend support to the median voter theorem since the sign on the coefficient is *negative*, and hence opposite of what the rational choice model predicts.

The negative coefficient of the initial level of private sector inequality is partly an artefact of the way the dependent variable—relative redistribution—has been constructed. Recall that it is obtained by dividing absolute redistribution (i.e. the difference between the Gini coefficients for disposable incomes and private sector incomes) by the Gini coefficient for private sector incomes. Hence, the same level of absolute redistribution will result in lower value for relative redistribution if initial, private sector inequality were higher. The variable 'pi_gini' (private sector inequality) can therefore best be thought of as control variable that is necessitated by nature of dependent variable. As robustness test, models (7) and (8) re-estimate the previous regressions with *absolute* redistribution as a dependent variable (Table 4). As expected, the coefficient on the initial level of inequality loses its significance in all specifications. While the size of the coefficients changes (the mean and standard deviation of the dependent variable are now smaller), the results generally replicate those of the previous analysis and the coefficient on 'support for redistribution' remains significant at the 0.01- and 0.05-level when the reduced sample of OECD countries is used. It also becomes (marginally) significant at the 0.10-level in the pooled cross-section time-series random effects model on the full sample.

What do these results imply? One interpretation would be that governments in the old OECD countries and in Eastern Europe are, to some extent, responsive to public demands to reduce inequality through the tax and transfer system. However, this conclusion comes with two caveats. The first is that the finding is based primarily on the between-country variation; due to data limitations, this paper has not exploited the time-series element if the dataset. Kenworthy and McCall (2008) study over-time variation in support for redistribution and changes in actual redistribution for 15 countries and find no consistent pattern. Over-time variation in support for redistribution is relatively small when compared to between-country differences, and does not necessarily match redistributive outcomes (which are heavily influenced by other factors, such as the business cycle and unemployment).

**Table 4: Robustness tests for models with 'support for redistribution' as an explanatory variable (dependent variable: absolute redistribution)**

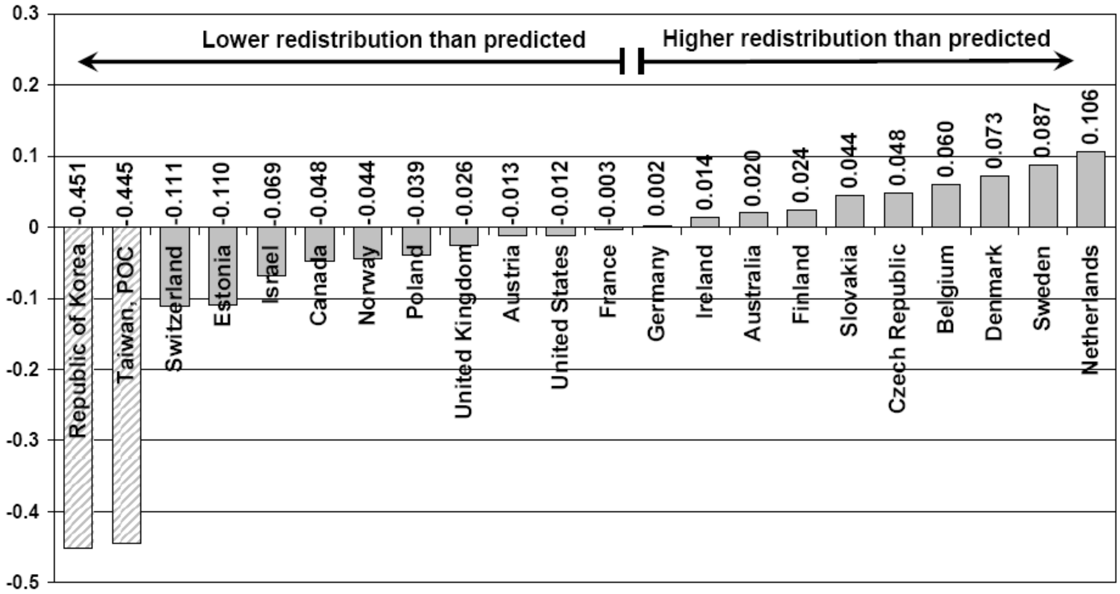
| Variable / Model | (7) Random effects, robust cluster SE | | (8) Between-country effects | |
|---|---|---|---|---|
| | (7a) full sample | (7b) OECD | (8a) full sample | (8b) OECD |
| pi_gini (Gini, private sector incomes) | 0.263<br>(0.191) | -0.019<br>(0.179) | 0.325<br>(0.288) | -0.218<br>(0.268) |
| support (support for redistribution) | 0.089*<br>(0.045) | 0.151***<br>(0.039) | 0.026<br>(0.078) | 0.165**<br>(0.072) |
| unemp (unemployment rate) | 0.276<br>(0.175) | 0.173<br>(0.136) | 0.579<br>(0.441) | 0.243<br>(0.356) |
| oldage (population 65+ years) | 0.983***<br>(0.271) | 0.714**<br>(0.308) | 1.073**<br>(0.425) | 0.553<br>(0.359) |
| Constant | -0.169<br>(0.92) | -0.031<br>(0.087) | -0.191*<br>(0.102) | 0.071<br>(0.107) |
| n = | 58 | 56 | 58 | 56 |
| Number of clusters / groups | 22 | 20 | 22 | 20 |
| $R^2$ | 0.540 | 0.529 | 0.622 | 0.473 |

Notes: $R^2$ refers to $R^2$ (overall) for the random effects model (7) and to $R^2$ (between) for the between-effects model (8).
Standard errors are given in parenthesis; those in model (7) refer to robust cluster standard errors.
***, **, and * denote significance at risks levels 0.01, 0.05 and 0.10, respectively.
OECD refers to old OECD member countries and those in Europe; the Republic of Korea is excluded.

Source: Based on Luxembourg Income Study (pi_gini and dependent), ISSP and related sources (support), ILO (unemp), World Bank (oldage) and Statistics Bureau of Taiwan, Province of China (unemp and oldage for Taiwan, POC). For details see Appendix 2.

The second caveat is that the aggregate finding may obscure a lack of government responsiveness in some countries, or in fact hide a bias in one direction across all countries. In the United States, the corrosive effects of inequality on democracy itself have become an issue of debate, as highlighted by the influential study of the APSA Task Force on Inequality and American Democracy (see Jacobs and Skocpol 2005). To summarize a complex literature, it appears that a government is responsive to citizens' demands, but more so to the views of affluent voters (Gilens 2005). This matters since the poor and the rich differ in their preferences when it comes to welfare spending and other policies with redistributive consequences (Gilens 2009). At the level of individual United States senators, Bartels (2005) shows that they are more responsive to the views of affluent constituents in their home state than to those held by middle-class voters; the preferences of the bottom strata have no statistically significant impact on senators' voting behaviour in Congress. Effectively, this literature gives support to the argument that the hypothetical median voter of Meltzer's and Richard's model world is not the decisive voter in the real world.

**Figure 4: Departures from predicted extent of relative redistribution, by country (Residuals from regression model 6b)**




Note: Grey bars refer to residuals from regression model (6b). The same regression equation was also applied to Korea and Taiwan, Province of China, to predict the extent of redistribution under the counterfactual assumption that these two economies displayed the same characteristics as the advanced countries (light grey bars).

Source: Based on Luxembourg Income Study (pi_gini and dependent), ISSP and related sources (support), ILO (unemp), World Bank (oldage) and Statistics Bureau of Taiwan, Province of China (unemp and oldage for Taiwan, POC). For details see Appendix 2.

Does the United States stand out for ignoring redistributive preferences of their voters? To approach this question, it is useful to look at the unexplained departure from the extent of redistribution that one would expect to find, given public support for redistribution and demographic factors. Figure 4 therefore displays the residuals from the between-effects regression model (6b). The striking finding is that redistribution in the United States is almost exactly in line with the model prediction (residual: -0.012). When compared to France (residual: -0.003) or Germany (residual: 0.002), the lower level of redistribution in the United States largely reflects difference in (measured) public opinion,[28] initial inequality, unemployment and demographic structure—and not a fundamental difference in how the political system translates inputs into outputs. This finding, however, leaves open to debate whether all of these countries share the same elite-bias. It is informative that some European welfare states (Denmark, Sweden, and the Netherlands) offer greater redistribution than expected—and somewhat counterintuitive that Switzerland, with its strong tradition of direct democracy, redistributes substantially less. In both cases, the historical evolution of the welfare state might offer an explanation. By far the greatest disconcordance can be observed in the two East Asian economies: Both the Republic of Korea (residual: -0.451) and Taiwan, Province of China, (residual: -0.445) have less government redistribution than one would expect if they behaved like the old OECD countries.

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[28] An argument could be made that measured public opinion differs from actual public opinion, given disproportionately high non-response to the ISSP item described in Section 5 among poorer respondents (see Jacobs and Skocpol 2005: 217). However, this effect should also apply in other countries.

## 6. Conclusions and discussion

Explanations for the extent to which governments redistribute income through the tax and transfer system provide for an interesting example to contrast rational choice and behavioural perspectives, and how they differ in understanding human motivation. In a classical paper, Meltzer and Richard (1981) provide a theoretical 'proof' that individual utility maximization and the vote-seeking behaviour of politicians under majority rule produce greater redistribution when inequality is high. The model exemplifies the deductive reasoning of rational choice, and applies the median-voter theory of Schumpeter (1942) and Downs (1957) to a tangible question. As even critics would concede, the model is elegant and parsimonious and its logic is intuitively compelling. Yet, it suffers from the shortcomings of its very foundations that behavioural economics has found wanting. By portraying humans as 'rational fools' (to use Amartya Sen's term), rational choice ignores that people are embedded in a society and share values and perceptions of fairness and social justice.

While a host of papers has tested the relationship between inequality and redistribution, one unresolved issue in the literature has been how best to define redistribution in empirical tests. While some authors have used 'absolute redistribution' (measured as the difference between the Gini for private sector incomes and the Gini for disposable incomes), others have chosen a relative concept of redistribution (i.e. the reduction of the Gini coefficient relative to its initial level). To resolve this question, the present paper has re-visited the original article by Meltzer and Richard and shown that their model assumptions imply that *relative* redistribution should rise in line with initial inequality. An increase in *absolute* redistribution should arise from the automatic stabilization properties of welfare states and would not provide conclusive evidence.

The empirical analysis in this paper has—in line with previous findings—shown that the simple mechanism of rational choice is a poor guide to reality. The 'pooled' analysis of cross-section time-series data with 110 observations from 26 countries revealed no significant relationship between inequality of private sector incomes and subsequent relative distribution. The approach also failed to account for differences between countries, but at first appeared to have some utility in explaining within-country changes over time. However, the explanatory power of the model remained poor and the regression coefficient on inequality became insignificant once control variables were added. As it turned out, changes in unemployment and an increasing share of the elderly population offer a simple and more powerful alternative explanation for the observed over-time changes in relative redistribution since the 1970s.

The failure of the real world to behave in line with model prediction puts into question the two underlying premises: that voters' support for redistribution strictly depends on what they personally have to gain from it (i.e. their utility maximization), and that the political system produces outputs that are aligned to the median voter's interests. While a large body of literature has concentrated on the second point, the present paper has adopted the perspective of behavioural economics that has challenged the axiom of rational utility maximization and explored the role of altruism, social norms and values and fairness in explaining people's choices. These have of course been central to political sociology and comparative welfare state research long before they entered mainstream economic analysis, as evident from a rich body of literature that has previously studied voters' actual views on inequality and redistribution. As it turns out, observed support for redistribution—measured as the share in the population who thinks that it is the government's role to reduce income differences—is a far better predictor for actual redistribution, at least in the old industrialized countries.

One caveat needs to be added to this analysis. People's views on what is just and fair, and on how the government should intervene in market outcomes, are of course shaped (but not fully conditioned) by their socialization in a political system. Hence, the institutions that redistribute income—primarily the tax system and social security institutions—will influence voters' views on redistribution and can often generate their own legitimacy through performance. The direction of causality is therefore open to debate, and it may well run in both ways. In Easton's (1965) terminology, the output of a political system is evaluated by the population and, through a feedback loop, influences the input that feeds into the system through elections or other forms of political participation. The central argument made in this paper is that, when analyzing inputs into the system, observing and measuring what people want is a better guide to reality than simply deducting what they want on the basis of assumptions about their rational utility maximization.

This type of micro-level analysis has some justification and utility. Beliefs in the fairness of the market can help to explain why support for redistribution has remained lukewarm among the American electorate despite the surge in inequality in the United States (see Shapiro and Young 1989; Osberg and Smeeding 2006). The failure of voters to demand sharp tax increases for the top 1 per cent of income earners who control almost a quarter of incomes (Atkinson et al. 2011) could be pinned down to plain 'irrationality', but historical explanations are richer and more useful to understand why Americans are so adverse to government redistribution (see e.g. Lipset and Marks 2001). Having established that beliefs in fairness matter, behavioural economics has broadened the perspective of economics to include such considerations. The limitation of this analysis is that it often remains ahistorical, and abstracts from power relations within society. The danger is to simply replace the supposedly 'rational' choice of individuals with another simplistic explanation of human choice that follows universal behavioural dispositions and has no space for human agency or historical context (see Streeck 2010).

For the analysis of distribution and redistribution, an approach in the tradition of classical political economy might offer deeper insights. The historical evolution of wealth and income inequality in Latin America is a case in point. The lack of welfare states in the region needs a complex explanation, and is certainly not due to lack of public support for redistribution. Opinion survey data from Latin America in fact show overwhelming support for reducing income inequality, but the historical legacy and wealth concentration mean that inclusive social security institutions are largely absent. However, the upshot is that leaders such as President Lula da Silva in Brazil can ride on public opinion to expand social security schemes like as *bolsa familia*—and that even conservative opposition parties extend their support to them once they have become popular. Redistribution might not follow automatically where it is most needed to reduce poverty, but democracy opens up the space for human agency to affect policy outcomes (Huber et al. 2006).

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## **Appendix 1**: Alternative proof for equation (1')

The Gini coefficient for market incomes, $G_m$, can be calculated using the Brown (1994) formula:

$$G_m = 1 - \sum_{k=1}^{n} (X_k - X_{k-1})(Y_k + Y_{k-1}) \quad \text{(A 1)}$$

where $X_k$ is the cumulated proportion of the population variable, for k = 0,...,n, with $X_0 = 0$, $X_n = 1$; and $Y_k$ is the cumulated proportion of the income variable, for k = 0,...,n, with $Y_0 = 0$, $Y_n = 1$. The equation can be transformed into:

$$\sum_{k=1}^{n} (X_k - X_{k-1})(Y_k + Y_{k-1}) = 1 - G_m \quad \text{A 1')}$$

When redistribution takes place, the cumulated proportion of the income variable, $Y_k$, is first reduced by the tax rate *t* (and therefore multiplied by *1-t*). Total tax receipts equal *t* and are redistributed in equal lump-sums across the entire population. Hence, each cumulated proportion *k* receives lump sums proportional to its population share, or $t \times X_k$ . The same applies to the cumulated proportion for the preceding class, *k-1*. One can thus calculate the Gini coefficient for disposable incomes, $G_d$, as follows:

$$G_d = 1 - \sum_{k=1}^{n} (X_k - X_{k-1}) \times \left[(1-t) \times Y_k + t \times X_k + (1-t) \times Y_{k-1} + t \times X_{k-1}\right] \quad \text{(A 2)}$$

$$G_d = 1 - \sum_{k=1}^{n} (X_k - X_{k-1}) \times \left[(1-t)(Y_k + Y_{k-1}) + t(X_k + X_{k-1})\right] \quad \text{(A 3)}$$

$$G_d = 1 - \sum_{k=1}^{n} (1-t)(X_k - X_{k-1})(Y_k + Y_{k-1}) + t(X_k - X_{k-1})(X_k + X_{k-1}) \quad \text{(A 4)}$$

$$G_d = 1 - \sum_{k=1}^{n} (1-t)(X_k - X_{k-1})(Y_k + Y_{k-1}) + t({X_k}^2 - {X_{k-1}}^2) \quad \text{(A 5)}$$

$$G_d = 1 - \left[\sum_{k=1}^{n} (1-t)(X_k - X_{k-1})(Y_k + Y_{k-1})\right] - \left[\sum_{k=1}^{n} t({X_k}^2 - {X_{k-1}}^2)\right] \quad \text{(A 6)}$$

$$G_d = 1 - (1-t)\left[\sum_{k=1}^{n} (X_k - X_{k-1})(Y_k + Y_{k-1})\right] - t\left[\sum_{k=1}^{n} ({X_k}^2 - {X_{k-1}}^2)\right] \quad \text{(A 7)}$$

$$G_d = 1 - (1-t)\left[\sum_{k=1}^{n} (X_k - X_{k-1})(Y_k + Y_{k-1})\right] - t\left[{X_n}^2 - {X_0}^2\right] \quad \text{(A 8)}$$

$$G_d = 1 - (1-t)\left[\sum_{k=1}^{n} (X_k - X_{k-1})(Y_k + Y_{k-1})\right] - t\left[1^2 - 0^2\right] \quad \text{(A 9)}$$

$$G_d = 1 - (1-t)\left[\sum_{k=1}^{n} (X_k - X_{k-1})(Y_k + Y_{k-1})\right] - t \quad \text{(A 10)}$$

Substituting (A 1') into (A 10) leads to:

$$G_d = 1 - (1-t)(1-G_m) - t \qquad \text{(A 9)}$$

$$G_d = (1-t) \times G_m \qquad \text{(1')}$$

## **Appendix 2**: Variable definitions and sources

| Variable | Definition | Source |
|---|---|---|
| pi_gini | Gini coefficient for private sector incomes (defined as sum of LIS variables MI, PRIVATI and V36). Standard LIS routines on equivalence scale, missing observations and top- and bottom-coding are used. | Luxembourg Income Study Database (LIS); analysis of micro-data completed between February and May 2011. |
| dpi_gini | Gini coefficient for disposable incomes (defined as LIS variable dpi). Standard LIS routines on equivalence scale, missing observations and top- and bottom-coding are used. | |
| absolute | Absolute difference between Gini coefficient for private sector incomes and Gini coefficient for disposable incomes (pi_gini – dpi_gini). | |
| relative | Difference between Gini coefficient for private sector incomes and Gini coefficient for disposable incomes, expressed as a fraction of the Gini coefficient for private sector incomes ([pi_gini – dpi_gini] / pi_gini). | |
| OECD | Dummy variable that takes the value of 1 for the old OECD countries and countries in Eastern Europe, and the value of 0 for all other countries and territories (Brazil, Colombia, Guatemala, Republic of Korea, Taiwan [Province of China]). | Organisation for Economic Co-operation and Development. |
| unemp | Unemployment rate, as a fraction of the total labour force. | ILO Laborsta and Statistics Bureau of Taiwan, Province of China. |
| oldage | Share of the population aged 65 years and above, as a fraction of the total population. | World Bank, World Development Indicators, and Statistics Bureau of Taiwan, Province of China. |
| support | Share of respondents who either support or strongly support the statement ‘It is the responsibility to reduce the differences in income between people with high incomes and those with low incomes’, expressed as a fraction.<br>If no survey data are available for the year to which the LIS observation refers, the data-point is estimated based on simple interpolation (in case two neighbouring observations are available) or based on the assumption that support for redistribution has remained stable (in case only one neighbouring observation is available).<br>The observations for post-reunification Germany are a population-weighted average of the ISSP observations for East and West Germany. | ISSP modules in Social Inequality (1987, 1992, 1999 and 2009) and Role of Government (1990 and 1996) from Gesis (archive numbers ZA1680, ZA1950, ZA2310, ZA2900, ZA 3430 and ZA5400).<br>Additional observations for Switzerland (1999) from SIDOS (archive number SID6396), for Ireland (1987 and 1999) from the SSRC, Dublin (no archive number supplied), and for Denmark (1999) from Alborg University (no archive number supplied). |

## **Appendix 3**: Basic descriptive statistics of the dataset

| Country | Observations | | | pi_gini | relative | unemp | oldage | support |
|---|---|---|---|---|---|---|---|---|
| | Number | First | Last | mean | mean | mean | mean | mean |
| Australia | 6 | 1981 | 2003 | 0.440 | 0.311 | 0.069 | 0.114 | 0.458 |
| Austria | 1 | 2004 | 2004 | 0.458 | 0.414 | 0.049 | 0.160 | 0.741 |
| Belgium | 2 | 1992 | 1997 | 0.465 | 0.491 | 0.083 | 0.159 | 0.738 |
| Brazil | 1 | 2006 | 2006 | 0.555 | 0.124 | 0.084 | 0.063 | n/a |
| Canada | 10 | 1971 | 2004 | 0.407 | 0.269 | 0.082 | 0.110 | 0.464 |
| Colombia | 1 | 2004 | 2004 | 0.507 | -0.001 | 0.128 | 0.050 | n/a |
| Czech Republic | 3 | 1992 | 2004 | 0.437 | 0.442 | 0.052 | 0.134 | 0.652 |
| Denmark | 5 | 1987 | 2004 | 0.415 | 0.440 | 0.065 | 0.152 | 0.503 |
| Estonia | 1 | 2004 | 2004 | 0.493 | 0.310 | 0.097 | 0.164 | 0.768 |
| Finland | 5 | 1987 | 2004 | 0.432 | 0.476 | 0.091 | 0.142 | 0.736 |
| France | 4 | 1979 | 1994 | 0.469 | 0.377 | 0.094 | 0.141 | 0.711 |
| Germany | 9 | 1973 | 2004 | 0.421 | 0.367 | 0.070 | 0.155 | 0.611 |
| Guatemala | 1 | 2006 | 2006 | 0.519 | 0.023 | 0.018 | 0.043 | n/a |
| Ireland | 2 | 1987 | 2004 | 0.500 | 0.360 | 0.106 | 0.109 | 0.749 |
| Israel | 6 | 1979 | 2005 | 0.461 | 0.289 | 0.079 | 0.095 | 0.748 |
| Korea | 1 | 2006 | 2006 | 0.334 | 0.068 | 0.035 | 0.097 | 0.751 |
| Netherlands | 5 | 1983 | 2004 | 0.445 | 0.424 | 0.078 | 0.129 | 0.651 |
| Norway | 6 | 1979 | 2004 | 0.387 | 0.383 | 0.037 | 0.154 | 0.585 |
| Poland | 2 | 1999 | 2004 | 0.499 | 0.390 | 0.164 | 0.126 | 0.833 |
| Romania | 2 | 1995 | 1997 | 0.372 | 0.255 | 0.070 | 0.123 | n/a |
| Slovakia | 1 | 1992 | 1992 | 0.399 | 0.526 | 0.114 | 0.106 | 0.745 |
| Sweden | 8 | 1967 | 2005 | 0.429 | 0.466 | 0.040 | 0.165 | 0.574 |
| Switzerland | 5 | 1982 | 2004 | 0.395 | 0.272 | 0.026 | 0.151 | 0.475 |
| Taiwan, POC | 7 | 1981 | 2005 | 0.292 | 0.039 | 0.025 | 0.072 | 0.663 |
| United Kingdom | 9 | 1969 | 2004 | 0.449 | 0.299 | 0.067 | 0.151 | 0.593 |
| United States | 7 | 1979 | 2004 | 0.452 | 0.229 | 0.057 | 0.121 | 0.340 |
| *Total* | *110* | *1967* | *2006* | *0.426* | *0.323* | *0.066* | *0.135* | *0.584* |

Note: Data for the variables pi_gini, relative, unemp and oldage are available for all data points listed in the first columns; data for the variable support are sometimes only available for a sub-set of observations. For details, please see Appendix 4 and use the full dataset for any replication of between-effects models.

Source: Based on Luxembourg Income Study (pi_gini and relative), ISSP and related sources (support), ILO (unemp), World Bank (oldage) and Statistics Bureau of Taiwan, Province of China (unemp and oldage for Taiwan, POC). For details see Appendix 2.

## **Appendix 4**: Documentation of the dataset

| country | pi_m | dpi_m | absolute | relative | OECD | unemp | oldage | support |
|---|---|---|---|---|---|---|---|---|
| Australia | 0.396 | 0.281 | 0.115 | 0.290 | 1 | 0.058 | 0.098 | -1.000 |
| Australia | 0.417 | 0.292 | 0.124 | 0.298 | 1 | 0.083 | 0.103 | 0.429 |
| Australia | 0.428 | 0.304 | 0.124 | 0.290 | 1 | 0.062 | 0.110 | 0.428 |
| Australia | 0.463 | 0.308 | 0.154 | 0.334 | 1 | 0.084 | 0.121 | 0.425 |
| Australia | 0.474 | 0.317 | 0.157 | 0.332 | 1 | 0.068 | 0.126 | 0.501 |
| Australia | 0.460 | 0.312 | 0.148 | 0.322 | 1 | 0.059 | 0.127 | 0.506 |
| Austria | 0.458 | 0.269 | 0.190 | 0.414 | 1 | 0.049 | 0.160 | 0.741 |
| Belgium | 0.449 | 0.224 | 0.226 | 0.502 | 1 | 0.077 | 0.154 | -1.000 |
| Belgium | 0.481 | 0.250 | 0.231 | 0.480 | 1 | 0.089 | 0.164 | 0.738 |
| Brazil | 0.555 | 0.486 | 0.069 | 0.124 | 0 | 0.084 | 0.063 | -1.000 |
| Canada | 0.395 | 0.316 | 0.079 | 0.199 | 1 | 0.064 | 0.080 | -1.000 |
| Canada | 0.375 | 0.289 | 0.086 | 0.229 | 1 | 0.069 | 0.085 | -1.000 |
| Canada | 0.370 | 0.284 | 0.086 | 0.233 | 1 | 0.075 | 0.096 | -1.000 |
| Canada | 0.387 | 0.283 | 0.104 | 0.270 | 1 | 0.088 | 0.107 | -1.000 |
| Canada | 0.405 | 0.281 | 0.124 | 0.306 | 1 | 0.103 | 0.114 | 0.482 |
| Canada | 0.419 | 0.284 | 0.135 | 0.322 | 1 | 0.104 | 0.118 | 0.457 |
| Canada | 0.417 | 0.291 | 0.126 | 0.301 | 1 | 0.091 | 0.122 | 0.446 |
| Canada | 0.441 | 0.311 | 0.130 | 0.295 | 1 | 0.083 | 0.124 | 0.460 |
| Canada | 0.429 | 0.315 | 0.115 | 0.267 | 1 | 0.068 | 0.126 | 0.475 |
| Canada | 0.432 | 0.318 | 0.113 | 0.263 | 1 | 0.072 | 0.130 | -1.000 |
| Colombia | 0.507 | 0.508 | -0.001 | -0.001 | 0 | 0.128 | 0.050 | -1.000 |
| Czech Republic | 0.406 | 0.207 | 0.198 | 0.489 | 1 | 0.033 | 0.129 | 0.669 |
| Czech Republic | 0.438 | 0.259 | 0.179 | 0.409 | 1 | 0.039 | 0.134 | 0.603 |
| Czech Republic | 0.467 | 0.267 | 0.200 | 0.428 | 1 | 0.083 | 0.141 | 0.684 |
| Denmark | 0.398 | 0.254 | 0.143 | 0.361 | 1 | 0.061 | 0.154 | -1.000 |
| Denmark | 0.426 | 0.236 | 0.189 | 0.445 | 1 | 0.090 | 0.155 | -1.000 |
| Denmark | 0.421 | 0.218 | 0.203 | 0.483 | 1 | 0.070 | 0.153 | -1.000 |
| Denmark | 0.412 | 0.225 | 0.188 | 0.455 | 1 | 0.046 | 0.148 | 0.493 |
| Denmark | 0.419 | 0.228 | 0.190 | 0.455 | 1 | 0.056 | 0.150 | 0.513 |
| Estonia | 0.493 | 0.340 | 0.153 | 0.310 | 1 | 0.097 | 0.164 | 0.768 |
| Finland | 0.393 | 0.209 | 0.184 | 0.469 | 1 | 0.050 | 0.128 | -1.000 |
| Finland | 0.407 | 0.210 | 0.197 | 0.485 | 1 | 0.066 | 0.136 | -1.000 |
| Finland | 0.438 | 0.217 | 0.222 | 0.506 | 1 | 0.152 | 0.142 | -1.000 |
| Finland | 0.459 | 0.246 | 0.213 | 0.464 | 1 | 0.097 | 0.149 | -1.000 |
| Finland | 0.463 | 0.252 | 0.211 | 0.456 | 1 | 0.088 | 0.157 | 0.736 |
| France | 0.447 | 0.294 | 0.154 | 0.344 | 1 | 0.059 | 0.141 | -1.000 |
| France | 0.469 | 0.298 | 0.171 | 0.365 | 1 | 0.098 | 0.131 | -1.000 |
| France | 0.474 | 0.287 | 0.187 | 0.395 | 1 | 0.094 | 0.139 | -1.000 |
| France | 0.485 | 0.288 | 0.197 | 0.405 | 1 | 0.123 | 0.152 | 0.711 |
| Germany | 0.356 | 0.271 | 0.085 | 0.240 | 1 | 0.012 | 0.144 | -1.000 |
| Germany | 0.387 | 0.264 | 0.123 | 0.317 | 1 | 0.039 | 0.155 | -1.000 |
| Germany | 0.388 | 0.245 | 0.143 | 0.369 | 1 | 0.042 | 0.155 | -1.000 |
| Germany | 0.385 | 0.260 | 0.125 | 0.324 | 1 | 0.082 | 0.150 | -1.000 |
| Germany | 0.442 | 0.265 | 0.177 | 0.400 | 1 | 0.087 | 0.147 | -1.000 |
| Germany | 0.429 | 0.258 | 0.171 | 0.398 | 1 | 0.080 | 0.148 | 0.643 |

| country | pi_m | dpi_m | absolute | relative | OECD | unemp | oldage | support |
|---|---|---|---|---|---|---|---|---|
| Germany | 0.450 | 0.270 | 0.180 | 0.401 | 1 | 0.103 | 0.153 | 0.622 |
| Germany | 0.464 | 0.266 | 0.198 | 0.427 | 1 | 0.079 | 0.164 | 0.568 |
| Germany | 0.488 | 0.278 | 0.210 | 0.430 | 1 | 0.110 | 0.184 | -1.000 |
| Guatemala | 0.519 | 0.507 | 0.012 | 0.023 | 0 | 0.018 | 0.043 | -1.000 |
| Ireland | 0.500 | 0.328 | 0.171 | 0.343 | 1 | 0.169 | 0.108 | 0.716 |
| Ireland | 0.500 | 0.312 | 0.188 | 0.376 | 1 | 0.044 | 0.110 | 0.781 |
| Israel | 0.411 | 0.303 | 0.109 | 0.264 | 1 | 0.029 | 0.085 | -1.000 |
| Israel | 0.448 | 0.308 | 0.140 | 0.313 | 1 | 0.071 | 0.088 | -1.000 |
| Israel | 0.442 | 0.305 | 0.137 | 0.309 | 1 | 0.112 | 0.094 | 0.658 |
| Israel | 0.473 | 0.336 | 0.137 | 0.290 | 1 | 0.077 | 0.099 | 0.736 |
| Israel | 0.502 | 0.346 | 0.156 | 0.311 | 1 | 0.094 | 0.099 | 0.804 |
| Israel | 0.490 | 0.370 | 0.120 | 0.246 | 1 | 0.090 | 0.101 | 0.793 |
| Korea | 0.334 | 0.311 | 0.023 | 0.068 | 0 | 0.035 | 0.097 | 0.751 |
| Netherlands | 0.470 | 0.260 | 0.210 | 0.447 | 1 | 0.134 | 0.118 | -1.000 |
| Netherlands | 0.475 | 0.256 | 0.219 | 0.461 | 1 | 0.100 | 0.124 | 0.651 |
| Netherlands | 0.448 | 0.266 | 0.182 | 0.406 | 1 | 0.070 | 0.129 | -1.000 |
| Netherlands | 0.372 | 0.231 | 0.142 | 0.381 | 1 | 0.036 | 0.135 | -1.000 |
| Netherlands | 0.459 | 0.263 | 0.196 | 0.427 | 1 | 0.050 | 0.140 | -1.000 |
| Norway | 0.364 | 0.223 | 0.141 | 0.387 | 1 | 0.020 | 0.145 | -1.000 |
| Norway | 0.352 | 0.233 | 0.119 | 0.339 | 1 | 0.020 | 0.159 | -1.000 |
| Norway | 0.374 | 0.231 | 0.142 | 0.381 | 1 | 0.055 | 0.163 | 0.583 |
| Norway | 0.400 | 0.238 | 0.162 | 0.405 | 1 | 0.049 | 0.159 | 0.577 |
| Norway | 0.403 | 0.250 | 0.153 | 0.379 | 1 | 0.034 | 0.150 | 0.609 |
| Norway | 0.430 | 0.256 | 0.174 | 0.405 | 1 | 0.045 | 0.146 | 0.569 |
| Poland | 0.474 | 0.289 | 0.185 | 0.390 | 1 | 0.139 | 0.120 | 0.849 |
| Poland | 0.524 | 0.320 | 0.204 | 0.389 | 1 | 0.190 | 0.131 | 0.818 |
| Romania | 0.373 | 0.277 | 0.095 | 0.256 | 1 | 0.080 | 0.120 | -1.000 |
| Romania | 0.371 | 0.277 | 0.094 | 0.254 | 1 | 0.060 | 0.126 | -1.000 |
| Slovakia | 0.399 | 0.189 | 0.210 | 0.526 | 1 | 0.114 | 0.106 | 0.745 |
| Sweden | 0.390 | 0.260 | 0.130 | 0.333 | 1 | 0.019 | 0.130 | -1.000 |
| Sweden | 0.400 | 0.215 | 0.185 | 0.462 | 1 | 0.016 | 0.151 | -1.000 |
| Sweden | 0.411 | 0.197 | 0.214 | 0.521 | 1 | 0.025 | 0.166 | -1.000 |
| Sweden | 0.428 | 0.218 | 0.210 | 0.491 | 1 | 0.021 | 0.180 | -1.000 |
| Sweden | 0.461 | 0.229 | 0.232 | 0.503 | 1 | 0.052 | 0.177 | 0.537 |
| Sweden | 0.459 | 0.221 | 0.237 | 0.518 | 1 | 0.077 | 0.175 | 0.579 |
| Sweden | 0.446 | 0.252 | 0.195 | 0.436 | 1 | 0.047 | 0.172 | 0.593 |
| Sweden | 0.440 | 0.237 | 0.203 | 0.462 | 1 | 0.060 | 0.172 | 0.586 |
| Switzerland | 0.397 | 0.309 | 0.088 | 0.221 | 1 | 0.004 | 0.140 | -1.000 |
| Switzerland | 0.403 | 0.307 | 0.096 | 0.238 | 1 | 0.028 | 0.146 | 0.487 |
| Switzerland | 0.387 | 0.280 | 0.107 | 0.276 | 1 | 0.027 | 0.154 | 0.422 |
| Switzerland | 0.392 | 0.274 | 0.119 | 0.302 | 1 | 0.029 | 0.156 | 0.471 |
| Switzerland | 0.395 | 0.268 | 0.128 | 0.323 | 1 | 0.043 | 0.158 | 0.521 |
| Taiwan, POC | 0.272 | 0.267 | 0.005 | 0.018 | 0 | 0.014 | 0.044 | -1.000 |
| Taiwan, POC | 0.275 | 0.269 | 0.007 | 0.024 | 0 | 0.027 | 0.053 | -1.000 |
| Taiwan, POC | 0.281 | 0.271 | 0.009 | 0.034 | 0 | 0.015 | 0.065 | -1.000 |
| Taiwan, POC | 0.295 | 0.277 | 0.019 | 0.063 | 0 | 0.018 | 0.076 | -1.000 |
| Taiwan, POC | 0.297 | 0.287 | 0.010 | 0.032 | 0 | 0.027 | 0.081 | -1.000 |
| Taiwan, POC | 0.302 | 0.289 | 0.013 | 0.044 | 0 | 0.030 | 0.086 | -1.000 |

| country | pi_m | dpi_m | absolute | relative | OECD | unemp | oldage | support |
|---|---|---|---|---|---|---|---|---|
| Taiwan, POC | 0.324 | 0.305 | 0.019 | 0.058 | 0 | 0.041 | 0.097 | 0.663 |
| United Kingdom | 0.353 | 0.267 | 0.085 | 0.242 | 1 | 0.033 | 0.128 | -1.000 |
| United Kingdom | 0.350 | 0.268 | 0.081 | 0.233 | 1 | 0.026 | 0.138 | -1.000 |
| United Kingdom | 0.396 | 0.270 | 0.126 | 0.318 | 1 | 0.053 | 0.148 | -1.000 |
| United Kingdom | 0.476 | 0.303 | 0.173 | 0.363 | 1 | 0.118 | 0.153 | 0.582 |
| United Kingdom | 0.475 | 0.336 | 0.139 | 0.293 | 1 | 0.084 | 0.158 | 0.614 |
| United Kingdom | 0.502 | 0.339 | 0.163 | 0.324 | 1 | 0.096 | 0.159 | 0.614 |
| United Kingdom | 0.503 | 0.344 | 0.158 | 0.315 | 1 | 0.086 | 0.159 | 0.569 |
| United Kingdom | 0.497 | 0.347 | 0.150 | 0.302 | 1 | 0.060 | 0.159 | 0.587 |
| United Kingdom | 0.490 | 0.345 | 0.146 | 0.297 | 1 | 0.047 | 0.160 | 0.595 |
| United States | 0.402 | 0.301 | 0.101 | 0.250 | 1 | 0.058 | 0.111 | -1.000 |
| United States | 0.432 | 0.335 | 0.097 | 0.224 | 1 | 0.070 | 0.119 | 0.292 |
| United States | 0.439 | 0.338 | 0.101 | 0.229 | 1 | 0.068 | 0.124 | 0.362 |
| United States | 0.465 | 0.355 | 0.110 | 0.236 | 1 | 0.061 | 0.125 | 0.359 |
| United States | 0.475 | 0.372 | 0.103 | 0.216 | 1 | 0.049 | 0.124 | 0.335 |
| United States | 0.473 | 0.368 | 0.105 | 0.222 | 1 | 0.040 | 0.124 | 0.353 |
| United States | 0.481 | 0.372 | 0.108 | 0.225 | 1 | 0.055 | 0.123 | -1.000 |

Note: See Appendix 2 for variable definition. The value -1 refers to missing data. Due to space constraints, all figures reproduced here are rounded to three digits behind the decimal point. Replication results might therefore differ slightly from the results presented in this paper.

Source: Based on Luxembourg Income Study (pi_gini and relative), ISSP and related sources (support), ILO (unemp), World Bank (oldage) and Statistics Bureau of Taiwan, Province of China (unemp and oldage for Taiwan, POC). For details see Appendix 2.