

An Asset Theory of Social Policy Preferences

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We present 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 weigh more prominently. We test the theory on public opinion data for eleven advanced democracies and suggest how differences in educational systems can help explain cross-national differences in the level of social protection.

It is well known 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 only the cognitive aspects of education have received systematic attention in explaining political preferences (Duch and Taylor 1993; Kitschelt 1991; Klingemann 1979).

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 similar to the way transaction cost economics explains the use of nonmarket 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 it is difficult to withdraw these assets in response to adverse market conditions, firms want protection against the effects of such adversity. 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 nonmarket protection. Skills that are portable, by contrast, do not require extensive nonmarket protection, just as the exchange of homogeneous goods does not require elaborate nonmarket governance structures.³

Our theory does not necessarily contradict a long tradition in the literature that emphasizes redistribution as a key political motive behind the welfare state (e.g., Esping-Andersen 1990; Korpi 1983). Indeed, Meltzer and Richard's (1981) influential work on the median voter and 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 are workers with lower income. Our argument parts ways with the Meltzer-Richard model, however, because we explicitly recognize that social protection also has an insurance aspect (Moene and Wallerstein 2001; Sinn 1995) and that demand for insurance varies among workers according to their degree of exposure to labor market risks (Baldwin 1992). A critical point is that in our model exposure to risk is inversely related to the portability of skills. We show that this proposition has strong support in public opinion data from eleven advanced democracies, with important implications for explaining differences in the level of social protection across countries.

The remainder of this article is divided into three sections. In the first we present the model and its main

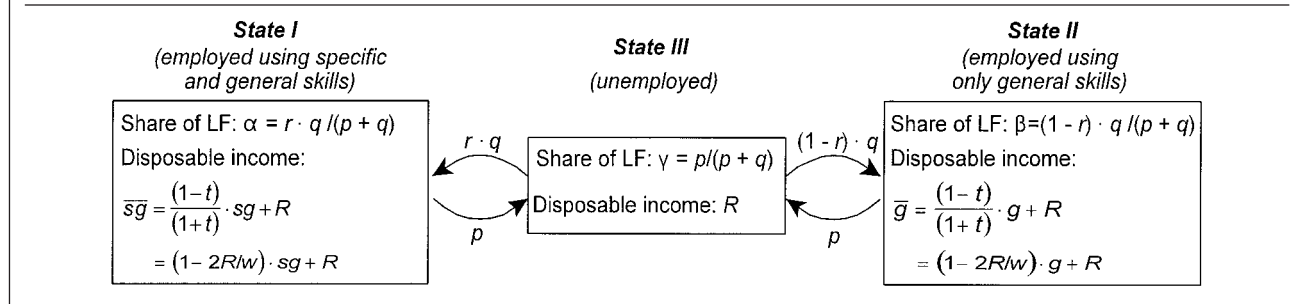
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¹ The trade policy argument treats skilled labor as a more or less abundant factor, and relative abundance is hypothesized to favor free trade. See Rogowski 1987, Frieden 1991, and especially Scheve and Slaughter 1999 for different versions of this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

² Alt and his colleagues (1999) provide 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.

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

empirical implications. In the second we test these implications on public opinion data from eleven countries. In the third we discuss 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 valuable only to a single firm or a group of firms (whether an industry or a sector), whereas general skills are portable across all firms. We distinguish three 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 uses both her specific and general skills; in State II the worker is employed in a firm that only uses his general skills; and in State III the worker is unemployed (i.e., none of her skills are being used).

We define by g the market value of a worker's general skills in State II when his 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 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 skill specificity).

In addition to market income, workers receive transfer income from the government, hereunder unemployment benefits, health care benefits, pensions, and other forms of nonwage compensation. Although some of these benefits are received by people outside the labor market, what matters to our argument is that they are viewed by workers as part of their compensation (in the neocorporatist literature, sometimes termed the "social wage"). Those 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, also will 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, Iversen, and Soskice (2001), one can distinguish between transfers that go to support the income of employed workers, or *Wage Protection*, and transfers that go to the unemployed, or *Unemployment Protection*.⁵ In developing our model we will discuss what happens if R only goes to unemployment protection, but ultimately we assume that all workers receive the same flat-rate subsidy, which may simply be termed *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:

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

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 pretax earnings. Then, tax income per capita is

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

Figure 1 illustrates the three states of the labor market and shows the disposable (aftertax) income associated with each: I: \bar{sg} , III: R , and II: \bar{g} . For a given period, there is a probability, p , of losing one's job and another probability, q , of reemployment. In equilibrium $p \cdot e = q \cdot \gamma$, where e is the share of employed workers, and γ is the share of the workforce that is unemployed ($e = 1 - \gamma$). 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 that uses both general and specific skills), then the equilibrium share of the labor force employed in State I is

⁴ This is also a realistic assumption, since aftertax income is distributed significantly more equally than pretax income (see Gottschalk and Smeeding 2000; Huber and Stephens 2001).

⁵ A third type identified by Estevez-Abe, Iversen, and Soskice (2001) is employment protection, which refers to legal and other barriers to layoffs. This type could be modeled as the probability of keeping a job that fully uses a worker's skills.

$$\alpha = r \cdot e = r \cdot q/(p + q). \quad (3)$$

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

$$\beta = (1 - r) \cdot e = (1 - r) \cdot q/(p + q). \quad (4)$$

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

$$\gamma = p/(p + q). \quad (5)$$

For any individual worker with both specific and general skills, the proportions α , β , and γ 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. If we ignore the discounting of future income (which makes no substantive difference to our results), this is captured by the following utility function:

$$V = \alpha \cdot u(\bar{s}g) + \beta \cdot u(\bar{g}) + \gamma \cdot u(R), \quad (6)$$

where $u(\cdot)$ 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 :

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

A number of the results below hold for this general form of utility function (notably the Meltzer-Richard results). Yet, 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,

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

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

Optimizing Social Preferences

We first consider a simple baseline model with no insurance effects, no tax disincentives, only general skills, and no unemployment (case 1). We then introduce tax disincentives to get the Meltzer-Richard result (case 2) and subsequently add insurance effects (and unemployment) to explore the effects of risk aversion (case 3). Finally, we show what happens to the demand for social protection when the composition of skills is allowed to vary (case 4). To keep the presentation simple, all proofs are in Appendix A.

Case 1: 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 ($e = 1$). In the simplest case 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:

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

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

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

and because $0 \leq R \leq w$, $g > w$ implies that V_R is uniformly negative for all values of R and hence t . Therefore, maximization of V requires $t = 0$. Thus, voters with income above average will choose a zero tax rate. An analogous argument for $g < w$ implies that voters with income below average 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 $g > w$. 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 2A.

Case 2: Disincentive Effects, No Insurance Effects: The Meltzer-Richard Model. If we now include the tax disincentive effect that $l(t) = 1/(1+t)$, we have:

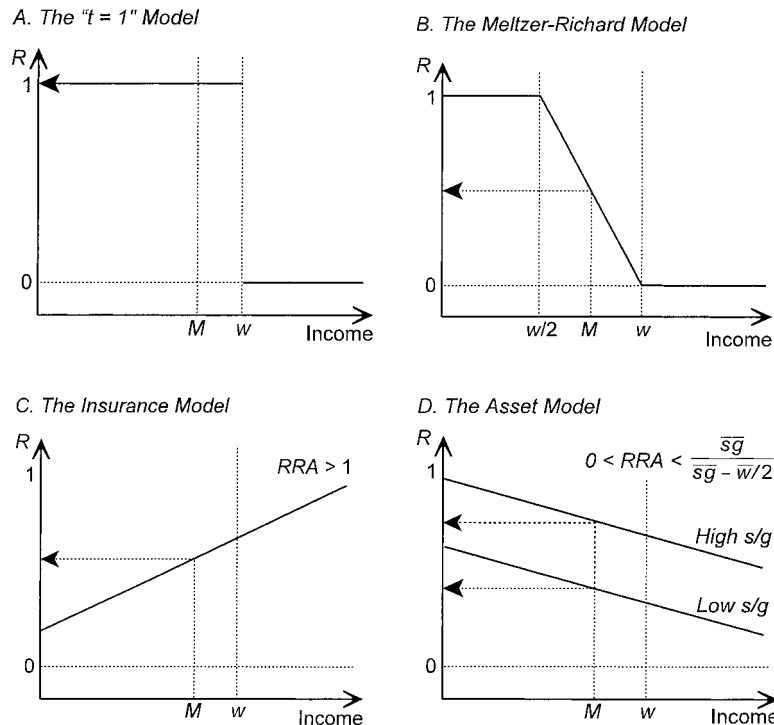
$$\begin{aligned} V &= u\left(\frac{1-t}{1+t}g + \frac{tw}{1+t}\right) = u(g(1 - 2R/w) + R) \\ &= \frac{1}{1-a} \cdot (g(1 - 2R/w) + R)^{1-a}. \end{aligned} \quad (9)$$

This implies

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

so in the Meltzer-Richard model, only voters with a g below that of half the average hourly wage ($g = w/2$) will vote for a maximum tax rate. As illustrated in Figure 2B, if the median voter has a g above $w/2$ he will not vote for the maximum tax rate. Because of the simplicity of our tax disincentive function, voters with g below $w/2$ will vote for $t = 1$, and voters with g above $w/2$ will vote for $t = 0$.⁶ With the more complex tax disincentive function used by Meltzer and Richard, workers with income in the range $[w/2, w]$ will prefer taxation up to the point at which 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,

⁶ The Meltzer-Richard model has a more general tax disincentive function than that used here. In consequence the tax rate that maximizes tax revenue can be less than 1.

FIGURE 2. Four Models of Social Policy Preferences

Note: R = preferred level of social spending; M = income of median voter; w = mean income; RRA = relative risk aversion; s/g = specific to general skills ratio. Arrows indicate preferred level of social spending by the median voter.

then she may vote for a positive tax rate less than 1 (as illustrated in Figure 2).

One implication of the Meltzer-Richard model is that voter turnout will be positively related to the level of government transfers because nonvoting tends to be concentrated among low-income people (Lijphart 1997). There is some cross-national evidence in support of this proposition (see Franzese 1998), but there is little empirical support for another key implication of the model, developed by Alesina and Rodrik (1994): Relatively inequalitarian 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).

Case 3: Disincentive Effects, Insurance Effects: The Insurance Model. Moene and Wallerstein (2001) offer one possible explanation for this puzzle by introducing insurance effects (see also Sinn 1995). For insurance effects to matter, we need at least two states of the world. In contrast to the Moene and Wallerstein (2001) model, in which there are high- and low-wage groups in addition to the unemployed, we now assume that workers can either be employed at a gross wage equal to their “tax-incentivized” skill level $g/(1+t)$ or be unemployed. There are no specific skills ($s = 1$), so equation 6 becomes:

$$V = \beta \cdot u(\bar{g}) + \gamma \cdot u(R). \quad (11)$$

In this model, if relative risk aversion is constant and greater than unity, then 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. Thus, the relationship between income and preferred level of spending is positive, as can be seen in Figure 2C.⁷

To get this result, we assume that R is only paid to those who are unemployed, so that $\bar{g} = g \cdot (1 - 2R/w)$. It can then be shown (as we do in Appendix A) that

$$dR/dg < 0 \quad \text{iff} \quad RRA < 1, \quad (12)$$

and

$$dR/dg > 0 \quad \text{iff} \quad RRA > 1. \quad (13)$$

A key implication of the result for $RRA > 1$ is that, contrary to the Meltzer-Richard model, provided that the income distribution is skewed to the right, a means-preserving increase in inequality will reduce the median voter’s preferred level of social protection. The

⁷ This differs from Moene and Wallerstein (2001) because in their model workers with high income are assumed not to be exposed to labor market risks and hence prefer zero spending and taxation. Since we do not make any distinctions in the risk exposure of different income groups, we refer to case 3 as the insurance model rather than the Moene-Wallerstein model. In the empirical section we test this insurance model, not the Moene-Wallerstein model (which, like our own argument, is consistent with both positive and negative effects of income when tested across the entire income scale).

reason is that 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 is thus one potential explanation for the empirically puzzling link between income equality and redistributive spending.

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

Case 4: 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:

$$V = \alpha \cdot u(\bar{s}g) + \beta \cdot u(\bar{g}) + \gamma \cdot u(R), \quad (6)$$

where

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

and

$$\bar{g} = g \cdot (1 - 2R/w) + R.$$

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

$$y \equiv \alpha \cdot sg + \beta \cdot g.$$

We then proceed to ask three questions of the model. First, up to what value of y is the chosen R maximal, that is, $t = 1$? Second, under what RRA conditions does R fall or rise as y rises above this value? 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 1: *Given the assumptions of the asset model, $t = 1$ iff $y \leq w/2$.*

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 2: *Given the assumptions of the asset model, holding s constant and with $y > w/2$,*

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

What this equality says is that the direction of the relationship between R and income (the sign, sgn , of $\partial R / \partial y$) depends on the level of risk aversion, just as in

the simple insurance model. In order for income to be positively related to support for spending, however, the RRA requirement is more stringent ($RRA > \bar{s}g/(\bar{s}g - \bar{w}/2)$) 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 insurance 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 (2001).

Now we come to the critical result that differentiates our approach from previous ones. Central to our argument 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 3: *Assuming a constant relative risk aversion utility function and $RRA > 0$, $\partial R / \partial s > 0$, 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. Workers with specific skills who are laid off face the risk of being reemployed in a sector in which 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 is the demand for income protection (R). The logic, as illustrated in Figure 2D, implies that the median voter's support for social protection depends on the composition of her skills.

Summary. 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 or class politics. 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 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, 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, then rising income may be associated with increased demand for protection, since high-

income workers have more to lose than low-income workers. This is the simple insurance model. If some transfers go to the employed, however, then the threshold of risk aversion for which this relationship holds rises, since transfers to the employed only serve redistributive purposes.

Finally, when skill specificity is introduced, which requires at least two employment states (State I and State 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 by empirical evidence from public opinion surveys in eleven advanced democracies.

TESTING THE MODEL

Statistical Model

In our asset model the relationship between the “preferred” level of R and the two exogenous variables, y (expected income) and s (skill specificity), is given by two implicit equations:

$$V_R(R, s, g) = 0, \text{ and } y = \alpha \cdot sg + \beta \cdot g. \quad (14)$$

From equation 14 we derived result 2,

$$\frac{\partial R}{\partial y} < 0 \quad \text{if} \quad 0 < RRA < \frac{\bar{sg}}{\bar{sg} - \bar{w}/2},$$

and result 3,

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

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 equation 14. Thus, our regressions take the form

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

By implication, if our estimate of b is significantly different from zero and negative, we can infer that $0 < RRA < \bar{sg}/(\bar{sg} - \bar{w}/2)$. 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.⁸

More generally, the asset model encompasses the other three. Hence, we can test for these models as

well. The first model (Meltzer-Richard without tax disincentives) implies that $b = c = 0$. The Meltzer-Richard model, with tax disincentives, implies that $b < 0$ and $c = 0$. The insurance model, with

$$RRA > \frac{\bar{sg}}{\bar{sg} - \bar{w}/2},$$

implies that $b > 0$ and $c = 0$.⁹

The Data

We use data from eleven advanced democracies obtained from two sets of national mass surveys conducted under the auspices of the International Social Survey Program (ISSP), one in 1996 and the other in 1997 (ISSP 1999, 2000).¹⁰ These surveys offer by far the best individual-level data on skills and preferences for social protection. We supplement this information with economywide unemployment data. The next two sections 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 appear in a cluster that asks whether the respondent would like to see more or less government spending on unemployment benefits, health care, and pensions (see Appendix C for details). Pertinent to an assumption in our model is the fact that respondents were warned that more spending may require higher taxes. The fourth variable is based on an item 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 those with specific skills to be more concerned than those with general skills about keeping their present job and income. Moreover, although respondents were 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 about 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. It is often claimed that better general education is associated with increased support for spending on “postmaterialist” activities, whereas our theory says that it is related to reduced support for spending in

⁹ Any general insurance model is, of course, consistent with ours. The hypothesis about the insurance model we are testing is the particular version (case 3) together with the assumption of $RRA > \bar{w}g/(\bar{sg} - \bar{w}/2)$. This solves the inequality-spending puzzle by implying a positive relationship between income and support for redistributive spending.

¹⁰ The countries are: Australia, Britain, Canada, France, Germany, Ireland, Italy, the Netherlands, Norway, Sweden, and the United States. Japan, covered in both ISSP surveys, could not be included because of missing data on a key occupational variable (explained below).

⁸ 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.

the social policy area (cf Kitschelt 1991).¹¹ Because one might object that our findings for skills reflect general ideological opposition to government spending among those with more years of formal education, it is useful to show that the relationship between skills and support for spending varies by policy area.

For economy of presentation we use confirmatory factor analysis (CFA) to create two 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 is 0.94 and varies little by country (the range is 0.90–0.98).¹²

Independent Variables. Our two different approaches to the measurement of skill specificity reflect different aspects of the theoretical model. The first classifies 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 in which one's skills are needed is proportional to their specificity. This is an attempt to gauge s indirectly through rq : the probability of reemployment into State I.

The first approach is based on the latest detailed classification of occupations by the International Labor Office: the International Standard Classification of Occupations (ISCO-88). ISCO-88 bases classification on two criteria: the level of skills required for an occupation and the degree of specialization of those skills. A distinction is made among four broad skill levels, which are a function of "the range and complexity of the tasks involved" and explicitly depend on informal as well as formal training (ILO 1999, 6). Skill level thus corresponds to $(s + g)$ in our model. All other distinctions 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" (p. 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.¹³ At the most disaggregated level, called the unit level, there are 390

occupational categories (unit groups) with highly specific job descriptions.¹⁴

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 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 (1) the size of the labor market segment captured by that class and (2) the degree of skill specialization of occupations found in that particular labor market segment. For example, 8% of the workforce across our countries is classified as "plant and machine operators and assemblers" (major group 8), but this group accounts for 70 of the 390 unit groups, or 18%. If occupations at the unit level are, on average, equally homogeneous in terms of skills, then the disproportionate share of unit groups in major group 8 will reflect a high degree of skill specialization within that major group. By dividing the share of unit groups (.18) by the share of the labor force (.08), we can generate a measure of the average skill specialization within that particular major group (2.1). This calculation also can be done at the lower submajor level, and we have used the mean of these calculations to get proxy for s .¹⁵ The resulting variable has 27 values, ranging from 0.4 to 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.¹⁶ This gives us a proxy for $s/(s + g)$ which we will refer to as *Skill Specificity 1* or 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 *Skill Specificity 2* or 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

¹¹ Duch and Taylor (1993) make a similar argument concerning postmaterialist attitudes (although they do not directly discuss spending).

¹² We used LISREL version 8.5 for the CFA and the resulting factor loadings to construct the two indexes. The model was estimated from the covariance matrix for six spending items, assuming that these are indicators of two latent spending variables: social and postmaterialist spending. The factor loadings for the latent variable are as follows: social spending—.52 (subsidies to protect jobs), .48 (health insurance), .58 (pensions), and .55 (unemployment insurance); postmaterialist spending—.58 (environment) and .51 (culture and the arts). Alternatively, the indexes can be constructed from the results of fitting CFA to the covariance 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.

¹³ 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 will tend to have more minor and unit groups.

¹⁴ Unit group 3144, for example, represents "air traffic controllers," within the minor group "ship and aircraft controllers and technicians," itself one of five categories in the major group "technicians and associate professionals."

¹⁵ 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. Yet, the greater variance of the measure at lower levels of aggregation helps reduce the standard error on the estimated parameter for the skill variable.

¹⁶ 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.

TABLE 1. Summary of Independent Skill Variables

Variable Name	Definition	Intercorrelations				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 ^a	—				
s_2	(Share of ISCO-88 level 4 groups)/(share of labor force) divided by level of general education ^a	.82	—			
s_3^b	Response to question about difficulty of finding an acceptable job	.38	.53	—		Not clear whether this is a measure of absolute or relative share of skills
s_4^b	s_3 divided by level of general education	.66	.59	.62	—	Assumes that s_3 measures absolute skills (although s_4 will always be a relative measure)

^aShares are calculated at both the first and second ISCO-88 level and then averaged.

^bBefore calculating the intercorrelations, the number of categories for s_3 and s_4 were reduced to the same number as for s_1 and s_2 .

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 that precisely taps the assessment of outside options: "If you were looking actively, how easy or difficult do you think it would be for you to find an acceptable job?" The response choices were: "very easy," "fairly easy," "neither easy nor difficult," "fairly difficult," or "very difficult." The difficulty of finding an acceptable job is likely to be related to skill portability. High specificity means that relatively few jobs use those skills, and the number of openings is also likely to be small 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 home, which is also a likely component of what people consider "acceptable," depends on the number of openings in a given geographical area.¹⁷ All else equal, the item about finding an acceptable job is thus 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 about as good a measure of rq as is possible. We refer to it as *Skill Specificity 3* or s_3 .

There is an ambiguity in the relationship of s_3 to the theoretical concept of s . We cannot know for certain whether survey responses reflect a person's absolute level of specific skills or the *relative* share of his skills that are 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 *Skill Specificity 4* or 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 intercorrelations are listed in Table 1. Not surprisingly, the correlations are higher between measures that use either the survey question or the ISCO classification. The lowest correlations are between s_3 and either 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 coworkers) 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.

One final methodological issue needs to be addressed. Because the item 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 always be calculated from s_3). For this purpose we calculated averages for s_3 at the three-digit ISCO-88 level in the 1997 survey and then assigned these values to individuals in the 1996 survey based on their three-digit ISCO classification in that survey.¹⁸ Since the classification of occupations is based on the skills required, it is reasonable to expect that 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,¹⁹ we show in Appendix D that s_3 , averaged by

¹⁷ In a 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 not recorded by the ISSP survey.

¹⁸ There are 116 unique groups at the three-digit level. The more fine-grained four-digit level is not available for some countries and sometimes contains a large number of empty categories.

¹⁹ With the minor qualification that people who turned 18 in the interim were not eligible for the 1996 survey.

ISCO three-digit groups, is an unbiased estimator for the original variable.

In addition to the skill variables (s_1 – s_4), we used self-reported pretax and pretransfer income as a proxy for y (converted into dollars at 1996 exchange rates and square root transformed), as well as the following set of controls.

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

Gender. As argued by Orloff (1993) and Estevez-Abe, Iversen, and Soskice (2001), women may demand more protection than men in comparable jobs because they need to not only leave but also return to the labor market for the purpose of child rearing.

Union Membership. A main function of unions is to insure members against labor market risks, so it is reasonable to expect that union members are particularly concerned with social protection (see, e.g., Korpi 1989).

Part-Time Employment. On the one hand, 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, they depend more on flexible labor markets to generate nonstandard jobs, which suggests a countervailing effect.

Nonemployed. Esping-Andersen (1999) argues that some outsiders 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 nonemployed have very different attitudes from the employed.

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

Self-Employment. The self-employed are expected to favor free markets and a low level of social protection because they 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 can be argued that better informed people may have adopted the predominant view, which tended to see budget cuts as necessary on efficiency grounds (corresponding to a higher cost of distortionary taxation in our model).²⁰ Information is measured by respondents' subjective understanding of politics (see Appendix C for details).

Left-Right Position. Attitudes toward social protection may partly reflect ideological predisposition or perhaps the socializing effects of political parties.²¹ We control for this possibility by including positions on a Left-Right scale based on the respondent's declared party support (see Appendix C for details).

National Unemployment. Although our theory implies that people discount cyclical unemployment, it still may affect 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 $U_j \cdot y_{ij}$ and $U_j \cdot s_{ij}$ 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 effect of national macroeconomic conditions).²² To cope with problems of missing observations we used a multiple imputation technique developed by Honaker and others (1999). This strategy is superior to the traditional approach of listwise deletion, which is both inefficient and potentially biased (King et al. 2001).²³ The following presentation is divided into a section on the key results and a section on their robustness, which also discusses potential objections to the way we interpret the results.

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. Because the Italian survey was conducted in 1990 and lacks information on several of the control variables, it was not included in the calculation. In the next section we show that the results for Italy are consistent with those presented in Table 2.

The model in column 1 uses the average of the four measures of skills, *Skill Specificity Composite* or $s_{\text{composite}}$, 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

²⁰ This idea was suggested by an anonymous reviewer.

²¹ 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 .

²² All data analysis was done using Stata 6.0 for Windows.

²³ 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 that method, but the standard errors are also larger.

TABLE 2. Support for Social Spending among the Publics of Ten Countries, 1996

	(1)	(2)	(3)	(4) ^a	(5) ^a	(6) ^b
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)
$S_{\text{composite}}$	0.233** (0.014)	—	—	—	—	0.219** (0.013)
Skill Specificity 1 (s_1)	—	0.148** (0.010)	—	—	—	—
Skill Specificity 2 (s_2)	—	—	0.150** (0.010)	—	—	—
Skill Specificity 3 (s_3)	—	—	—	0.105** (0.013)	—	—
Skill Specificity 4 (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 employment	-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)
Nonemployed	-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)
Left-Right 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)
$U_j \cdot y_{ij}$	-0.0002** (0.0001)	-0.0002** (0.0001)	-0.0002** (0.0001)	-0.0003** (0.0001)	-0.0004** (0.0001)	-0.0003** (0.0001)
$U_j \cdot s_{ij}$	-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^2	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

Note: All regressions include a full set of country dummies (not shown). Standard errors are in parentheses. * $p = .05$, ** $p = .01$.

^aExcludes Australia, Ireland, and Italy, for which data are not available.

^bExcludes Australia, for which union membership data are not available.

risk aversion is not sufficiently high to make public demand for transfers rise with income. Technically speaking, $RRR < \bar{s}g/(\bar{s}g - \bar{w}/2)$, which means that the Meltzer-Richard redistribution logic dominates the insurance logic. As expected, the relationship is little affected by differences in national unemployment rates, despite their considerable variation in the survey year. Thus, an increase of one standard deviation 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 $RRR > 0$). 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 $U_j \cdot s_{ij}$.²⁴ The parameter for s_3 is lower than for the other measures, which is not entirely unexpected, because this variable may capture absolute rather than

²⁴ For example, an increase of one standard deviation in U_j only reduces the effect of $s_{\text{composite}}$ from .23 to .22.

TABLE 3. Estimated Magnitude of the Effects of Independent Variables

	Proportion of Explained Variance ^a		Effect of One SD Change ^b	
	Lower Bound ^c	Upper Bound ^d	95% Confidence Interval Upper Bound	Lower Bound
Income	11	51	-0.22	-0.19
$s_{\text{composite}}$	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
Nonemployed	0	8	-0.03	-0.01
Self-employed	2	8	-0.07	-0.05
Informed	1	8	-0.05	0.04
Left-Right party support	5	5	-0.09	-0.07
$U_j \cdot y_{ij}$	1	13		
$U_j \cdot s_{ij}$	0	8		
Income + $s_{\text{composite}}$	38	73	0.38	0.44
All controls combined	27	52	0.36	0.47

^aIncrease in explained variance by each variable as a proportion of the total explained variance of all (nondummy) variables (based on model 6 in Table 2).

^bThe change in support for social spending (measured in standard deviations) as a result of an increase of one standard deviation in each independent variable (in the cases of income and $s_{\text{composite}}$, 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).

^cIncrease in explained variance (compared to model with only country dummies) when each variable is included as the last variable.

^dIncrease in explained variance when a variable is included as the first variable.

relative endowment of specific skills (or a combination of both). For all correlations between s and g that are greater than -1 , absolute measures of s will yield lower parameter estimates than relative measures.²⁵

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 estimate the portion of the explained variance for each independent variable as well as the effect on preferences of a change in each variable of one standard deviation. The estimates are based on the results of model 6 in Table 2, which includes all the relevant variables.

Although we cannot precisely attribute the proportion of explained variance to each independent variable, we can calculate the likely ranges. The upper bound is found by recording the increase in explained variance (measured as a percentage of the total) 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 bound is calculated as the increase in explained variance (as a percentage of the total) when a variable is entered as the last predictor. This proce-

sure eliminates all hypothesized individual-level spurious effects of the variable, but it 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 factors in explaining social policy preferences among the variables included in this analysis. Income accounts for between 11% and 51% of the total explained variance, and skills account for between 26% and 38%. Jointly, they capture between 38% and 73%, with the rest accounted for by the controls.

The key role of income and skills is confirmed when we consider the effect of a change of one standard deviation in these variables (column 3). Such a change in either variable is associated with about 20% 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 influence of income and skills is as great as the joint effects 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% confidence interval).

The results for the controls also generally confirm our expectations. Individuals who are particularly ex-

²⁵ 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.

posed to labor market risks—the unemployed, women, and older workers—are more favorably disposed to an increase in social spending than are 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 likely to oppose spending, which may reflect a political environment at the time that was hostile to the welfare state. Supporters of rightist parties, not surprisingly, also express less support for social spending than supporters of leftist parties. Finally, the attitudes of part-time employees and those outside the labor market are indistinct from the attitudes of others. These groups are evidently too heterogeneous to share any common interest in social policies.

Gender stands out among the control variables, accounting for between 8% and 17% of the total explained variance, and it has the greatest effect among the controls. As argued by Estevez-Abe, Iversen, and Soskice (2001), women require more protection than men in comparable jobs because of maternity leave, but 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. Because women know they are likely to leave their job before they can reap the full returns on specific skill investments, they are dissuaded from making such investments (Estevez-Abe, Iversen, and Soskice 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 for men ($t = 14$). In other words, when women invest in specific skills they are more likely than men to support a high level 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. We first note that the findings for y and s stand up to any combination of the controls we included, and they are robust to the inclusion of any other variable used in the survey (region, public sector employment, urbanization, and supervisory position) in any combination.²⁶ Although 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 estimated parameters may even reverse in particular cases. In addition, pooling usually yields exaggerated t -scores compared to those found for individual countries.²⁷ Therefore, we ran our regressions on each of the countries individually. The results for the theoretical variables are shown in Table 4.

²⁶ None of these was used in every survey, so rather than clutter the presentation with several additional columns, we omitted them from the main analysis.

²⁷ 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 , because a squared term is added on the top, and 1 is added to the bottom (although it does not have to be so).

Note that every regression yields results that are consistent with the pooled analysis; each of the 60 parameters records the correct sign, and most are significant at the .01 level or better. The composite skill variable is always significant at .01 or better, and for nine of the countries the parameter estimates for s vary in a fairly narrow range between 0.16 and 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 $s_{\text{composite}}$ in both cases is 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, which potentially dilutes the skill distinctions among categories.

As in 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 United States) is there 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 capture an ideological aversion to government spending among those with higher education. Two measures of s have formal education in the denominator, and the other two implicitly assume it. In quantitative terms, general education accounts for roughly one-third of the variance in $s_{\text{composite}}$. It is therefore conceivable that the proxies for skills may partly capture an ideological effect of higher education. For example, much of the economic theory taught to university students during the 1990s emphasized the efficiency of free markets over state intervention.²⁸

To some extent we control for this possibility through the variables for self-assessed level of information and party support. If the highly educated consider themselves better informed about the costs of generous social spending, this is likely to show up in the information variable. Likewise, those who are ideologically committed to a small welfare state are presumably more likely to support rightist 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.

There still may 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, which means that $s_{\text{composite}}$ will only pick up the effects of specific skills. In this setup we would expect formal education to have the opposite effect of 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. We can be certain, however, that what-

²⁸ In terms of the formal model, this can be captured by different assessments of the distortionary effects of taxation.

TABLE 4. Income, Skills, and Support for Social Spending in Eleven Countries

	Income ^a	S _{composite}	S ₁	S ₂	S ₃	S ₄	N
Australia	-0.0030** (-0.0004)	0.156** (0.027)	0.129** (0.023)	0.127** (0.026)	n.a ^b	n.a	2,151
Britain	-0.0029** (-0.0006)	0.219** (0.042)	0.105** (0.026)	0.121** (0.027)	0.135** (0.039)	0.181** (0.046)	989
Canada	-0.0054** (-0.0008)	0.219** (0.053)	0.102* (0.042)	0.140** (0.045)	0.087* (0.041)	0.220** (0.047)	1,182
France	-0.0055** (-0.0007)	0.235** (0.038)	0.158** (0.035)	0.147** (0.028)	0.097** (0.037)	0.111** (0.026)	1,312
Germany	-0.0027** (-0.0007)	0.255** (0.030)	0.182** (0.028)	0.155** (0.023)	0.115** (0.035)	0.212** (0.028)	2,361
Ireland	-0.0030** (-0.0008)	0.116** (0.028)	0.092** (0.029)	0.122** (0.027)	n.a	n.a	994
Italy	-0.0021 (-0.0016)	0.105** (0.037)	0.104* (0.040)	0.095** (0.034)	n.a	n.a	983
Norway	-0.0026** (-0.0006)	0.257** (0.033)	0.139** (0.024)	0.165** (0.027)	0.076** (0.027)	0.250** (0.030)	1,344
New Zealand	-0.0049** (-0.0007)	0.176** (0.038)	0.125** (0.030)	0.094** (0.030)	0.134** (0.041)	0.131** (0.034)	1,198
Sweden	-0.0055** (-0.0010)	0.282** (0.039)	0.130** (0.030)	0.140** (0.028)	0.156** (0.033)	0.278** (0.030)	1,238
United States	-0.0024* (-0.0010)	0.294** (0.052)	0.192** (0.030)	0.189** (0.037)	0.042 (0.036)	0.272** (0.058)	1,332

Note: All regressions include the same set of controls as in Table 2, column 1. Standard errors are in parentheses. * $p = .05$, ** $p = .01$.

^aThe effect of income is only shown for S_{composite} but varies little across the four measures of s .

^bData are not available to estimate this parameter.

ever effect remains of s , it cannot be attributed to general education.

The first column of Table 5 shows the results of reestimating 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, which is consistent with the skill asset argument. More important, the parameter on the specific skills variable remains positive and statistically significant. It is not surprising that the effect of s falls from 0.23 to 0.14, but this is still a very considerable influence. Even if we completely discount the effect of general education as a measure of general skills, therefore, the results lend unambiguous support to our argument.

Yet, 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. Surely, if highly educated individuals believe in the efficiency of free markets and the waste of government spending, then they also should oppose public spending on the environment, culture, and the arts.²⁹ 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 in these areas than are 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. Workers with general skills demand little social protection but are enthusiastic consumers of a clean environment and state-subsidized culture. Workers with specific skills are deeply concerned about social protection but not enthusiastic about state subsidization for the environment and the arts. There is no blanket support for, or opposition to, government spending among any particular group of workers.

CONCLUSION

It is well known 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,

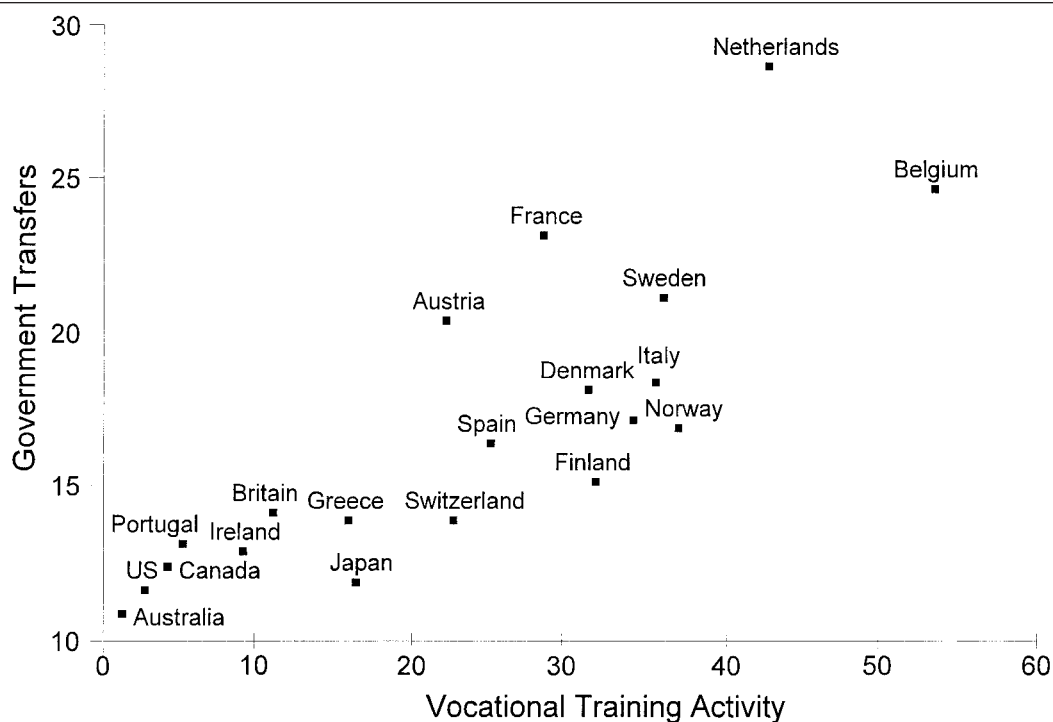
internalized a general aversion to government spending through their educational experience.

²⁹ "The environment" may be conceived as a collective good that improves overall welfare (it is a little harder to argue this with respect to fine arts subsidies), 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); rather, they may have

TABLE 5. Formal Education and Support for Two Types of Spending in Ten Countries, 1996

	Support for Social Spending	Support for Postmaterialist Spending	
Formal education	-0.105** (0.008)	0.130** (0.007)	—
$S_{\text{composite}}$	0.143** (0.015)	—	-0.097** (0.014)
Income	-0.0027** (0.0002)	-0.0004 (0.0002)	0.0004 (0.0002)
Age	0.0015* (0.0006)	-0.0064** (0.0006)	-0.009** (0.001)
Gender (female)	0.203** (0.018)	0.092** (0.018)	0.087** (0.019)
Part-time employment	-0.025 (0.028)	0.104** (0.029)	0.112** (0.030)
Unemployed	0.303** (0.040)	0.085* (0.043)	0.089* (0.044)
Nonemployed	-0.067** (0.025)	0.077** (0.024)	0.093** (0.025)
Self-employed	-0.243** (0.028)	0.021 (0.025)	0.011 (0.026)
Informed	-0.031** (0.008)	0.069** (0.009)	-0.083** (0.009)
Left-Right party support	-0.049** (0.004)	-0.060** (0.005)	-0.059** (0.005)
$U_j \cdot Y_{ij}$	-0.0002 (0.0001)	0.0000 (0.0001)	0.0001 (0.0001)
$U_j \cdot S_{ij}$	-0.006 (0.005)	0.008 (0.005)	0.007 (0.006)
Adj. R^2	0.22	0.09	0.07
N	14,101	14,101	14,101

Note: Regressions include a full set of country dummies. Standard errors are in parentheses. * $p = .05$, ** $p = .01$.

FIGURE 3. Vocational Training Intensity and Government Transfers in Twenty Countries

Sources: For government transfers, Cusack 1991 and OECD, *National Accounts*, various years. For vocational training, UNESCO 1999.

Note: Government transfers are all government payments to the civilian household sector (including social security, government grants, public employee pensions, and transfers to nonprofit institutions serving the household sector) as a percent of GDP. Vocational training activity is the number of people in secondary vocational training as a percentage of all those in the secondary school age cohort plus the number of people in postsecondary (ISCED5) vocational training as a percentage of all those in the postsecondary school age cohort.

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 do workers who have portable skills. For this reason, workers with specific skills have a strong 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. At any given level of income, however, workers with specific skills are more inclined to support a high level of protection than are those with general skills. This tendency can help us understand cross-national variance in social protection, because the profile of skills is likely to vary in accord with the structure of a nation's educational system. Some countries provide much more training than others in vocational skills that are specific to particular jobs, firms, or industries. In Germany, for instance, more than one-third of any age cohort goes through extensive vocational training (3–5 years), whereas in the United States the comparable figure is only a few percentage points (even counting those in the junior college system). If these differences are reflected in political preferences, either through the electoral system or the system of interest representation, then a new explanation of the welfare state based on differences in national skill profiles may be appropriate.

The empirical plausibility of such an explanation is suggested by Figure 3, which uses the vocationally trained share of an age cohort 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 ($r = .82$), although future research is needed to eliminate the effects of potentially confounding variables.

Cross-national differences in skill composition also may explain why income equality is linked to higher social spending in cross-national comparisons. Since the extent of vocational training is strongly positively related to pretax income equality—the correlation coefficient is .73 using the earnings of a worker in the bottom decile of the earnings distribution divided by the earnings of a worker in the top decile of the earnings distribution (d9/d1 ratios) as a measure of equality (OECD n.d.)—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 income distribution.

Finally, our model points to an important source of cross-time variance in support for social protection: unanticipated shocks to the occupational structure. When workers invest heavily 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 explain 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, in the training system, and 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.

APPENDIX A: MATHEMATICAL PROOFS

Derivation of Equations 12 and 13

The choice of the optimal R requires that:

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

Totally differentiating both sides we get:

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

Since the denominator is negative,

$$\frac{dR}{dg} > 0 \quad \text{iff} \quad [\bar{g}u''(\bar{g}) + u'(\bar{g})] < 0.$$

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 equations 12 and 13 follow directly.

Proof for Result 1

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

From equation 6 the necessary condition for optimal R is

$$\alpha \cdot u'(\bar{s}g) \cdot \left(1 - \frac{2sg}{w}\right) + \beta \cdot u'(\bar{g}) \cdot \left(1 - \frac{2g}{w}\right) + \gamma \cdot u'(R) \geq 0. \quad (\text{A-1})$$

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

Proof for Results 2 and 3 of the Asset Model

The necessary condition for optimal choice of R is $V_R(R, s, g) = 0$. This is given by A-1. Totally differentiating V_R gives:

$$\begin{aligned} & \alpha \cdot \left[u''(\bar{s}g) \cdot \left(\frac{2sg}{w} - 1\right) \cdot \left(1 - \frac{2R}{w}\right) \cdot g + u'(\bar{s}g) \cdot g \cdot \frac{2}{w} \right] \cdot ds \\ & + \alpha \cdot \left[u''(\bar{s}g) \cdot \left(\frac{2sg}{w} - 1\right) \cdot \left(1 - \frac{2R}{w}\right) \cdot s + u'(\bar{s}g) \cdot s \cdot \frac{2}{w} \right] \cdot dg \\ & + \beta \cdot \left[u''(\bar{g}) \cdot \left(\frac{2g}{w} - 1\right) \cdot \left(1 - \frac{2R}{w}\right) + u'(\bar{g}) \cdot \frac{2}{w} \right] \cdot dg \end{aligned}$$

$$= \left\{ \alpha \cdot u''(\bar{s}g) \cdot \left(\frac{2\bar{s}g}{w} - 1 \right)^2 + \beta \cdot u''(\bar{g}) \cdot \left(\frac{2\bar{g}}{w} - 1 \right)^2 + \gamma \cdot u''(R) \right\} \cdot dR. \quad (\text{A-2})$$

Note that the term in braces on the right-hand side, which we will call B , is negative. Also, we can write $(\bar{s}g - w/2) \cdot (1 - 2R/w) = \bar{s}g - \bar{w}/2$. Furthermore,

$$\begin{aligned} & [u''(\bar{s}g) \cdot (\bar{s}g - \bar{w}/2) + u'(\bar{s}g)] \\ &= u'(\bar{s}g) \cdot \left[1 - RRA \cdot \frac{\bar{s}g - \bar{w}/2}{\bar{s}g} \right] \equiv u'(\bar{s}g) \cdot L(\bar{s}g), \end{aligned} \quad (\text{A-3})$$

so equation A-2 can be written:

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

Since $dy = \alpha \cdot g \cdot ds + \alpha \cdot s \cdot dg + \beta \cdot dg$, we can further rewrite equation A-4 as

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

To prove results 3 and 4, note that in terms of equation A-5 $\partial R/\partial y = dR/dy$ and $\partial R/\partial s = dR/ds$. We show first that $L(\bar{s}g) < L(\bar{g})$. From the definition in equation A-3, this follows if $s > 1$ —as is the case apart from purely general skills—and if $RRA > 0$. Result 3 is that $\text{sgn } \partial R/\partial y < 0$ if $RRA < \bar{s}g/(\bar{s}g - \bar{w}/2)$. Since $B < 0$, $L(\bar{s}g) < L(\bar{g})$, and $u'(x) > 0$, this follows from equation A-5 if $L(\bar{s}g) > 0$. This requires that $RRA < \bar{s}g/(\bar{s}g - \bar{w}/2)$. This is a sufficient condition; a necessary and sufficient condition is that the numerator of the second term in square brackets on the right-hand side of equation A-5 is positive.

Result 4 is that $\text{sgn } \partial R/\partial s > 0$. Since $B < 0$, this requires that the numerator in the first square bracket on the right-hand side of equation A-5 is negative. Since $u'(\bar{s}g) < u'(\bar{g})$ from diminishing marginal utility, a sufficient condition is that $L(\bar{s}g) < L(\bar{g})$, which is true so long as $RRA > 0$ and $s > 1$. So result 4 follows from the existence of risk aversion and specific skills.

APPENDIX B: DERIVING THE ESTIMATING EQUATION

We show that the estimating equation we use,

$$R = k + b \cdot y + c \cdot s, \quad (\text{B-1})$$

is equal to

$$R = k + \frac{\partial R}{\partial y} \cdot y + \frac{\partial R}{\partial s} \cdot s, \quad (\text{B-2})$$

where equation B-2 is a first-order Taylor expansion of $V_R(R, s, g) = 0$ and $y = \alpha \cdot s \cdot g + \beta \cdot g$ evaluated around $(R, s, g) = (\bar{R}, \bar{s}, \bar{g}) \equiv \bar{x}$.

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

$$R = K + \frac{V_{R,s}}{V_{R,R}} s + \frac{V_{R,g}}{V_{R,R}} g. \quad (\text{B-3})$$

In terms of equation A-5:

$$\frac{V_{R,s}(\bar{x})}{V_{R,R}(\bar{x})} = \frac{u'(\bar{s}g) \cdot L(\bar{s}g) \cdot \alpha \cdot \bar{g}}{(w/2) \cdot B}, \quad (\text{B-4a})$$

and

$$\frac{V_{R,g}(\bar{x})}{V_{R,R}(\bar{x})} = \frac{u'(\bar{s}g) \cdot L(\bar{s}g) \cdot \alpha \cdot \bar{s} + u'(\bar{g}) \cdot L(\bar{g}) \cdot \beta}{(w/2) \cdot B}. \quad (\text{B-4b})$$

The first-order Taylor expansion of y is:

$$y = k(\bar{x}) + [\alpha \cdot \bar{s} + \beta] \cdot g + [\alpha \cdot \bar{g}] \cdot s. \quad (\text{B-5})$$

Rewrite equation B-5:

$$g = \frac{y - k(\bar{x}) - \alpha \bar{g}}{\alpha \bar{s} + \beta}$$

and substitute into equation B-3, using equations B-4a and 4b. This yields equation B-2.

APPENDIX C: DETAILED INFORMATION ABOUT VARIABLES

Dependent variables

The spending variable, R , is measured by 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." The options are: (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 an earlier version (ISCO-68) of occupational classifications was used, but these 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 a 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. 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. 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

TABLE C-1. Party Classification for Ten Countries

	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, PRG	Socialist Party	UDF	RPR	National Front
Germany	PDS	Social Democratic Party, Greens	Free Democrats	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, Socialist Party	Center Party, Liberals, Christian Democrats, Greens	Conservatives	
United States		Democratic Party	Independent	Republican Party	

Note: NDP = New Democratic Party; PC = Progressive Conservative Party; PRG = Parti Radical de Gauche; RPR = Rassemblement pour la République; UDF = Union pour la Démocratie Française; PDS = Partei des Demokratischen Sozialismus; CDU/CSU = Christlich Demokratische Union/Christlich-Soziale Union. Party affiliation is not available for Italy.

education levels enabled us to map the Italian codes to the one-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 proxy for g and has five levels: (1) completed primary degree or lower; (2) incomplete secondary; (3) completed secondary; (4) incomplete and completed semihigher 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 do not use this distinction). Alternatively we could have used years of formal schooling as a measure of g , but the results are very similar.

Information. An item asked the degree of agreement with the following statement: "I feel that I have a pretty good understanding of the important political issues facing our country." The options were: (1) strongly agree; (2) agree; (3) neither agree nor disagree; (4) disagree; (5) strongly disagree. The coding was reversed so that higher values represent 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. Parties are classified as shown in Table C-1 (data are not available for Italy).

National Unemployment. The standardized rate of unemployment at the time of the national surveys (1996 unless noted otherwise) minus the overall 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. (Because it is the same in both cases, we will simply refer to s .) 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, so it was necessary to "translate" the 1997 information on s for use in the 1996 survey. For this purpose we calculated averages for s at the three-digit ISCO-88 level in the 1997 survey and then assigned these values to individuals in the 1996 survey, based on their three-digit ISCO classification in that survey. We show here that the estimated coefficient of b is consistent but has an approximate small sample bias that biases the estimated coefficient downward toward zero if $b > 0$ and upward toward zero if $b < 0$.

The structural model is

$$R_{i,j}^{96} = k + b \cdot y_{i,j}^{96} + c \cdot s_{i,j}^{96} + \epsilon_{i,j}^{96}, \quad (\text{D-1})$$

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

$$s_{i,j}^{96} = s_j + \eta_{i,j}^{96}, \quad (\text{D-2})$$

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

$$s_{i,j}^{97} = s_j + \eta_{i,j}^{97};$$

$$E\eta_{i,j}^{9x} = 0 \quad \forall i, j, x;$$

$$E\eta_{i,j}^{96}\eta_{r,j}^{97} = 0 = E\eta\varepsilon \quad \forall i, r, j;$$

$$E\eta^2 = \sigma_\eta^2; \quad E\varepsilon^2 = \sigma_\varepsilon^2; \quad (\text{D-3})$$

that is η_{ij}^{96} and η_{ij}^{97} can be thought of as random drawings from the same distribution. We now run the regression

$$R_{i,j}^{96} = k + by_{i,j}^{96} + c\bar{s}_j + \varepsilon_{i,j}^{96} + v_{i,j}^{96},$$

where

$$\bar{s}_j \equiv \frac{\sum s_{i,j}^{97}}{N_j}$$

and

$$v_{i,j}^{96} \equiv c[s_{i,j}^{96} - \bar{s}_j], \quad (\text{D-4})$$

and N_j is the number of individuals in ISCO category j in the 1997 survey. From equations D-3 and D-4,

$$\eta_{i,j}^{96} = \eta_{i,j}^{96} - \frac{\sum \eta_{i,j}^{97}}{N_j}. \quad (\text{D-5})$$

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

$$E\hat{b} = a,$$

$$E\hat{c} = c \left(1 - E \frac{\sum_j \bar{s}_j \cdot \bar{\eta}_j}{\sum_j \bar{s}_j^2} \right). \quad (\text{D-6})$$

If we make 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

$$\frac{\sum \bar{s}_j \cdot \eta_j}{\sum \bar{s}_j^2}$$

around the expected values of numerator and denominator:

$$\begin{aligned} E \frac{\sum \bar{s}_j \eta_j}{\sum \bar{s}_j^2} &\approx E \left[\frac{E \sum \bar{s}_j \eta_j}{E \sum \bar{s}_j^2} + \frac{1}{E \sum \bar{s}_j^2} \cdot (\sum \bar{s}_j \eta_j - E \sum \bar{s}_j \eta_j) \right. \\ &\quad \left. - \frac{E \sum \bar{s}_j \eta_j}{(E \sum \bar{s}_j^2)^2} \cdot (\sum \bar{s}_j^2 - E \sum \bar{s}_j^2) \right] \\ &= \frac{E \sum \bar{s}_j \eta_j}{E \sum \bar{s}_j^2}. \end{aligned}$$

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

$$E\hat{c} = c \left(1 - \frac{\sigma_\eta^2}{\sigma_\eta^2 + N \cdot \frac{\sum s_j^2}{J}} \right). \quad (\text{D-7})$$

Since J is constant, equation D-7 tells us first that as N increases the approximate bias goes to zero. More important, it implies that for a small sample

$$c < E\hat{c} < 0 \quad \text{if} \quad c < 0; \quad (\text{D-8})$$

$$c > E\hat{c} > 0 \quad \text{if} \quad c > 0.$$

Finally, it implies that

$$\text{If } c = 0 \Rightarrow E\hat{c} = 0, \quad (\text{D-9})$$

and the standard significance tests hold.

APPENDIX E: MULTILEVEL MODEL

Write the level 2 observation on individual i in economy j as $R_{ij} = \gamma + \eta y_{ij} + \mu s_{ij} + X'_{ij} \delta + \varepsilon_{ij}$, where X_{ij} is the vector of controls, and define level 1 by the fixed effects model $\eta = \bar{\eta}_1 U_j$ and $\mu = \bar{\mu} + \mu_1 U_j$. The single-stage regression model is derived by substituting the level 1 model into level 2: $R_{ij} = \gamma + \bar{\eta} Y_{ij} + \eta_1 y_{ij} U_j + \bar{\mu} s_{ij} + \mu_1 s_{ij} U_j + X'_{ij} \delta + \varepsilon_{ij}$. If we make the assumption that all U_j are exogenous and that this is a nonrandom effects model, then the implication is that the single-stage multilevel equation conforms to the standard ordinary least-squares conditions.

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