MEMORANDUM

No 02/2006

Do the rich vote Conservative because they are rich?

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Do the rich vote Conservative because they are rich?*

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February 3, 2006

Abstract

Political economy models predict that the rich oppose redistribution, and hence vote for conservative parties. Although this seems to fit the data well, I show that this is not true when we control for unobservable characteristics. Using Norwegian survey data, I study to what extent voting is caused by income. Unobserved characteristics correlated with income are handled by using fixed effects panel data discrete choice models. Although a positive association between income and conservative voting persists when controlling for unobservables, the magnitude of the effect is reduced by a factor of five. To correct for measurement error, I instrument income with average income by profession. The magnitude of the coefficients is increased, but the main conclusions remain.

Keywords: Political economy, redistribution, voting, multinomial logit, panel data
JEL Classification: C23, C25, D31, D72, H11, H53

1 Introduction

Workhorse models in political economy and public choice entail that a rich person prefers a lower tax rate than a poor. The reason is simple: The richer you are, the more taxes you pay whereas the benefits remain almost the same. This is a crucial element of much of political economics, such as the median voter approaches to tax determination of Romer (1975), Roberts (1977), and Meltzer and Richard (1981), probabilistic voting approaches, such as Dixit and Londregan (1996) and Lindbeck and Weibull (1987), and models of special interest groups as discussed by e.g. Grossman and Helpman (2001). If income does not determine voting, this literature is simply flawed.

For instance the Meltzer-Richard model implies that there is more redistribution the more unequal a society is. Empirically, however, there is no general tendency that unequal societies redistribute more than equal societies. A proper understanding of the relationship between income, preferences for redistribution, and actual voting is a key to understanding this "redistribution puzzle".

Apparently, the standard political economy prediction of the rich opposing redistribution and voting for conservative parties seems to fit the data well. Figure 1 shows opinions on tax cuts for high incomes and support for the Conservative party by income quintile for a sample of Norwegian respondents.[1] Respondents in the higher income groups are much more positive to cuts in taxes they do support the Conservative party. Preferences for tax cuts are doubled and conservative voting almost tripled in the 5th quintile relative to the 1st and 2nd.

However, this may be overly simplistic. First, variables such as gender or education might have an impact. The conclusion is almost unchanged if we control for such observable characteristics. What we should worry more about is unobservable characteristics of the agents, such as social background and experience-induced beliefs. Social background captures a set of value norms and views on the working of society learned at young age. Political sociology suggests that voting is determined by location in social structure and party identification acquired during childhood and adolescence. This means that voting is determined by social background. Income is also highly correlated with social background, and hence income and

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[1]A description of the data is given in Section 4. Income before 1993 is based on the simulation technique of Section 3.
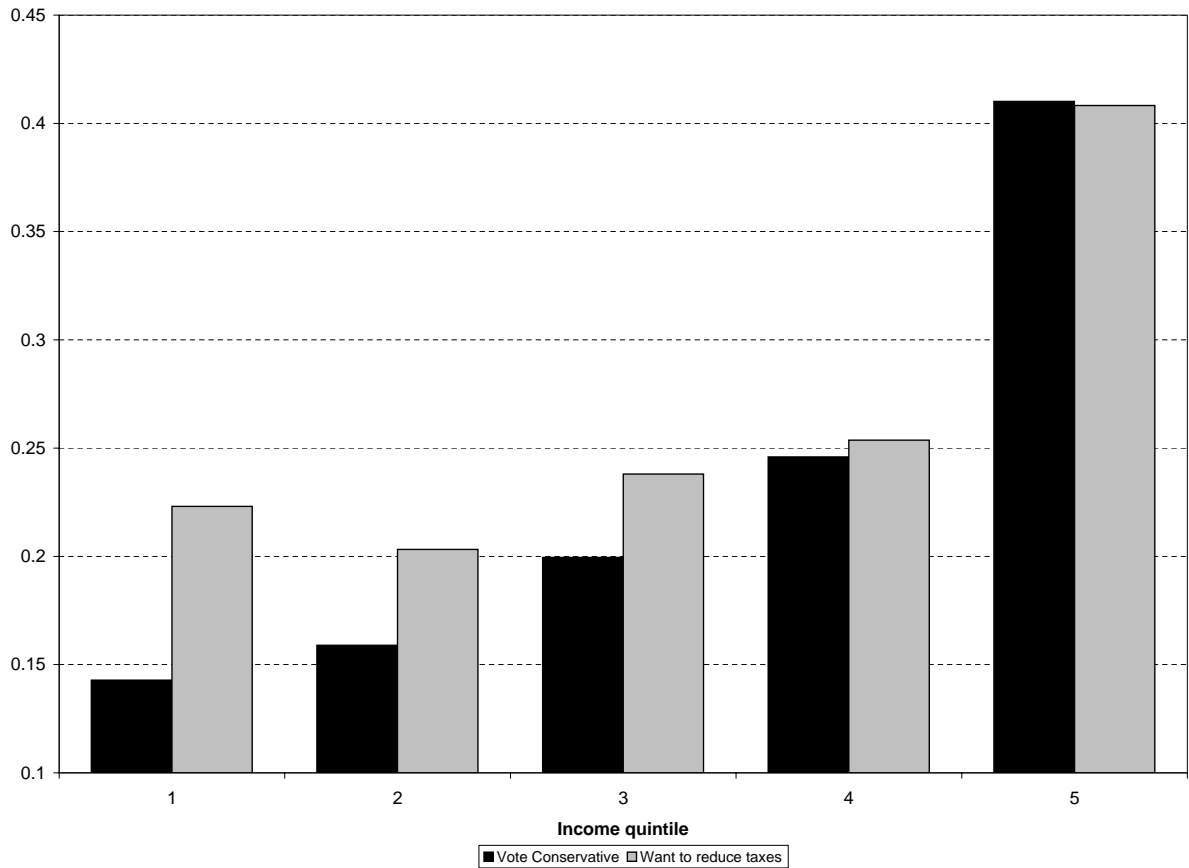
Figure 1: *Stated preferences for tax cuts for high incomes and fraction voting Conservative by income quintile.*

voting will be correlated. But this relationship is not causal, so a change in income will not lead to a change in voting behaviour. People's beliefs about the working of the economy also influences their political preferences. As highlighted by e.g. Thomas Piketty (1995), beliefs about the economy are likely to be influenced by the agent's own experience. If he has experienced upward economic mobility, he tend to believe that effort is the main ingredient of success, and hence that state intervention is not necessary. On the opposite, if he has experienced downward mobility, he attributes this to unfortunate circumstances, and hence believe that state intervention is necessary. Hence there is correlation between high income and the belief in hard work rather than luck as the source of mobility. We could use proxies to control for these, but this is unlikely to remove more than a small fraction of the induced bias.

To the best of my knowledge, there are no studies that estimate the effect of personal income on voting controlling for unobserved social background in a satisfactory manner. That is what I do in this paper, and it turns out that controlling for unobservables has a major impact on the estimated effect of income on voting behaviour. To estimate this, I assume that during a limited period of time, unobserved characteristics remain virtually constant, so they are individual specific. If a change in income over a limited period of time has an impact on expressed political preferences, we may quite reasonably say that the income change is the true cause of this change in preferences.

To estimate the effect of income on voting, I use panel data from the Norwegian Election Study for the seven elections between 1977 and 2001 where each respondent is interviewed after two consecutive elections. There are several reasons why Norway is a suitable country for this study. First, as it has a proportional electoral system, there are more than two parties and more than one "active" cleavage. Hence there is not the same tendency for convergence towards the middle that is found in most majoritarian polities. Second, there is a clearly right-wing party, the Conservative party, which could potentially attract voters from the whole electorate independently of religious or ethnic belongings.

I first formulate a random utility model for party choice, which leads me to a logit type choice structure. Unobserved characteristics are modelled as individual specific effects. I consider three main cases: First an ordinary random effects panel data logit model. The disadvantage of this model is that we have to assume that the individual effects, which contain the unobserved characteristics, are uncorrelated with the regressors, most importantly income. As I argued above, this is probably not the case. However, this model gives a benchmark to compare against when we later allow for unobserved effects that are correlated with income. To see whether income has an impact on those who actually change their party preferences, I secondly restrict the sample to respondents who vote for different parties in the two periods of interview. Finally, I introduce a fixed effects panel specification. Here, one preference parameter is estimated for each party for each voter. The incidental parameter problem is solved by using a version of conditional maximum likelihood. This method is robust with regard to correlation between the individual effects and the regressors.

When I assume that the individual specific effects are uncorrelated with income, I find strong support for the hypothesis that high incomes induce Conservative voting. However,

once I introduce estimators that are more robust to correlations, these results become insignificant. The magnitude of the effect is also reduced by a factor of five. This seems to indicate that the rich vote Conservative, but only to a limited extent because they are rich.

As the study is based on stated income, one explanation of this finding may simply be measurement error in income. It is well known that variables of interest, such as income, are more persistent than measurement error. This leads fixed effects estimators to have a stronger attenuation bias than other estimators, and hence constitute a rival explanation for my findings. To solve the problem of measurement error, I use an instrumental variables technique for limited dependent variables where income is instrumented with the average income of the respondent's profession. This yields numerically higher estimates, so it seems measurement error is a matter of concern. However, the estimates are still much smaller when the fixed effect is included, so the hypothesis that the correlation between income and voting is largely driven by unobserved characteristics is corroborated.

Although voting behaviour is what finally determines policy choices, it is also useful to look at opinions on economic policy. To do this, I study the answer to questions about whether the respondents favour tax cuts for high incomes and extensions of social security. This analysis confirms my previous findings. In the benchmark random effects model, income has a significantly positive effect on preferences for tax cuts and negative effect on preferences for extensions of social security. When we either focus only on those who give different answers in the two periods or introduce fixed effects, the effect of income almost totally disappears. This is also the case when we use instrumental variables to correct for measurement error.

The existence of these unobservables implies that there are two different types of comparative statics we can perform. On the one hand, we can condition on background and experience and change the agents' income. This is what we associate with such experiments as mean preserving spreads, and gives the short run effect of a change in income. On the other hand, we can change both income and background characteristics. Then we would expect to find the same effect as is found in cross sectional analyses ignoring unobservable characteristics of the agents. This corresponds to the effects of a change in income in the long run, where long run may be several generations. These analyses are fundamentally different, and it is crucial to match the right analysis to the right question.

2   Income and voting

Before explaining the estimation strategy in detail, I review some of the literature on income and voting. I first consider why we should expect a rise in income to lead to preferences for lower taxes before I review other literature that may explain a correlation between high incomes and conservative preferences and voting without creating a causal relationship.

The simplest, and probably best known, framework for analysing how income determines voting behaviour is the simple median voter models of Romer (1975), Roberts (1977), and Meltzer and Richard (1981). Essentially, the economy is populated by a number of agents with different incomes. They vote over a linear tax rate that is used for a uniform transfer to all agents. Under general conditions, the tax rate preferred by the agent with median income, which is the median voter, is a Condorcet winner, i.e. there is a majority of voters who prefers this tax rate to every other tax rate. In these models, an agent pays a constant share of his income in taxes but receive a lump sum in transfers. Hence a tax cut makes him richer if his income is high and poorer if his income is low.[2] The richer he is, the stronger are his preferences for low taxes and vice versa.

One can envisage a large number of extensions to this basic model. For instance Moene and Wallerstein (2001) extend public policy to consist of both ordinary redistribution and social insurance. As insurance is a normal good, the rich may have a higher demand for it than the poor. However, as the rich also usually have better access to private insurance, the should still be a negative relationship between income and support for social security should. Lind (2005) considers various forms of social conscience and altruism towards society as a whole and members of say one's ethnic group. This may entail that agents with identical income have different preferences for redistribution. If we control for group belonging, however, the basic relationship should remain.

A class of somewhat more advanced models are the models of probabilistic voting. The seminal paper is Lindbeck and Weibull (1987); see also Dixit and Londregan (1996) as well as the overview given by Persson and Tabellini (2000). Here, one allows that some rich agents vote for politicians favouring high taxes and vice versa. In particular, each agent has preferences consisting of a deterministic term similar to the preferences described above,

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[2]The turning point is located somewhere below the average income. The exact location depends on the nature of deadweight losses.

but also a stochastic term to accommodate differing views on the particular politicians or to accommodate for other dimensions to politics. Within this class of models, we should also expect to see a positive correlation between high income and preferences for low taxes.

The relationship between income and opinions about redistribution or voting has been studied before. There are survey-based studies exploring how income affects preferences for redistribution (Husted 1989), taxes (Lewis 1979, Furnham 1984), and provision of public goods (Bergstrom et al. 1982, Gramlich and Rubinfeld 1982, Preston and Ridge 1995). Another approach is to see whether those who benefit from a government program are more positive to it than others. In an early study, Mueller (1963) uses US survey data on opinions about a number of government programs and possible tax cuts or raises. She finds some support for the hypothesis that those who benefit from a program are more positive to it than others. A somewhat more recent study is Sanders (1988). He mostly confirms her findings, using a different methodology.

It has also been found that persons who believe in upward economic mobility tend to oppose redistributive policies and vice versa. In a large international survey, Corneo and Grüner (2002) find strong support for the hypothesis that those who expect to benefit from redistributive programs are more supportive of them than others. Ravallion and Lokshin (2000) find similar results on Russian data. They also find that respondents who believe in upward economic mobility tend to oppose redistributive policies and vice versa. Alesina and La Ferrara (2002) confirm this result using US survey data. A problem with this literature is that stated opinions on redistribution may not be decisive for the voters once they are to cast their vote.

McCarty et al. (2003) find a positive relationship between income and Republican partisanship which has been strengthening over time. They argue that this is due to increased polarization as income inequalities have been increasing. However, they make no real attempts at controlling for unobserved heterogeneity in their study. In parts of the literature on voting, income is also used as a control variable, but again no attempts for controlling for unobservables is done, and most of the time little emphasis is put on income.

Another strand of literature investigates how the macroeconomic situations influence the incumbent party's popularity. Douglas Hibbs (2000) has coined the term "bread and peace voting" for such behaviour. On time series data, Kramer (1971) had one of the

first contributions. Using cross sectional data, one can distinguish between pocketbook (or egotropic) voting, i.e. how your own economic situation influences support for the incumbent party, and sociotropic voting, how the macroeconomic situation affects voting. There seems to be strong support for sociotropic voting but almost no support for pocketbook voting. Kinder and Kiewiet (1979, 1981) use cross sectional data to distinguish between pocketbook voting, i.e. how your own economic situation influences support for the incumbent party, versus sociotropic voting, how the macroeconomic situation affects voting. On US data, they find strong support for sociotropic voting but almost no support for pocketbook voting. Lewis-Beck (1986) confirms these findings on data from Britain, France, Germany, and Italy. However, in the recent symposium on economic voting in *Electoral Studies*, edited by Lewis-Beck and Paldam (2000), the overall conclusion is somewhat moderated. For most countries there is little support for pocketbook voting, but there are exceptions, most notable Denmark and the UK.

A different approach taken by a number of researchers is to test the key outcome of the standard political economy models that there should be more redistribution in inegalitarian societies. The most popular approach is to regress some measure of the amount of redistribution on a measure of inequality on cross sections of countries; see Bénabou (1996) for a survey of these. Generally, if there is a relationship, it is a negative correlation between inequality and redistribution. However, most of the time, no significant relationship is found.

There is also an enormous number of non-economic theories of voting, and some of these can explain why there is a correlation between income and conservative voting without the relationship being causal. A group of theories posit that party preferences to a large extent are inherited from your parents. Furthermore, an agent's income is highly correlated with his parent's income. Hence if a person comes from a high income home where the parents used to vote conservative, he will also tend to both have a high income and vote conservative. However, the conservative voting is not necessarily driven by income.

Lazarsfeld et al. (1948) emphasize the importance of socioeconomic status. Level is assigned by interviewers, who are "trained to assess the homes, possessions, appearance, and manner of speech of the respondents and to classify them into their proper stratum in the community according to a set quota. The people with the best homes, furniture, clothes etc., i.e., the ones with most money, would be classed as A's; and the people at the other

extreme would be D's" (p. 17). They find a strong relationship between this measure of socioeconomic status and voting behaviour, and it is clear that it is also correlated with income. However, few objects to this being a less than perfect measure of socioeconomic status, and it is probably impossible to construct a measure that fully controls for status.

Campbell et al. (1960) also find social background as a strong predictor of party preferences, but they emphasize party identification as an important determinant of voting behaviour. Identification is defined as "the individual's affective orientation to an important group-object in his environment" (p. 121), where the political party serves as group for party identification. They find strong correlation between an individual's party identification and his parent's political orientation: "[A]n orientation toward political affairs typically begins before the individual attains voting age and [...] this orientation strongly reflects his immediate social milieu, in particular his family" (p. 146f). Niemi and Jennings (1991) confirm the finding that parents play a major role in shaping the offspring's party identification using parent-offspring panel data. Party identification is usually measured by self classification. This does probably not give a complete measure of both the direction and strength of party identification, among others because the question is rather abstract and different respondents interprets it differently. Hence using the self reported measures to control for party identification does not solve the omitted variable bias, but it does indicate that party identification is an important determinant of voting behaviour. This has also been confirmed by a large number of studies. See e.g. Berglund (2004) for a recent study of the Norwegian context.

There is also a voluminous literature on class voting. Very generally, working class background is a good predictor of socialist vote. For instance Nieubeerta and Ultee (1999) confirm this on a large number of countries, and the relationship also holds over time, although the strength of class has been declining in most countries. Class background is clearly correlated with income. Some research has actually used income to operationalize class. However, this is in my view a misinterpretation of the concept of class, and it also seem that both class and income have separate effects on voting (Weakliem and Heath 1994). Again, a perfect measure of class is difficult, so this is also likely to be an omitted variable.

A last explanation for a correlation between income and tax preferences that does not go directly through income is the one presented by Piketty (1995). Is model is to some

extent a model of party identification through learning. He depicts an economy where there is disagreement about the disincentive effects of taxation. Through their own experiences as well as that of their ancestors, agents learn about the working of the economy. Then we can get left-wing dynasties who don't believe in a large degree of social mobility, and hence favour more redistribution than right-wing dynasties, which in general has experienced more upward mobility. This then generates a positive relationship between income and preferences for low taxes.

3   Empirical model and estimation

To estimate a model taking into account some of the features set out above, I equip each voter with a set of preference functions for each party available. We may interpret this as the agent's utility of the given party coming to power. As there is little scope for strategic voting in parliamentary elections, it is natural that the agent also vote for the party he most prefer to see in power. We may also construct other models of how the agent decides on which party to vote. Generally, I will talk about the utility he expects to get from voting for a party, rather than from seeing the party in power, and leave the precise interpretation to the reader.

I assume that there is a stochastic component to utility, closely related to the assumptions of probabilistic voting models. Utility for a party depends on agent and time specific effects, his income, and other background variables. Assuming a linear structure, agent i gets utility

$$v_{itj} = \alpha_{ijt} + \beta_j y_{it} + \gamma_j z_{it} + \varepsilon_{ijt} \tag{1}$$

from voting party j at time t. In the present work, I only study the effects of the current income y_{it}. In a companion paper I study the effect of permanent versus transitory income on voting behaviour (Lind 2004). α_{ijt} is a latent party, period, and individual specific term, z_{it} individual characteristics other than income, and ε_{ijt} a standard extreme value distributed taste shifter. Throughout the analysis, I use the specification $\alpha_{ijt} = \alpha_{ij}^I + \alpha_{jt}^T + \alpha_j^P$ where α_{ij}^I is the individual specific term, which I give different specifications outlined below, α_{jt}^T a time-party specific fixed effect, and α_j^P a parameter which applies if the agent is in his second period of interviewing. Since the model contains the period specific effects α_{jt}^T, all effects related to incumbency, macroeconomic conditions, popularity of specific politicians, and so

forth are controlled for. To test whether the conventional approach to political economy is legitimate, we need to test whether $\beta_j = 0$ for all j. The appropriate alternative hypothesis is that β_j is high for right-wing parties and low for socialist or left-wing parties.

Let us first consider a standard random effects model with normally distributed individual effects, which is used as a benchmark model. The individual effects are here modelled as $\left(\alpha_{i1}^I, \ldots, \alpha_{iN}^I\right)' \sim NID\left(0, \Sigma_\alpha\right)$, i.e. independent of income y_i. Computationally, the model is estimated by Monte Carlo integration, hence by drawing αs from the appropriate distribution and averaging. Appendix B.1 gives further details. This model requires the random effects α_{ij}^I to be independent of y_{ijt} and z_{ijt}, a strong and probably unrealistic assumption: In the previous section I argued that the non-economic utility of voting for right-wing parties should be higher for those with high incomes. As only a part of this non-economic utility is picked up by the observables z_{ij}, it is likely that y_{it} and α_{ij}^I are correlated hence making estimation based on (3) invalid.

To solve this problem, we can let the α_{ij}^I's be time independent individual specific taste parameter. However, as the sample size increases keeping the number of periods each individual is observed fixed, the number of α_{ij}^I's to estimate increases proportionally. This is Neyman and Scott's (1948) incidental parameter problem. Andersen (1970), however, shows that if we can find a sufficient statistic for the incidental parameters, we may use conditional maximum likelihood, so we maximize the likelihood of the data conditional on a sufficient statistic for the incidental parameters. See also Cox (1975). Chamberlain (1980) has derived a conditional maximum likelihood estimator for fixed effects logit models and show that knowledge of what parties an agent vote at some time is a sufficient statistic for the individual specific effects. We then use the order in which he votes these parties to estimate the parameters. I limit the exposition to the case of two periods as this is the case we encounter below. Also, for simplicity of notation, I omit the period specific and second period effects. Assume that individual i votes for party p_1 and p_2 in period 1 and 2. Conditional on having voted for parties p_1 and p_2, i.e. on one of the two voting sequences (p_1, p_2) or (p_2, p_1), the probability of this observation is

$$\frac{\Pr\left(p_1, p_2\right)}{\Pr\left(p_1, p_2\right) + \Pr\left(p_2, p_1\right)}.$$

As the individual effects enter similarly into both $\Pr\left(p_1, p_2\right)$ and $\Pr\left(p_2, p_1\right)$, they cancel out in this expression. Details are provided in Appendix B.2. This estimator is extremely

robust to correlation between the unobserved individual effect an the explanatory variables. However, it is also a costly estimator as we only use those observations where the individual has changed the party he voted for and also loose information as we only consider within-variation in regressors and ignore between-variation. I have run most of the estimations shown below using a less costly, but also less robust estimator proposed by Chamberlain (1980). Here, the individual effects are supposed to be correlated with the average income over the two periods. It turns out that the results are similar to those obtained from the random effects estimator, so I have not included these results in the paper.

A final complication is that for the first periods in the sample, we do not have the precise value of the income y_{it}, only an interval in which it lies. We could then use a dummy variable for each income group. However, these groups change quite considerably over time, so this would not be possible on the whole panel. Rather, I have used simulated maximum likelihood. I assume that a log normal distribution is a reasonable approximation to the income distribution in each period. Using the number of respondents in each income interval, we can then estimate the two parameters μ and σ, the mean and standard deviation of log income, for each period.[3] The estimated parameters for the relevant years are found in Table 1. One could estimate the parameters of the income distribution jointly with the policy preference parameters, but for computationally simplicity I use the two step procedure of first estimating the income distribution and then estimating the preference parameters. Some respondents have reported zero income. They are left out of the estimation of μ and σ, and their income is kept at zero throughout the simulations.

Let $\mathcal{L}(y_i, z_i, \Theta)$ denote the appropriate log likelihood function for individual i who has income y_i, other characteristics z_i, and where Θ denotes the parameters to be estimated. We only know that y_i is in the interval $(\underline{y}, \overline{y})$. Then the log likelihood given his income group, but ignoring his precise income, is

$$\frac{1}{F(\underline{y}) - F(\overline{y})} \int_{\underline{y}}^{\overline{y}} \mathcal{L}(y, z_i, \Theta) \ dF(y),$$

where F is the CDF of the income distribution, in this case the log normal. To calculate

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[3]As $\ln y_i \sim N(\mu, \sigma^2)$, we have $\Pr(\underline{y} \leq y_i \leq \overline{y}) = \Phi\left(\frac{\ln \overline{y} - \mu}{\sigma}\right) - \Phi\left(\frac{\ln \underline{y} - \mu}{\sigma}\right)$ where Φ is the standard normal CDF. μ and σ are then estimated from the log likelihood function $\sum_{g=1}^{G} N_g \ln\left[\Phi\left(\frac{\ln \overline{y}_g - \mu}{\sigma}\right) - \Phi\left(\frac{\ln \underline{y}_g - \mu}{\sigma}\right)\right]$, where G is the number of income groups, N_g is the number of individuals in group g and $\overline{y}_g$ and $\underline{y}_g$ the upper and lower incomes in the group.

this integral, I simulate the income of the agent by making a number of draws from the log normal distribution with the appropriate parameters, conditional on the agent being in his income group.

4   The Norwegian polity and the data set

The model is estimated on data from the Norwegian Election Study, made available through the Norwegian Social Science Data Services. The years 1977 to 2001, a total of seven elections, are used.[4] The study is a rotating panel where most of the respondents participate in the survey in two consecutive elections. Each survey contains a large number of questions on the respondent's opinion about political question as well as some background variables. Although the questions vary over time, there are some key questions that are asked each time. For our purposes, the most interesting are the party the respondent voted for and a statement of household income. Responses to opinions on economic policy was asked in a different way in 2001, so analyses of political opinions are restricted to a sample from 1977 to 1997.

Income is measured by recall, and is a measure of last year's income. With minor modifications, it has been measured by the question

> *What gross income did you (and your spouse/co-habitee) have [last year]? With gross income, we mean total income before any deductions or tax.*

From 1977 to 1989 the respondents were only asked to identify their income group and not the precise income. As explained in Section 3, this may be solved by using simulated maximum likelihood. Descriptive statistics on the simulated incomes as well as the stated incomes from 1993 to 2001 are found in Table 1. As expected, real incomes are increasing over the period. Also, we can see a slight increase in inequality, particularly in 2001.

Political economy models don't usually tell us whether it is the agents' permanent or transitory income that should matter for policy preferences. If politics are chosen independently after each election, then what should matter is the agent's expected permanent

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[4]An overview of the whole data set is found in Kiberg et al. (2000). See Statistics Norway (1978) for documentation of the 1977 wave, Statistics Norway (1982) for the 1981 wave, Statistics Norway (1986) for the 1985 wave, Valen et al. (1990: Appendix B) for the 1989 wave, Aardal et al. (1995) for the 1993 wave, Aardal et al. (1999) for the 1997 wave, and Aardal et al. (2003) for the 2001 election.

income during the election period. In this case, using the income for say the first year could be a reasonably good approximation. However, it is likely that there is a degree of status quo bias in politics: Major changes of policy usually takes a while. Hence it can be rational for the agents to look beyond the present electoral period. Consequently, what should matter for policy preferences is a mix of today's income and the agent's expected permanent income.

In the present work, I have included some results where I use CPI adjusted income, but I have relied mostly on incomes relative to annual averages, as this is what determine whether an agent gains or looses from a redistributive scheme with a linear tax. Also, I have only looked at the effects of income, ignoring the effect of wealth. The reason is that my data set does not have respondent's wealth. Also, reliable data on wealth are hard to obtain. Estate taxation, which is an important part of taxation of wealth is in Norway determined at the local level, whereas I study (national) parliamentary elections. Finally, the two are usually reasonably highly correlated, so results that hold on income are likely to be rather similar for wealth.

Table 1 about here

Throughout the paper, I group the parties into five group. Each party group's vote share in the sample and at the elections are given in Table 2. The *Socialist parties* consist of the Social Leftist party as well as the Norwegian Communist Party and the Workers' Communist Party. The second group is the *Labour party*. As this is the largest group, I use it as the reference group. Then follows the *Centrist parties* which consist of the Centre Party, the Liberal party (Venstre), the Liberal Popular Party, and the Christian Popular Party. The *Conservative party* (Høyre) is grouped alone as is the *Progress party*. There seems to be a slight under reporting of voting Progress relative to the outcomes of elections, but otherwise the data seem to reflect the actual outcomes fairly well.

Table 2 about here

The best overview of Norway's political history and political landscape up to the 1960s is still Rokkan (1967). In his path breaking study, he distinguished between five dimensions of conflict in Norwegian politics. These are (1) the territorial (center vs. periphery), (2) the sociocultural, (3) the religious, (4) an economic conflict on the commodity market (producers vs. buyers of agricultural products), and finally (5) an economic conflict on the

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labour market. For our purposes, the last cleavage, which corresponds to the traditional division between (economic) right and left, is the important one. Over the last 30 years, this has also probably been the most important one. Rasch (2003) reviews a number of studies who attempt to order the Norwegian parties from left to right using survey data on electors, interview of politicians and political experts, studies of party programs, and studies of parliamentary voting. Quite generally, these studies support the ordering Socialist-Labour-Centrist-Conservative-Progress. The Progress Party, however, is somewhat unclear and is placed between Labour and Centrist by some studies. For this reason, I consider the Conservative party as the advocate for low taxes and study to what extent Conservative voting is associated with high incomes.

This is also confirmed by respondents' response to questions regarding their preferences over policy issues. Figure 2 show the fraction of respondents who oppose tax cuts and favour extensions of social security by party preference. As both of these answers relate to an extended scope of government, we should expect left wing voters to be more positive to the proposed changes. The figure shows that respondents who vote for left wing parties also oppose tax cuts and favour social security extensions.

5   Results from the basic models

The results estimation of the baseline models described above are given in Table 3. These estimations include income and a vector of control variables. I have reported results from random effects models on the whole sample and restricting the sample to respondents who vote for different parties in the two waves of interviews as well as from the fixed effects models. All the models contain year specific dummies and a dummy for the agent being in his second period of observation. The latter, however, is redundant in the fixed effects model.[5] Income relative to annual averages is the main measure of income throughout the paper. Appendix Table A1 reports estimates using income in absolute terms (CPI adjusted to 2001). To facilitate reading of the results, I only include the estimated parameter on income; the full estimation results as well as the estimated covariance matrices of the random effects

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[5]When we both have an individual fixed effect and a year dummy, introducing a dummy for the second period would result in perfect multicollinearity. This is similar to the difficulty of observing individual, year, and cohort effects.
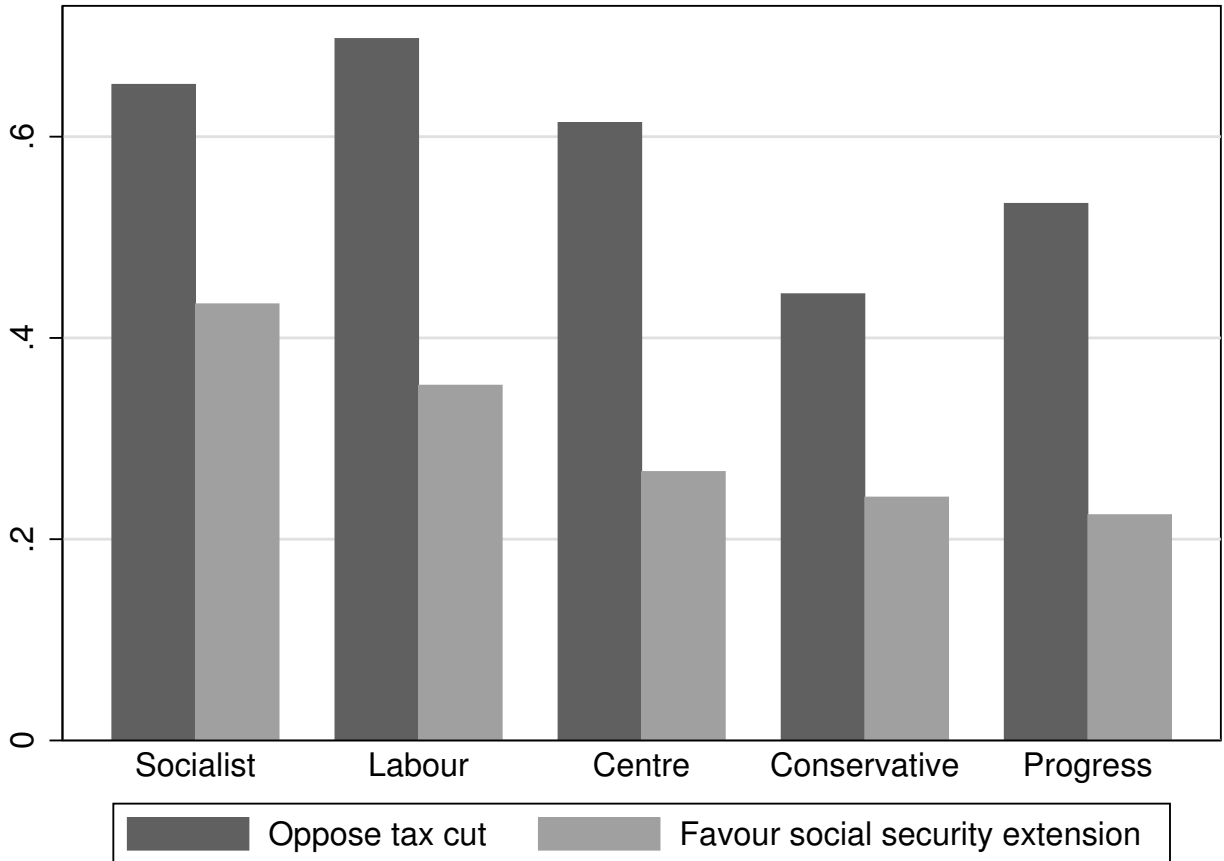
Figure 2: *Preferences for tax cuts and extensions of social security by party preference. Sample is 1977-97. "Oppose tax cut" are respondents who answer that they oppose or oppose strongly reductions in the tax rate of high incomes versus those who favour it, favour it strongly, or indifferent. "Favour social security" are respondents who say they favour extensions versus oppose it or indifferent.*

estimators are available upon request.

Table 3 about here

To test the validity of the political economy approach to voting behaviour, the most interesting estimates are the income coefficients on the Conservative party. The coefficient is positive and significantly different from zero in both random effects models. This is similar to the evidence in Figure 1. However, the estimated parameter is much higher for the full sample than for the sample restricted to party changers. Already, we see a tendency for conservative voters being rich, but that party changes only to a limited degree can be explained by income. The results from the fixed effects estimations confirm these

results: Here the estimated parameter on the Conservative party is even lower, and it is now insignificant. Hence when we control for unobserved characteristics, income has very little explanatory power on Conservative voting.

Another interesting feature of Table 3 is that the socialist parties, which according to the theories of Section 2 should attract the poorest voters, rather seem to attract voters with a relatively high income. But this effect is much smaller than for the Conservative party in all specifications, and in the fixed effects specification it even has a positive sign. Also, the parameter is not significantly different from zero at any conventional level of confidence in any of the specification. This may, however, be due to the reference party being Labour, which also is located towards the left wing.

The Progress party, which is often judged as a populist right-wing party, is slightly troublesome as the estimated parameters do not have a consistent sign. These estimates are also generally not significantly different from zero. This may be explained by the fact that the party to a large extent gets "protest votes". Throughout the period, it was also not expected that the party would be in position. Finally, particularly at the end of the period I study, the party's ideology started to include some more left-wing components such as increased emphasis on the public responsibility for the welfare of senior citizens.

Estimation using income in price adjusted levels rather than relative income reported in Appendix Table A1 gives qualitatively very similar consolations. The figures reported in Table 3 should be interpreted as parameters in a linearized utility function for voting a given party. It may also be of interest to look at the effect of income on choice probabilities. If β_j is the coefficient on party j and P_j the probability of voting for the party, then we see that

$$\frac{\partial P_j}{\partial y} = P_j \left(\beta_j - \bar{\beta} \right) \tag{2}$$

where $\bar{\beta} = \sum P_j \beta_j$. Estimated choice probabilities corresponding to the estimates of Table 3 are reported in Appendix Table A2, where the P_js used are the sample means. Increasing the income by an amount equal to the mean income is seen to increase the probability of voting Conservative by a substantial amount, about 25%, in the random effects models on the whole sample. Restricting the sample to party changers and by using the fixed effects specification, however, the effect is reduced to about 7% and 5%.

The reason for not including number of children in the control variables is that this variable first appeared in the survey in 1985. I have ran estimations on the last five periods

including number of children; the estimates do not change much (estimates not reported). I have also tried to use income per capita or income adjusted by the OECD equivalence scale. The effect of income relative to period averages on preferences for the conservative party do not change much.

To conclude this section, it seems that income has a positive effect on Conservative voting. The effect is reduced strongly by focusing on party changers and even more when introducing individual fixed effects. In the latter case, the estimate is reduced by a factor of about 5 relative to the standard random effects specification. I interpret this as income having a smaller effect on Conservative voting when we control for unobserved effects correlated with income. Consequently, by just observing a positive correlation between income and Conservative voting, we tend to overstate the causal effect of income.

6   Model specification and measurement error

In the previous section, a common parameter on income was estimated for the whole sample. One may worry that as party platforms have changed over the years, the effect of income on party preferences have also changed. To see whether there is a problem of parameter instability, I have rerun the estimations allowing the parameters on income to vary over time. The upper panel of Table 4 shows results from estimations where the sample was split between 1985 and 1989. It is seen that there is a slight tendency for the parameter on the Conservative party to decline over time. However, for the estimators using party changers and fixed effects, we cannot reject the LR test of no break at a 5% level of confidence.

If we allow for a separate coefficient in each year, we still cannot reject LR tests of no structural change for the fixed effects model, but now both random effects models rejects no structural break at the 5% level. Figure 3 show the estimated parameters for each year. We see that almost all parameters have a tendency to fall, indicating increased Labour voting among high income respondents. Focusing on the Conservative party, we also see that we have the same ranking of estimates: Full sample random effects indicates a large, albeit falling, effect of income on conservative voting, the changers sample random effects a smaller effect, and the fixed effects the smallest effect.
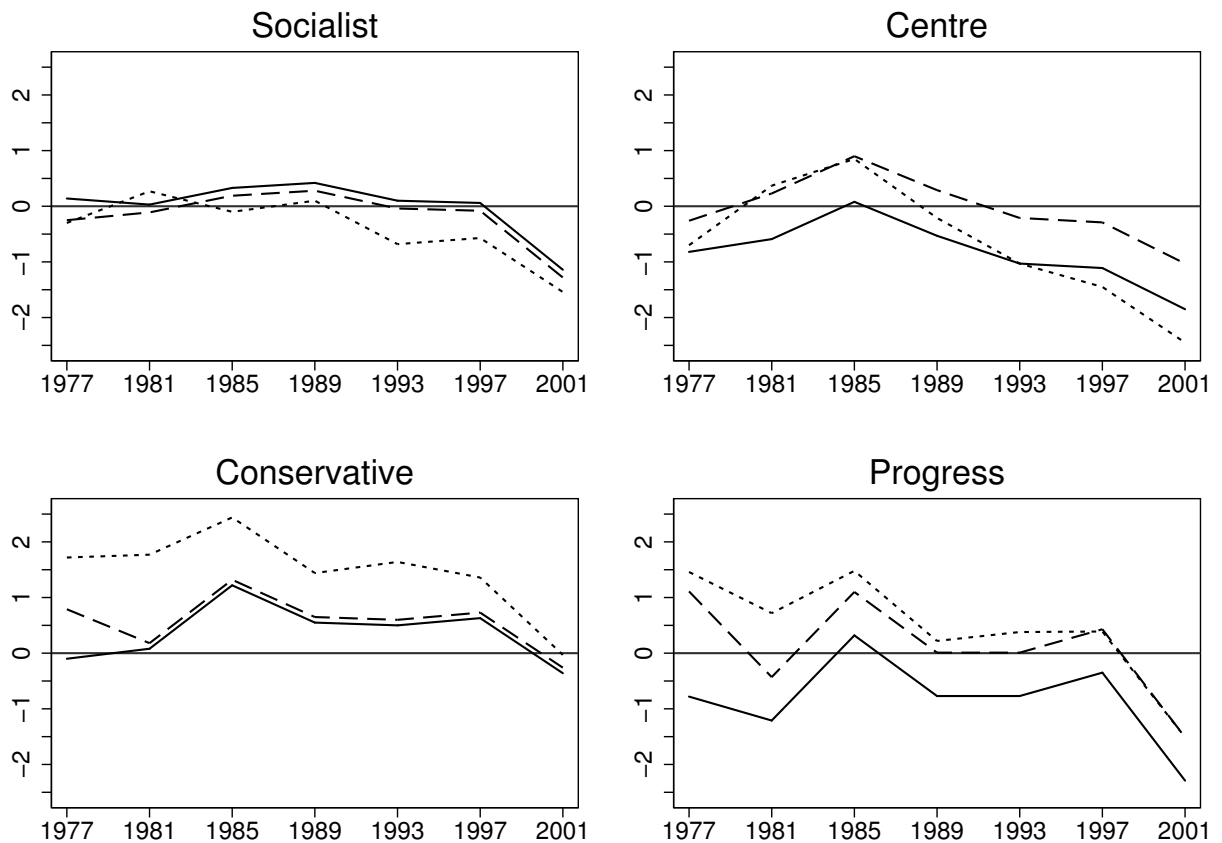
Table 4 about here

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Figure 3: *Year specific coefficients on income in the three specifications. Fixed effects estimator as solid line, random effects estimator on party changers as dashed line, and random effects estimator on full sample as dotted line.*

As estimation is carried out using panel data techniques I only use data on respondents who voted in both periods of interview. We could fear that this leads to a selection bias. To see whether this is that case, I have rerun the estimations including abstention as a separate "party ". This has very little effect on the other estimates, so I have not included these estimates.

The fixed effects estimator is essentially based on studying whether a change in income tends to push voters towards or away from a party. We may, however, expect an increase and a decrease to have non-symmetrical effects. Table 5 show results from estimations that study this. Here I have interacted income with a dummy variable for increasing income from the first to the second period of interviews. The first thing to notice is that the coefficients on the interaction terms are not jointly significant in any of the specifications, so there

is no clear sign of asymmetrical effects of income changes. However, the interaction with increased income seems to have a negative coefficient in many specifications, particularly for the Conservative and the Progress parties. This indicates that an income increase has less of an effect of pulling voters to the Conservative party than a decline has of pushing them away. Overall, there also seems to be a slight tendency for voters to go towards Labour when the income declines. This could be motivated by Labour being seen as a safer party than other choices.

<div align="center">Table 5 about here</div>

For the fixed effects estimator, we only use data on the agents that change party during the two periods. It is possible that those who change their party have preferences that differ from those who stay with the same party in both periods. Obviously, we cannot test this with results from the fixed effects estimator, but results from the random effects estimator give us some insight. Table 6 reports results from an estimation where income has been interacted with a dummy for voting for different parties in the two periods of observation. The first thing to notice is that an LR test of all the interaction terms being zero is rejected at a very high level of significance, so there seems to be differences between party changers and party stayers. Furthermore, it seems that income is a better predictor of Conservative voting for party changers than for party stayers. For the other three party groups, the effect seems to be the contrary.

<div align="center">Table 6 about here</div>

To get en idea of who the party changers and party stayers are, Tables 7 and 8 report some summary statistics on party choices and demographic characteristics of changers and stayers. Table 7 show that there is higher volatility in voting for parties at the extremes, i.e. Socialist or Progress. Hence party changers have a higher tendency for having voted for one of these party groups in one of the periods of observation. Table 8 shows the average of certain demographic characteristics of respondents who voted for the same party and different parties in the two periods as well as results from a logit estimation of the probability of changing party. We see that the income level and the change in income between the two periods is almost the same, and does not turn out to be significant in the logit estimation. Large changes in income in absolute value, however, tends to increase the probability of changing party. Age also seem to have an impact on the probability of changing party:

Party changers are generally younger whereas older voters seems to have found "their" party to a larger extent. Men and women seems to have approximately the same propensity to change parties, but there is some tendency that higher education increases the probability of changing party. However, the effect of education disappears in the multivariate logit analysis, so this may to a large extent be due to highly educated voters being younger on average than less educated voters.

<div align="center">Tables 7 and 8 about here</div>

An important caveat with the results presented above is that income is likely to be measured with error as the data are based on stated recall of last year's income. As usual, this induces an attenuation bias in the results. Furthermore, it is possible that measurement error is more marked in the fixed effects model as income is persistent. Hence the lower estimate of the effect of income in the fixed effects model than the random effects model may simply be the result of measurement error. As the variance of the measurement error is unknown, we have to attempt some instrumental variables-type estimator.

Assume that we have an instrument that is correlated with the true income, uncorrelated with the measurement error, and that doesn't influence party preferences. For the probit model with normally distributed measurement errors, a number of consistent estimators have been developed. See e.g. Nelson and Olson (1978) and Newey (1986). However, as I want to use a fixed effects model, the probit framework is not usable.

I suggest to use a two stage estimator closely related to the ones suggested by Rivers and Vuong (1988) and Stefanski and Buzas (1995), which in their words is approximately consistent, i.e. it converges in probability to a value approximately equal to the true value. The estimator is described in Appendix B.3. The main ingredients are to run a first stage where the explanatory variable measured with error is regressed on the instruments and then estimate the logit model using the predicted values from the first stage. The residuals from the first stage are also included to improve the accuracy of the estimator.

To instrument for income, I use two instruments. The first is the average income of agents of the respondent's profession, measured at the two-digit level using the NYK classification (which is almost identical to the ISIC-58 standard, see Arbeidsdirektoratet (1995) for details), using the whole sample to calculate the means. The second is year specific averages of the one digit NYK classification to get more variability over time. The coding of profession was

changed in the 2001 round of the survey, so I only use data from 1977 to 1997 for the IV estimations.

<div align="center">Table 9 about here</div>

The results are shown in Table 9. The most striking result is that the estimated coefficients are numerically much larger indicating the presence of measurement error in the reported incomes. Furthermore, the increase in the parameter values seem to be as large for the random effects models as the fixed effects model, so the lower value in the fixed effects model was not driven by measurement error. Otherwise, the pattern is similar to above: Income has a large and significantly positive effect on preferences for the Conservative party in the random effects model, but the parameter becomes smaller and insignificant once fixed effects are introduced.

One objection to these results may be that the instrument is invalid as average salary of your profession may be a signal of future income. This would then be an omitted variable in both the baseline estimation and the instrumental variables approach. Call the true coefficient on salary by profession π and the bivariate regression coefficient between salary by profession on income b. In a linear model, the coefficient on income from an OLS estimation would converge to $(\beta + \pi b)\, \sigma_y^2 / \left(\sigma_y^2 + \sigma_u^2\right)$, a combination of the well known results of the effects of omitted variables and errors in variables. If we still include income as the sole explanatory variable, but instrument it with average income by salary, the estimated coefficient converge in probability to $(\beta + \pi b)$. It is probably possible to prove similar results for logit models. This show that even if the instruments should have been included as an explanatory variable, the increase in the estimated parameter values stems from the removal of measurement error. As Table 9 shows that the increase is larger for the random effects models than the fixed effects model, this indicates that the drop in the estimate for the fixed effects model in Table 3 cannot be caused by measurement error in the income variable.

7 Public policy opinions and party choice

So far I have only considered the determinants of voting. However, Section 2 dealt mostly with preferences for low taxes or much redistribution. The main reason for concentrating on party choice is a revealed preference-argument. It is probably easier to say that you favour

redistribution than to vote for a party that favours redistribution in a way that may be costly to you. Norway being an egalitarian social democracy, there may also be social norms that makes it harder for some to admit that they want low taxes and don't care about supporting the poor.

Nevertheless, it is also interesting to see how income and indicators of expectations of income changes affect opinions on economic policy. Table 10 reports results from binomial panel data logit estimations with random and fixed effects. Panel A studies the response to a question on whether the respondent favour a tax cut for those with high incomes. I distinguish between those who answer that they oppose it or oppose it strongly against those who favour it, favour it strongly, or are indifferent. Panel B reports results from a question on whether the respondent believe that social security should be expanded, versus leaving it as it is today or reducing it. Panels C and D study the same outcomes using the instrumental variables technique from Section 6.

Table 10 about here

From column (1) of Panel A, we see that in the whole sample, high income is associated with preferences for tax cuts. The coefficient is both numerically relatively large and strongly significantly different from zero. This seems to support the models outlined in Section 2. However, we notice that for the sample of agents that change their opinion on tax cuts, reported in column (2), income has an insignificant, but slightly positive effect on opposing tax cuts, contrary to the predictions of the models outlined in Section 2. This is also clear when we use the fixed effects estimator reported in column (3). Hence it seems that richer agents favour tax cuts, but an income rise is not a cause of starting to favour tax cuts, so there does not seem to be a causal relationship.

Panel B confirms to a large extent the conclusions above. For the whole sample, income has the predicted negative effect on preferences for social security extensions. The coefficient is also significantly different from zero. For the opinion changing sample, income has an negative effect, but it is now much smaller and insignificant. The same is true when we introduce a fixed effects or the average income of the Chamberlain estimator. Hence again, rich agents tend to oppose social security extensions, but probably not because they are rich in itself.

In Panel C we return to the opinions on tax cuts for the rich, now using instrumental

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variables to correct for measurement error in stated income. As in Table 9, estimated coefficients are numerically considerably larger, indicating the presence of measurement error. For the full sample, income is seen to have a strong negative effect on opposition to tax cuts. Again, if we look at those changing their opinion, the effect becomes much smaller, but the reductions seems to be less strong than above. This may indicate that a part of the result stated above was due to more measurement error in the income of those who change their opinion (in the sense of their error having larger variance). Nonetheless, it again turns out that changes in income only has a limited effect on changes in opinion on tax cuts. The conclusions seem to be about the same for Panel D where the instrumental variables estimator is employed on the stated opinion on social security extensions.

It is also interesting to see to what extent these stated opinions affects party choice, and to what extent the effect income has on party choice is through these opinions. Table 11 reports estimations when opinions on expansions of social security and tax cuts are included as explanatory variables. First we notice that the parameters on the opinion variables seem to have reasonable values. People who want tax cuts and oppose social security expansions tend to vote for the Conservative party, those with the opposite view Labour or Socialist. In most of the cases, the estimated parameters are numerically large and significantly different from zero in most cases, also when individual fixed effects are introduced. The LR-test of joint significance of all the opinion variables being irrelevant is rejected at extremely high levels of significance, even in the fixed effects model. Hence we can quite safely conclude that opinions on economic policy has an effect on voting behaviour.

Table 11 about here

From the standard political economy approaches of Section 2, a high income should be associated with favouring tax cuts and reductions of social security. Hence when we control for these, the direct effect of income should be drastically reduced. Comparing with Table 3, we see that if there is a change in the magnitude of the income term, it is rather in the opposite direction. Hence including proxies for political opinions does not reduce the effects of income on voting so it seems that income has other effects than influencing opinions on economic policy.

There may be a problem of reverse causality in these estimations. Although we usually think of agents first making up their mind on a set of issues and then deciding which party

to vote for, other scenarios are possible. It may be that some agents first choose a party, for instance because of the party's charismatic leader or the voter's upbringing. Then once the party preferences are chosen, the voter rationalizes his choice by also adhering to the party's main policies.[6] It would be very difficult to empirically distinguish this from the opinion-based voting theory. However, it is less plausible that opinion changes are due to party changes, so the fixed effects estimator should be more robust to this criticism. This may be one explanation for the numerically smaller estimates on the fixed effects estimator than the random effects estimators.

To conclude, it seems that controlling for opinions about economic policy tends to reduce the direct effect of income slightly, but less then we could expect. Hence although stated opinions have an impact on voting, income also seems to have an effect on its own.

8 Conclusion

Although there may seem to be a positive causal relationship between high income and Conservative voting as predicted by most political economy models, this may be due to factors like social background and learning about the working of the economy. In this paper I have used panel data discrete choice models to extract the pure causal effect of income by looking at income changes over a limited period of time and their impact on changes in voting behaviour. The result is that income has a causal effect on Conservative voting, but the effect is smaller in magnitude than what we find using traditional tools. The reduction in magnitude seems to be in the order of a magnitude of five.

The same is true for stated opinions on tax cuts and extension of social security. High income respondents tend to favour tax cuts and oppose social security extensions, but changes in these opinions are only to a limited degree explained by changes in income and does not survive inclusion of individual fixed effects. Opinions in favour of tax cuts and reductions in social security seems to induce Conservative voting as we would expect, but the effect of income on party choice is virtually unchanged from including these opinion variables.

Thus income has most effect on party choice and political preferences in the long run, possibly over several generations. As income levels are quite persistent, this can explain why

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[6]This line of reasoning is closely related to the concept of cognitive dissonance, see e.g. Akerlof and Dickens (1982).

high income respondents tend to vote Conservative and favour a small welfare budget. There could be several explanations for this. One is that permanent income is more relevant for preferences than transitory income (Lind 2004). As a change in annual income may not have a large impact on lifetime income, we may not expect it to have a large impact on voting. However, it is not obvious that permanent income is more important than current income, given that politics are quite volatile over time. We could also envisage a more sociological impact where a high income tends to socialize people into different groups or classes with different views and beliefs, so people vote in accordance with who they think they are. These perceptions are only changing slowly, which explains why income only has an effect in the long run. It seems that the traditional political economy models, although not proven to be wrong, only tell a part of the story. Unobserved characteristics correlated with income has almost the same impact on voting as income.

A   Additional estimation results

Table A1-A2 about here

B   Further details on the empirical model and estimation

B.1   The random effects estimator

Preferences over different parties may be described by the utility function (1). For ease of notation, let $X_{it} := \left(\begin{array}{cc} y_{it} & z'_{it} \end{array} \right)'$ and $\Theta_j := \left(\begin{array}{cc} \beta_j & \gamma \end{array} \right)$, so $v_{itj} = \alpha_{ijt} + \Theta_j X_{it} + \varepsilon_{ijt}$. From this specification, it follows that the probability of agent i choosing party j at date t is

$$p_{ijt} = \frac{\exp\left(\alpha_{ijt} + \Theta_j X_{it}\right)}{\sum_k \exp\left(\alpha_{ikt} + \Theta_k X_{it}\right)}.$$

In the random effects model, the individual effects satisfy $\left(\alpha_{i1}^I, \ldots, \alpha_{iN}^I\right)' \sim NID\left(0, \Sigma_\alpha\right)$, and the probability of choosing party j becomes

$$p_{ijt} = \int \cdots \int \frac{\exp\left(\alpha_{jt}^T + \alpha_j^P + \alpha_{ij}^I + \Theta_j X_{it}\right)}{\sum_k \exp\left(\alpha_{kt}^T + \alpha_k^P + \alpha_{kj}^I + \Theta_k X_{it}\right)} \phi\left(\alpha_{i1}^I, \ldots, \alpha_{iN}^I; \Sigma_\alpha\right) \, d\alpha_{i1}^I \ldots d\alpha_{iN}^I, \quad (3)$$

where $\phi\left(\cdot, \Sigma_\alpha\right)$ is the PDF of a multinomial normal distribution with mean zero and co-variance matrix Σ_α. This expression may be computed by simulation of the individual specific effects, i.e. using Monte Carlo integration. For each individual, we make n drawings $\left\{\tilde{\alpha}_{i1}^I \ldots \tilde{\alpha}_{iN}^I\right\}_{l=1}^n$ from the appropriate normal distribution and replace (3) by

$$\tilde{p}_{ijt} = \frac{1}{n} \sum_{l=1}^n \frac{\exp\left(\alpha_{jt}^T + \alpha_j^P + \tilde{\alpha}_{ij}^{Il} + \Theta_j X_{it}\right)}{\sum_k \exp\left(\alpha_{kt}^T + \alpha_k^P + \tilde{\alpha}_{kj}^{Il} + \Theta_k X_{it}\right)}. \tag{4}$$

As n grows, this approximation converges towards the exact value of p_{ijt}. To enhance the efficiency of the simulation, antithetic variates are used (see e.g. Train (2003, Section 9.3.1)). Adding the likelihoods over individuals and periods, we get the likelihood for the whole sample. Let v_{ijt} be a dummy taking the value 1 if individual i voted party j at time t, and 0 otherwise, and let T_i be the set of time periods in which individual i is observed. An approximation of the log likelihood for the whole sample is then

$$\sum_i \ln \frac{1}{n} \sum_{l=1}^n \prod_{t \in T_i} \prod_{j=1}^N \left(\frac{\exp\left(\alpha_{jt}^T + \alpha_j^P + \tilde{\alpha}_{ij}^{Il} + \Theta_j X_{it}\right)}{\sum_k \exp\left(\alpha_{kt}^T + \alpha_k^P + \tilde{\alpha}_{kj}^{Il} + \Theta_k X_{it}\right)} \right)^{v_{ijt}}.$$

This expression is maximized to obtain estimates of the parameters of interest Θ_j, but also the period dummies α_{jt}^T and α_j^P and the values of the Cholesky decomposition of Σ_α. As in all models of discrete choice, we can only identify the effect of a variable on the choice of an alternative relative to another alternative. Hence we restrict $\Theta_j = 0$ for one party and imposing the individual and time effects for the same alternative to be zero, so Σ_α, the covariance matrix of the individual effects, is in reality of dimension $N - 1$.

B.2 The fixed effects estimator

Now consider Chamberlain's (1980) conditional maximum likelihood estimator for fixed effects logit models. I limit the exposition to the case of two periods as this is the case we encounter below. Also, for simplicity of notation, I omit the period specific and second period effects. Assume that individual i votes for party p_1 and p_2 in period 1 and 2. The likelihood of this observation is

$$\frac{\exp\left(\alpha_{ip_1}^I + \Theta_{p_1} x_{i1}\right)}{\sum_k \exp\left(\alpha_{ik}^I + \Theta_k x_{i1}\right)} \frac{\exp\left(\alpha_{ip_2}^I + \Theta_{p_2} x_{i2}\right)}{\sum_k \exp\left(\alpha_{ik}^I + \Theta_k x_{i2}\right)}. \tag{5}$$

If $p_1 = p_2$, we can perfectly predict the agent's actions by letting $\alpha_{ip_1}^I = \alpha_{ip_2}^I \to \infty$ and $\alpha_{ik}^I \to -\infty$ for $k \neq p_1$. Hence only those who vote for different parties give any information

about the parameters of interest Θ_j. Conditional on one of the two voting sequences (p_1, p_2) or (p_2, p_1), the probability becomes

$$\frac{\Pr(p_1, p_2)}{\Pr(p_1, p_2) + \Pr(p_2, p_1)}.$$

For parties k for which he does not vote in any of the two periods, $\alpha_{ik}^I \to -\infty$. We then get the likelihood for individual i as

$$\frac{\exp(\Theta_{p_1} x_{i1} + \Theta_{p_2} x_{i2})}{\exp(\Theta_{p_1} x_{i1} + \Theta_{p_2} x_{i2}) + \exp(\Theta_{p_2} x_{i1} + \Theta_{p_1} x_{i2})}, \tag{6}$$

which is easily maximized to obtain ML estimates of the parameters.

B.3 The instrumental variables estimator

Let y_{it}^* denote agent i's (unobserved) true income and $y_{it} = y_{it}^* + u_{it}$ denote the observed income with error. The error u_{it} is assumed to be iid $(0, \sigma_u^2)$ across both i and t. The instrument w_{it} is assumed to satisfy $y_{it}^* = \pi w_{it} + v_{it}$ where $\pi \neq 0$ and v_{it} is iid $(0, \sigma_v^2)$. In the first stage we regress y_{it} on w_{it}. This gives a consistent estimate of π, so the predicted value $\hat{y}_{it}$ converges in probability to $y_{it}^* - v_{it}$ when either the sample size or the number of time periods grows. Also, the residuals $\widehat{u_{it} + v_{it}}$ converge in probability to $u_{it} + v_{it}$. The utility of voting party j is

$$\upsilon_{itj} = \alpha_{ijt} + \beta_j y_{it}^* + \gamma_j z_{it} + \varepsilon_{ijt}.$$

First, replacing y_{it}^* by $\hat{y}_{it}$ yields the equation

$$\upsilon_{itj}^{(1)} = \alpha_{ijt} + \beta_j \hat{y}_{it} + \gamma_j z_{it} + \varepsilon_{ijt} \xrightarrow{p} \upsilon_{itj} + \beta_j v_{it}. \tag{7}$$

Here, $\beta_j v_{it}$, is an unobserved nuisance term. Although it is uncorrelated with $\hat{y}_{it}$, its variance tends to give a downward bias in the estimates, particularly if the correlation between y_{it}^* and w_{it} is low. To partially remedy this, I suggest to include $\widehat{u_{it} + v_{it}}$ as an additional regressor. This gives the model

$$\upsilon_{itj}^{(2)} = \alpha_{ijt} + \beta_j \hat{y}_{it} + \check{\beta}_j(\widehat{u_{it} + v_{it}}) + \gamma_j z_{it} + \varepsilon_{ijt} \xrightarrow{p} \upsilon_{itj} - \check{\beta}_j u_{it} + \left(\beta_j - \check{\beta}_j\right) v_{it}. \tag{8}$$

Now the nuisance term becomes $-\check{\beta}_j u_{it} + \left(\beta_j - \check{\beta}_j\right) v_{it}$, which has a lower variance than $\beta_j v_{it}$. To see this, notice that $\check{\beta}_j \xrightarrow{p} \frac{\sigma_v^2}{\sigma_u^2 + \sigma_v^2} \beta_j$ so the variance of the nuisance term in (8) approaches $\frac{\sigma_u^2 \sigma_v^2}{\sigma_u^2 + \sigma_v^2} \beta_j^2$ whereas the variance of the nuisance term in (7) approaches $\sigma_v^2 \beta_j^2 \geq \frac{\sigma_u^2 \sigma_v^2}{\sigma_u^2 + \sigma_v^2} \beta_j^2$.

Particularly, the nuisance term in (8) disappears if either there is no measurement error, i.e. $\sigma_u^2 = 0$, or the instrument is perfect, i.e. $\sigma_v^2 = 0$. Nevertheless, the unobserved nuisance term leads to a downward bias in this case as well.

Notice that we can estimate $\sigma_u^2 + \sigma_v^2$ as the variance of $(\widehat{u_{it} + v_{it}})$ and $\frac{\sigma_v^2}{\sigma_u^2 + \sigma_v^2}$ as the ratio $\beta_j / \check{\beta}_j$ and hence derive estimates of σ_u^2 and σ_v^2. This may help us to get an idea of the bias of the estimates. However, knowledge of the variance of the nuisance term in (8) is probably not sufficient to get the exact bias. Stefanski and Buzas (1995) discuss some possible adjustments based on knowledge of these variances, but their simulations indicate that the improvement in the accuracy of the estimator is not worth the increase in complexity of estimation, so I omit this from the present work.

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Table 1
Descriptive statistics for income

Year	Number of obsevations	1000 1997 kroner	Relative to average	μ	σ
1977	473	251.62	1.07	4.13	0.52
		(145.41)	(0.62)	(0.01)	(0.01)
1981	1125	265.38	1.08	4.52	0.55
		(159.65)	(0.65)	(0.01)	(0.01)
1985	1342	267.18	1.09	4.89	0.51
		(156.75)	(0.64)	(0.01)	(0.01)
1989	1270	322.42	1.10	5.26	0.61
		(214.17)	(0.73)	(0.01)	(0.01)
1993	1137	354.35	1.00	5.44	0.77
		(196.03)	(0.55)	(0.002)	(0.001)
1997	1114	401.51	1.00	5.62	0.78
		(221.77)	(0.55)	(0.002)	(0.001)
2001	557	451.72	1.00	5.85	0.73
		(374.97)	(0.83)	(0.003)	(0.001)
Total	7018	325.93	1.05		
		(218.21)	(0.65)		

Sample averages of income and income relative to year average by year, with standard deviations in parenthesis. Numbers up to 1989 are based on 100 simulations. Parameters μ and σ are estimated parameters based on a log normal distribution, used to simulate incomes throughout the paper. Results for 1993 to 2001 are only included for reference. Standard errors of the estimates in parenthesis.

Table 2
Descriptive statistics for party choice

Year	Socialist	Labour	Centre	Conservative	Progress	Total
1977	29	208	110	122	4	473
	6.1	44.0	23.3	25.8	0.9	100
	4.6	42.3	21.2	24.5	1.9	94.5
1981	63	437	233	353	39	1,125
	5.6	38.8	20.7	31.4	3.5	100
	6.0	37.1	17.0	31.8	4.5	96.4
1985	90	505	286	419	42	1,342
	6.7	37.6	21.3	31.2	3.1	100
	6.1	40.8	18.5	30.4	3.7	99.5
1989	163	446	257	280	124	1,270
	12.8	35.1	20.2	22.1	9.8	100
	10.1	34.3	18.2	22.2	13.0	97.8
1993	103	470	322	192	50	1,137
	9.1	41.3	28.3	16.9	4.4	100
	9.0	36.9	28.2	17.0	6.3	97.4
1997	95	404	328	181	106	1,114
	8.5	36.3	29.4	16.3	9.5	100
	7.8	35.0	26.1	14.3	15.3	98.5
2001	84	132	148	145	48	557
	15.1	23.7	26.6	26.0	8.6	100
	13.8	24.3	21.9	21.2	14.6	95.8
Total	627	2,602	1,684	1,692	413	7,018
	8.9	37.1	24.0	24.1	5.9	100

First row is number of observations in sample, second year-specific percentage, and third result of elections.

Table 3
Effect of relative income on party choice

	(1) Random effects	(2) Random effects	(3) Fixed effects
Socialist	-0.269	-0.106	0.014
	(0.202)	(0.175)	(0.357)
Centre	-0.615	-0.077	-0.114
	(0.207)	(0.150)	(0.158)
Conservative	1.090	0.361	0.191
	(0.184)	(0.145)	(0.251)
Progress	0.123	-0.033	-0.305
	(0.233)	(0.190)	(0.378)
Log likelihood	-7560.0	-2556.1	-463.2
Observations	3509	890	890
Time frame		1977-2001	
Sample	Full	Changers	Changers

Values are coefficient on income measured relative to period averages. Standard errors in parenthesis. Control variables are age, two categories of education, and dummy variables for sex, marital status, student, retired, staying at home, employed in the public sector, having zero income, and year. Sex and age are omitted for the fixed effects model. Estimation is by simulated maximum likelihood with 100 simulations per observation.

Table 4
Stability of the income parameters over time

		(1) Random effects	(2) Random effects	(3) Fixed effects
One cut				
Socialist	Income	-0.467	-0.157	0.183
		(0.247)	(0.193)	(0.453)
	Income ×	0.439	0.140	-0.242
	before '85	(0.363)	(0.343)	(0.619)
Centre	Income	-1.250	-0.323	-0.220
		(0.270)	(0.178)	(0.229)
	Income ×	1.462	0.732	0.803
	before '85	(0.355)	(0.280)	(0.436)
Conservative	Income	0.625	0.162	0.032
		(0.233)	(0.160)	(0.319)
	Income ×	1.218	0.608	0.680
	before '85	(0.329)	(0.270)	(0.463)
Progress	Income	-0.358	-0.210	-0.520
		(0.282)	(0.209)	(0.447)
	Income ×	1.306	0.606	0.818
	before '85	(0.427)	(0.388)	(0.629)
log L		-7549.79	-2551.67	-460.59
$\chi^2(4)$		20.35	8.95	5.19
		$[4.27 \times 10^{-4}]$	[0.06]	[0.27]
Full specification				
Log likelihood		-7528.57	-2533.92	-448.02
$\chi^2(24)$		62.79	44.46	30.33
		$[2.55 \times 10^{-5}]$	[0.01]	[0.17]
Time frame			1977-2001	
Sample		Full	Changers	Changers
Observations		3509	890	890

Figures in the first panel are coefficients on income relative to period average and income interacted with a dummy for being observed between 1977 and 1985. Control variables are as in Table 3. $\chi^2(4)$ is a LR test for structural break, comparing estimates with Table 3. Standard errors are in parenthesis, p-values in square brackets.
The second panel reports log likelihood for the specification with separate income parameters for each year. $\chi^2(24)$ is a LR test for the parameters being different, comparing estimates with Table 3. Estimation is by simulated maximum likelihood with 100 simulations per individual.

Table 5
Differences between increasing and decreasing income

		(1) Random effects	(2) Random effects	(3) Fixed effects
Socialist	Income	-0.170	-0.183	-0.485
		(0.230)	(0.191)	(0.629)
	Income × Income up	-0.180	0.172	1.027
		(0.212)	(0.159)	(0.945)
Centre	Income	-0.517	-0.135	0.228
		(0.246)	(0.164)	(0.509)
	Income × Income up	-0.156	0.124	-0.392
		(0.261)	(0.137)	(0.571)
Conservative	Income	1.246	0.394	0.621
		(0.240)	(0.156)	(0.550)
	Income × Income up	-0.240	-0.072	-0.613
		(0.240)	(0.140)	(0.736)
Progress	Income	0.427	0.016	0.345
		(0.273)	(0.204)	(0.656)
	Income × Income up	-0.647	-0.114	-1.227
		(0.281)	(0.189)	(1.040)
$\chi^2(4)$		6.39	3.52	3.42
		[0.17]	[0.48]	[0.49]
Log likelihood		-7556.8	-2554.3	-461.5
Observations		3509	890	890
Time frame			1977-2001	
Sample		Full	Changers	Changers

Figures are coefficients on income relative to period average and this variable interacted with a dummy for income increasing between the two periods of observation. Control variables are as in Table 3. $\chi^2(4)$ is an LR test on all four interaction terms being zero. Standard errors are in parenthesis, p-values in square brackets. Estimation is by simulated maximum likelihood with 100 simulations per individual.

Table 6
Differences between party changers and party stayers

		Random effects
Socialist	Income	-0.856
		(0.201)
	Income × changer	1.522
		(0.199)
Centre	Income	-0.995
		(0.216)
	Income × changer	1.084
		(0.234)
Conservative	Income	0.856
		(0.183)
	Income × changer	0.566
		(0.202)
Progress	Income	-0.829
		(0.277)
	Income × changer	2.219
		(0.304)
$\chi^2(4)$		61.0
		$[1.82\times10^{-12}]$
Log likelihood		-7529.5
Observations		3509
Time frame		1977-2001
Sample		Full

Figures are coefficients on income relative to period average and this variable interacted with a dummy for voting for different parties in the two periods of observation. Control variables are as in Table 3. χ^2 is an LR test on all four interaction terms being zero. Standard errors are in parenthesis, p-values in square brackets. Estimation is by simulated maximum likelihood with 100 simulations per individual.

Table 7
Fraction changing party

Party	Fraction changing from	Fraction changing to
Socialist	0.41	0.42
Labour	0.18	0.17
Centre	0.23	0.26
Conservative	0.26	0.23
Progress	0.51	0.57
Total	0.25	0.25

Fraction of voters who report voting for different parties in the two periods by voting pattern in the first and second period.

Table 8
Characteristics of party changers and party stayers

	Averages			Logit
	Stayers	Changers	Difference	
Income	1.05	1.06	0.004	-0.093
	(0.014)	(0.025)	(0.029)	(0.077)
Change in income	-0.01	0.02	-0.02	-0.029
	(0.011)	(0.024)	(0.024)	(0.074)
Absolute value of	0.36	0.40	-0.04	0.135
change in income	(0.008)	(0.020)	(0.019)	(0.088)
Age	48.79	43.62	5.17	-0.022
	(0.291)	(0.477)	(0.571)	(0.003)
Female	0.48	0.46	0.01	-0.004
	(0.010)	(0.016)	(0.019)	(0.079)
Low education	0.42	0.37	0.06	-0.105
	(0.009)	(0.016)	(0.019)	(0.170)
Middle education	0.23	0.20	0.03	-0.064
	(0.008)	(0.013)	(0.016)	(0.185)
High education	0.35	0.43	-0.09	0.119
	(0.009)	(0.016)	(0.018)	(0.156)

Average values for certain descriptive statistics for voters reporting voting for the same party and different parties in the two periods. Variables are measured in the last period of observation, except changes in income. Forth column is results from a logit model where all variables are combined to explain party change. Standard errors in parenthesis.

Table 9
Instrumental variables estimation of the effect of income on party choice

	(1) Random effects	(2) Random effects	(3) Fixed effects
Socialist	0.861	0.149	-0.086
	(0.740)	(0.775)	(1.503)
Centre	0.143	0.443	-0.125
	(0.826)	(0.644)	(1.309)
Conservative	5.366	1.976	0.529
	(0.872)	(0.673)	(1.385)
Progress	2.741	1.236	-1.357
	(0.934)	(0.799)	(1.746)
Log likelihood	-5932.6	-1943.5	-353.9
Observations	2848	681	681
Time frame		1977-1997	
Sample	Changers	Changers	Changers

Figures are coefficients on income relative to period average. Instruments for income are average income of profession at two-digit NYK level and yearly averages at one-digit NYK. Control variables are the same as for Table 3. Standard errors are in parenthesis.
Estimation is by two step simulated maximum likelihood with 100 simulations per observation

Table 10
Determinants of opinions on economic policy

		(1) Random effects	(2) Random effects	(3) Fixed effects
A	Against tax cuts	-0.700	-0.066	-0.071
	Ordinary logit	(0.080)	(0.097)	(0.125)
	Log likelihood	-3400.6	-1201.9	-592.8
	Observations	2846	896	896
B	For social security extension	-0.261	-0.023	-0.045
	Ordinary logit	(0.088)	(0.109)	(0.151)
	Log likelihood	-3076.9	-1029.6	-496.7
	Observations	2846	853	853
C	Against tax cuts	-1.470	-0.427	-0.687
	Instrumental variables	(0.295)	(0.453)	(0.650)
	Log likelihood	-3078.28	-1097.28	-539.096
	Observations	2642	812	812
D	For social security extension	-0.382	-0.220	-0.301
	Instrumental variables	(0.325)	(0.479)	(0.712)
	Log likelihood	-2977.16	-952.654	-459.98
	Observations	2642	816	816
	Time frame		1977-1997	
	Sample	Full	Changers	Changers

Estimates are coefficient on income measured relative to period averages. Standard errors in parenthesis.
Control variables are as in Table 3, instruments as in Table 9.
Variable "Against tax cut" is one for respondents who answered that they oppose or oppose strongly tax cuts for high incomes and zero for respondents who answer that they favour it, favour it strongly, or don't know.
Variable "For social security extension" is one for respondents who answered that they favour or favour strongly extensions of social security, zero for respondents who answered that they oppose it, oppose it strongly, or don't know.
Estimation is by simulated maximum likelihood with 100 simulations per observation

Table 11
Effect of opinions on economic policy on voting

		(1) Random effects	(2) Random effects	(3) Fixed effects
Socialist	Income	0.131	0.221	0.354
		(0.224)	(0.197)	(0.419)
	Extend social security	0.812	0.418	0.076
		(0.204)	(0.206)	(0.325)
	Against tax cut	0.452	0.413	-0.241
		(0.230)	(0.242)	(0.362)
Centre	Income	0.132	0.179	0.259
		(0.250)	(0.173)	(0.334)
	Extend social security	-0.425	-0.153	-0.034
		(0.226)	(0.188)	(0.298)
	Against tax cut	-1.071	-0.504	-0.814
		(0.227)	(0.184)	(0.306)
Conservative	Income	1.821	0.665	0.442
		(0.257)	(0.172)	(0.340)
	Extend social security	-0.941	-0.461	-0.201
		(0.249)	(0.209)	(0.359)
	Against tax cut	-2.122	-0.941	-0.741
		(0.237)	(0.192)	(0.326)
Progress	Income	0.497	0.028	-0.203
		(0.281)	(0.219)	(0.484)
	Extend social security	-0.989	-0.287	0.212
		(0.286)	(0.247)	(0.456)
	Against tax cut	-1.903	-1.098	-0.975
		(0.264)	(0.226)	(0.417)
$\chi^2(8)$		130.34	69.68	0.16
		[2.41×10^{-24}]	[5.68×10^{-12}]	[1.00]
logL		-5869.4	-1933.9	-349.4
Observations		2846	689	689
Time frame			1977-1997	
Sample		Stayers	Changers	Changers

Numbers are coefficients on income relative to period average and dummies for favouring or strongly favouring expanding social security and opposing or opposing strongly tax cuts for high incomes. Control variables are as in Table 3. Standard errors in parenthesis.
χ^2 is the test of all parameters on opinions being zero. p-values in brackets.
Estimation is by simulated maximum likelihood with 100 simulations per observation.

Table A1
Effect of income in constant prices on party choice

	(4) Random effects	(1) Random effects	(3) Fixed effects
Socialist	-0.273	-0.131	-0.107
	(0.094)	(0.074)	(0.166)
Centre	-0.577	-0.159	-0.065
	(0.091)	(0.061)	(0.084)
Conservative	0.110	0.005	0.034
	(0.038)	(0.035)	(0.137)
Progress	-0.215	-0.096	-0.234
	(0.081)	(0.074)	(0.186)
Log likelihood	-7575.0	-2557.8	-462.5
Observations	3509	890	890
Time frame		1977-2001	
Sample	Full	Changers	Changers

Values are coefficient on income measured in 100 000 1997-NOK. Standard errors in parenthesis. Control variables are as in Table 3. Estimation is by simulated maximum likelihood with 100 simulations per observation.

Table A2
Marginal effect of income on choice probabilities

	(1) Random effects	(2) Random effects	(3) Fixed effects
Socialist	-0.040	-0.015	0.001
	(0.016)	(0.013)	(0.029)
Centre	-0.178	-0.032	-0.030
	(0.032)	(0.023)	(0.028)
Conservative	0.258	0.073	0.050
	(0.032)	(0.022)	(0.041)
Progress	0.002	-0.006	-0.017
	(0.011)	(0.009)	(0.019)
Log likelihood	-7560.0	-2556.1	-463.2
Observations	3509	890	890
Time frame		1977-2001	
Sample	Full	Changers	Changers

Numbers are the estimated marginal effect of income relative to period averages on choice probabilities. Marginal effects are calculated at sample means. Standard errors in parenthesis. Control variables are as in Table 3. Estimation is by simulated maximum likelihood with 100 simulations per observation.