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Abstract

A common assumption in political economy is that there exists a consistent and well defined policy space. Often, this space is assumed to be adequately represented by a single 'left' - 'right' dimension. This paper makes the case that it is not only convenient but also meaningful to talk of the left and the right. Motivated, in part, by recent work in political psychology, this paper compares how individuals place themselves on a left-right scale with their answers to substantive policy questions, to provide evidence that the left-right scale has a consistent meaning across time and place. It is also finds consistent differences in how different demographic groups perceive the 'left'-'right' continuum. In particular, it finds important differences associated with ageing, gender, income and education. It provides evidence that this is true for both abstract alternatives and concrete choices, questions of redistribution and broader conceptions of social justice. Heterogeneity is taken seriously, analysing variation within cohorts defined by country, date of birth, and gender - a variety of different forms are hypothesised, tested for, and rejected. Finally, it provides evidence that increases in income may lead to increased levels of political polarisation.

Keywords: Ideology, Voter Preferences, Voting, Polarization

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Reducing the complexity of differences in political opinion to a single left-right scale is commonplace in the media, public discourse, and academic inquiry. It is also often a convenient modelling assumption in Economics, Political Science, and other social-sciences. It is thus important to be clear about whether this is a reasonable assumption. This paper provides evidence that the terms left and right have a consistent meaning across countries and time - vitiating this assumption. The evidence further suggests that there is also a similarly consistent relationship between individuals' characteristics and their own ideological positions. But, as the title suggests, both findings are subject to an important caveat – certain characteristics, for example gender, consistently predict differences between individuals' subjective view of their ideological position compared to a range of objective measures. For example, across country, time, and generation, the better educated are found to consistently be more likely to consider themselves liberal or left-wing whilst being more likely to favour increased inequality. This results seems driven by differences in self-perception as it holds for a range of different objective questions. Whilst, this suggests that we should be careful when analysing the political preferences of different groups using a left-right measure, it also suggests we can improve the usefulness of this simple scale by conditioning on demographic variables. The results are of relevance beyond Economics, and the qualified support for a left-right scale has implications for how ideological position should be quantified to maximize external validity in Political Science and Political Psychology.

Politicians, parties, and voters are often described as being on the left or the right. Similarly, when economists model political processes we often assume political actors can be satisfactorily ordered on some left-right continuum. This reduction of the vast range of different issues, opinions, and groups to a single dimension is of course only ever an approximation. It is only a useful approximation, if the terms 'left' and 'right' can be said to mean something. That is they have an interpretation that is consistent beyond a specific election, poll, or study. This paper analyses data from the World Values Survey covering just over 280,000 people in 80 countries for nearly 30 years. The evidence suggests that there is, broadly speaking, a consistent meaning of the terms 'left' and 'right'. This paper presents evidence that not only are some groups more likely to be more left wing than others, but their perceptions of their position also vary in a consistent manner. Thus, the usefulness of the left-right continuum can be improved by understand how the population distribution over varies by observable characteristics and allowing for these systematic variations. Of course, there is a great deal of variation some of which may be attributed to the local and temporary issues that may gain and lose salience, but nevertheless it is case that some of this variation should be attributed to differences in how different demographic groups see the relationship between policies and position.

To find such consistency not only in the meaning of left and right, who feels they are left and right, and for whom the definition varies may seem surprising at first. It could seem odd to claim that, say, Swedish politics in the 1980s had much in common with politics in, say, Mexico in 2010. But, there are several good reasons not to be surprised. Firstly, the size and scope of government, and the extent of redistribution is likely to be an important issue in most democracies, as most voters may be expected to have much to lose or gain. (cf. Meltzer and Richard (1981)). This explanation suggests that the terms left-wing or right-wing retain meaning over time because the core issue remain the same. An alternative explanation is offered by a growing literature investigating the psychological, physiological, and genetic correlates of ideological position. One reading of this literature, elaborated below, is that there is a consistent left-wing right-wing dimension because this represents the distribution of physiological types in a society. One example of these competing, although perhaps complimentary, explanations is gender differences in political preferences. As discussed below, this has been argued to represent differences in the benefits of a welfare-state (Edlund and Pande (2002)) but also attributed to differences in empathy (Dodd et al. (2011)) or neurology (Amodio et al. (2007)).

The remainder of this paper is organised as follows. The next section provides a necessarily brief overview of the relevant literature in Economics, Political Science, and Political Psychology. The results are then organised into two sections, the first compares who identifies as left wing and who may reasonably judged as being so. The second, presents evidence that this these relationships are consistent across time, place, and generations. The final section concludes.

2 Defining Left and Right-Wing

2.1 The left-right scale in Economics

Taken together the celebrated results of Black (1948) and Arrow (1963) demonstrate that stable democratic equilibria will only exist if voters preferences are suitably restricted. Black (1948) showed that majoritarian equilibria exist (don't exhibit Condorcet cycles) if the distribution of voters preferences are uni-dimensional and single-peaked. Subsequent work has relaxed these requirements to one of single crossing. An alternative approach to the cycling problem has been probabilistic voting, but such models require similarly restrictive assumptions. The model of Besley and Coate (1997) provides for equilibria in the case of multi-dimensional preferences that are not single-peaked but these equilibria are often in mixed strategies and non-unique. Whilst, this certainly was a major improvement it is limiting from the point of view of generating empricial predictions. It is perhaps unsurprising, therefore, that many more applied papers focus on the uni-dimensional case such as Meltzer and Richard (1981), Wittman (1977), Alesina (1988), Morelli (2004), and more recently, Pickering and Rockey (2011, 2012), Bjørnskov and Paldam (2012), Ensley (2012), Funk and Gathmann (2013), and Potrafke (2012),.[1]

This paper argues, with some caveats, that the evidence from many countries over a period of nearly 30 years lends credibility to these assumptions. The remainder of this section will argue that whilst previous research has demonstrated that a unidimensional model can rarely if ever provide a full account of political preferences in any given country at any given time, that there is evidence for a left-right dimension with a consistent definition. This section also suggests that the this view is not only analytically convenient and supported by evidence but may well have physiological underpinnings.

[1] An exception is Roemer (1999) who models political competition as taking place between parties themselves composed of competing factions and develops a new equilibrium concept, Party Unanimity Nash Equilibrium, which as discussed in Roemer (2001) exists in multi-dimensional policy spaces. However, this clearly represents a major contribution but is limited to situations in which the particular conception of political parties employed is appropriate.

There is also an emerging body of evidence on the the relationship between gender and physiology more generally and political attitudes and behaviour. One interpretation of this work is that it suggests, that in a fixed population, there should be some consistency in the distribution of preferences.[2] However, the potential implications run in both directions, if the meaning of 'left' and 'right' is not well defined over time then the appropriate inference drawn from the results of these studies is more complicated.

Indeed recent work has also suggested how macroeconomic factors can alter political preferences. Di Tella and MacCulloch (2009) argue that individuals who perceive corruption as being widespread are more likely to be left-wing.[3]

Alesina and Giuliano (2009) review the literature on 'Preferences for Redistribution'. Analysing survey data, they ask: What makes individuals desire more or less inequality? Broadly, they consider two classes of explanations. Individuals may care about inequality because of its impact in some way upon themselves or because they desire an income distribution compatible with a preferred concept of social justice. In particular their results suggest that within the US, Women, Blacks, and the more educated all tend to have a greater preference for redistribution. Di Tella and MacCulloch (2005) consider the possibility that causation runs in the opposite direction, and provide evidence for how political preferences may affect macroeconomic outcomes.

These papers also serve to highlight the importance of survey-based research for the economic study of ideology, and the widespread use of the left-right spectrum. This in turn highlights the importance of being clear about the reliability of answers to survey questions and the meaning of a left-right dimension.

The seeming (albeit perhaps reluctant) acceptance of a uni-dimensional model in Economics was at odds until recently with trends in Political Psychology and related fields. Jost et al. (2006) describes how the notion that there was a well defined, stable, and empirically reliable left-right measure of political preference was from Converse (1962) and until quite recently, intellectually, beyond the pale. As he further describes this conclusion is at the very least at odds with the reporting of politics in the media. Despite the arguments of Converse (1962) and others, and despite the political upheaval subsequent to the Civil Rights Movement, the terms remained in widely used. This would be less remarkable if the terms, as employed, referred to very different policies and instincts, but again this does not appear to be the case. The distinction between those on the left and right (then literally) in Revolutionary France has much in common with the broad distinctions between left and right today. Such similarities are not confined to beliefs about more or less redistribution, but are reflected in broader views about the organisation of society, the role of the church, the importance of tradition, etc. (Graham et al. (2009)). In such an unequal, class based, society these views may be expected to represent the common interests of each class.[4] Yet, in modern societies we observe similar patterns of ideological preference that

[2]Presumably, such preferences aren't entirely pre-determined and so we should observe some variation.

[3]In common with this paper they use the WVS to obtain estimates for a cross-section of countries on individual beliefs. These results are complemented by country-level panel data estimates of the relationship between corruption and the ideology of different branches of government. They present a model in which an externality of corruption is reduced incentives for entrepreneurs. This leads individuals to become more left-wing, that is to advocate greater redistribution, and hence impedes the flow of capitalism to poor countries.

[4]For evidence on French income inequality at the time of the revolution see Morrisson (2000) and Bourguignon and Morrisson (2002).

can not be as easily rationalised on the basis of direct economic interest.

This result is perhaps less surprising if considered in the context of the emergent literature emphasising the role of physiological and genetic traits as determinants of political behaviour. Oxley et al. (2008) argue that political views are associated with sensitivity to perceived threats, those with lower physical sensitivity were more likely to favour 'liberal' policies, and vice-versa. The central conclusion of the literature is arguably that those on the right are more anxious (Blair et al. (2003)), more sensitive to loud noises and flashing lights and pain (Oxley et al. (2008)). They are also found to be less curious about the world and new experiences (Carney et al. (2008), Settle et al. (2010)); less altruistic (Zettler and Hilbig (2010)); more responsive to threats (Vigil (2010)); less empathetic (Dodd et al. (2011)); and even less intelligent (Hodson and Busseri (2012)). This litany of negative correlates of being right-wing is perhaps troubling. Despite all of this, right-wingers are somehow happier (Napier and Jost (2008), Vigil (2010)). Amodio et al. (2007) provides evidence for a neurological basis for ideological differences. More recently other neuroscientists, Kanai et al. (2011), have related this basis to differences in brain anatomy. In particular, to differences in the size of the anterior cingulate cortex and right amygdala. The former is associated with being more left-wing and is the area of the brain responsible for threat monitoring, the latter with being more right-wing and is the area of the brain associated with emotional processing. Petersen et al. (2013) finds evidence that relationship between income and preferences for redistribution in men depends on physical strength, with the effect of income more pronounced in stronger men. Fowler et al. (2008) and Alford et al. (2005) use twin studies to measure the effects of genetic variation as a determinant of the tendency to vote and political ideology respectively. They find that greater genetic similarity leads to more similar political behaviour.[5] However, as Alford et al. (2005) note, this is not an argument for genetic determinism.

Thus, one interpretation is that the common use in economics of a left-right scale has solid psychological and physiological underpinnings. Yet, this conclusion would seem to be at odds with the main thrust of research in political science. The next section presents a (necessarily) partial review of this literature and seeks to shed some light on why the different approaches reach opposed conclusions.

2.2 Political Science

There is a significant body of work in Political Science that is relevant for this paper. Most relevant is the prominent literature on the measurement of political preferences and beliefs. This is a vast and complex literature and any proper account would fill several volumes. Here, attention is necessarily restricted to a small sample of recent contributions. Ordeshook (1993) argued that:

> "The idea of spatial preferences – of representing the set of feasible alternative as a subset of an m-dimensional Euclidean space, of labeling the dimensions "issues," of assuming that people (legislators or voters) have an ideal preference on each

[5] Both studies compare the difference in the variance of turnout rate/political attitudes between monozygotic twins (who share 100% of their DNA) and dizygotic twins (who share on average 50%).

issue, and of supposing that each person's preference (utility) decreases as we move away from his or her wi-dimensional ideal policy is now commonplace and broadly accepted as a legitimate basis for modeling electorates and parliaments."

The increasing focus of scholars on measuring evaluating the substantive interpretation of these dimensions can only vitiate a claim that the spatial model is by now even further entrenched in our thinking. Concomitant with this, is a body of work evaluating and discussing each of the constituent assumptions of the model. An important initial question is the extent to which the form of preferences must be restricted if they are to be satisfactorily represented by a spatial model. Put differently, are there important types of preferences which cannot be represented spatially? Austen-Smith and Banks (2000, 2005) discuss these issues in detail, and show that few restrictions are necessary. Ordeshook (1993) argues persuasively that even though much of the assemblage of the spatial model, such as Euclidean preferences, aren't necessary for some of its key results that they are important in helping the theory marry with intuition. Benoit and Laver (2012) make a similar point, that relative comparison between individuals, policies, and parties is so entrenched in political discourse that it is hard to do without it.

Yet, whilst the spatial model need not mean the Euclidean metric. There is more at stake here then may first appear. A key result is that of Plott (1967), who showed that generically majoritarian voting equilibria are unstable, Kramer (1973), and Rubinstein (1979) showed that the core, the set of Pareto optimal Social Choice Function is empty if preferences are continuous. That is, there is almost never the possibility that voting will lead to optimal outcomes. On the other hand, the Manhattan or taxi-cab norm is an example of a non-continuous norm for which stable equilibria are known to exist. Humphreys and Laver (2009) show that if preferences are described by the city-block or l_1 norm then majoritarian equilibria exist. Eguia (2009) shows that if voters are risk-neutral then their utility functions are decreasing in the city-block distance from their ideal point. Risk neutrality, however, itself implies what may be a less palatable assumption. Risk neutrality, and the l_1 norm imply that voters have zero trade-off between different issue dimensions. Whilst, this provides for majoritarian equilibria as the dimensions can be treated independently, it rules out any possibility that an individual may be increasingly (or decreasingly) upset by every further move in policy away from that which they prefer prefer, or that they can trade-off, say a candidates views on the EU with those on Environmental issues.

The next issue, and most relevant to this paper, concerns the measurement of the number of dimensions. Benoit and Laver (2012) emphasize that this can be done either inductively or deductively. The deductive approach is a data-driven process in which some attempt is made to find the number of dimensions which can best rationalize the data. As Benoit and Laver (2012) note the dimensions and content are inherently unobservable, and that any quantification is thus only ever an approximation. However, there a variety of sophisticated methods have been used to generate estimates useful for a variety of purposes. These include, Expert judgement (cf. Benoit and Laver (2006)); the ideological content of manifestos (cf. Budge (2001); Klingemann et al. (2006); speeches Schwarz et al. (2013); votes (cf. Poole and Rosenthal (1997, 2006) for the US, or Hix (2001) for the EU). Other work has focused more explicitly on the number of dimensions of the issue space Lijphart (1984, 1999), Taagepera and Grofman (1985) and more recently Stoll

(2011). Other work, has considered the extent to which political beliefs are stable Epstein et al. (1998), Krosnick (1991). Finally, an alternative view is that political preferences maybe better understood as partisan attachment Goren (2005), Gerber et al. (2010). This literature, of which the cited papers are necessarily only a small sample, routinely reaches a conclusion that there is more than one policy dimension. Moreover, they often find that this number is changing, and that the composition of these dimensions is also changing. Benoit and Laver (2006) provide a summary of this view when they write:

> The pessimistic conclusion, then, is that we may well be treading on thin ice methodologically when comparing left-right policy positions across space or time. Because the substantive meaning of the left-right dimension is so context-dependent, it may be impossible for any single scale to measure this dimension in a manner than can be used for reliable or meaningful cross-national comparison.

They reach this conclusion by analysing expert assessments of parties' positions on different issues (and the importance ascribed to them) and expert judgements of the relative left-right position of each party. Others, see Stoll (2011) for a recent example, have often reached similar conclusions with alternative methods and data. Yet, these conclusions are at odds with those of Poole and Rosenthal (1997, 2006). They estimate the ideological positions of all US congressmen up to the 109th congress. They demonstrate that these positions could be compared over any of the "stable two party periods"). Importantly, they find that 81 percent of the 13 million roll call votes made in the congress since 1789 can be explained with one liberal - conservative ideological dimension.[6] Hix et al. (2006) employs a similar algorithm and concludes that '[...]the main dimension of politics in the European Union's only elected institution is the classic left-right dimension found in domestic politics.' This result also rules out of a lack of consistency over time. How is it then that they reach such a different conclusion? One interpretation is that this seeming contradiction reflects a difference of focus. Benoit and Laver (2006) are concerned with national parties' (estimated) substantive policies (implemented or not) on a range of topics. Jost et al. (2003, 2006), Jost (2009), Jost et al. (2009) and Poole and Rosenthal (1997, 2006), Hix et al. (2006) are concerned with individuals. Perhaps, as argued by Ladley and Rockey (2014) multiple dimensions are important for the formation and functioning of parties but not individuals. The policy platform of parties is both a strategic choice and the result of intra-party compromise (e.g. Roemer (1999)), that is easier with more dimensions. But, this is contrary to the suggestion of Jost (2009) who argues that the policy space of unsophisticated voters is more likely to be irreducible to a single dimension. Thus, there is something of an impasse as to how many dimensions there are at any time or context. Following, Benoit and Laver (2012) this may be a natural consequence of attempts to measure an unobservable and imperfectly definable construct. In particular, it might reflect methodological differences or differences in intent. Like this paper work such as Poole and Rosenthal (1997, 2006) and Hix (2001) in some sense is searching for the minimal number of necessary preference dimensions whereas others seek to develop a more complete, albeit parsimonious, representation. Further, along this spectrum and not discussed here are studies of specific elections or parties which further

[6] 85 percent of variation can be explained using two dimensions, with no meaningful improvement with three or more.

prioritize understanding the detail, subtly, and nuance of a particular political context.

This paper approaches these issues from a different perspective and a different motivation. It does not have modelling the distribution of preferences as its focus. Rather, it asks whether there is a sufficiently well-defined and widely present left-right dimension that to support the assumptions made by much of the Political Economy literature. As noted above, a common assumption is that preferences can be represented with a single dimension. Moreover, due to the infrequency of elections the data considered are often for many countries or states and regularly for the majority of the post-1945 period. The next two sections provide evidence for the existence of a left-right dimension, but emphasise the importance of controlling for demographic differences.

3 Methodology and Results

If the left-right scale is meaningful then it should have a consistent relationship with substantive policy questions over time and place. That is, there should be a consistent pattern in between individuals' views and where they place themselves on this scale. The empirical analysis will therefore test for the existence of such a consistent relationship. Thus, in statistical terms, this paper is concerned with the (polychoric) correlation between variables measuring these views and where they position themselves on a left-right scale. If this correlation is consistently positive across specifications and observations then it suggests that there is indeed a common component of the left-right dimension across place and time. The larger this correlation is, the larger the average share of this common component. Thus, if the correlation is close to 1 then it would suggest that the substantive question at hand predicts, at all times and in all places, every individual's positioning on the left-right scale. If it is zero, then it has no predictive power. Thus, high values of this correlation are a tall order - at the very least we should expect significant noise. The appropriate critical value will vary with context. Yet, clearly values substantially below 1 will sometimes be sufficient evidence that the left-right spectrum is an adequate simplification of the political space. Lesser values, but still significantly greater than zero may be interpreted as evidence that there exists a consistent meaning of the left-right scale. This unconditional correlation in the data at hand, $\rho = 0.11$, can only be suggestive. Further analysis is needed to understand how much this relatively low value is due to noise versus substantive instability in the definition of left and right. Thus, the empirical work begins with a further observation that, if the left-right dimension is meaningful not only should we expect this unconditional correlation to be positive, but also that similar people (as defined by various demographic characteristics) should given similar answers to both questions. Furthermore, if the conditional correlation is positive but there are consistent discrepancies in the answers to the two questions then understanding these discrepancies is not only interesting, but also important in understanding the potential usefulness of the left-right dimension. I provide evidence that there is a consistent and quantitatively important relationship and several interesting, and equally consistent and quantitatively important, discrepancies. The remainder of this section first briefly introduces the data before outlining the results from regression estimates that allows us to analyse the correlation conditional on demographics and or including country and time fixed effects. These estimates also reveal important demographic patterns in ideological preference. It then shows that the results are largely unchanged using a number of alternative questions about substantive views. Section 4 considers various forms of heterogeneity that may be driving the results. It provides evidence that these results are consistent both across and within cohorts. Finally, it considers whether the relationship is conditional on an individual's ideology.

As a first step I analyse the demographic correlates of individuals describing themselves as more or less left-wing –who is leftwing – and also their answers to a substantive policy question – who just thinks they are. To do this I use survey questions from the World Values Survey (WVS) which ask individuals to place themselves on a ten point left-right scale, *leftright*, and their substantive views are captured by *moreineq* which asks them to locate their view as to whether inequality should be increased to provide greater incentives or wages should made more equal on the same ten point scale.The WVS contains questions concerning the beliefs and opinions of over 280,000 individuals in 84 countries on a wide variety of issues along with basic demographic information. Whilst, the sample includes many countries that are autocracies or at best imperfect democracies given the expanding literature on the distribution of power in autocratic regimes (cf. Besley and Kudamatsu (2008) and Robinson et al. (2006)) it is worthwhile understanding if political discourse – however muted – remains organised on the same underlying basis.[7] *leftright* measures individuals' self-perception of their political beliefs, that is how left or right wing they consider themselves to be. It does not define what the different values mean: for example, what 2 means compared to 3 is left as a judgement for the individual. referred to here was used. It asks:

> In political matters, people talk of "the left" and "the right." How would you place your views on this scale, generally speaking?
>
> 1: 'Left'
> 2: '2'
> :
> :
> 9: '9'
> 10: 'Right'

The decision as to which variable represents best the *actual* political preferences of individuals is further complicated by, as discussed above, the possibility that the nature of political debate and the ideological cleavages that motivate it vary substantially between countries. This additional concern necessitates using a variable that both represents as much as possible of the

[7] The data are taken from all five waves of the World Values Survey conducted in 1981-1984, 1989-1991, 1994-1999, 1999-2004, and 2005-2008 respectively. Data for the variables used in this paper were available for 82 countries. These were: Albania Algeria, Andorra, Azerbaijan, Argentina, Australia, Austria, Bangladesh, Armenia, Belgium, Bosnia And Herzegovina, Brazil, Bulgaria, Belarus, Canada, Chile, Taiwan, Colombia, Cyprus, Czech Republic, Dominican Republic, El Salvador, Ethiopia, Estonia, Finland, Georgia, Germany, Ghana, Guatemala, Hong Kong, Iceland, India, Indonesia, Iran, Ireland, Italy, Japan, Jordan, South Korea, Kyrgyzstan, Latvia, Lithuania, Luxembourg, Mali, Mexico, Moldova, Morocco, New Zealand, Nigeria, Norway, Pakistan, Peru, Philippines, Poland, Puerto Rico, Romania, Russian Federation, Rwanda, Slovakia, Viet Nam, Slovenia, South Africa, Zimbabwe, Spain, Sweden, Switzerland, Thailand, Trinidad And Tobago, Turkey, Uganda, Ukraine, Macedonia, Egypt, Tanzania, United States, Burkina Faso, Uruguay, Venezuela, Serbia And Montenegro, Zambia, Northern Ireland, and Serbia.

variation in individuals' ideological position, whilst remaining consistent in its interpretation across countries. [8] It is based upon the following question:

> "Incomes should be made more equal vs We need larger income differences as incentives. How would you place your views on this scale?
> 1: 'Incomes should be made more equal'
> 2: '2'
> :
> :
> 9: '9'
> 10:'We need larger income differences as incentives'

Hence, *moreineq* can be seen to represent what Immervoll et al. (2007) refer to as the ""old debate", that is, the traditional conflict between equality and efficiency. The relative merits of the arguments that increased inequality improves efficiency or that greater equality is the more ethical outcome are ignored here. It suffices to assert that a great deal of current and historical political debate has centred around arguments like this or those that embody similar ideological principles. For example, a debate on how health care or education should be provided is in essence very similar: a conflict between ideas of equality of provision and the mooted greater efficiency of the market. Section 3.1 expands the definition by considering alternative dependent variables designed to measure other differences in political beliefs.

Since, *moreineq* is related to desire for change in income inequality it is comparable across countries. More generally, since the question is focused on what explains the variation in perceived and underlying ideological positions within countries, differences in the average ideological position between countries are not important. In sections 3.1, and 4 I present evidence that the results are robust to a variety of other measures of individuals' substantive views.

The regression specification needs to capture the possibility that *incomeequal* and *leftright* might be jointly determined and that there may be important information in the ordinal structure of the data. Our starting point is therefore a bivariate ordered probit model. Let the latent variables $y^*_{i,1}$ and $y^*_{i,2}$ denote individuals' true self-perceived and actual ideological positions and are measured using the ordered dependent random variables $y_{i,1}$ and $y_{i,2}$ respectively. The disturbance terms $\varepsilon_{i,1}$, and $\varepsilon_{i,2}$ y^2_i are assumed to be jointly normally distributed with correlation parameter ρ and $\mathbf{x_i}$ is the vector of independent variables. Then following Greene and Hensher (2009) the estimation may be stated as a 'seemingly unrelated regressions (SUR) model for the latent regressions':

$$\begin{aligned} y^*_{i,1} &= \beta_1'\mathbf{x}_{i,1} + \varepsilon_{i,1}, \quad y_{i,1} = j \text{ if } \mu_{j-1} < y^*_{i,1} < \mu_j,\ j = 0, ..., J_1 && (1) \\ y^*_{i,2} &= \beta_2'\mathbf{x}_{i,2} + \varepsilon_{i,2}, \quad y_{i,2} = j \text{ if } \delta_{j-1} < y^*_{i,2} < \delta_j,\ j = 0, ..., J_2 && (2) \\ \begin{pmatrix}\varepsilon_{i,1} \\ \varepsilon_{i,2}\end{pmatrix} &\sim N\left[\begin{pmatrix}0 \\ 0\end{pmatrix}, \begin{pmatrix}1 & \rho \\ \rho & 1\end{pmatrix}\right] && (3) \end{aligned}$$

[8] In particular *leftright* is WVS variable *E033* and *moreineq* WVS variable *E035*.

A key advantage of explicitly modelling the dependence between y_1^* and y_2^* is that the estimate of ρ apportions any simultaneity into the error terms, providing consistent estimates. As a first step in addressing heterogeneity, country-wave dummies are included, and the error terms are similarly clustered.

Table 1 contains the results of this bivariate ordered probit estimation.[9] Column 1, includes no demographic controls only country and wave fixed effects. Reassuringly, the estimated correlation coefficient is not qualitatively different. Column 2 reports results including demographic variables, but not country and wave fixed effects. Column 3 includes reports the same specification but including fixed effects. A first-glance at these results suggests that men are both more right-wing than women, and perceive themselves as being so. This would seem to complement the results of previous work such as that of Edlund and Pande (2002) and Aidt and Dallal (2008). Edlund and Pande (2002) argue that the decline in US marriage rates (and the increase in divorce rates) has made women less well-off and men better-off. They provide evidence that this decline is associated with the rise of a difference in political allegiance between men and women. This gap was around 15 percentage points by 1996, whilst in general prior to 1980 there was no significant gap. Moreover, they suggest that around three percentage points of the Political Gender Gap, can be attributed to the impact of divorce on the voting intentions of women. Aidt and Dallal (2008) exploit the variation in when Women gained the franchise in Europe to obtain results that "support the hypothesis that countries experienced an increase in social spending after women were given voting rights. In the short-run, the effect is a 0.6-1.2% increase [...] the long run impact beeing three to eight times larger." Cavalcanti and Tavares (2006) argue that as income per capita increases women substitute remunerated employment for household production, and that this leads to the demand for government services. Their headline empirical result, coinciding with the predictions of their theoretical analysis, is that "an increase in female labour force participation of 10% leads to a increase of government spending of about 2.5 percent as a share of GDP". Funk and Gathmann (2008) suggest using data from Swiss Referenda that men and women favour different policies, *ceteris paribus.* In particular, women seem to "*care more about the environment, public health, social welfare and are more sceptical towards nuclear energy or the military. Regarding the fiscal consequences of female policy makers, we find a bigger impact on the composition rather than the size of government.*"There is also evidence for a 'Political Gender Knowledge Gap', for example Frazer and Macdonald (2003) provide evidence that all else equal British women have less 'political knowledge than men'. More generally, they find that the younger and less educated are less knowledgeable about politics based on answers to a set of questions. Mondak and Anderson (2004) argue that the size of the gap is inflated but otherwise suggest that young, uneducated, and female Americans are more likely to be politically uninformed. Interestingly, when the analysis is repeated separately (results available upon request) for men and for women, the estimated coefficients are qualitatively similar for the other independent variables. This suggests differences between men and women are not contingent on other demographic factors. Thus we are unable to provide additional evidence of such a political gender knowledge gap.

[9]As noted by Hoetker (2007), it cannot be assumed that the unobserved variation is the same in the two equations. This precludes comparison of the size of coefficients between the two equations, instead the discussion is restricted to the significance and sign of the coefficients.

Age seems to be an important factor for individuals perception of their ideology although has no significant effect on views on redistribution versus efficiency. In this specification the effect of generation is confounded with age, but this result potentially compliments with that of Sørensen (2013) who analyses how age effects relative preferences for different public goods. He finds that consistent with a life-cycle model individuals increasingly prefer health care spending to education provision as they get older. Thus we should not necessarily expect to find here a preference over the overall degree of redistribution.

How does education affect ideology? It would seem that the better educated, if anything, are less accurate in how they perceive their ideology. Higher levels of education are associated with being less likely to believe oneself to be right-wing, whilst simultaneously associated with being in favour of increased inequality. This result contrasts with those for income: higher levels of income are associated with both believing oneself to be more right-wing as well as considering more inequality to be necessary. The effects of income on ideological self-position are interesting. Whilst, there remains too much unobserved heterogeneity to make a causal claim, one interpretation is that, *à la* Alesina and Angeletos (2005), on average the rich believe themselves to be richer due to their effort and thus policy should incentivise effort. That there are opposing effects of education on ideological self-position and substantive views might be indicative of peer-group effects, or a changes in individual's perception of the range of possible views. Some evidence for the latter is found in Section 4.1 where evidence is found that the better educated are more likely to judge themselves to be less extreme for any given ideological position. Similarly, evidence is found that greater income is associated with greater polarization.

It is not obvious why living with one's parents is associated with an increase in the probability that an individual considers themselves to be right wing. But, it is notable that the positive coefficient remains significant in all of the specifications considered here.

I now consider the variables that describe individuals' occupations. The coefficient on being self-employed is large, and positive, for both *incomeequal* and *selfpospolit*. Although it is only significant in Table 1 for *incomeequal*. Like the results for gender, education, and income this result seems stable. Having full-time or part-time employment is also associated with a preference for more inequality but the estimates are statistically significant and the coefficients for *selfpospolit* are negative. This difference has at least two possible interpretations. Either, those who are exposed to what is often considered as the greater risks and rewards of self-employment are come to understand better the importance of income (inequality) as an incentive. Alternatively, those who believe in the power of individual effort are more likely to become self-employed.

Those with non-labour market occupations, that is; students, what the WVS refers to as 'housewives', and the retired, seem to be more likely to be right-wing. The retired seem to perceive themselves as being slightly more left-wing than average, but whilst the coefficient on their estimated actual ideological position is also negative it is small and insignificant.[10]

[10] The results are robust to including measures of respondents' marital status. The preferred specification ignores these as such decisions as to to whether to get married or divorced could partly reflect cultural conservatism and hence lead to biased estimates. However, the results reported below are robust to the inclusion of these variables.

Comparing columns 2 and 3 suggests that the results are broadly speaking, robust to the exclusion of fixed effects. This may be interpreted as some evidence that the results reflect long-term, cross-national, patterns.

3.1 Defining Left and Right Wing

If the single left-right measure is at all adequate then it should also relate consistently to a variety of other plausible substantive questions. Whilst I argue above that beliefs about inequality – as measured by *incomeequal* – represent a key aspect of the left-wing continuum it is worthwhile to consider other measures that capture beliefs about how society should be organised. To this end Table 2 reports using the same specification as in Table 1 but employing alternative measures of individuals' beliefs. These are; *basicatt*- a trinary question recording whether the respondent believes society needs radical change, reform, or defending; *stateown* -whether state ownership of industry should be increased; *govtresp* - whether it is an individuals or their governments responsibility to ensure they are provided for; hardsuccess - captures whether hard work or luck brings success; and *wealthaccum* measures whether individuals believe that 'people can only get rich at the expense of others' or if 'wealth can grow so there's enough for everyone'? An alternative, related concern, it that the other questions maybe too abstract. To allay such concerns, the analysis was also repeated for the binary variable *secfair*. This variable records the answer to whether the respondent feels the following scenario is fair or not fair.

> Imagine two secretaries, of the same age, doing practically the same job. One finds out that the other earns considerably more than she does. The better paid secretary, however, is quicker, more efficient and more reliable at her job. In your opinion, is it fair or not fair that one secretary is paid more than the other?

This question is felt to be meaningfully different to *moreineq*, in particular because it describes a scenario rather than a conceptual choice, whilst also measuring the respondents' views on the same equality-incentives dimension. Inspection of Table 2 suggests that the results in Table 1 are not artefacts of the choice of ideology measure. More specifically, there is a good deal of consistency across all six measures. Older respondents continue to see themeselves as more right-wing whilst their views on the substantive questions mostly don't change. The pattern for gender also remains largely unchanged. Finally, and perhaps of greatest interest is that the effects of education and income found previously seem to be repeated here. The exceptions to these patterns are largely for the *basicattsoc*. Analysis suggests that these are driven by the statement 'society must be valiantly defended' as the rightmost of the three answers. Repeating the analysis using instead binary variables defined as choosing or not this option or 'society must be radically changed' suggests the defence of society alternative picks up something different to all of the other questions. In an effort to capture the common underlying variable, *f1*, the first principal component of the measures of beliefs is computed and used as a regressand. The results of this analysis are again consistent with those obtained for the individual dependent variables.

4 Taking heterogeneity seriously

The purpose of this paper is to identify consistent demographic patterns in preferences and preference mis-perception across time and place. In this context, one may reasonably ask how consistent the estimated relationships are across countries or generations. In the remainder of this paper, using a variety of alternative estimation strategies, I consider three possible forms of heterogeneity and argue that the results support the use of a single left-right dimension. To the extent that aggregate experiences are one source of ideological preference as suggested by Alesina and Giuliano (2009) and Di Tella and MacCulloch (2009) then one might expect ideology to vary by generation within countries for reasons separate to age. One way to address is this is to control for such generation or cohort fixed-effects, and focus on variation within cohorts. As, membership needs to be fixed over time, I define cohorts by decade of birth $D \in \{pre1930, `40s, \ldots, `90s\}$ and gender $G \in \{Male, Female\}$, and index these cohorts by $j \in J = G \times D$. The key forms of heterogeneity considered are:

1. That the 'left-right' spectrum may mean different things across time and place.

$$corr(y_1^*, y_2^*) = \rho_j \neq \rho_{-j}$$

2. Variation across cohorts in how demographic characteristics affect ideology and its perception.[11]

$$E[y_{1,2}^*|X_j'\beta_j] \neq E[y_{1,2}^*|X_{-j}'\beta_{-j}]$$

3. That how demographics preducts self-reported ideological position varies with actual ideological position.

$$E[y_1^*|X] \neq E[y_1^*|X, y_2^*]$$

4.0.1 How consistent is the meaning of 'left' and 'right-wing'

If what is meant by 'left-wing' and 'right-wing' means is not consistent then there will be little consistency in the correlation in the residuals from the regression residuals for the *incomeequal* and *selfpospolit* equations. That it is positive and significantly different from zero in all of the regressions reported above does not imply that there isn't a significant proportion of countries or periods for which it isn't true. Of particular concern, following Pop-Eleches and Tucker (2010) and Alesina and Fuchs-Schndeln (2007) is that individuals living in post-communist countries may interpret the left and the right differently. They find this discrepancy, depends on the extent to which a person lived under communism during their formative years. To investigate this hypothesis, the analysis in Column 3 of Table 1 was repeated separately for each cohort, the solid line in Figure 5 is a kernel density plot of the resulting distribution of $\hat{\rho}$ The dotted line is

[11] One further possibility is a non-linear relationship between the dependent and the independent variables: $y_{1,2}^* \sim f(X, \beta) \neq X'\beta$. The categorical nature of most of the left hand side variables means this is less of a concern here.

the distribution excluding ex-communist countries.[12] It suggests that while the strength of the correlation is variable, it is almost always positive as is hoped. À la Pesaran (2007) I calculate the average of the Z-statistics for $\hat{\rho}$ over the cohort specific regressions.[13] The average z-statistic is 2.12, excluding ex-communist countries, the number is only slightly higher at 2.26. Excluding all cohorts for which $\hat{\rho} \leq 0$ does not change the results. This analysis suggests that whilst there is inevitably variation in what 'left' or 'right' wing mean, that that for the vast majority (88%) of observations there is a consistent correspondence.

4.0.2 Do the results reflect variation across cohorts

The WVS is, unfortunately, not a panel but rather a repeated cross-section of nationally representative polls. Deaton (1985) argued that whilst it is impossible to track an individual with such data, that if the unit of analysis was instead a 'cohort' —a group of individuals exogenously determined with fixed membership— and representative survey data are available then consistent estimates for the linear fixed-effects model can be straightforwardly obtained. The intuition is that as each poll is representative (and there are no problems of attrition) then the evolution of the population as whole is described by the series of cross-sections. By suitably defining cohorts as a group of individuals that can be identified in each survey, then the each cohort's fixed effect is the sum of the individual fixed effects providing for consistent estimates. Writing the standard within-estimator as:

$$y_{it} = x_{it}'\beta + \alpha_i + u_{it}, \qquad t = 1, \ldots, T;\; i = 1, \ldots, N \tag{4}$$

Here, and in what follows I employ the notation of Verbeek (2008). Deaton (1985) shows that taking averages by cohort provides the following population model, with asterisks denoting population quantities:

$$\overline{y}_{ct}^{*} = \overline{x}_{it}^{*\prime}\beta + \alpha_c^{*} + u_{ct}^{*}, \qquad t = 1, \ldots, T;\; c = 1, \ldots, C \tag{5}$$

As this is a population identity it will in practice be subject to sampling error. Deaton (1985) therefore proposes the resulting errors-in-variables estimator, which was subsequently modified by Verbeek and Nijman (1993) to be consistent if the number of time periods is small (as is the

[12] Following Pop-Eleches and Tucker (2010) these countries are: Armenia, Azerbaijan, Belarus, Bosnia and Herzegovina, Croatia, Estonia, Georgia, Kazakhstan, Kyrgyzstan, Latvia, Lithuania, Macedonia, Moldova, Russian Federation, Serbia and Montenegro, Slovenia, Tajikistan, Turkmenistan, Ukraine, Uzbekistan.

[13] Specifically, I perform a Fisher Z-transformation (inverse hyperbolic tangent transformation) of the estimated values of $\hat{\rho}$ so that the transformed sampling distribution is approximately normally distributed.

case here).[14] Thus, assuming the measurement error is distributed as follows:

$$\begin{pmatrix} \overline{y}_{ct} - y^*_{ct} \\ \overline{x}_{ct} - x^*_{ct} \end{pmatrix} \sim i.i.d. \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_{00} & \sigma' \\ \sigma & \Sigma \end{pmatrix} \right]$$

Then, the estimator employed is:

$$\hat{\beta} = (M_{xx} - \tau\hat{\Sigma})^{-1}(m_{xy} - \tau\hat{\sigma}) \tag{6}$$

Where:

$$M_{xx} = \frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}(\overline{x}_{ct} - \overline{x}_c)(\overline{x}_{ct} - \overline{x}_c)' \tag{7}$$

$$m_{xy} = \frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}(\overline{x}_{ct} - \overline{x}_c)(\overline{y}_{ct} - \overline{y}_c) \tag{8}$$

This estimator, ignoring measurement error also has a straight-forward Instrumental Variables representation proposed by Moffitt (1993) in which x_{it} is instrumented for using the interaction of cohort and time dummies.[15] This interpretation makes the requirements for identification clear; the cohorts must be defined such that the resultant instruments are both relevant and exogenous. In practice this means that the mean of each independent variable should vary over time within each cohort, and defining these cohorts based on variables that don't vary over the period. Taking these concerns into account, cohorts are defined on the basis of decade of birth and gender.

Table 3 contains estimates corresponding to the specification in Table 1 although now *age*100 and *male* are subsumed into the cohort fixed effect. Columns 1 & 2 suggest that conditioning on cohorts, those who are better educated still judge themselves as more leftwing than their policy preferences suggest. Similarly, the more affluent continue to judge themselves more right wing than implied by their policy preferences. Indeed the results are broadly speaking consistent with those obtained previously. This is an important result as it means that the results are not being driven by generational and gender differences or national trends. Most importantly,

[14]Of course, as the sample size approaches the limit then the measurement error, given a suitable sampling scheme, should converge to zero and the application of OLS to Equation 5 will provide consistent estimates. This is an assumption that has often been made in practice. But, the asymptotic analysis presented by Verbeek (2008) emphasizes that consistency of the OLS estimator requires the the number of observations per cohort to be large, while the finite - sample analysis of Devereux (2007) suggests that even when the number of observations per cohort is several thousand substantial biases remain if we do not allow for measurement error. Given the minimum cohort size here is restricted to be at least 200 with an average of 452 ignoring measurement error would still seem ambitious. For the bivariate ordered probit results in Section 4.0.1 I focus on cohorts with at least 200 members, although the results aren't sensitive to the specific number chosen, and the focus there is on the sign not the magnitude of the coefficients.

[15]In the notation above, and again following Verbeek (2008) this estimator is simply

$$\hat{\beta}_{IV} = (\frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}(\overline{x}_{ct} - \overline{x}_c)x'_{it})^{-1}\frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}(\overline{x}_{ct} - \overline{x}_c)y_{it})$$

. Equally, Verbeek (2008) notes the standard within-estimator applied to the cohort means also corresponds to the case where $\tau = 0$.

is that it allows for stronger inferences. We can credibly claim that the effects of education, income, or employment on individuals' political views and their perception of those views are consistent across the wide range of countries studied. Thus, these coefficients can be understood to measure the average effect of being, for example, better educated, rather than describing the views of more educated, for example, Germans or Argentinians.

Columns 3 and 4 repeat this analysis but in the absence of well developed asymptotics for clustered standard errors in pseudo-panels report bootstrapped standard errors. The overall interpretation is largely unchanged, although as expected the errors are now larger - most notably the coefficient on education in the *selfpospolit* equation is no longer significant.[16]

4.1 Does the relationship between ideological self-placement and demographics vary with ideology?

Intuitively, there are several reasons why the relationship between demographics and (mis-)perception may itself vary with ideology. For example, as discussed above the better educated tend to believe they are more leftwing whilst favouring more right-wing policies. But, is this effect uniform across the population? Or, instead does education magnify or ameliorate this effect? To investigate this possibility Equation 5 was augmented with pairwise interaction terms to give the following population identity:

$$\overline{selfpospolit}^*_{ct} = \overline{x}^{*\prime}_{it}\beta + (\overline{x}^*_{it} \times \overline{incomeequal}^*_{ct})'\gamma + \alpha^*_c + u^*_{ct}, \qquad t = 1, \ldots, T;\ c = 1, \ldots, C \quad (9)$$

The results of this analysis are reported in Table 4. Now the conditional marginal effect of x on $selfpospolit$ conditional on *incomequal* is simply: $E(\beta|incomeequal) = \beta + \gamma \times incomeequal$. Thus, $E[y^*_1|X] = E[y^*_1|X, y^*_2] \leftrightarrow H_0 : \gamma = 0$ However, *student* and *senior(non)manual*, are the only variables for which both β and γ are insignificant. There is also no evidence that the effect of being *retired* varies with ideology, whilst being childless (*nokid*) has no significant effect at average levels of ideology but there is evidence it does at either end of the distribution. But, for the majority of regressors the average effect, given by β, is now positive whilst the interaction term, γ, is negative. This implies that in most cases the distribution of $selfpospolit$ is moderated for extreme values of *incomeequal* compared to the counterfactual of the omitted category.

Through inspection of β and γ alone it is often difficult to quickly work out the conditional effect at any given level interacted variable, here *incomeequal*, and even more so its error. Therefore, as in Braumoeller (2004), I plot the conditional marginal effect, and its error, over the domain of *incomeequal*. A further obstacle to inference is that it is difficult to compare the relative homogeneity of different variables with different standard deviations. As such, I report instead $\frac{\sigma_x}{\sigma_y}E[\beta|incomeequal]$. These graphs are reported in Figure 5. If a the effect of a

[16] One disadvantage of this approach is that the ordinal nature of the dependent variables is not modelled. Results, not reported, using an ordered - logit estimator with cohort-fixed effects and clustered standard errors support similar conclusions.

particular variable does not vary with *incomeequal* then we should expect the estimated curve to be approximately flat. Conversely, the more heterogeneity there is in the effect of a given variable the steeper the gradient of the curve. Figure 5 reveals that that for some variables there is evidence of heterogeneity. The effect of *parttime* varies most with *incomeequal*, but the negative slope implies that part-time workers are less likely to judge themselves to hold extreme ideological positions. That is, a positive effect at low values of *incomeequal* and a negative effect at high values reduces the dispersion in predicted values of *selfpospolit*. A similar pattern is true, although to a lesser extent, for for the *selfemployed* and *fulltime* workers, and those who live with their parents. Note however, that this shrinkage does not imply shrinkage towards the average *selfpospolit*, rather the average effect is given by the intersection of the curve and the y-axis. Thus, for example, for the *parttime* the average effect is positive, implying that part-time workers (judge themselves to be) more rightwing on average. The displayed confidence interval is also important for inference. It suggests that whilst there is probably no significant effect of associated with *retired* that the similarly homogenous effects of *(senion)nonmanual* clearly are negative and significant at all points. But, for *nokids*, *retired*, *housewife*, and *skilledmanual* when the estimated curve is relatively flat and close to the y-axis a reasonable interpretation is of no quantitatively important effect.

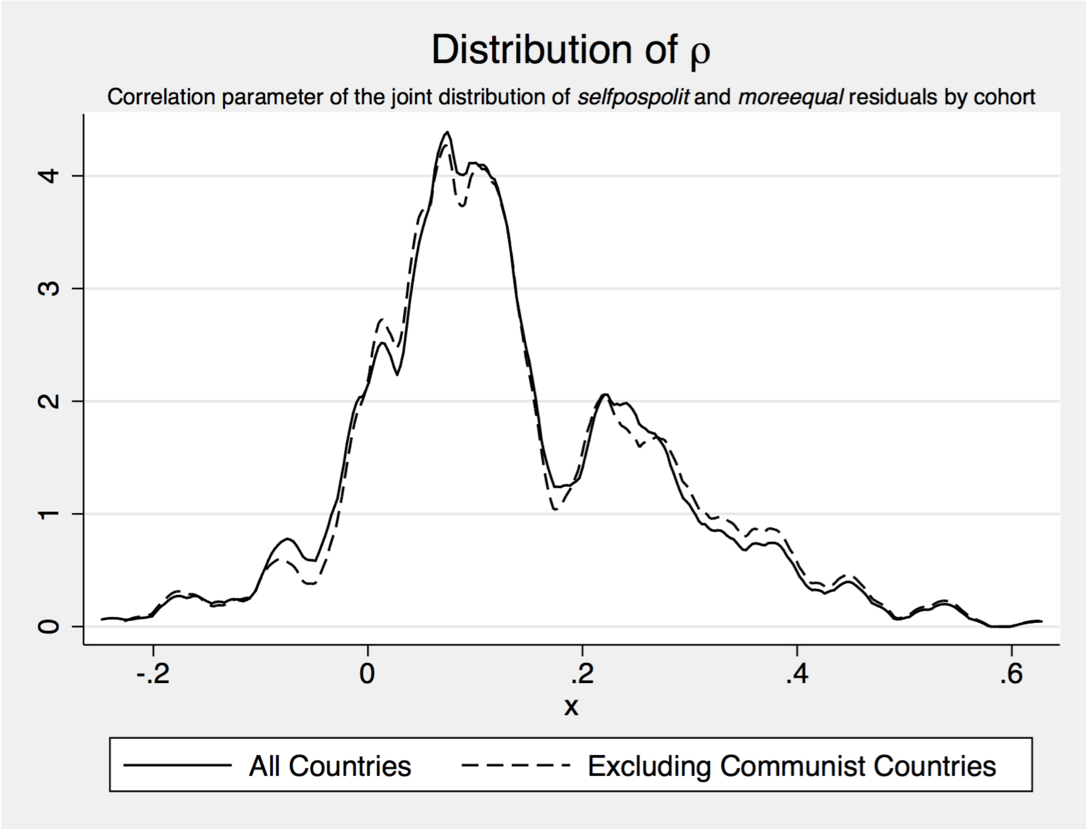
The coefficients of *highesteduc* and *scaleofy* continue to be amongst the most interesting results and they merit especial attention here. Figure 1(a) displays the conditional marginal effect of additional education for different levels of *incomeequal*, while Figure 1(b) plots the effect of moving up an income decile by answer to *incomeequal*. The graph for *highesteduc* is similar to *parttime* in interpretation and supports a conclusion that greater education is associated with believing oneself to be less extreme, other things equal. Income has the opposite effect, and is associated with judging oneself to be more extreme. This effect is different, but similar in spirit, to that reported by Gelman et al. (2008) who writes "[. . .] *being in a red or blue state matters more for rich than for poor voters*". This result is also interesting in the context of the literature on political polarization. Ensley (2012) documents the increasing polarization of the US electorate. and McCarty et al. (2003) suggests that this is related to income growth. Our results suggest that this effect is consistent across countries, and suggests that in general at higher levels of income there is a greater distinction between policy preferences and self-reported preferences. This distinction is suggestive of one potential mechanism for the polarization Gelman finds associated with richer voters – an increasing willingness or need to self-describe oneself in more polarised terms. For those that fret about the implications of continued growth incomes for the future practice of politics, our results suggest that growth in education may be a key moderating mechanism. In terms of the results of the rest of the paper, taken as a whole Figure 5 does not give any cause to suggest that that the results in Tables 1 and 2 should be seen as being contingent on certain values of incomeequal.

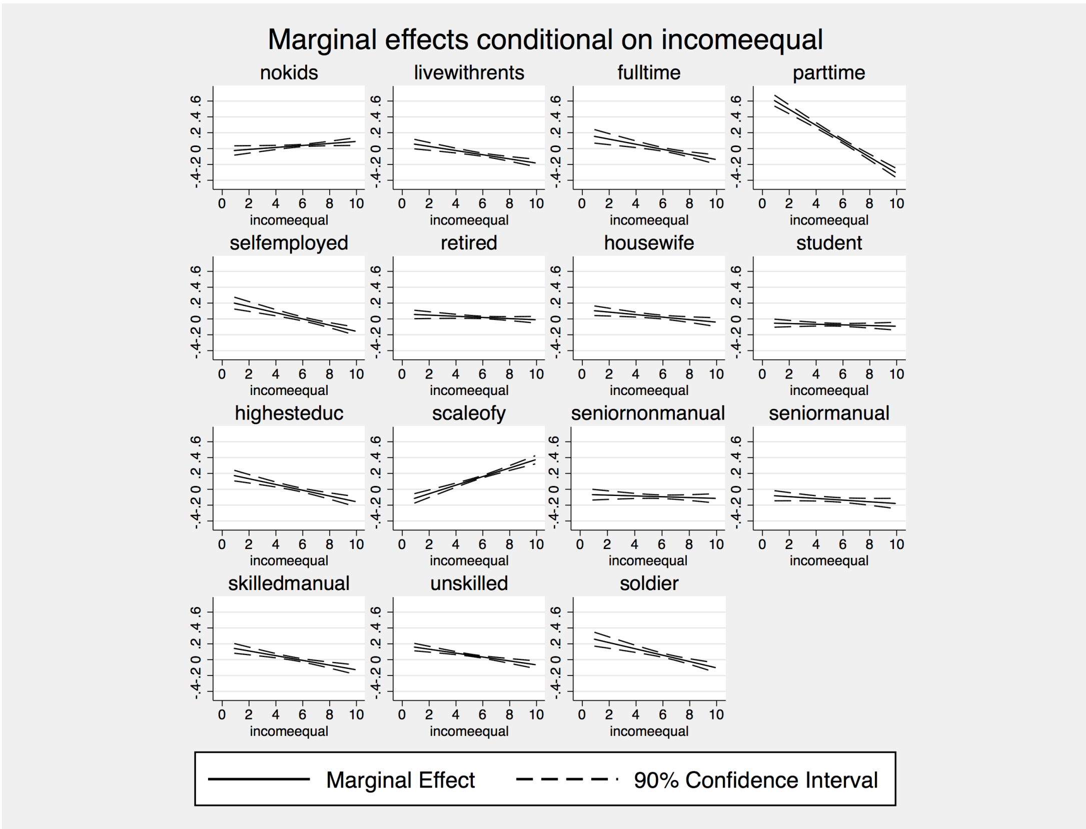
Thus, whilst there is evidence of polarising effects of income and 'anti-polarizing' effects of education, the three forms of heterogeneity considered can be largely rejected. Further, there is little evidence that the results are contingent on the measure of ideology used.

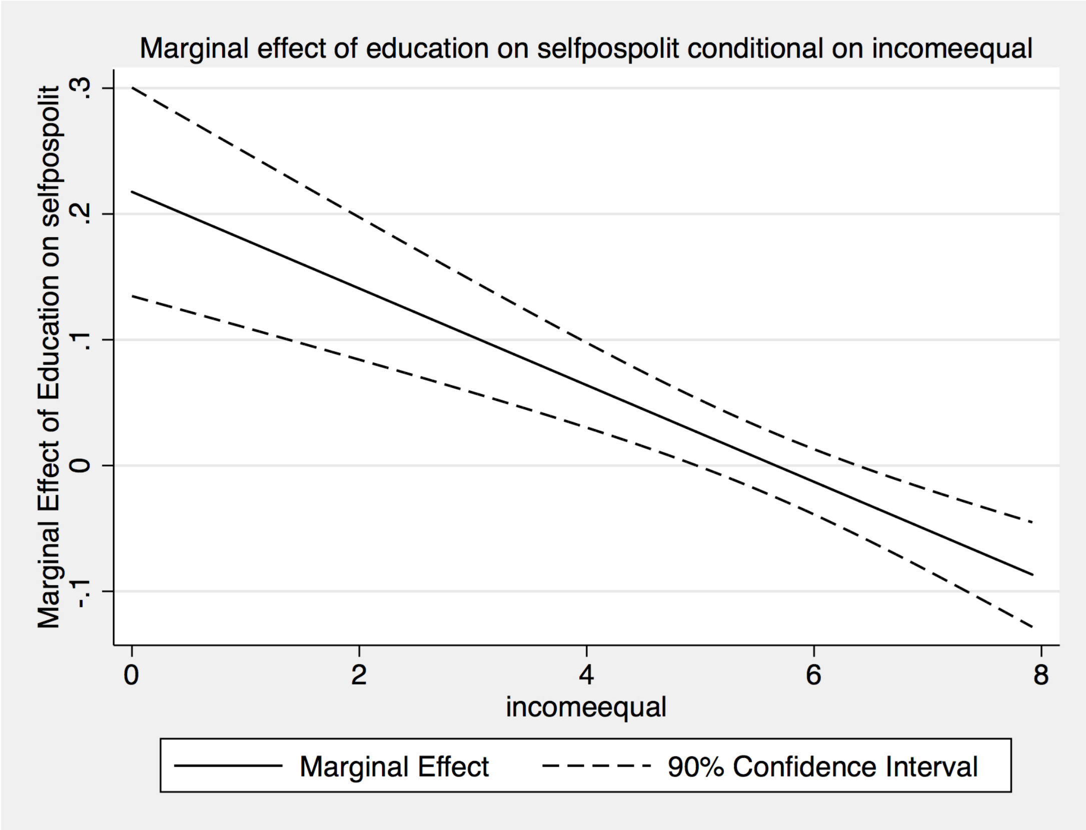
5 Conclusion

This paper makes the case that is meaningful to talk of the left and the right. It argues that whilst different demographics perceive left and right differently, that both these differences and the meaning left and right are consistent across time and place. It provides evidence that this is true for both abstract alternatives and concrete choices, questions of redistribution and broader conceptions of social justice. Finally, it suggests that whilst income seems to lead the rich tgo think of themselves as more polarized than they are, in general differences across demographics are predominantly in mean only and thus easily accommodated in regression analyses.

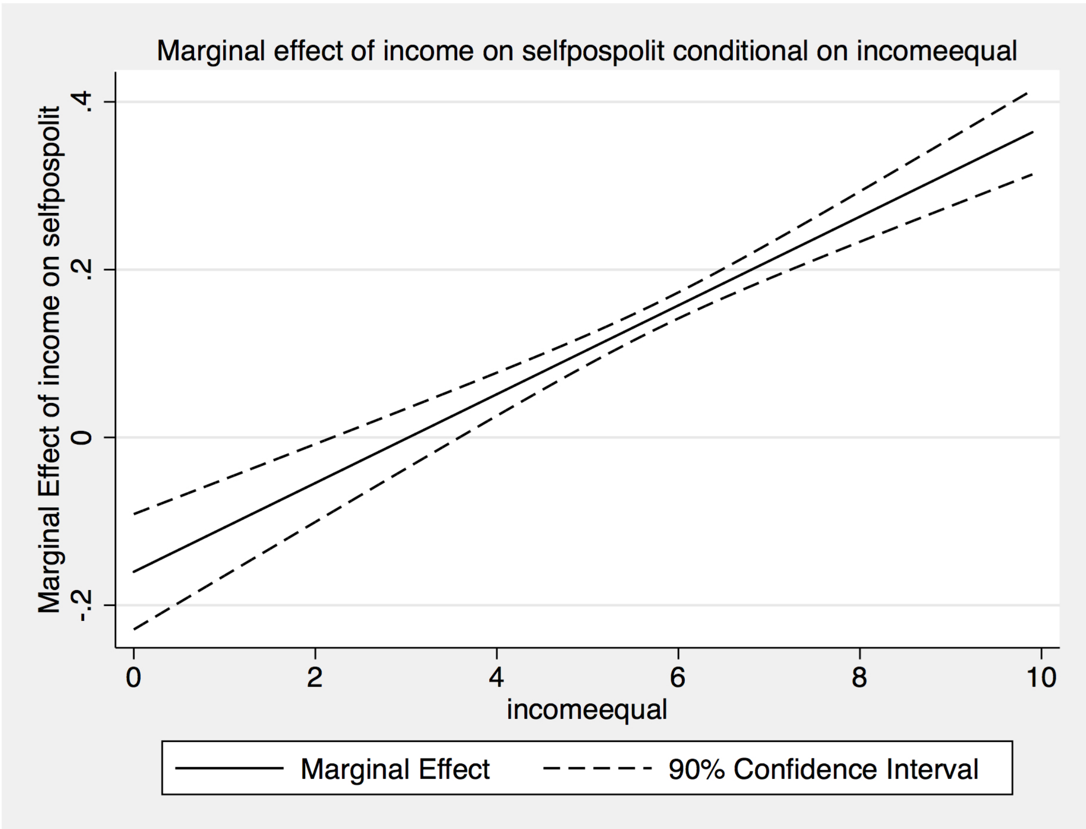
None of this is to argue that the results preclude many other important factors at any given time, or the possibility of further and important issue dimensions. Again, at each election other dimensions and factors will also be important. Other studies have had more to say about these other factors, but the evidence for the broad consistency in the interpretation of left and right wing presented here will hopefully make inference easier in such studies.


Distribution of ρ
Correlation parameter of the joint distribution of *selfpospolit* and *moreequal* residuals by cohort
0
1
2
3
4
-.2
0
.2
.4
.6
x
All Countries
Excluding Communist Countries


Marginal effects conditional on incomeequal
nokids
livewithrents
fulltime
parttime
selfemployed
retired
housewife
student
highesteduc
scaleofy
seniornonmanual
seniormanual
skilledmanual
unskilled
soldier
incomeequal
Marginal Effect
90% Confidence Interval



(a) The effect of education conditional on incomeequal



(b) The effect of income conditional on incomeequal

Table 1: Bivariate Ordered Probit Estimates

	1	2		3	
		incomeequal	selfpospolit	incomeequal	selfpospolit
male	–	0.026	0.041^{**}	0.032^{***}	0.041^{***}
		(0.021)	(0.019)	(0.010)	(0.012)
age100	–	−0.150	0.114	−0.032	0.180^{**}
		(0.110)	(0.087)	(0.055)	(0.072)
nokids	–	−0.015	−0.039	−0.000	−0.015
		(0.028)	(0.027)	(0.016)	(0.013)
livewithrents	–	0.032	0.083^{***}	0.002	0.046^{***}
		(0.030)	(0.021)	(0.010)	(0.011)
fulltime	–	−0.072	-0.105^{*}	0.029	−0.012
		(0.063)	(0.058)	(0.018)	(0.021)
parttime	–	−0.099	−0.057	−0.001	−0.018
		(0.074)	(0.046)	(0.031)	(0.021)
selfemployed	–	0.038	0.035	0.053^{**}	0.042^{*}
		(0.073)	(0.047)	(0.025)	(0.024)
retired	–	-0.110^{*}	-0.099^{**}	0.002	0.002
		(0.065)	(0.045)	(0.021)	(0.028)
housewife	–	−0.024	0.024	0.039	0.069^{**}
		(0.072)	(0.076)	(0.031)	(0.028)
student	–	−0.045	−0.100	0.011	−0.036
		(0.059)	(0.065)	(0.029)	(0.029)
highesteduc	–	0.034^{***}	-0.014^{*}	0.031^{***}	-0.018^{***}
		(0.007)	(0.008)	(0.005)	(0.006)
scaleofy	–	0.022^{***}	0.019^{**}	0.033^{***}	0.018^{***}
		(0.009)	(0.009)	(0.006)	(0.004)
seniornonmanual	–	0.075	0.034	0.066^{***}	0.010
		(0.048)	(0.041)	(0.019)	(0.015)
seniormanual	–	0.034	0.051	0.017	0.007
		(0.062)	(0.053)	(0.023)	(0.025)
skilledmanual	–	−0.008	−0.022	−0.016	-0.037^{**}
		(0.044)	(0.048)	(0.019)	(0.018)
unskilled	–	−0.046	0.084	−0.030	0.034^{**}
		(0.046)	(0.063)	(0.020)	(0.017)
soldier	–	0.111^{*}	0.198^{**}	0.026	0.042
		(0.058)	(0.100)	(0.041)	(0.044)
ρ	0.114^{***}	0.123^{***}		0.116^{***}	
	(0.012)	(0.017)		(0.012)	
Country & Wave Effects	Yes	No		Yes	
N	136046	136046		136046	

Estimates are from a bivariate ordered probit model. Standard Errors clustered by country in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 2: Other measures of political preferences

	basicattsoc		incomeequal		stateown		govtresp	
	selfpospolit	basicattsoc	selfpospolit	secfair	selfpospolit	stateown	selfpospolit	govtresp
male	0.03^{**}	-0.05^{***}	0.04^{***}	0.06^{***}	0.04^{***}	0.13^{***}	0.04^{***}	0.04^{***}
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
age100	0.23^{***}	0.21^{***}	0.18^{**}	0.36^{***}	0.16^{**}	0.05	0.18^{**}	0.01
	(0.09)	(0.08)	(0.08)	(0.07)	(0.07)	(0.07)	(0.07)	(0.05)
nokids	−0.01	0.02	−0.02	0.04^{**}	−0.01	0.02	−0.01	0.02
	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)
livewithrents	0.06^{***}	0.03^{*}	0.05^{***}	−0.01	0.04^{***}	−0.02	0.04^{***}	−0.00
	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)	(0.01)
fulltime	−0.02	0.01	−0.02	0.05	−0.01	0.00	−0.01	0.04^{***}
	(0.02)	(0.02)	(0.02)	(0.03)	(0.02)	(0.03)	(0.02)	(0.01)
parttime	−0.04	0.02	−0.02	−0.00	−0.02	−0.00	−0.02	0.02
	(0.02)	(0.02)	(0.02)	(0.04)	(0.02)	(0.04)	(0.02)	(0.02)
retired	−0.04	0.11^{***}	−0.00	0.03	−0.00	−0.02	0.00	0.02
	(0.03)	(0.02)	(0.03)	(0.04)	(0.03)	(0.04)	(0.03)	(0.02)
housewife	0.05	0.09^{***}	0.07^{**}	0.07^{*}	0.07^{**}	0.05	0.07^{**}	0.06^{***}
	(0.03)	(0.03)	(0.03)	(0.04)	(0.03)	(0.03)	(0.03)	(0.02)
student	−0.04	0.03	−0.03	0.07	−0.03	0.02	−0.04	0.03
	(0.04)	(0.02)	(0.03)	(0.04)	(0.03)	(0.03)	(0.03)	(0.02)
highesteduc	-0.02^{***}	-0.03^{***}	-0.02^{***}	0.04^{***}	-0.02^{***}	0.03^{***}	-0.02^{***}	0.02^{***}
	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.00)	(0.01)	(0.00)
scaleofy	0.01^{***}	−0.00	0.02^{***}	0.03^{***}	0.02^{***}	0.03^{***}	0.02^{***}	0.03^{***}
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.01)	(0.00)	(0.01)
seniornonmanual	−0.01	0.03	0.01	0.08^{***}	0.01	0.01	0.01	0.03^{**}
	(0.02)	(0.02)	(0.02)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)
seniormanual	−0.01	0.05^{*}	0.01	0.03	0.01	−0.03	0.01	0.01
	(0.03)	(0.03)	(0.03)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)
skilledmanual	-0.06^{**}	0.04^{**}	-0.04^{**}	-0.05^{*}	-0.04^{**}	-0.06^{***}	-0.04^{**}	−0.02
	(0.02)	(0.02)	(0.02)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)
unskilled	0.01	0.03^{*}	0.03^{*}	−0.06	0.03^{*}	-0.07^{**}	0.03^{**}	−0.03
	(0.02)	(0.02)	(0.02)	(0.04)	(0.02)	(0.03)	(0.02)	(0.02)
soldier	0.00	0.12^{***}	0.04	0.00	0.04	-0.09^{**}	0.04	0.06^{*}
	(0.05)	(0.03)	(0.04)	(0.08)	(0.04)	(0.04)	(0.04)	(0.03)
N	82613		128782		129605		135167	

Estimates are from a bivariate ordered probit model, also included were country and wave fixed effects. Clustered standard errors in parentheses. Dependent variables are *basicatt*- a trinary question recording whether the respondent believes society needs radical change, reform, or defending; *stateown* -whether state ownership of industry should be increased; *govtresp* - whether it is an individuals or their governments responsibility to ensure they are provided for. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 2: Other measures of political preferences - Continued

	hardwork		wealthaccum		f1	
	selfpospolit	hardwork	selfpospolit	wealthaccum	selfpospolit	$f1$
male	0.05^{***}	0.06^{***}	0.05^{***}	-0.03^{**}	0.11^{***}	0.13^{***}
	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)
age100	0.19^{**}	0.27^{***}	0.19^{**}	0.20^{***}	0.61^{***}	0.19^{***}
	(0.09)	(0.05)	(0.09)	(0.05)	(0.10)	(0.05)
nokids	0.00	-0.00	0.01	0.00	0.08^{***}	-0.01
	(0.02)	(0.01)	(0.02)	(0.01)	(0.03)	(0.01)
livewithrents	0.03^{**}	-0.02	0.03^{***}	0.01	0.19^{***}	-0.02^{*}
	(0.01)	(0.02)	(0.01)	(0.01)	(0.03)	(0.01)
fulltime	0.01	-0.00	0.01	0.00	-0.19^{***}	-0.08^{***}
	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.01)
parttime	0.01	-0.04^{*}	0.01	-0.01	-0.19^{***}	-0.06^{***}
	(0.02)	(0.02)	(0.02)	(0.02)	(0.04)	(0.02)
retired	0.03	0.05	0.03	0.02	-0.27^{***}	-0.11^{***}
	(0.03)	(0.03)	(0.03)	(0.02)	(0.05)	(0.02)
housewife	0.12^{***}	0.01	0.11^{***}	0.01	0.23^{***}	0.13^{***}
	(0.03)	(0.03)	(0.03)	(0.02)	(0.04)	(0.02)
student	-0.00	0.07^{**}	-0.01	0.01	-0.05	0.03
	(0.03)	(0.03)	(0.03)	(0.03)	(0.05)	(0.03)
highesteduc	-0.02^{***}	-0.00	-0.02^{**}	0.00	-0.03^{***}	0.05^{***}
	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.00)
scaleofy	0.02^{***}	0.01^{***}	0.02^{***}	0.01^{***}	0.02^{***}	0.05^{***}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
seniornonmanual	0.03	0.04^{*}	0.02	0.03	0.20^{***}	0.06^{***}
	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.02)
seniormanual	0.01	-0.00	0.00	-0.01	0.10^{*}	0.06^{**}
	(0.03)	(0.03)	(0.03)	(0.03)	(0.05)	(0.03)
skilledmanual	-0.04^{**}	-0.01	-0.05^{**}	-0.02	-0.07^{**}	-0.18^{***}
	(0.02)	(0.02)	(0.02)	(0.03)	(0.04)	(0.02)
unskilled	0.04^{**}	0.01	0.04^{*}	-0.04	0.09^{**}	-0.14^{***}
	(0.02)	(0.04)	(0.02)	(0.04)	(0.04)	(0.02)
soldier	0.10^{**}	0.02	0.10^{**}	-0.04	0.29^{***}	-0.10^{**}
	(0.05)	(0.05)	(0.05)	(0.05)	(0.10)	(0.05)
N	94658		92510		45667	

Dependent variables are *hardwork* - captures whether hard work or luck brings success; *wealthaccum* measures whether individuals believe that people can only get rich at the expense of others or if wealth can grow so theres enough for everyone? *f1* the first principal component of the all of the other dependent variables. Other details as previous page.

Table 3: Pseudo-Panel Estimates

	(1) incomeequal	(2) selfpospolit	(3) incomeequal	(4) selfpospolit
nokids	-0.24^{***}	0.18^{**}	-0.24	0.18
	(0.09)	(0.07)	(0.20)	(0.18)
livewithrents	-0.50^{***}	-1.05^{***}	-0.50^{*}	-1.05^{***}
	(0.12)	(0.09)	(0.27)	(0.21)
fulltime	0.10	-0.26^{***}	0.10	-0.26
	(0.10)	(0.08)	(0.23)	(0.17)
parttime	-2.18^{***}	2.75^{***}	-2.18^{***}	2.75^{***}
	(0.33)	(0.25)	(0.68)	(0.49)
selfemployed	-2.17^{***}	-0.35^{*}	-2.17^{***}	-0.35
	(0.24)	(0.18)	(0.50)	(0.36)
retired	-0.99^{***}	-0.12	-0.99^{***}	-0.12
	(0.12)	(0.09)	(0.25)	(0.21)
housewife	-0.64^{***}	-0.22	-0.64^{*}	-0.22
	(0.19)	(0.15)	(0.38)	(0.41)
student	-0.19	-0.86^{***}	-0.19	-0.86^{**}
	(0.16)	(0.12)	(0.34)	(0.38)
highesteduc	0.21^{***}	-0.06^{**}	0.21^{***}	-0.06
	(0.03)	(0.02)	(0.05)	(0.06)
scaleofy	-0.13^{***}	0.09^{***}	-0.13^{***}	0.09^{***}
	(0.02)	(0.01)	(0.03)	(0.03)
seniornonmanual	0.66^{***}	-0.31^{**}	0.66^{**}	-0.31
	(0.17)	(0.13)	(0.26)	(0.28)
seniormanual	-3.54^{***}	-2.78^{***}	-3.54^{***}	-2.78^{***}
	(0.57)	(0.43)	(1.01)	(0.73)
skilledmanual	-1.29^{***}	-0.18	-1.29^{***}	-0.18
	(0.16)	(0.12)	(0.34)	(0.26)
unskilled	-0.38^{***}	0.61^{***}	-0.38^{**}	0.61^{***}
	(0.07)	(0.06)	(0.18)	(0.17)
soldier	2.06	4.79^{***}	2.06	4.79^{***}
	(1.62)	(1.24)	(2.34)	(1.58)
N	66408	66408	66408	66408

Columns 1 and 2 report the results of a pseudo-panel estimator correcting for measurement error, standard errors calculated under the assumption of homoskedasticity. Columns 3 and 4 repeats the analysis in Columns 7 and 8 but report bootstrapped standard errors. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Allowing for the interaction of *incomeequal* and demographic variables.

	selfpospolit	
nokids	−0.20	(0.24)
livewithrents	0.42*	(0.22)
fulltime	0.90***	(0.30)
parttime	6.16***	(0.43)
selfemployed	1.78***	(0.40)
retired	0.56*	(0.33)
housewife	0.85***	(0.31)
student	−0.48	(0.35)
highesteduc	0.22***	(0.05)
scaleofy	−0.16***	(0.04)
seniornonmanual	−0.35	(0.27)
seniormanual	−0.75	(0.48)
skilledmanual	1.04***	(0.27)
unskilled	1.03***	(0.19)
soldier	6.20***	(1.30)
incomeequal	3.18***	(0.38)
incomeequal × nokids	0.07*	(0.04)
incomeequal × livewithrents	−0.14***	(0.04)
incomeequal × fulltime	−0.16***	(0.05)
incomeequal × parttime	−0.89***	(0.07)
incomeequal × selfemployed	−0.30***	(0.06)
incomeequal × retired	−0.07	(0.05)
incomeequal × housewife	−0.12**	(0.05)
incomeequal × student	−0.04	(0.06)
incomeequal × highesteduc	−0.04***	(0.01)
incomeequal × scaleofy	0.05***	(0.01)
incomeequal × seniornonmanual	−0.03	(0.04)
incomeequal × seniormanual	−0.11	(0.08)
incomeequal × skilledmanual	−0.18***	(0.05)
incomeequal × unskilled	−0.14***	(0.04)
incomeequal × soldier	−0.84***	(0.20)
N	66408	

Estimates are obtained using the pseudo-panel fixed-effects estimator described in Equation 9. Standard Errors in Parenthesis * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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