An Asset Theory of Social Policy Preferences

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Abstract

This paper presents a theory of social policy preferences that emphasizes the composition of people's skills. The key to our argument is that individuals who have made risky investments in skills will demand insurance against the possible future loss of income from those investments. Because the transferability of skills is inversely related to their specificity, workers with specific skills face a potentially long spell of unemployment, or a significant decline in income, in the event of job loss. Workers deriving most of their income from specific skills therefore have strong incentives to support social policies that protect them against such uncertainty. This is not the case for general skills workers for whom the costs of social protection will weigh more prominently. We test the theory on public opinion data for 11 advanced democracies, and suggest how differences in educational systems can help explain cross-national differences in the level of social protection.

Introduction

It is a well known fact that human capital rivals physical capital as a source of personal and national wealth. Indeed, it is the single most important determinant of personal income in advanced industrialized countries. Yet, whereas physical assets -- buildings, machinery, goods, and money -- have long been recognized as essential for understanding the political interests of their owners, surprisingly little is known about the role of human capital in explaining public policy preferences. With the exception of trade policy,[1] it is only the *cognitive* aspects of education that have received systematic attention in explaining political preferences (Klingemann 1979; Kitschelt 1991; Duch and Taylor 1993).

Following Becker (1964), we conceptualize human skills as an *investment*, and ask how the character of this investment affects workers' preferences for social protection. We approach this question in a fashion that is similar to the way transaction cost economics explains the use of non-market institutions to overcome market failures (Williamson 1985). In a political version of this logic, endogenous trade theory hypothesizes that investments in physical assets that are specific to a particular location or economic transaction lead firms to lobby the state for protection against uninsurable risks (see Alt et al. 1999).[2] Since pulling out assets in response to adverse market conditions is difficult, firms will want protection against the effects of such conditions. We start from the similar idea that investment in skills that are specific to a particular firm, industry or occupation exposes their owners to risks for which they will seek non-market protection. Skills that are portable, by contrast, do not require extensive non-market protection, just as the exchange of homogeneous goods does not require elaborate non-market governance structures.[3]

Our theory does not necessarily contradict a long tradition in the study of the welfare state that emphasizes redistribution as a key political motive behind the welfare state (e.g., Korpi 1983; Esping-Andersen 1990). Indeed, Meltzer and Richard's (1981) influential median voter result for government spending, which focuses on the redistributive aspect of social protection, emerges as a special case in our model. Given a particular composition of skills, workers with higher income are likely to demand less social protection than workers with low income. Our argument parts ways with the Meltzer-Richard model, however, because we explicitly recognize that social protection also has an insurance aspect (Sinn 1995, Moene and Wallerstein 1999), and that demand for insurance varies between workers according to their degree of exposure to labor market risks (Baldwin 1992). Critically, in our model exposure to risk is inversely related to the portability of workers' skills. We show that this proposition has strong support in public opinion data from 11 advanced democracies, with important implications for explaining differences in the level of social protection across countries.

The remainder of this paper is divided into three sections. In the first we present the model and its main empirical implications. In the second we test these implications on public opinion data from 11 OECD countries. The concluding section discusses the broader implications of the model for explaining differences in social protection across countries.

The Model

Assumptions

Workers derive their income from skills that can be either general or specific. Specific skills are skills that are valuable only to a single firm or to a group of firms (whether an industry or a sector), whereas general skills are portable across all firms. We distinguish between three different employment situations, or states of the world, each associated with distinct levels of income. In *State I* a worker is employed in a firm that utilizes both his specific and general skills; in *State II* the worker is employed in a firm that only utilizes his general skills; and in *State III* the worker is unemployed (i.e., none of his skills are being utilized).

We define by g the market value of a worker's general skills in *State 2* when his or her specific skills are not being used. In *State I* (when the specific skills are being used as well) the worker is paid sg, the value of his or her combined specific and general skills. If a worker has no specific skills, then $s=1$,and she is always employed at the market value of her general skills. The key assumption is that general skills are marketable in all sectors of the economy, whereas specific skills are only marketable in one sector (the size of which is defined by the specificity of skills).

In addition to market income, which includes both wage and non-wage compensation, workers receive transfer income from the government, hereunder unemployment benefits, health care benefits, pensions, and other forms of non-wage compensation. Although some of these benefits are received by people who are outside the labor market, what matters to our argument is that they are viewed by workers as part of their compensation (what in the neo-corporatist literature is sometimes referred to as the "social wage"). Those who are most fearful of losing the labor market power of their skills, and hence their ability to secure good health and pension plans through their employer, will also be most concerned about guaranteeing a high level of benefits, even if the benefits are "deferred" to the future.

We assume that transfers come in the form of a flat-rate payment, *R*, which incorporates the idea in the Meltzer-Richard model that there is a redistributive aspect to social protection.[4] Following the terminology in Estevez-Abe et al. (2001), one can distinguish between transfers that go to support the income of employed workers, *wage protection*, and transfers that go to the unemployed, *unemployment protection*.[5] In the development of the model we will discuss what happens if *R* only goes to unemployment protection. But in the main model we will assume that all workers receive the same flat-rate subsidy, which may simply be referred to as *income protection*.

Transfers are paid out of a flat-rate tax (t) on all wages. Total per capita receipts are T, and all receipts are spent on transfers (i.e., we assume balanced budgets). As in the Meltzer-Richard model, taxation is assumed to create work disincentives, captured here by the following simple labor supply function:

(1) $$l(t) = 1/(1+t),$$

where $l(t)$ is the number of hours worked or the intensity of effort (the particular form of this function is chosen for mathematical convenience). Define w as average hourly pre-tax earnings. Then tax income per capita is

(2) $$T = t \cdot w \cdot l(t) = \frac{t \cdot w}{1+t} = R.$$

Figure 1 illustrates the three states of labor market, and shows the disposable (after tax) income associated with each state: $I\colon \overline{s}\overline{g},\ III\colon R,\ and\ II\colon \overline{g}.$

[Figure 1 about here]

For a given period of time, there is a probability, *p*, of losing one's job, and another probability, *q*, of re-employment. In equilibrium $p \cdot e = q \cdot g$, where *e* is the share of employed workers and g is the share of the workforce that is unemployed $(e = 1 - g)$. This implies that in equilibrium *e*=*q*/(*p*+*q*). Furthermore, if *r* is the probability that an employed worker is in *State I* (i.e., is in a job where both general and specific skills are utilized), then the equilibrium share of the labor force employed in *State I* is

(3) $$a = \mathrm{r} \cdot \mathrm{e} = \mathrm{r} \cdot \mathrm{q}/(\mathrm{p} + \mathrm{q}).$$

Likewise, the share of the labor force employed in *State II* is

(4) $$b = (1 - \mathrm{r}) \cdot \mathrm{e} = (1 - \mathrm{r}) \cdot \mathrm{q}/(\mathrm{p} + \mathrm{q}),$$

while the share of the labor force in *State III* (unemployment) is

(5) $$g = \mathrm{p}/(\mathrm{p} + \mathrm{q}).$$

For any individual worker with both specific and general skills, the proportions a, b, and g can be interpreted as probabilities in a lottery with three possible outcomes. An employed *sg*-worker will therefore seek to maximize the expected utility of income across all three states. Ignoring the discounting of future income (which makes no substantive difference to our results), this is captured by the following utility function:

(6) $$V = a \cdot u(\overline{s}\overline{g}) + b \cdot u(\overline{g}) + g \cdot u(R),$$

where *u*(.) is the worker's utility from net income, which for simplicity we assume is spent on consumption. Using standard assumptions, we impose the following constraints on *u*:

(7) $$u_c > 0, \quad u_{cc} < 0, \text{ and} \quad \lim_{c \to 0} u'(c) = \infty.$$

A number of the results below hold for this general form of utility function (notably the Meltzer-Richard results). However, since the insurance function of the social wage will play an important role and since we then need specific conditions on risk aversion, we use the standard assumption of a constant Arrow-Pratt relative risk aversion (*RRA*) utility function. Specifically,

$$\text{(7a)}\qquad \begin{aligned} u(c) &= \frac{c^{1-a}}{1-a} \quad \forall\, a > 0, \neq 1 \\ &= \log c \quad \mathit{for} \quad a = 1 \end{aligned}$$

With these assumptions in mind, we can now determine workers' utility-maximizing preferences for social protection.

Optimizing social preferences

The logic of the presentation in this section is as follows. We first consider a simple base-line model with no insurance effects, no tax disincentives, only general skills, and no unemployment (section i). We then introduce tax disincentives to get the Meltzer-Richard result (section ii), and subsequently add insurance effects (and unemployment) to get the Moene-Wallerstein result (section iii). Finally, we show what happens to the demand for social protection when the composition of skills is allowed to vary (section iv). To keep the presentation simple, all proofs are put in an appendix.

(i) *No insurance effects, no disincentive effects: the "t=1" model*. In solving workers' maximization problem we begin by assuming a labor force with only general skills (s=1), and no unemployment (e=1). The simplest case is where there are no tax disincentive effects on the number of hours supplied (so that $l(t)=1$ rather than $l(t)=1/(1+t)$ as we shall subsequently assume).

When s=1, e=1, and $l(t)$=1 equation (6) reduces to:

$$\text{(8)}\qquad \begin{aligned} V &= u\big((1-t)g + tw\big) = u\big(g(1-R/w) + R\big) \\ &= \frac{1}{1-a}\cdot\big(g(1-R/w)+R\big)^{1-a} \end{aligned}$$

We want to choose R to maximize V, where R is bounded between R=0, corresponding to t=0, and R=w, corresponding to t=1. Since

$$V_R = (g(1-R/w)+R)^{-a}\cdot(1-g/w)$$

and, since $0\le R\le w$, $g>w$ implies that V_R is uniformly negative for all values of R and hence t: maximization of V therefore requires $t=0$. Thus voters with above average income will choose a zero tax rate. And analogous argument for $g<w$ implies that voters with below average income will choose the maximum tax rate of 100%. This is the standard result that in the absence of insurance functions and tax disincentives, voters will want the maximum R (i.e., t=1) if $g<w$ and a zero R (t=0) if $g>w$. Thus, if the median voter, M, has an income less than the average income of w, the median voter will always vote for a maximum tax rate. The result is illustrated in Figure 2, *panel a*.

[Figure 2 about here]

(ii) <u>*Disincentive effects, no insurance effects: the Meltzer-Richard model*</u>. If we now include the tax disincentive effect that $l(t)=1/(1+t)$, we have:

(9)
$$\begin{aligned} V &= u\left(\frac{1-t}{1+t}g+\frac{tw}{1+t}\right)=u\big(g(1-2R/w)+R\big)\\ &=\frac{1}{1-a}\cdot\big(g(1-2R/w)+R\big)^{1-a}\end{aligned}$$

This implies

(10) $$V_R = (g(1-2R/w)+R)^{-a}\cdot(1-2g/w)$$

so that in the Meltzer-Richard model, only voters with a g-level below that of half the average hourly wage ($g=w/2$) will vote for a maximum tax rate. As illustrated in Figure 2, *panel b*, if the median voter has a g-level above $w/2$ he will therefore not vote for the maximum tax rate. Because of the simplicity of our tax disincentive function, voters with g levels below $w/2$ will vote for t=1 and voters with g levels above $w/2$ will vote for t=0.[6] With the more complex tax disincentive function used by Meltzer-Richard, workers with income in the range [w/2, w] will prefer taxation up to the point where the benefits to them from redistribution are exactly outweighed by the efficiency costs of tax disincentives. If the median voter is in this range, as the Meltzer-Richard model assumes, then she or he may vote for a positive tax rate less than 1 (as illustrated in Figure 2).

One of the implications of the Meltzer-Richard model is that voter turnout will be positively related to the level of government transfers because non-voting tends to be concentrated among low-income people (Lijphart 1997). There is some cross-national evidence for this proposition (see Franzese 1998). On the other hand, there is little empirical support for another key implication of the model, developed by Alesina and Rodrik (1994), namely that relatively inegalitarian societies will exhibit greater pressures for redistributive spending than relatively egalitarian ones (see Perotti 1996 for a review of the evidence). Among advanced countries the relationship is actually the reverse (Bénabou 1996).

(iii) *Disincentive effects, insurance effects: the Moene-Wallerstein model*. The Moene-Wallerstein model offers one possible explanation for this puzzle, which results from the introduction of insurance effects. For insurance effects to matter, we need at least two states of the world. In a simple version of the Moene-Wallerstein model (Moene and Wallerstein 1999), it is assumed that workers can either be employed at a gross wage equal to their “tax-incentivised” skill level $g/(1+t)$ or be unemployed. There are no specific skills (s=1), so equation (6) becomes:

(11) $$V = b \cdot u(\overline{g}) + g \cdot u(R)$$

Moene-Wallerstein show that if relative risk aversion is greater than unity, workers will choose a higher tax rate as they become wealthier: in other words, their aversion to risk outweighs the increased cost to them of insurance as their income increases.

To get the Moene-Wallerstein results on risk aversion, we assume that $\overline{g} = g \cdot (1 - 2R / w)$ so that R is only paid to those who are unemployed. It can then be shown (as we do in Appendix A) that

(12) $dR/dg < 0 \quad iff \quad RRA < 1$, and

(13) $dR/dg > 0$ *iff* $RRA > 1.$

Put differently, when risk-aversion is high (*RRA*>1), and if all transfers go to the unemployed, the relationship between income (*g*) and the preferred level of social protection is positive (see Figure 2, *panel c*). A key implication of this result is that, contrary to the Meltzer-Richard model, a means-preserving increase in inequality will *reduce* the median voter's preferred level of social protection (provided that the income distribution is skewed to the right). The reason is that such a rise in inequality lowers the income of the median voter, and since the insurance motive dominates the redistribution motive (*RRA*>1), demand for social protection will decline. In the Meltzer-Richard model there is no insurance motive, so a fall in the income of the median voter always leads to a rise in the demand for social protection. Risk aversion thus poses one potential solution to the empirical puzzle of why income equality is linked to redistributive spending.

Yet, despite the neatness of this result, our econometric estimations clearly reject its implication that people prefer less redistribution at lower levels of income. This leaves the negative relationship between redistribution and inequality as an important unsolved puzzle for comparative political economy. We discuss below how our distinction between specific and general skills permits an alternative and more plausible interpretation.

(iv) *Disincentive effects, insurance effects, specific and general skills: the asset model.* This is the most general model and requires us to consider all three states in Figure 1. We therefore return to the present value of utility given by equation (6):

(6) $$V = a \cdot u(\overline{sg}) + b \cdot u(\overline{g}) + g \cdot u(R)$$

where

$$\overline{sg} = s \cdot g \cdot (1 - 2R/w) + R$$
$$\overline{g} = g \cdot (1 - 2R/w) + R$$

In addition, we now introduce an important variable, expected hourly income before taxes and transfers, *y*. This is simply defined as:

$y \equiv a \cdot sg + b \cdot g$

We then proceed to ask, first, up to what value of *y* is the chosen *R* maximal, i.e. $t = 1$; second, under what *RRA* conditions does *R* fall or rise as *y* rises above this value; and, third, what happens to the choice of *R* as the balance of general and specific skills changes, holding *y* constant?

In the first result we show that a worker will only choose the maximum tax rate if $y \leq w/2$. Stated formally (the proof is in Appendix A):

Result I: *Given the assumptions of Model (iv),* $t = 1$ *iff* $y \leq \frac{w}{2}$

With this property established, we now consider what happens to *R* when *y* increases. As we show in Appendix A, this yields the following result:

Result II: *Given the assumptions of Model (iv), holding s constant and with y > w/2,*

$$\operatorname{sgn}\frac{\partial R}{\partial y} = \operatorname{sgn}\left[RRA(\bar{s}\bar{g}) - \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2} \right].$$

What this equality says is that the direction of the relationship between *R* and income (the sign, sgn, of $\frac{\partial R}{\partial y}$) depends on the level of risk-aversion, just as in the simple Moene-Wallerstein model. However, for income to be positively related to support for spending the *RRA* requirement is more stringent ($RRA > \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$) than before (*RRA*>1). The reason is that *R* now goes to the employed as well as to the unemployed, and since employed workers in the Moene-Wallerstein model only have an insurance incentive in relationship to unemployment, *RRA* must be higher for the insurance motive to dominate the redistribution motive. This implication is also demonstrated by Moene and Wallerstein.

Now we come to the critical result which differentiates our approach from previous ones. Central to the argument of the paper is the proposition that an increase in specific skills relative to general skills, holding constant the level of expected income, implies an increase in preferred *R*: put broadly, workers with specific skills will prefer higher taxes and social protection than workers with general skills. The following result is also proved in Appendix A:

Result III: *Assuming a constant relative risk aversion utility function and RRA>0,* $\frac{\partial R}{\partial s} > 0$ *holding y constant.*

In other words, as *s* rises, the preferred level of *R* also rises. The intuition behind this key result is that workers with specific skills have more to fear if they lose their job than workers with general skills. This is because specific skill workers who are laid off face the risk of being re-employed in a sector where their skills are not needed. If this happens they will lose some of their previous income, including employer-provided insurance against illness and old age. General skill workers do not face this problem because they are always compensated at the value of their general skills. Hence, the more income derived from specific as opposed to general skills – that is, the higher the ratio *s*/*g* – the greater the demand for income protection (*R*). The logic is illustrated in Figure 2, *panel d*, and implies that the median voter's support for social protection depends on the composition of his or her skills.

Summarizing the results in this section: with the simplest set of assumptions – only one state of the world (employment), only general skills, and no tax disincentives – the politics of social spending is all about redistribution (class politics if you will): those with a wage below the mean will want a maximum rate of taxation (*t*=1) whereas those above the mean will want zero taxation. If we add tax disincentives, however, the cost of redistribution may deter those low-income workers closest to the mean from demanding confiscatory taxation, and the median voter is likely to be among those workers. This is the Meltzer-Richard model.

When an unemployment state is added to the model, an entirely new motive enters into workers' calculations of their interests: insurance against loss of income. If workers are sufficiently risk-averse, and if all transfers go to the unemployed, rising income may in fact be associated with *higher* demand for protection since high-income workers have more to lose than low-income workers. This is the Moene-Wallerstein model. If some transfers go to the employed, however, the threshold of risk-aversion for which this relationship holds goes up since transfers to the employed only serve redistributive purposes.

Finally, when differences in the specificity of skills are introduced, which require at least two employment states (Sector *I* and *II* in our model), the insurance motive plays a crucial role *even* when workers are only moderately risk-averse (0<*RRA*<1) and *even* when transfers are distributed to both employed and unemployed workers. The reason is that employed workers risk losing the income from their specific skills, regardless of their exposure to unemployment. This coupling between skills and demand for insurance thus transforms the relationship between income and social policy preferences. The next section explores whether this proposition is supported using empirical evidence from public opinion surveys in 11 advanced democracies.

Testing the Model

Statistical Model

In the last section we showed in our Model IV that the relationship between the "preferred" level of *R* and the two exogenous variables *y* (expected income) and *s* (skill specificity) is given by the implicit equations:

$$(14) \qquad \begin{aligned} &V_R(R,\ s,\ g) = 0 \\ &y = a \cdot s \cdot g + b \cdot g \end{aligned}$$

And from (14) we derived Result II that

$$\frac{\partial R}{\partial y} < 0 \quad if \quad 0 < RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$$

and Result III

$$\frac{\partial R}{\partial s} > 0 \quad if \quad 0 < RRA.$$

We show in Appendix B that

$$R = K + \frac{\partial R}{\partial y} \cdot y + \frac{\partial R}{\partial s} \cdot s$$

is the first-order Taylor expansion of (14). Thus our regressions take the form

(15) $$R = k + b \cdot y + c \cdot s.$$

By implication if our estimate of *b* is significantly different from zero and negative, we can infer that $0 < RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$. If *c* is significantly different from zero and positive, 0 < *RRA*, so that skill specificity increases the demand for social protection. This is our main argument and hypothesis.[7]

More generally, Model IV encompasses Models I, II and III. Hence, we can test for these models as well. Model I (Meltzer-Richard without tax disincentives) implies that $b=c=0$. Model II (Meltzer-Richard with tax disincentives) implies that $b<0$ and $c=0$. And Model III (Moene-Wallerstein) implies that $b>0$ and $c=0$.

The data

We use individual-level data from 11 advanced democracies obtained from two sets of national mass surveys conducted under the auspices of the International Social Survey Program (ISSP), one in 1996 the other in 1997 (ISSP 1999; 2000).[8] These surveys offer by far the best individual-level data on skills and preferences for social protection. We supplement these individual-level variables with economy-wide unemployment data. In the following two sections we describe the operationalization of the dependent and independent variables.

Dependent variables

The 1996 survey contains four spending questions that closely match our theoretical emphasis on income protection (*R*). Three of the four are used in a cluster of questions that asks whether the respondent would like to see more or less government spending on a) unemployment benefits, b) health care, and c) pensions (see Appendix C for details). Reflecting an assumption in the model, respondents were warned that more spending may require higher taxes. The fourth variable is based on a question that asks whether the respondent favors government spending on declining industries for the purpose of protecting jobs (see Appendix C for details). This question is as much about job security as it is about income security, but the two are obviously closely related, and we expect specific skills workers to be more concerned than general skills workers with keeping their present job and income. Moreover, although the respondent was not explicitly told about the potential costs of government subsidies, such subsidies are widely acknowledged to be problematic for economic efficiency.

The survey also asked people whether they favored more or less spending on "culture and the arts" and "the environment." These policy areas are clearly unrelated to social protection, but they are nevertheless relevant to our argument because general education is often argued to *increase* support for spending on "post-materialist" activities, whereas our theory says that it reduces support for spending in the social policy area (cf Kitschelt 1991).[9] Since one might object that our findings for skills reflect general ideological opposition to government spending among those with long formal educations, it is useful to be able to show that the relationship between skills and support for spending varies by policy area.

In order to economize on the presentation we used confirmatory factor analysis (CFA) to create two spending indexes: one for social spending and one for postmaterialist spending (both constructed to have a standard deviation of one). The Adjusted Goodness of Fit Index of the CFA model applied to all 11 countries is 0.94 and varies little by country (the range is 0.90-0.98).[10]

Independent variables.

We use two different approaches to the measurement of skill specificity, reflecting different aspects of the theoretical model. The first is to classify workers' skills, or the skills required to perform certain jobs, according to their degree of specialization or specificity. This is an attempt to gauge s directly. The second starts from the model assumption that the difficulty of finding a job where one's skills are needed is proportional to their specificity. This is an attempt to gauge s indirectly through rq: the probability of re-employment into *state I*.

The first approach is based on the ILO's detailed classification of people's occupations: the International Standard Classification of Occupations (ISCO-88). ISCO-88 classifies workers in "occupations" based on two criteria: the *level* of skills required for an occupation, and the *degree of specialization* of those skills. ISCO-88 distinguishes between four broad skill levels, which are a function of "the range and complexity of the tasks involved" and explicitly dependent on informal as well as formal training (ILO 1999, 6). Skill level thus corresponds to $(s+g)$ in our model. All other distinctions between occupations are based on the specialization of skills required to carry out particular jobs, reflecting “the type of knowledge applied, tools and equipment used, materials worked on, or with, and the nature of the goods and services produced” (ILO 1999, 6). Guided by this logic, the subdivision of skills proceeds through four levels of aggregation until a high degree of skill homogeneity is reached within each group.[11] At the most disaggregated level, called the unit level, there are 390 occupational categories with highly specific job descriptions.[12]

Since the occupation of every respondent in the ISSP surveys was classified according to ISCO-88 at either the most detailed or second most detailed level (for exceptions, see Appendix C), we can exploit the skill-based hierarchical structure of ISCO-88 to the capture the specialization of workers’ skills. We accomplish this by comparing the share of unit groups in any higher-level class to the share of the workforce in that class. The logic is that the number of unit groups in any higher-level class will be a function of the size of the labor market segment captured by that class, *and* of the degree of skill specialization of occupations found in that particular labor market segment. For example, 8 percent of the workforce across our countries are classified as "plant and machine operators and assemblers” (major group 8), whereas this group accounts for 70 out of the 390 unit groups, or 18 percent of all unit groups. If occupations at the unit group level are, on average, equally homogeneous in terms of skills, the disproportionate share of unit groups in major group 8 will reflect a greater degree of specialization of skills found within that major group. By dividing the share of unit groups (.18) by the share of the labor force (.8), we can therefore generate a measure of the average skill specialization within that particular major group (2.1). This calculation can also be done at the lower sub-major level, and we have used the mean of these calculations to get proxy for s.[13] The resulting variable has 27 values ranging from 0.4-4.7.

Because the theoretical concept of skill specificity is a *relative* variable, the final step is to divide the absolute skill specialization measure, s, by the ISCO measure of the *level* of skills.[14] This gives us a proxy for $s/(s+g)$ that we will refer to as s_1. Alternatively we can divide s by a proxy for peoples’ general skills, g, which gives us a measure for s/g. We call this alternative measure s_2. The proxy for g that we use is the respondent’s highest academic degree as recorded by the respondent (see Appendix C for details).

The second approach to measuring skill specificity is based on the observation that the probability of moving from any particular job into one that makes use of a worker's skills *(state I)* is rq for specific skills workers and q for general skills workers, where r<1. If we conceive of rq as an element in the continuum [0, q], r would then be a measure of the asset-specificity of a worker's skills. At the heart of the concept of job specificity is the idea that outside options are more limited for workers with specific skills than for workers with general skills.

The 1997 ISSP survey contains a question is that precisely taps workers' assessment of their outside options. The question reads as follows:

> If you were looking actively, how easy or difficult do you think it would be for you to find an acceptable job?

The respondent could answer "very easy", "fairly easy", "neither easy nor difficult", "fairly difficult", and "very difficult." The key here is that the difficulty of finding an acceptable job is likely to be related to how portable a person's skills are. High skill specificity means that there are fewer jobs where these skills are used, and the number of job openings is also likely to be smaller because asset-specific investments by employers and employees tend to lengthen tenure and limit turnover. In addition, the probability of finding an appropriate job close to a person's current residence, which is also a likely component of what an individual considers "acceptable," falls with the number of job openings in a given geographical area.[15] Asking people about the probability of finding an acceptable job is therefore likely to generate answers that are systematically related to a person's skills. In the absence of extensive information about individual work histories, and employment conditions in particular labor market niches, the question is therefore about as good a measure of rq as one could hope for. We refer to it as s_3.

There is however an ambiguity in the relationship of s_3 to the theoretical concept of s. The reason is that we cannot know for sure if peoples' responses reflect their *absolute* level of specific skills or the *relative* share of their skills that is specific. To make sure that the skill measure is a relative measure, as required by the theoretical model, we can divide s_3 by g. We call this alternative measure s_4. If s_3 is already a relative measure, we simply get another relative measure that should also be positively related to preferences for social spending.

The different skill measures, and their inter-correlations, are listed in Table 1. Not surprisingly, the correlations are higher between measures using either the survey question *or* the ISCO classification. The lowest correlations are between s_3 and s_1 or s_2. To some extent this may reflect that s_3 is an absolute rather than a relative measure, but the main reason is simply that s_3 is influenced by a number of factors (such as how much people like their current co-workers) that are unrelated to either skills or social policy preferences. These factors will wash out in the regression, but they reduce the correlation with the other measures. To facilitate comparison of the effects of the different variables in the subsequent regression analysis, all proxies for *s* have been divided by their standard deviation.

[Table 1 about here]

One final methodological issue needs to be addressed. Because the question used as the basis for s_3 and s_4 was asked only in the 1997 survey, whereas all the questions about spending were asked only in the 1996 survey, it was necessary to "translate" the 1997 information on s_3 so it could be used in the 1996 survey (s_4 can be always be calculated from s_3). For this purpose we calculated averages for s_3 at the 3-digit ISCO-88 level in the 1997, and then assigned these values to individuals in the 1996 survey based on their 3-digit ISCO classification in that survey.[16] Since the classification of occupations is motivated by the skills required in these occupations, it is reasonable to expect the original information about *s* is preserved to a considerable extent in this translation. Moreover, because the 1996 and 1997 samples are drawn from the same populations,[17] we show in Appendix D that s_3, averaged by ISCO level-3 groups, is an unbiased estimator for the original variable.

In addition to the skill variables (s_1-s_4) we used self-reported pre-tax and transfer income as a proxy for *y* (converted into dollars at 1996 exchange rates), and the following set of controls:

Age. Older workers are likely to be more concerned with job security and income than younger workers since their time to retirement is shorter and since their ability to find new employment is likely to be more limited.

Gender. As argued by Orloff (1993) and Estevez-Abe (2001), women may demand more protection than men in comparable jobs because they need to be able to leave, *and return*, to the labor market for the purpose of child rearing.

Union membership. Since one of the main functions of unions is to insure their members against labor market risks, it is reasonable to expect that union members are particularly concerned with social protection (see, for example, Korpi 1989).

Part-time employment. Part-time employees are often in vulnerable labor market positions, and this may cause particular concern for job security and income protection. On the other hand, part-time employees depend more on flexible labor markets to generate non-standard jobs, which suggests a countervailing effect.

Non-employed. Esping-Andersen (1999) has argued that some outsider groups may share an interest in social and economic policies that maximize their ability to enter employment. But this is an extremely heterogeneous group that may not have common policy preferences. We need to include the variable to control for the possibility that the non-employed have very different attitudes than the employed.

Unemployed. The expectation is obviously that the unemployed, relying as they do on transfers, will support high levels of income protection.

Self-employment. The self employed are expected to favor free markets and low levels of social protection because the self-employed depend on flexible labor markets and often on relatively low-paid workers.

Information. It is conceivable that better information about the economy yields particular views on the desirability of social spending. There was an intense public debate about the proper role of the state in the 1990s, and it could be argued that better informed people may reflect the predominant view in this debate, which tended to see cut-backs as necessary on efficiency grounds (corresponding to a higher cost of distortionary taxation in our model).[18] Information is measured by respondents' subjective understanding of politics (see Appendix C for details).

Left-right position. Attitudes to social protection may in part be a reflection of people's ideological predispositions, or perhaps the socializing effects of political parties.[19] We control for this possibility by including positions on a left-right scale based on the respondent's declared support for parties that are ranked from far left to far right (see Appendix C for details).

National unemployment. Although our theory implies that individuals discount cyclical unemployment, it has been suggested that such unemployment could impact individual-level social preferences. Testing this requires a multilevel modeling procedure, with countries as Level 1 and individuals as Level 2. Collapsing both levels into a single equation (as shown in Appendix E) implies the inclusion of the product variables $U_j.y_{ij}$ and $U_j.s_{ij}$ in the regression model, where U_j is the rate of unemployment in country j (see Appendix C for details on measurement).

Findings

We estimated the regression model in equation (15) on all countries (technically speaking as a single stage multilevel procedure to incorporate the possible impact of national macroeconomic conditions).[20] To cope with problems of missing observations we used a multiple imputation technique developed by Honaker et al. (1999). This strategy is superior to the traditional approach of "listwise deletion", which is both inefficient and potentially biased (King et al 2000).[21] The following presentation is divided into a section with the key results, and a section where we test the robustness of these results and discuss potential objections to the way we interpret them.

The basic results

To give a sense of the central tendency of the estimates, Table 2 shows the results from a pooled analysis, including a full set of country dummies. Since the Italian survey was conducted in 1990 and lacks information on several of the control variables, it was not included in the calculation of these pooled results. In the next section we show that the results for Italy are consistent with those presented in Table 2.

[Table 2 about here]

The model in column (1) uses the average of the four measures of skills, called $s_{\text{composite}}$, as a summary variable for skill composition. The next four columns show the results for each of the component measures (s_1-s_4). Model (6) is identical to (1) except that the regression now includes union membership as an independent variable. Since union membership was not recorded in Australia, the estimation of model (6) excludes this country.

In interpreting the results, first note that the parameters for income, *y*, and the four measures of skill, s_1-s_4, are in the predicted direction, and highly statistically significant. The negative effect of income implies that people's risk aversion is not sufficiently high to make their demand for transfers rise with income. Technically speaking, $RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$, which means that the Meltzer-Richard redistribution logic dominates the Moene-Wallerstein insurance logic. As expected, the relationship is little affected by differences in national unemployment rates, despite considerable variation in unemployment in the survey year. Thus, a one standard deviation increase in unemployment would only change the parameter on *y* from .0033 to .0038.

Yet, for our purposes the key finding is the positive effect of specific skills on preferences for spending (which implies that $RRA>0$). Each of the four (standardized) skill variables is associated with significantly higher support for spending, and three of the four measures exhibit similar magnitudes of effects. Again, these relationships hold for all levels of unemployment as can be seen from the negligible parameters for $U_j \cdot s_{ij}$.[22] The parameter for s_3 is lower than for the other measures, but this is not entirely unexpected given that this variable may capture absolute rather than relative endowments of specific skills (or a combination of absolute and relative endowments). For all correlations between *s* and *g* that are greater than -1, absolute measures of *s* will yield lower parameter estimates than relative measures.[23]

Considering the very different approaches to measuring skills, it is reassuring that the results are consistent across definitions. Yet, statistically significant effects do not necessarily imply large substantive effects. In Table 3 we have therefore estimated the portion of the explained variance accounted for by each of the independent variables, as well as the impact on preferences of a one standard deviation change in each of the independent variables. The estimates are based on the results of model (6) in Table 2, which includes all the relevant variables.

[Table 3 about here]

Although we cannot precisely attribute the proportion of explained variance to each of the independent variables, we can calculate the likely ranges. The *upper* bounds of these ranges are found by recording the increase in explained variance (measured as a percentage of the total explained variance) when a variable is included as the *first* predictor (apart from the country dummies). This number encompasses every direct, indirect, and spurious effect of the variable. The *lower* bounds are calculated as the increase in explained variance (as a percent of the total explained variance) when a variable is entered as the *last* predictor. This procedure eliminates all hypothesized individual-level spurious effects of the variable, but also discounts all possible indirect effects. The true explanatory power of any variable is likely to be somewhere in between these bounds.

Using this method, Table 3 shows that income and skills are unambiguously the most important variables in explaining social policy preferences among the ones included in this analysis. Thus, income accounts for between 11 and 51 percent of the total explained variance, whereas skills account for between 26 and 38 percent. Jointly, income and skills capture between 38 and 73 percent of the explained variance, with the rest accounted for by the controls.

The key role of income and skills in explaining social policy preferences is confirmed when we consider the impact of a one standard deviation change in these variables (column 3). A standard deviation change in either variable is associated with about 20 percent of a standard deviation change in preferences (since the dependent variable is standardized, the recorded effects can be interpreted directly in terms of standard deviations). Together, the impact of income and skills is as great as the joint effect of a standard deviation change in all controls simultaneously. Note also that the effects of both variables are estimated very precisely, varying in a narrow range between (-)0.19 and (-)0.22 (95-percent confidence interval).

The results for the controls also generally confirm our expectations. Individuals who are particularly exposed labor market risks – the unemployed, women, and older workers – are more favorably disposed to increasing social spending than others. The same is the case for union members, whereas the self-employed are more likely to oppose social spending. Those who consider themselves well-informed about politics are also more likely to oppose spending, perhaps reflecting a political reality at the time that was hostile to the welfare state. Supporters of right parties, not surprisingly, also express less support for social spending than supporters of left parties. Finally, we note that the attitudes of part-time employees and those outside the labor market are indistinct from the attitudes of others. These groups are evidently too heterogenerous to share any common interest in social policies.

Gender stands out among the control variables, accounting for between 8 and 17 percent of the total explained variance, and having the greatest impact among the controls. As argued by Estevez-Abe (2001), women require more protection than men in comparable jobs because they need to be able to leave, *and return*, to the labor market for the purpose of child rearing. Yet, almost half of this effect disappears if the skill variable is removed from the equation. The reason, we believe, is closely related to our theoretical argument. Since women know that they are likely to leave their jobs before they can reap the full returns on specific skill investments, they are dissuaded from making such investments (Estevez-Abe et al. 2001). This shows up in our data as a negative effect of (female) gender on s. Thus, our skill specificity variable is 0.27 standard deviations lower for women than it is for men (t=14). In other words, *if* women invest in specific skills they are more prone than men to support high levels of social protection, but they are somewhat less likely to invest in these skills in the first place.

Robustness tests

In this section we test the robustness of the results, and address some potential objections to our interpretation of the results. We first note that the findings for *y* and *s* stand up to any combination of the controls included above, and they are robust to the inclusion of any other variable used in the survey, hereunder region, public sector employment, urbanization, and supervisory position -- in any combination.[24] But while income and skills are powerful explanatory variables in the pooled analysis, pooling can disguise considerable cross-national variation in the strength of the results, and sometimes estimated parameters can even reverse in particular cases. In addition, pooling usually yields exaggerated t-scores compared to those found for individual countries.[25] We therefore ran our regressions on each of the 11 countries individually. The results for the theoretical variables are shown in Table 4.

[Table 4 about here]

Note that every regression yields results that are consistent with the pooled analysis, with each of the 60 parameters recording the correct sign, and most being significant at the .01 level or better. The composite skill variable is always significant at a .01 level or better, and for 9 of the 11 countries the parameter estimates for *s* vary in a fairly narrow range between 0.16-0.29 (the parameter in the pooled analysis is .23). Only Ireland and Italy fall slightly out of the pattern with parameters just below .12. Yet, the effects for these countries are still statistically highly significant, and it should be noted that $s_{composite}$ in both cases are based on only two proxies for *s*. In the case of Italy, these proxies also use a crude occupational variable that maps rather poorly onto ISCO-88, potentially diluting the skill distinctions between categories.

As in the case of the pooled analysis, we also note that the results for s_3 are somewhat weaker across all cases than for the other skill measures, but only in one instance (the US) do we get a statistically insignificant result. Given the variety of countries and the differences in measurements, the combination of results is clear support for our theory.

Another objection that can be raised to our findings for skills is that they may in part be capturing an ideological aversion to government spending among those with higher education. Two of the measures of *s* have formal education in the denominator, and the other two implicitly assume that such skills are part of the denominator. In quantitative terms, general education accounts for roughly one third of the variance in $s_{composite}$. It is therefore conceivable that the proxies for skills may in part capture an ideological effect of higher education. For example, much of the economic theory taught to university students during 1990s emphasized the efficiency of free markets over state intervention.[26]

To some extent we have already controlled for this possibility by including variables for people's assessment of their own level of information, as well as their support for parties on the left-right scale. If the highly educated consider themselves better informed about the costs of generous social spending, this is likely to show up in the variable measuring information. Likewise, those ideologically committed to a small welfare state are presumably more likely to support right parties. The fact that a large effect of skills persists after we control for these variables suggests that our conception of skills as assets is correct.

But there may still be unmeasured aspects of formal education that somehow confound the effects of our skill variable. One way to address this issue is simply to include general education as a separate variable. That way the effect of $s_{composite}$ will only pick up the effects of specific skills. In this setup we would expect formal education to have the opposite effect than the specific skills variable, and the separating out of general skills will necessarily weaken the effect of the original variable if general education is indeed a measure of general skills. However, we can be certain that whatever effect remains of s, it cannot be attributed to general education.

The first column of Table 5 shows the results of re-estimating model (1) in Table 2, using formal education as a separate independent variable. Formal education has a strong *negative* effect on support for social protection. This is consistent with the skill asset argument. But more importantly, the parameter on the specific skills variable remains positive and statistically significant. Not surprisingly, the effect of s falls from 0.23 to 0.14, but this is still a very considerable impact. Even if we were to completely discount the effect of general education as a measure of general skills, therefore, the results lend unambiguous support to our argument.

[Table 5 about here]

Yet, we think it would be a mistake to treat general education as a proxy for unmeasured ideological effects, and we can support this claim with results for the "postmaterialist" spending index explained above. Surely, if highly educated individuals believe in the efficiency of free markets and the waste of government spending, they should also oppose public spending on the environment, culture, and the arts.[27] But the exact opposite is true as shown in column (2) of Table 5. People with high general education are much *more* likely to support government spending on these areas than others. Conversely, if we use our composite measure of specific skills (column 3), the effect of skills is reversed: specific skill workers want less "postmaterialist" spending even though they support more social spending. Evidently people prefer government spending in areas that are particularly conducive to their personal welfare. General skills workers demand little social protection but are enthusiastic consumers of a clean environment and state-subsidized culture. Specific skill workers are deeply concerned with social protection but not enthusiastic about state subsidization of environmental causes and the arts. There is no blanket support for, or opposition to, government spending among any particular group of workers.

Conclusions and future research

It is a well-known fact that a substantial portion of both national and personal income can be attributed to human capital, broadly conceived. It is therefore not surprising that the asset-specificity of this capital matters a great deal for the amount of social insurance demanded by individual workers. Like physical capital, human capital can be more or less mobile, and workers who have made heavy investments in asset-specific skills stand a greater risk of losing a substantial portion of their income than workers who have invested in portable skills. For this reason, specific skill workers have a greater incentive to support policies and institutions that protect their jobs and income.

Because social protection tends to benefit low-income people more than high-income people, position in the income distribution also divides public opinion. However, at any given level of income, workers with specific skills are more inclined to support high levels of protection than those with general skills. This can help us understand cross-national variance in social protection since the profile of skills is likely to vary depending on the structure of the educational system. It is a well-documented fact, for example, that some countries train much more than others in vocational skills that are specific to particular jobs, firms, or industries. In Germany, for instance, more than a third of an age cohort goes through a longer vocational training (3-5 years), whereas in the US the comparable figure is only a few percentage points (even counting those in the American junior college system). If these differences are reflected in political preferences, either through the electoral system or the system of interest representation, it suggests a new explanation of the welfare state based on differences in national skill profiles.

The empirical plausibility of such an explanation is suggested by Figure 3, which uses the share of an age cohort going through vocational training as an indicator for the composition of labor force skills, and government transfers as a share of GDP as the proxy for the income transfer variable (*R*). As expected, there is a strong positive association between the two variables (r=.82), although future research is needed to eliminate the effects of potentially confounding variables.

Taking account of cross-national differences in skill composition also suggests a solution to the puzzle that income equality is linked to higher social spending in cross-national comparisons. Since vocational training activity is strongly positively related to pre-tax income equality – the correlation coefficient is .73 using d9/d1 earnings ratios as a measure of equality (OECD, undated) – income equality and social spending tend to go hand in hand across countries. In the pure Meltzer-Richard model this is ruled out because the pressure for redistribution is always greatest in countries with the most skewed distribution of income.

[Figure 3 about here]

Finally, our model points to an important source of cross-time variance in support for social protection: unanticipated shocks to the occupational structure. If workers have sunk investments in skills that are not fully transferable, an increase in the risk of having to move across a skill boundary in the economy raises the level of demand for social insurance. This helps us understand why the dramatic decline of industrial employment in many countries over the past three decades is a very good predictor of welfare state expansion (Iversen and Cusack 2000). More generally, changes over time in the exposure to risks, changes in the training system, and technologically induced changes in the international division of labor all affect the political demand for social protection. Modeling these dynamics, and testing them empirically, are important tasks for future research.

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Appendix A:

Mathematical proofs.

Derivation of results (12) and (13) in section (iii):

The choice of the optimal *R* requires that:

$V_R \geq 0 \Leftrightarrow \; b \cdot u'(\bar{g}) \cdot 2g / w = g \;\cdot u'(R)$

Totally differentiating both sides we get:

$$\frac{dR}{dg} = \frac{\frac{2b}{w} \cdot [\bar{g} \cdot u''(\bar{g}) + u'(\bar{g})]}{b \cdot \left(\frac{2g}{w}\right)^2 \cdot u''(\bar{g}) + g \;\cdot u''(R)}$$

Since the denominator is negative,

$$\frac{dR}{dg} > 0 \;\; iff \;\; \left[\bar{g}u''(\bar{g}) + u'(\bar{g})\right] < 0$$

which implies

$$RRA(\bar{g}) \equiv -\frac{\bar{g}u''(\bar{g})}{u'(\bar{g})} > 1$$

where *RRA(x)* is the Arrow-Pratt definition of Relative Risk Aversion defined at *c=x*. The inequality conditions specified in (12) and (13) follow directly.

Proof for Result I in section (iv):

(i) Note first that *t*=1 maximizes *t*/(1+*t*) when $0 \leq t \leq 1$. Also, if *t*=1, *R*=*w*/2.

(ii) From (6) the necessary condition for optimal *R* is

$$(1\text{A}) \qquad a \cdot u'(\overline{sg}) \cdot (1 - \frac{2sg}{w}) + b \cdot u'(\overline{g}) \cdot (1 - \frac{2g}{w}) + g \cdot u'(R) \geq 0$$

If $R = w/2$, $\overline{sg} = \overline{g} = R$; hence the maximum combination of *sg* and *g* at which *R*=*w*/2, assuming it exists, requires that this condition holds with equality and that $u'(\overline{sg}) = u'(\overline{g}) = u'(R)$. These conditions imply directly that $a \cdot sg + b \cdot g = (a + b + g) \cdot \frac{w}{2}$ $= y = \frac{w}{2}$.

Proof for Results II and III of model (iv):

The necessary condition for optimal choice of *R* is V_R *(R, s, g) = 0.* This is given by 1A above.

Totally differentiating V_R gives:

$$(2\text{A}) \qquad \begin{aligned} & a \cdot \left[u''(\overline{sg}) \cdot \left(\frac{2sg}{w} - 1 \right) \cdot \left(1 - \frac{2R}{w} \right) \cdot g + u'(\overline{sg}) \cdot g \cdot \frac{2}{w} \right] \cdot ds \\ & + a \cdot \left[u''(\overline{sg}) \cdot \left(\frac{2sg}{w} - 1 \right) \cdot \left(1 - \frac{2R}{w} \right) \cdot s + u'(\overline{sg}) \cdot s \cdot \frac{2}{w} \right] \cdot dg \\ & + b \cdot \left[u''(\overline{g}) \cdot \left(\frac{2g}{w} - 1 \right) \cdot \left(1 - \frac{2R}{w} \right) + u'(\overline{g}) \cdot \frac{2}{w} \right] \cdot dg \\ & = \left\{ a \cdot u''(\overline{sg}) \cdot \left(\frac{2sg}{w} - 1 \right)^2 + b \cdot u''(\overline{g}) \cdot \left(\frac{2g}{w} - 1 \right)^2 + g \cdot u''(R) \right\} \cdot dR \end{aligned}$$

Note: (i) The term in curly brackets on the RHS, which we will call *B*, is negative. (ii) We can write $\left(\overline{sg} - w/2\right) \cdot \left(1 - 2R/w\right) = \overline{sg} - \overline{w}/2$. And (iii):

$$(3\text{A}) \qquad \begin{aligned} & \left[u''(\overline{sg}) \cdot \left(\overline{sg} - \overline{w}/2 \right) + u'(\overline{sg}) \right] \\ & = u'(\overline{sg}) \cdot \left[1 - RRA \cdot \frac{\overline{sg} - \overline{w}/2}{\overline{sg}} \right] \equiv u'(\overline{sg}) \cdot L(\overline{sg}) \end{aligned}$$

So (2A) can be written:

$$
\text{(4A)} \quad \begin{aligned} & u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) \cdot a \cdot g \cdot ds + u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) \cdot a \cdot s \cdot dg \\ & + u'(\bar{g}) \cdot L(\bar{g}) \cdot b\, dg = (w/2) \cdot B \cdot dR \end{aligned}
$$

Since $dy = a \cdot g \cdot ds + a \cdot s \cdot dg + b \cdot dg$, we can further rewrite (4A) as

$$
\text{(5A)} \quad \begin{aligned} dR = & \frac{2a \cdot b \cdot g}{wB} \cdot \left[\frac{u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) - u'(\bar{g}) \cdot L(\bar{g})}{a\,s + b} \right] \cdot ds \\ & + \frac{2}{wB} \cdot \left[\frac{u'(\bar{s}\bar{g}) \cdot L(\bar{s}\bar{g}) \cdot a\,s + u'(\bar{g}) \cdot L(\bar{g}) \cdot b}{a\,s + b} \right] \cdot dy \end{aligned}
$$

To prove **Results III** and **IV**, note that in terms of (5A) $¶R/¶y = dR/dy$ and $¶R/¶s = dR/ds$. We show first that $L(\bar{s}\bar{g}) < L(\bar{g})$. From the definition in (3A), this follows if $s > 1$ – as is the case apart from purely general skills – and if RRA > 0. **Result III** is that $\mathrm{sgn}\,¶R/¶y < 0 \;\; if \;\; RRA < \bar{s}\bar{g}/(\bar{s}\bar{g} - \bar{w}/2)$. Since $B < 0$, $L(\bar{s}\bar{g}) < L(\bar{g})$ and $u'(x) > 0$, this follows from (5A) if $L(\bar{s}\bar{g}) > 0$. This requires that $RRA < \frac{\bar{s}\bar{g}}{\bar{s}\bar{g} - \bar{w}/2}$. This is a sufficient condition: a necessary and sufficient condition is that the numerator of the second term in square brackets on the RHS of (5A) is positive.

Result IV is that $\mathrm{sgn}\,¶R/¶s > 0$. Since $B < 0$, this requires that the numerator in the first square bracket on the RHS of (5A) is negative. Since $u'(\bar{s}\bar{g}) < u'(\bar{g})$ from diminishing marginal utility, a sufficient condition is that $L(\bar{s}\bar{g}) < L(\bar{g})$, which is true so long as RRA > 0 and s > 1. So **Result IV** follows from the existence of risk aversion and specific skills.

Appendix B:
Deriving the estimating equation

In this appendix we show that the estimating equation that we use:

(6A) $$R = k + b \cdot y + c \cdot s$$

is equal to

(7A) $$R = k + \frac{\partial R}{\partial y} \cdot y + \frac{\partial R}{\partial s} \cdot s$$

where (7A) is a first-order Taylor expansion of $V_R(R,\ s,\ g) = 0$ and $y = a\ \cdot s \cdot g + b\ \cdot g$ evaluated around $(R,\ s,\ g) = (\tilde{R},\ \tilde{s},\ \tilde{g}) \equiv \tilde{x}$

Proof: The first-order Taylor expansion of V_R is given by:

(8A) $$R = K + \frac{V_{R,s}}{V_{R,R}} s + \frac{V_{R,g}}{V_{R,R}} g$$

In terms of (5A):

(9A.1) $$\frac{V_{R,s}(\tilde{x})}{V_{R,R}(\tilde{x})} = \frac{u'(\overline{sg}).L(\overline{sg}).a\ .\tilde{g}}{(w/2).B}$$

and

(9A.2) $$\frac{V_{R,g}(\tilde{x})}{V_{R,R}(\tilde{x})} = \frac{u'(\overline{sg}).L(\overline{sg}).a\ .\tilde{s} + u'(\overline{g}).L(\overline{g}).b}{(w/2).B}$$

The first-order Taylor expansion of y is:

(10A) $$y = k(\tilde{x}) + \left[a\ .\tilde{s} + b\ \right].g + \left[a\ .\tilde{g}\right].s$$

Rewrite (10A): $g = \dfrac{y - k(\tilde{x}) - a\,\tilde{g}}{a\,\tilde{s} + b}$ and substitute into (8A), using (9A.1) and (9A.2). This yields (7A).

Appendix C:
Detailed information about variables

Dependent variables:

The spending variable, *R*, is based on four issue items in the ISSP surveys. The first three are based on the following question:

"Listed below are various areas of government spending. Please show whether you would like to see more or less government spending in each area. Remember that if you say 'much more', it might require a tax increase to pay for it." The respondent is then presented with the different spending areas (unemployment, health, retirement) and the following range of possible responses: "1. Spend much more; 2. Spend more; 3. Spend the same as now; 4. Spend less; 5. Spend much less; 8. Can't choose, don't know."

The fourth variable is based on the following question:

"Here are some things the government might do for the economy. Please show which actions you are in favor of and which you are against. Please tick one box in each line." One of the actions is: "Support for declining industries to protect jobs: 1. Strongly in favor of; 2. In favor of; 3. Neither in favor of nor against; 4. Against; 5. Strongly against; 8. Can't choose, don't know; 9. NA, refused."

Independent variables:

s_1 and s_2:

In some countries individuals were classified using an earlier version of ISCO (ISCO-68). However, these classifications can be translated into ISCO-88 with considerable consistency using a coding scheme developed by Harry Ganzeboom at Utrecht University (see Ganzeboom and Treiman 1996 and http://www.fss.uu.nl/soc/hg/ismf for details). The Swedish occupational classification is based on an amended version of an older edition of ISCO. *Statistiska Centralbyrån* (Statistics Sweden) provided us with a conversion table to translate these codes into

ISCO-88 in reasonably consistent manner. Britain uses its own national classification system, but it is closely related to ISCO-88 and likewise uses skills as the basis for the classification. We received the British translation codes from *UK National Statistics*. The only problematic case is Italy where the few broad categories used in the 1996 ISSP survey are completely unrelated to the ISCO-88 categories. Instead we went back to an earlier 1990 ISSP study (ISSP 1993), which contains a somewhat more detailed occupational variable for Italy. Using this variable in conjunction with information on educational levels enabled us to map the Italian codes to the 1-digit ISCO-88 level in a fairly consistent manner. Yet, because of the lack of direct correspondence, the results for Italy must be viewed with caution.

General skills (used in the denominator of s_2 and s_4):
The variable is used as a prox for *g*, and has five levels: 1. completed primary degree or lower; 2. incomplete secondary; 3. completed secondary; 4. incomplete and completed semi-higher degree, or incomplete university degree; and 5. completed university degree (in some countries a distinction is made between incomplete and complete primary education, but we are not making use of this distinction). Alternatively we could have used years of formal schooling as a measure of *g*, but the results are very similar.

Information:
Gauged by a question that asked people to declare their degree of agreement with the following statement: “I feel that I have a pretty good understanding of the important political issues facing our country.” Respondents could indicate five levels of agreement: 1. Strongly agree; 2. Agree; 3. Neither agree nor disagree; 4. Disagree; 5. Strongly disagree. The variable was reversed so that higher values measure more information.

Left-right position:
This variable is based on the classification of parties from left to right developed by the International Social Survey Program to facilitate comparison of party support across countries. Individual parties are classified as follows (data on party support are not available for Italy):

	Far left (1)	Left, Center Left (2)	Center, Liberal (3)	Right, Conservative (4)	Far Right (5)
Australia	Greens	Labour	Democrats	Liberal Party	
Britain		Labour	Liberal Democrats	Conservatives	
Canada	Communists	NDP, Bloc Quebecois, Greens	PC, Liberal Party	Reform Party	
France	Communists, Far Left	Socialist Party	UDF	RPR	National Front
Germany	PDS	SPD, Greens	FDP	CDU/CSU	Republicans
Ireland		Worker's Party, Sinn Fein, Democratic Left	Fianna Fail, Fine Gael, Labour, Greens	Progressive Party	
Norway	Red Alliance	Labor, Socialist Left	Christian Democrats, Center Party, Liberal Party	Conservatives, Progress Party	
New Zealand	Alliance	Labour	New Zealand First	National Party	
Sweden		Labor, Socialists	Center Party, Liberals, Chr Democrats, Greens	Conservatives	
United States		Democrats	Independent	Republicans	

National unemployment: The standardized rate of unemployment at the time of the national surveys (1996 unless noted otherwise below) minus the OECD rate of unemployment at that time (the subtraction eliminates problems of multicollinearity while leaving the substantive results unaltered). *Source*: OECD (2000). Unemployment rates were: Australia, 8.5; Britain, 8.2; Canada, 9.6; France (1997), 12.3; Germany, 8.9; Ireland, 11.6; Italy, 11.7; Norway, 4.9; New Zealand (1997), 6.7; Sweden, 9.6; United States, 5.4.

Appendix D:
Statistical appendix

A problem arises in our use of s_3 and s_4 as explanatory variables. (Since it is the same in both cases we will simply refer to *s*.) Because the question used as the basis for *s* was asked only in the 1997 survey, whereas all the questions about spending were asked only in the 1996 survey, it was necessary to "translate" the 1997 information on *s* so it could be used in the 1996 survey. For this purpose we calculated averages for *s* at the 3-digit ISCO-88 level in the 1997, and then assigned these values to individuals in the 1996 survey based on their 3-digit ISCO classification in that survey. We show here that the estimated coefficient of *b* is consistent but has an approximate small sample bias which biases down the estimated coefficient towards zero if $b > 0$, and biases it upwards towards zero if $b < 0$.

The structural model is

(11A) $$R^{96}_{i,j} = k + b.y^{96}_{i,j} + c.s^{96}_{i,j} + e^{96}_{i,j}$$

where each observation is drawn from the 1996 survey and where *i* indexes the *i*th individual in the *j*th ISCO 3-digit level occupation group. We do not have data on $s^{96}_{i,j}$. Assume $s^{96}_{i,j}$ is generated by the process

(12A) $$s^{96}_{i,j} = s_j + h^{96}_{i,j}$$

where s_j is exogenous. s_j itself is unobservable but we have data from the 1997 survey generated by the same process:

(13A) $$\begin{aligned} &s^{97}_{i,j} = s_j + h^{97}_{i,j} \\ &Eh^{9x}_{i,j} = 0 \quad \forall\, i, j, x \\ &Eh^{96}_{i,j}.h^{97}_{r,j} = 0 = Ehe \quad \forall\, i, r, j \\ &Eh^2 = s^2_h; \; Ee^2 = s^2_e \end{aligned}$$

i.e., h_{ij}^{96} and h_{ij}^{97} can be thought of as random drawings from the same distribution. We now run the regression

$$(14A)\qquad \begin{aligned} &R_{i,j}^{96} = k + by_{i,j}^{96} + c\bar{s}_j + e_{i,j}^{96} + u_{i,j}^{96} \\ &where\ \ \bar{s}_j \equiv \frac{\sum s_{i,j}^{97}}{N_j} \\ &and\ \ u_{i,j}^{96} \equiv c\left[s_{i,j}^{96} - \bar{s}_j\right] \end{aligned}$$

where N_j is the number of individuals in ISCO category *j* in the 97 survey. From (A12, A13)

$$(15A)\qquad u_{i,j}^{96} = h_{i,j}^{96} - \frac{\sum h_{i,j}^{97}}{N_j}$$

We can simplify the exposition considerably by assuming that there is no correlation between $\bar{s}$ and *y*. This implies:

$$(16A)\qquad \begin{aligned} &E\hat{b} = a \\ &E\hat{c} = c\left(1 - E\frac{\sum_j \bar{s}_j.\overline{h}_j}{\sum_j \bar{s}_j^2}\right) \end{aligned}$$

Making the appropriate probability limit assumptions it is not difficult to show that $\hat{c}$ is a consistent estimator of *c*. We can get a better insight from the expectation of the exact first-order Taylor expansion of $\frac{\sum \bar{s}.h}{\sum \bar{s}^2}$ around the expected values of numerator and denominator:

$$E\frac{\sum \bar{s}h}{\sum \bar{s}^2} \approx E\left[\frac{E\sum \bar{s}h}{E\sum \bar{s}^2} + \frac{1}{E\sum \bar{s}^2}.\left(\sum \bar{s}h - E\sum \bar{s}h\right) - \frac{E\sum \bar{s}h}{\left(E\sum \bar{s}^2\right)^2}.\left(\sum \bar{s}^2 - E\sum \bar{s}^2\right)\right]$$
$$= \frac{E\sum \bar{s}h}{E\sum \bar{s}^2}$$

Let there be J ISCO categories and assume for convenience that $N_j = N \ \forall j$. Then this approximation produces:

(17A) $$E\hat{c} = c\left(1 - \frac{s_h^{\,2}}{s_h^{\,2} + N.\frac{\sum s_j^2}{J}}\right)$$

Since J is constant, (C.8) tells us first that as N increases the approximate bias goes to zero. Second and more important, it implies that for a small sample

(18A) $$\begin{array}{l} c < E\hat{c} < 0 \quad if \quad c < 0; \\ c > E\hat{c} > 0 \quad if \quad c > 0. \end{array}$$

Finally, it implies that

(19A) $$If \quad c = 0 \Rightarrow E\hat{c} = 0$$

and the standard significance tests hold.

Appendix E:
Multilevel model

Write the Level 2 observation on individual *i* in economy *j* as

$R_{ij} = g + hy_{ij} + ms_{ij} + \mathbf{X}'_{\mathbf{ij}}d + e_{ij}$ where $\mathbf{X}_{\mathbf{ij}}$ is the vector of controls, and define Level 1 by the fixed effects model $h = \bar{h} + h_1 U_j$ and $m = \bar{m} + m_1 U_j$. The single stage regression model is derived by substituting the Level 1 model into Level 2:

$$R_{ij} = g + \bar{h} y_{ij} + h_1 y_{ij} U_j + \bar{m} s_{ij} + m_1 s_{ij} U_j + \mathbf{X}'_{\mathbf{ij}} d + e_{ij}$$

Making the assumption that U_j s are exogenous, and that this is a non-random effects model, implies that the single stage multilevel equation conforms to the standard OLS conditions.

ENDNOTES

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1. The trade policy argument treats skilled labor as a more or less abundant factor, where a relatively abundant factor is hypothesized to favor free trade. See Rogowski (1987), Frieden (1991), and especially Scheve and Slaughter (1999) for different versions of this argument.

2. Alt et al. (1999) show empirical evidence that lobbying rises with the asset specificity of industries. See also Alt et al. (1996) for a more theoretical treatment of this and related arguments concerning the importance of asset specificity.

3. There is an analogy here to the argument that capital mobility increases the incentives of capital owners to oppose restrictions on trade and financial mobility. See Frieden (1991), and Bates, Brock, and Tiefenthaler (1991).

4. This is also a realistic assumption with after-tax income distributed significantly more equally than pre-tax income (see Gottschalk and Smeeding 2000 and Huber and Stephens 2001).

5. A third type of protection identified by Estevez-Abe et al. (2001) is called *employment protection* and refers to legal and other barriers to layoffs. This type of protection could be modeled as a probability of keeping a job where a worker's skills are fully utilized.

6. Meltzer-Richard have a more general tax disincentive function than that used here. In consequence the tax rate which maximizes tax revenue can be less than one.

7. The model also implies that the coefficients of *y* and *s* are independent of cyclical variations in the unemployment rate. This implication can be tested through multilevel modeling as discussed below.

8. The 11 countries are: Australia, Britain, Canada, France, Germany, Ireland, Italy, Netherlands, Norway, Sweden, and the United States. Japan, used in both ISSP surveys, could not be included because of missing data on a key occupational variable (explained below).

9. Duch and Taylor (1993) make a similar argument concerning postmaterialist attitudes (though they do not directly discuss spending).

10. We used LISREL Version 8.5 to conduct the confirmative factor analysis, using the resulting factor loadings to construct the two indexes from the individual spending variables. The model was estimated from the co-variance matrix for the six spending variables, assuming that the variables are indicators of two latent spending variables: social and postmaterialist spending. The factor loadings for each latent variable are as follows: i) *social spending*: .52 (subsidies to protect jobs), .48 (health insurance), .58 (pensions), and .55 (unemployment insurance); ii) *postmaterialist spending*: .58 (environment), and .51 (culture and the arts). Alternatively the indexes can be constructed from the results of fitting confirmatory factor models to the co-variance matrixes for individual countries, but the regression results are only marginally affected. These results, the results for each spending area separately, and the country-level CFA results are available from the authors.

11. We are not claiming that homogeneity is equivalent in every unit group. Yet, skills that are clearly distinct from one another are unlikely to be in the same group at the most disaggregated level, and major groups with a highly diverse skill structure therefore will tend to have more minor and unit groups.

12. Unit group 3144, for example, represents “air traffic controllers”, which is a member of the minor group “ship and aircraft controllers and technicians” -- itself one of five categories in the major group called “technicians and associate professionals”.

13. The sensitivity of *s* to small differences in the number of unit groups assigned to each higher-level group is greater at lower levels of aggregation, and these differences may not accurately

reflect differences in skill specificity. This source of error is minimized at the highest level of aggregation. However, the greater variance of the measure at lower levels of aggregation helps reduce the standard error on the estimated parameter for the skill variable.

14. Using an absolute measure of s generates results that are downward biased. At the limit, if the (unknown) correlation between *s* and *g* is 1, *s* will have no effect on preferences. It is therefore important to develop relative measures.

15. In an path-breaking analysis Scheve and Slaughter (2001) argue, and show empirically, that home ownership can be treated as a relatively immovable asset that affects people's preferences for trade protection. It would be interesting to interact home ownership with the question about the difficulty of finding an acceptable job, but residential status is unfortunately not recorded by ISSP.

16. There are 116 unique groups at the 3-digit level. The more fine-grained 4-digit level is not available for some countries, and contains a large number of empty categories where it is.

17. With the minor qualification that those who turned 18 between the 1996 and 1997 surveys were not part of the population in the former survey.

18. This idea was suggested by an anonymous reviewer.

19. This was suggested by an anonymous reviewer. We note that party support may in part be endogenous to skills. If so, the effect of skills will be underestimated by the parameter for *s*.

20. All data analysis was done using Stata 6.0 for Windows.

21. In practice, however, our results are very similar to those obtained by using listwise deletion. The effects of our theoretical variables tend to be slightly stronger when we use listwise deletion, but the standard errors are also larger.

22. For example, a one standard deviation increase in U_j only reduces the effect of $s_{composite}$ from .23 to .22.

23. In fact, the correlation between s_3 and a measure of *g* based on general education is close to 0 in our data, which implies an estimated effect of s_3 is half the "true" effect of skills.

24. None of these variables were used in every survey, so instead of cluttering the presentation with several additional columns, we left these variables out of the main analysis.

25. The reason is that the standard error has the form (s.e. of equation error)/(s.e. of variable). Since the denominator is the square root of the sum of squares of the explanatory variable divided by N, this normally increases with N since a squared term is added on the top and 1 is added to the bottom (though it does not have to be so).

26. In terms of the formal model this can be captured by different assessments of the distortionary effects of taxation.

27. It is true that “the environment” may be conceived as a collective good improving overall welfare (it is a little harder to argue this with respect to subsidies to the fine arts), but by the same token social protection may be conceived as welfare-improving insurance. The point is not that the highly educated are more informed about what is “good” and “bad” spending, we already control for this, but that they may have internalized a general aversion to government spending through their educational experience.

Tables and Figures

Table 1. Summary of Independent Skill Variables

Variable Name:	Definition:	Inter-correlations				Comment:
		s_1	s_2	s_3	s_4	
s_1	(Share of ISCO-88 level 4 groups)/ (share of labor force) divided by ISCO level of skills[1)]	1				
s_2	(Share of ISCO-88 level 4 groups)/ (share of labor force) divided by level of general education[1)]	.82	1			
s_3[2)]	Response to question about difficulty of finding an acceptable job.	.38	.53	1		Not clear whether this is a measure of absolute or relative skills
s_4[2)]	s_3 divided by level of general education	.66	.59	.62	1	Assumes that s_3 measures absolute skills (though s_4 will always be a relative measure)

[1)] Shares are calculated at both the first and second ISCO-88 level and then averaged. [2)] The number of categories on s_3 and s_4 have been reduced to the same number as on s_1 and s_2 before calculating the inter-correlations.

Table 2. Support for Social Spending Among the Publics of 10 OECD Countries, 1996 (Standard Errors in Parentheses)

	Dependent Variable: Support for Social Spending[1)]					
	(1)	(2)	(3)	(4)[2)]	(5)[2)]	(6)[3)]
Income	-0.0033** (0.0002)	-0.0036** (0.0002)	-0.0038** (0.0002)	-0.0044** (0.0002)	-0.0035** (0.0002)	-0.0036** (0.0002)
$s_{composite}$	0.233** (0.014)	-	-	-	-	0.219** (0.013)
s_1	-	0.148** (0.010)	-	-	-	-
s_2	-	-	0.150** (0.010)	-	-	-
s_3	-	-	-	0.105** (0.013)	-	-
s_4	-	-	-	-	0.218** (0.014)	-
Age	0.0029** (0.0006)	0.0043** (0.0006)	0.0034** (0.0005)	0.0042** (0.0006)	0.0018** (0.0006)	0.0027** (0.0006)
Gender (female)	0.215** (0.018)	0.208** (0.018)	0.205** (0.018)	0.124** (0.019)	0.148** (0.019)	0.198** (0.019)
Union membership	-	-	-	-	-	0.185** (0.023)
Part-time empl.	-0.029 (0.028)	-0.041 (0.028)	-0.033 (0.028)	-0.076* (0.031)	-0.058 (0.031)	-0.031 (0.029)
Unemployed	0.293** (0.041)	0.313** (0.041)	0.311** (0.042)	0.320** (0.047)	0.309** (0.046)	0.325** (0.043)
Non-employed	-0.079** (0.025)	-0.081** (0.025)	-0.086** (0.025)	-0.080** (0.026)	-0.074** (0.026)	-0.038 (0.026)
Self-employed	-0.232** (0.029)	-0.235** (0.028)	-0.250** (0.028)	-0.222** (0.028)	-0.221** (0.029)	-0.184** (0.027)
Informed	-0.041** (0.008)	-0.045** (0.008)	-0.047** (0.008)	-0.069** (0.010)	-0.050** (0.010)	-0.043** (0.009)
L-R party support	-0.050** (0.004)	-0.051** (0.004)	-0.050** (0.004)	-0.047** (0.005)	-0.047** (0.005)	-0.041** (0.005)
$U_j.y_{ij}$	-0.0002** (0.0001)	-0.0002** (0.0001)	-0.0002** (0.0001)	-0.0003** (0.0001)	-0.0004** (0.0001)	-0.0003** (0.0001)
$U_j.s_{ij}$	-0.008 (0.005)	-0.004 (0.004)	-0.002 (0.004)	-0.008 (0.005)	-0.012* (0.005)	-0.012* (0.004)
Adj. R-squared	0.21	0.20	0.20	0.18	0.20	0.22
N	14,101	14,101	14,101	10,956	10,956	11,950

Key: * significant at the .05 level; ** significant at the .01 level.

Notes: [1)] All regressions included a full set of country dummies (not shown); [2)] excludes Australia, Ireland, and Italy for which data are not available; [3] excludes Australia for which union membership data are not available.

Table 3. Estimates of the Magnitude of the Effects of Independent Variables

	Proportion of Explained Variance[1]		Impact of a One STD. Change[4]	
	Lower bound[2]	Upper bound[3]	95% confidence interval	
Income	11	51	-0.22	-0.19
$s_{composite}$	26	38	0.19	0.22
Age	1	2	0.03	0.05
Gender (female)	6	17	0.09	0.11
Union membership	1	4	0.07	0.09
Part-time employment	0	0	-0.02	-0.00
Unemployed	3	9	0.06	0.08
Non-employed	0	8	-0.03	-0.01
Self-employed	2	8	-0.07	-0.05
Informed	1	8	-0.05	0.04
L-R party support	5	5	-0.09	-0.07
$U_j \cdot y_{ij}$	1	13		
$U_j \cdot s_{ij}$	0	8		
Income *and* $s_{composite}$	38	73	0.38	0.44
All controls combined	27	52	0.36	0.47

Notes: [1] Increase in explained variance by each variable as proportion of the total explained variance of all (non-dummy) variables (based on model in Table 2, column 6). [2] Increase in explained variance (compared to model with only country dummies) when each variable is included as the last variable. [3] Increase in explained variance when a variable is included as the first variable. [4] The change in support for social spending (measured in standard deviations) as a result of a one standard deviation increase in each of the independent variables (in the cases of income and $s_{composite}$, unemployment is kept at its mean). The last two rows assume changes in the independent variables that raise support for spending (and take into account that some combinations of the employment variables are impossible).

Table 4. Income, Skills, and Support for Social Spending in 11 OECD Countries (t-Scores in Parentheses)[1)]

	Income[2)]	$s_{composite}$	s_1	s_2	s_3	s_4	N
Australia	-0.0030**	0.156**	0.129**	0.127**	n.a[3)]	n.a	2151
	(-0.0004)	(0.027)	(0.023)	(0.026)			
Britain	-0.0029**	0.219**	0.105**	0.121**	0.135**	0.181**	989
	(-0.0006)	(0.042)	(0.026)	(0.027)	(0.039)	(0.046)	
Canada	-0.0054**	0.219**	0.102*	0.140**	0.087*	0.220**	1182
	(-0.0008)	(0.053)	(0.042)	(0.045)	(0.041)	(0.047)	
France	-0.0055**	0.235**	0.158**	0.147**	0.097**	0.111**	1312
	(-0.0007)	(0.038)	(0.035)	(0.028)	(0.037)	(0.026)	
Germany	-0.0027**	0.255**	0.182**	0.155**	0.115**	0.212**	2361
	(-0.0007)	(0.030)	(0.028)	(0.023)	(0.035)	(0.028)	
Ireland	-0.0030**	0.116**	0.092**	0.122**	n.a	n.a	994
	(-0.0008)	(0.028)	(0.029)	(0.027)			
Italy	-0.0021	0.105**	0.104*	0.095**	n.a	n.a	983
	(-0.0016)	(0.037)	(0.040)	(0.034)			
Norway	-0.0026**	0.257**	0.139**	0.165**	0.076**	0.250**	1344
	(-0.0006)	(0.033)	(0.024)	(0.027)	(0.027)	(0.030)	
New Zealand	-0.0049**	0.176**	0.125**	0.094**	0.134**	0.131**	1198
	(-0.0007)	(0.038)	(0.030)	(0.030)	(0.041)	(0.034)	
Sweden	-0.0055**	0.282**	0.130**	0.140**	0.156**	0.278**	1238
	(-0.0010)	(0.039)	(0.030)	(0.028)	(0.033)	(0.030)	
United States	-0.0024*	0.294**	0.192**	0.189**	0.042	0.272**	1332
	(-0.0010)	(0.052)	(0.030)	(0.037)	(0.036)	(0.058)	

Key: * significant at the .05 level; ** significant at the .01 level.

Notes: [1)]All regressions included the same set of controls as in Table 2, column (1); [2)] The effect of income is only shown for $s_{composite}$, but varies little across the four measures of *s*; [3)] n.a.: data not available to estimate this parameter.

Table 5. Formal Education and Support for Two Types of Spending in 10 OECD Countries, 1996 (t-Scores in Parentheses)

	Support for Social Spending	Support for Postmaterialist Spending	
Formal education	-0.105**	0.130**	-
	(0.008)	(0.007)	
$s_{composite}$	0.143**	-	-0.097**
	(0.015)		(0.014)
Income	-0.0027**	-0.0004	0.0004
	(0.0002)	(0.0002)	(0.0002)
Age	0.0015*	-0.0064**	-0.009**
	(0.0006)	(0.0006)	(0.001)
Gender (female)	0.203**	0.092**	0.087**
	(0.018)	(0.018)	(0.019)
Part-time employment	-0.025	0.104**	0.112**
	(0.028)	(0.029)	(0.030)
Unemployed	0.303**	0.085*	0.089*
	(0.040)	(0.043)	(0.044)
Non-employed	-0.067**	0.077**	0.093**
	(0.025)	(0.024)	(0.025)
Self-employed	-0.243**	0.021	0.011
	(0.028)	(0.025)	(0.026)
Informed	-0.031**	0.069**	-0.083**
	(0.008)	(0.009)	(0.009)
L-R party support	-0.049**	-0.060**	-0.059**
	(0.004)	(0.005)	(0.005)
$U_j.y_{ij}$	-0.0002	0.0000	0.0001
	(0.0001)	(0.0001)	(0.0001)
$U_j.s_{ij}$	-0.006	0.008	0.007
	(0.005)	(0.005)	(0.006)
Adj. R-squared	0.22	0.09	0.07
N	14,101	14,101	14,101

Key: * significant at the .05 level; ** significant at the .01 level.

Note: Regressions included a full set of country dummies.

Figure 1: The Three States of an Individual in the Labor Market.

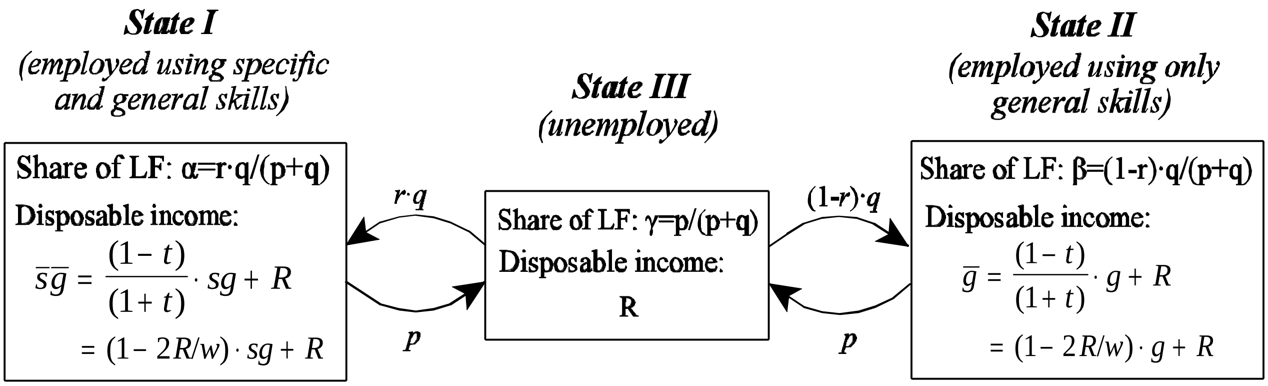
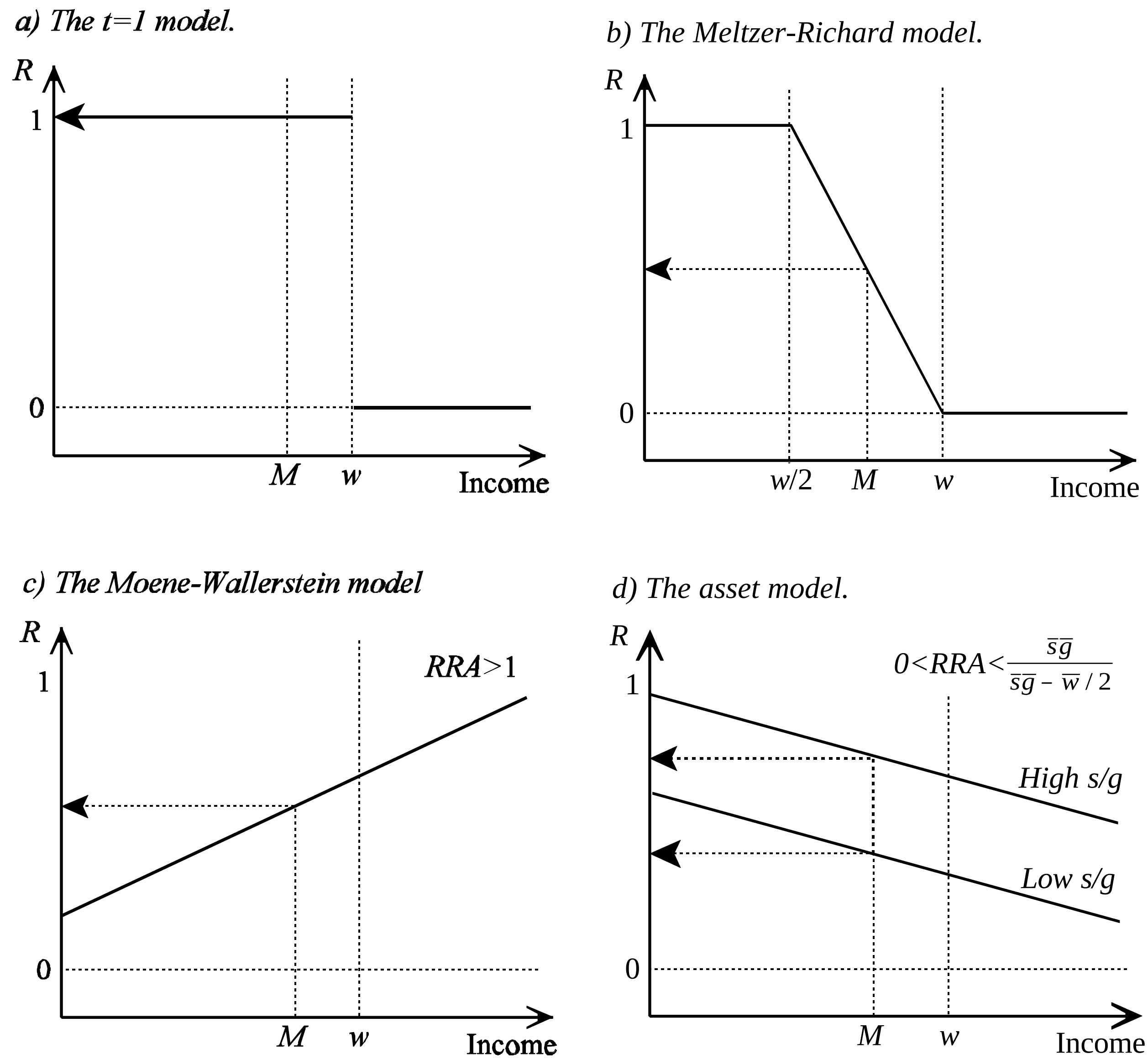
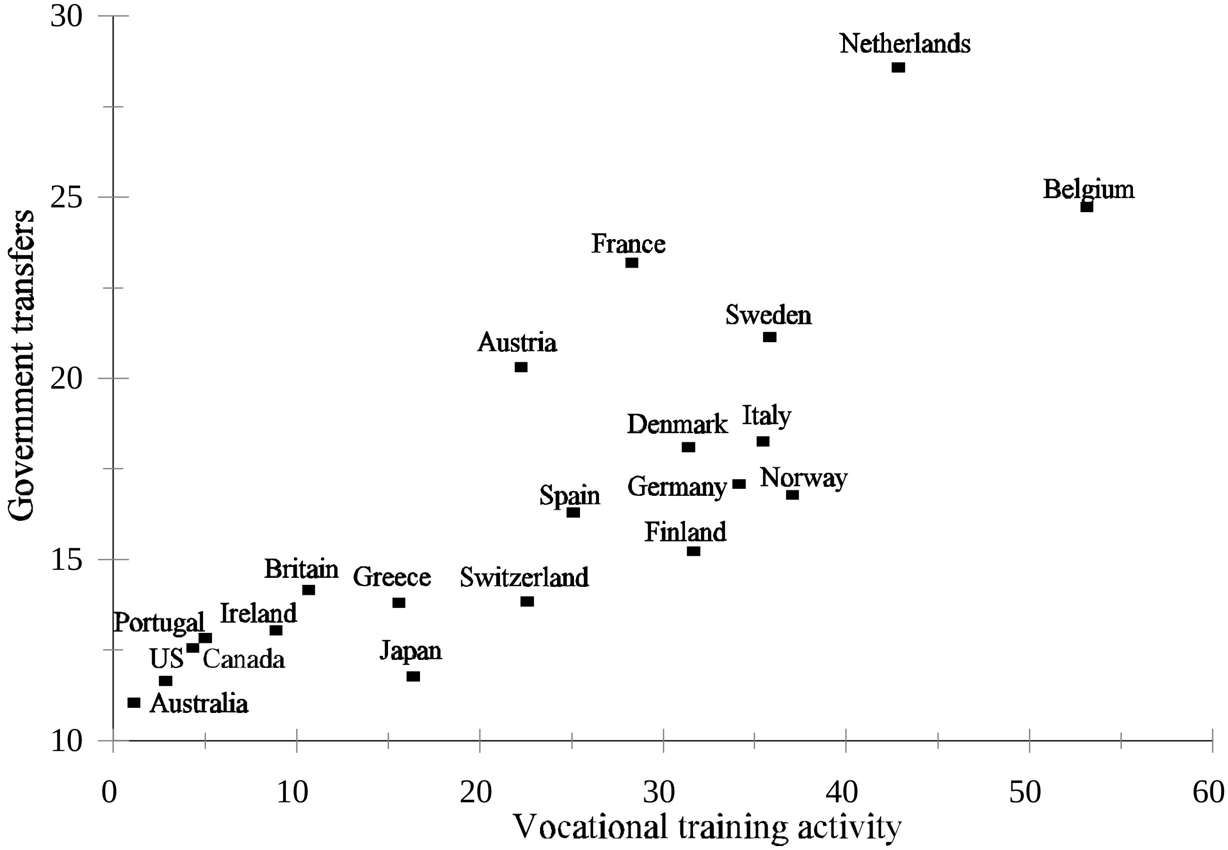


Figure 2. Four Models of Social Policy Preferences



Note: Arrows indicate preferred level of social protection by the median voter.

Figure 3. Vocational Training Intensity and Government Transfers



Notes: Government transfers is all government payments to the civilian household sector (including social security transfers, government grants, public employee pensions, and transfers to non-profit institutions serving the household sector) as a percent of GDP. *Sources*: Cusack (1991) and OECD, *National Accounts* (various years). Vocational training activity is the share of an age cohort in either secondary or post-secondary (ISCED5) vocational training. *Source*: UNESCO (1999).