<Online supplementary material>

**Online-Appendix**

**Online-Appendix Figure A**: Respondents’ preferences over public education spending across countries.



Source: Authors’ compilation, based on ISSP RoG 2006 data (see text for details).

**Online-Appendix Table A**: Descriptive statistics for variables used in survey analyses.

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
| Variable, Source, Operationalization | N | Mean | Std. Dev. | Min | Max |
| *Micro-level variables (Source: ISSP 2008):* |
| Education spending preference, (“V20”) | 19,372 | .7104 | .4536 | 0 | 1 |
| Unemployment spending preference, (“V23”) | 19,072 | .2775 | .4478 | 0 | 1 |
| Income, in country deciles | 19,890 | 3.3867 | 2.2551 | 1 | 10 |
| Gender (female) | 19,964 | .5057 | .5000 | 0 | 1 |
| Education (years) | 18,564 | 12.7671 | 4.1244 | 1 | 65 |
| Age (years) | 19,892 | 48.5462 | 16.5843 | 15 | 97 |
| Full-time worker, “WRKST” | 19,671 | .5251 | .4994 | 0 | 1 |
| Part-time worker, “WRKST” | 19,671 | .1059 | .3078 | 0 | 1 |
| Less than part-time worker, “WRKST” | 19,671 | .1095 | .3123 | 0 | 1 |
| Unemployed, “WRKST” | 19,671 | .0321 | .1762 | 0 | 1 |
| In education, “WRKST” | 19,671 | .0367 | .1880 | 0 | 1 |
| Retired, “WRKST” | 19,671 | .1907 | .3929 | 0 | 1 |
| No children, “HOMPOP” = 1 or 2 persons | 19,718 | .5359 | .4987 | 0 | 1 |
| *Micro-level variables included in robustness section (Source: ISSP 2008):* |
| No children (alternative operationalization), “HHCYCLE” = 1 or 2 persons | 19,573 | .5212 | .4996 | 0 | 1 |
| Party vote (left), “PARTY\_LR”, voted for a left or far-left party | 17,698 | .3370 | .4727 | 0 | 1 |
| Party vote (right), “PARTY\_LR”, voted for a right or far-right party | 17,698 | .2790 | .4485 | 0 | 1 |
| Should government cur spending?, “V11” | 18,878 | 2.6037 | 1.2169 | 0 | 4 |
| Offshore-ability index (Walter 2010a) | 17,817 | 18.4618 | 29.2760 | 0 | 100 |
| *Country-level variables (Source: OECD Stats, if not indicated otherwise)* |
| Level of trade openness (2000), (Imports + Exports) / GDP | 20,023 | 71.5490 | 32.7974 | 20.52 | 184.01 |
| Level of trade openness (2005), (Imports + Exports) / GDP | 20,023 | 68.8095 | 29.0180 | 26.49 | 150.70 |
| 5-year change in trade openness, Level of trade openness 2005 – level of trade openness 2000 | 20,023 | -2.7395 | 8.7638 | -33.31 | 10.52 |
| Level of inequality (2000), (SWIID, Version 3.1), net Gini  | 20,023 | 29.7010 | 4.1908 | 22.5 | 36.8 |
| Level of inequality (2005), (SWIID, Version 3.1), net Gini  | 20,023 | 28.9481 | 3.3786 | 23.60 | 35.93 |
| Level of public education spending all levels (2000) as a share of GDP | 20,023 | 5.0711 | .7998 | 3.50 | 6.40 |
| Level of public education spending all levels (2005) as a share of GDP | 20,023 | 4.9985 | .8659 | 3.40 | 6.80 |
| Level of public unemployment spending (2000) as a share of GDP | 20,023 | 1.2187 | .7077 | .233 | 3.006 |
| Level of public unemployment spending (2005) as a share of GDP | 20,023 | 1.2137 | .7667 | .275 | 2.829 |
| Level of deindustrialization (2000) | 20,023 | .6890 | .0609 | .53 | .77 |
| Change in deindustrialization (2005-2000) | 20,023 | .0267 | .0104 | .01 | .05 |
| Total foreign direct investment (2000) | 20,023 | 16.3672 | 10.8090 | .84 | 36.27 |
| Total foreign direct investment (2005) | 20,023 | 8.2206 | 5.7892 | 1.06 | 25.38 |
| Inwards foreign direct investment (2000) | 20,023 | 7.9177 | 6.2708 | .176 | 26.46 |
| Inwards foreign direct investment (2005) | 20,023 | 3.4538 | 3.0459 | .06 | 14.35 |
| Outwards foreign direct investment (2000) | 20,023 | 8.4495 | 6.6743 | .67 | 19.73 |
| Outwards foreign direct investment (2005) | 20,023 | 4.7668 | 4.2336 | .29 | 19.26 |
| Capital account transactions Index (2000), Armingeon et al. 2012 | 20,023 | 2.2875 | .4408 | 1.13 | 2.46 |
| Capital account transactions Index (2000), Armingeon et al. 2012 | 20,023 | 2.2875 | .4408 | 1.13 | 2.46 |
| KOF “economic globalisation index” (2000), Dreher et al. 2008 | 20,023 | 83.1185 | 8.5456 | 56.85 | 97.25 |
| KOF “overall globalisation index” (2000), Dreher et al. 2008 | 20,023 | 83.8081 | 6.5552 | 63.37 | 91.91 |

**Online Appendix: Robustness checks**

In order to corroborate our results, we conduct a number of robustness tests. Due to space limitations, we report only those for preferences on education spending, in which we are mainly interested. The reported results for unemployment spending are equally robust (results upon request).

First, we check whether our results depend on the inclusion or exclusion of certain control variables by estimating different models. For instance, we include respondents’ ideological predisposition on the left-right scale, their general preferences towards public spending, and an occupation-based estimation of the likelihood of future job-loss (“offshore-ability index”). As could be expected (Ansell 2010; Garritzmann 2015), LEFT-WING VOTERS favor more public education spending while RIGHT-WING VOTERS prefer less spending (Table B, models 1a, 2a, 3a). We abstain from including partisan variables in the main analyses above as this considerably decreases the number of cases due to missing values and because additional problems of endogeneity might occur (partisan identification is also associated with the other control variables in the model). Including party affiliation does *not* alter the reported main results.

Furthermore, one might argue that respondents’ preferences towards *education* spending might be partially explained by their PREFERENCES TOWARDS PUBLIC SPENDING IN GENERAL. While the fact that the effects of openness on preferences differed across policy-fields already puts some doubt on this, we nevertheless include a variable covering whether respondents think that government spending should be cut. Not surprisingly, those who favor retrenchment also favor cuts in education spending (Table B, models 1b, 2b, 3b). More importantly, however, the reported results remain unaffected by this control. Hence, we exclude this general spending variable, because of the theoretical and statistical complexities resulting from ‘explaining attitudes with attitudes’. In more general terms, multicollinearity does not affect our results, because the main effects remain robust irrespective of the variable selection and because the correlations between the included variables are relatively low.[[1]](#footnote-1)

Finally, we also try to control for respondents’ exposure to globalisation on the individual-level by utilizing Blinder’s (2007) “OFFSHORE-ABILITY INDEX”, which has been commonly used in the literature. The index tries to capture a job’s *potential* to be offshored. Blinder estimates this potential by rank-ordering occupations based on whether workers need to be present at a specific work location or not. This rank-ordered information is then transformed into a four-category variable, which is interpreted as the likelihood of future globalisation-related job-loss. Walter (2010a) transfers Blinder’s index from the U.S. to the European context using the respondents’ 4-digit ISCO-codes as provided in the ESS-data. We apply this index to our ISSP-data to control for within-country effects of globalisation.

We refrain from including this index in the main analysis for theoretical and methodological reasons. Theoretically, we are interested in the association between the macro-level context in terms of economic openness and individual-level policy preferences as argued above. This is because we expect that people’s preferences are shaped by the general openness of the economy above and beyond their individual exposure to globalisation. Moreover, the offshore-ability measure is methodologically questionable: First, the rank-ordering and categorization of occupations is rather arbitrary. Second, it is based on the case of the United States only and assumes that the rank-ordering can be equally applied in other countries, which is a very strong assumption. Applying the same procedure as Walter’s (2010a) for ESS data, the resulting variable is heavily right-skewed in our ISSP data: Across all countries, three-quarters of all workers are coded as entirely “sheltered”. On average only six percent are found in the highest offshore-ability-category. In some countries (e.g., in Canada, Finland, or France), fully 98 percent of the population are coded as being employed in jobs with no offshore-potential.[[2]](#footnote-2) If only a minority is directly affected by globalisation, as this data claims, it is hard to imagine how this minority could be responsible for the large degree of public support in favor of education policies. Also, the extremely low values of offshore-ability for many country cases fuel doubts about the principal usability of this measure in countries besides the U.S.

Nevertheless, we include the offshore-ability index despite these shortcomings both in addition to and instead of the labor market status variables to test the robustness of our results. The reported results remain unaffected: Trade openness on the macro-level is still found to increase demand for education, but not for unemployment spending. Moreover, the individual offshore potential does *not* have a significant effect on respondents’ attitudes towards education spending (Table B, models 1c, 1d, 1e). The same finding holds when we exclude those countries with the heaviest right-ward skew (Canada, Finland, France).

Moreover, we also tested another set of additional control variables:

1. Rural/urban (Table B, Models 1m, 2m, 3m): We added the URBRURAL variable (because it is better comparable across countries) and tested whether respondents from different communities have different preferences. The only significant finding is that respondents from “farms or homes in the countryside” are significantly less likely to support additional public education spending (which seems very plausible from the perspective of several theories, from both rational choice and sociological approaches).
2. Marital status (Table B, Models 1n, 2n, 3n): there are no significant differences between the groups, with the exception of “divorced” respondents being slightly more in favor of education spending (only at a 10% level, though). This finding might be due to the fact that among the divorced there might be more single parents, who might have higher demands for social investments (in order to be able to reconcile family and work life).
3. Employment (Table B, Models 1o, 2o, 3o): as could be expected, respondents working for the government are the strongest in favor of additional public education spending (which at least partly could be due to self-interest); workers in private firms and the self-employed are significantly less in favor.
4. Religious attendance (Table B, Models 1p, 2p, 3p) & denomination (Table B, Models 1q, 2q, 3q): compared to the “no religion” group, both Catholics and Protestants are significantly less likely to support additional public education spending; analogously, those who attend services less often are more in favor of education spending.
5. Different operationalization of income (Table B, Models 1r, 2r, 3r): We standardized the income variable in each country. Yet, this produces essentially the same findings.
6. Household composition/Having children (Table B, Models 1s, 2s, 3s): We also run additional models with a different coding of whether respondents live with children at home or not. The findings are essentially the same, i.e. irrespective of how we measure this, respondents without children in the household are less in favor of additional public education spending and respondents with children in the household are more in favor of additional education spending, in line with our expectation.
7. Educational degrees (Table B, Models 1t, 2t, 3t): using educational degrees instead of the education year variable yields highly similar findings. But the findings are a bit more nuanced: The results show that especially respondents who hold a higher education degree are much more in favor of additional education spending. Moreover, we find that – compared to respondents with no formal education – respondents in the second lowest category (i.e. respondents with a basic formal degree) are slightly more in favor of additional public education spending than those without any degrees.

The most important finding, however, is that the inclusion of these additional control variables does not alter the main findings reported in the body and does not challenge the argument made in this paper.

In addition to these supplementary micro-level controls, we also test Iversen’s (2001) and Iversen and Cusack’s (2000) claim that DEINDUSTRIALIZATION rather than globalisation affects governments’ (and we deduce: individuals’) preferences by including the level (or 5-year change) in deindustrialization. However, neither levels of nor changes in deindustrialization show significant effects in either model as soon as we control for inequality (Table B, models 2d, 3d, 2e, 3e), and only weakly significant results when deindustrialization is included as the only macro variable (Table B, models 1d and 1e). Most importantly, the effect of trade openness remains unaffected by these changes. In sum, the findings underpin that the results are not driven by the selection of variables and robust to including other controls.

Second (and besides the inclusion of additional controls), we check whether our results are sensitive to the inclusion of survey weights, or whether they are biased by possibly systematic patterns of missing values. The results (available on request) remain unaltered.

**Online Appendix Table B:** Testing rival explanations by including additional control variables.

|  |
| --- |
| *Note: The following table summarizes results of 15 separate models. The models replicate models 1 through 3 of Table 1 and include additional control variables. To keep the table readable, we only present results for the variables of interest here. Included, but not shown, are the same control variables as in Table 1.*  |
|  | (1a) | (2a) | (3a) |
| Respondent voted for a left party | 0.3216\*\*\* | 0.3284\*\*\* | 0.3292\*\*\* |
|  | (0.0486) | (0.0486) | (0.0486) |
| Respondent voted for a right party | -0.1917\*\*\* | -0.1851\*\*\* | -0.1850\*\*\* |
|  | (0.0476) | (0.0476) | (0.0476) |
|  | (1b) | (2b) | (3b) |
| Government should cut spending | -0.1212\*\*\* | -0.1211\*\*\* | -0.1212\*\*\* |
|  | (0.0162) | (0.0162) | (0.0162) |
|  | (1c) | (1d) | (1e) |
| “Offshore-ability index” (Blinder & Walter) | 0.0126 | 0.0133 | 0.0133 |
|  | (0.0197) | (0.0197) | (0.0197) |
|  | (1d) | (2d) | (3d) |
| Level of deindustrialization (2000) | -4.0502\* | -2.2313 | -2.1467 |
|  | (2.4209) | (2.2386) | (2.1781) |
|  | (1e) | (2e) | (3e) |
| Change in deindustrialization (2005-2000) | 16.1718 | -0.4532 | 0.9208 |
|  | (14.2727) | (13.9640) | (13.6325) |
| Type of community (self-assessed). Baseline category: urban/a big city | Suburb, outskirt of a big city | (1m)-0.0170(0.0551) | (2m)-0.0178(0.0551) | (3m)-0.0182(0.0551) |
| Town or small city | -0.0512(0.0529) | -0.0517(0.0528) | -0.0521(0.0528) |
| Country village, other type of community | -0.0324(0.0560) | -0.0301(0.0560) | -0.0303(0.0560) |
| Farm or home in the country | -0.2960\*\*\*(0.0754) | -0.2975\*\*\*(0.0754) | -0.2975\*\*\*(0.0754) |
| Marital status (baseline category: married) | Widowed | (1n)-0.0051(0.0849) | (2n)-0.0057(0.0849) | (3n)-0.0058(0.0849) |
| Divorced | 0.1280(0.0688) | 0.1284(0.0849) | 0.1285(0.0687) |
| Separated | -0.0023(0.1162) | -0.0043(0.1162) | -0.0039(0.1162) |
| Never married/single | 0.0072(0.0525) | 0.0074(0.0525) | 0.0072(0.0525) |
| Type of employment (baseline: work for government) | Public owned firm, national industry | (1o)0.0272(0.0843) | (2o)0.0263(0.0843) | (3o)0.0260(0.0843) |
| Private owned firm, others | -0.1003\*\*(0.0466) | -0.1007\*\*(0.0466) | -0.1014\*\*(0.0466) |
| Self-employed | -0.1883\*\*\*(0.0647) | -0.1903\*\*\*(0.0647) | -0.1911\*\*\*(0.0647) |
| GB: Other, charity, voluntary sector, ZA | -0.1535(0.6888) | -0.1727(0.6886) | -0.1750(0.6886) |
| Religious attendance (frequency) | (1p)0.0260\*\*\*(0.0097) | (2p)0.0269\*\*\*(0.0097) | (2p)0.0269\*\*\*(0.0097) |
| Religious denomination (baseline: no religion) | Roman catholic | (1q)-0.3719\*\*\*(0.0577) | (2q)-0.3730\*\*\*(0.0576) | (3q)-0.3724\*\*\*(0.0576) |
| Protestant | -0.2428\*\*\*(0.0534) | -0.2366\*\*\*(0.0533) | -0.2355\*\*\*(0.0534) |
| Christian Orthodox | 0.4849\*(0.2926) | 0.4880\*(0.2926) | 0.4888\*(0.2926) |
| Jewish | -0.2765(0.3557) | -0.2785(0.3558) | -0.2775(0.3558) |
| Islam | -0.0523(0.2075) | -0.0478(0.2074) | -0.0472(0.2075) |
| Buddhism | 0.1360(0.1453) | 0.1216(0.1452) | 0.1185(0.1455) |
| Hinduism | 0.4723(0.4881) | 0.4683(0.4881) | 0.4692(0.4881) |
| Other Christian religions | -0.4886\*\*\*(0.1142) | -0.4926\*\*\*(0.1142) | -0.4916\*\*\*(0.1142) |
| Other Eastern religions | -0.1248(0.4647) | -0.1376(0.4644) | -0.1400(0.4643) |
| Other religions | -0.4716\*\*(0.1848) | -0.4713\*\*(0.1847) | -0.4709(0.1847) |
| Income z-transformed (different operationalization, instead of the income variable used in the main body) | (1r)-0.0194(0.0238) | (2r)-0.0198(0.0238) | (3r)-0.0191(0.0238) |
| Having children in the household (different operationalization, instead of the variable used in the main models) | (1s)0.3357\*\*\*(0.0412) | (2s)0.3356\*\*\*(0.0412) | (3s)0.3358\*\*\*(0.0412) |
| Educational degrees (instead of education years variable) (baseline: no formal qualification, incomplete primary education) | Lowest formal qualification | (1t)-0.1992\*\*(0.0928) | (2t)-0.1953\*\*(0.0928) | (3t)-0.1962\*\*(0.0928) |
| Above formal qualification | -0.0068(0.0956) | -0.0013(0.0956) | -0.0023(0.0956) |
| Higher secondary completed | -0.0009(0.0952) | 0.0014(0.0952) | 0.0001(0.0952) |
| Above higher secondary level, other qualification | 0.0069(0.0929) | 0.0104(0.0928) | 0.0091(0.0929) |
| University degree completed, graduate studies | 0.2065\*\*(0.0965) | 0.2094\*\*(0.0965) | 0.2081\*\*(0.0965) |

\* *p*<0.1; \*\* *p*<0.05; \*\*\* *p*<0.01. Standard errors in parentheses.

**Online Appendix Table C:** Testing alternative operationalizations of globalisation.

|  |
| --- |
| *Note: The following table summarizes 15 separate models, replicating models 1 trough 3 of Table 1. Instead of “trade openness”, alternative operationalizations are used. To keep the table readable, we only present results for the variables of interest here. Included, but not shown, are the same control variables as in Table 1. (In models 1i, 2i, and 3i, we use changes in inequality and education spending instead of levels.)* |
|  | (1f) | (2f) | (3f) |
| FDI total (2000) | -0.0061 | 0.0239\* | 0.0228\* |
| (0.0144) | (0.0142) | (0.0139) |
|  | (1g) | (2g) | (3g) |
| FDI inwards (2000) | 0.0222 | 0.0528\*\*\* | 0.0506\*\*\* |
| (0.0234) | (0.0175) | (0.0179) |
|  | (1h) | (2h) | (3h) |
| FDI outwards (2000) | -0.0367\* | -0.0062 | -0.0032 |
| (0.0215) | (0.0242) | (0.0238) |
|  | (1i) | (2i) | (3i) |
| Capital account transaction “kaopen” index (2000) | -0.3243 | -0.1904 | -0.1481 |
| (0.5036) | (0.4213) | (0.4124) |
| KOF “economic globalisation index” (2000) | (1j)0.0106(0.0173) | (2j)0.0349\*\*(0.0137) | (3j)0.0397\*\*\*(0.0127) |
| KOF “overall globalisation index” (2000) | (1k)0.0117(0.02223) | (2k)0.0326\*(0.0180) | (3k)0.0373\*\*(0.0172) |
|  | (1l) | (2l) | (3l) |
| Change in trade openness (2000-2005) | 0.0251 | 0.0258\* | 0.0241 |
|  | (0.0153) | (0.0150) | (0.0153) |

\* *p*<0.1; \*\* *p*<0.05; \*\*\* *p*<0.01. Standard errors in parentheses.

Third, we test different operationalizations of economic globalisation to investigate whether the results depend on our focus on trade openness. To begin with, we use foreign direct investment (FDI inwards only, outwards only, and total) (Table C). Using total FDI as well as inward FDI clearly confirms the reported results: The higher the total or inward FDI, the higher the demand for public education spending (Table C, models 1f, 2f, 3f, 1g, 2g, 3g). *Outward* FDI, in contrast, are only significant when we do not control for inequality, and the effect is not very robust (Table C, models 1h, 2h, 3h). This seems reasonable because inward FDI are a lot more visible for respondents than outward FDI, although the latter might actually be more related to changes on the labor market. Moreover, we also use an index for the extent of openness in capital accounts (Armingeon et al. 2012). As the variable shows hardly any variance across space and time for our sample (see Table A in the Online-Appendix), it is not surprising that we do not observe any effect (Table C, models 1i, 2i, 3i). Furthermore, we used the KOF-index, which has become another standard measure of globalisation (Dreher et al. 2008). We tested both the KOF’s “economic globalisation index” (using data from actual economic flows as well as restrictions to trade and capital flows; cf. ibid.) and the “KOF overall index”, which also covers social and political aspects of globalisation above and beyond economic forces. The results (see Table C, models 1j, 2j, 3j, 1k, 2k, 3k) again reveal the same finding, i.e. globalisation is associated with higher demand for education spending.

We also test whether *changes* in trade openness instead of levels have an effect by using five-year differences. In two models (Table C, models 1l and 3l), the change in trade openness slightly misses conventional significant levels (it is significant at an 11-percent level), while it is significant at a 10-percent level in model 2l. Substantially, this indicates that not only the levels, but also the changes in trade openness might influence demand for public education spending, although this association is less robust than the association with levels of trade openness. This is strong additional proof for our claims. Finally, we *simultaneously* include the level of trade openness and the respective FDI measures, because it might be the case that they have distinct independent effects on attitudes because they cover different aspects of globalisation. Due to the high correlations between these measures, it is unsurprising that the significance levels drop slightly, while the reported effect sizes remain similar (results on request). In sum, we feel confident to conclude that the results do not seem to be driven by a specific operationalization of globalisation, but are observable for different measures covering several facets of globalisation.

Fourth, to make sure that the results are not driven by our decision to dichotomize the dependent variable, we use a different coding and dichotomize between those who are “strongly in favor” and all other respondents. Again, the findings hold, but the discriminating power of some variables is lower, i.e. the effect of trade openness remains highly similar, but suffers a slight loss in significance and size (results on request). Furthermore, we use different estimation techniques: While the theoretically most suitable models (multilevel ordered logit) do, unfortunately, not converge in several cases due to the heavy left-skewedness of the data in some countries (as discussed above), we use *single-level* ordered logit models with clustered standard errors as the second-best model choice. Moreover, the proportional lines assumption (i.e. that the effect of the independent variables is constant for each answer category of the dependent variable) was violated in some cases; therefore, we also estimate a “partial proportional odds model” (Williams 2006). In the Online-Appendix (Table B), we present and discuss this model and its findings in-depth. For the present purpose it suffices to summarize that the effect of trade openness remains robust: The coefficients of the openness variable are positive and significant for all answer categories, indicating that trade openness has a linear positive effect on education spending preferences.[[3]](#footnote-3) Thus, the results are not driven by our decision to dichotomize the dependent variable. We get the same findings when using probit or multinomial models.

In sum, the theorized effect of globalisation on demand for public spending seems highly robust as it holds for different variable selections, different operationalizations of globalisation, and across different model specifications, and is not affected by missing values or the inclusion of survey weights. This gives us confidence that the reported results are robust, reliable, and substantial.

**Online Appendix Table D**: Determinants of individual-level preferences towards public education spending in 17 countries in 2005/06, partial proportional odds model.

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
|  | Spend less, the same as now, more, or much more (=1). Much less (= 0). | Spend the same as now, more, or much more (=1). Less, or much less (= 0). | Spend more, or much more (=1). Much less, less, or the same as now (= 0). | Spend much more (=1). Much less, less, the same as now, or more (= 0). |
| Individual-level variables |
| Income | 0.2316\*\*\* | 0.0589 | 0.0040 | -0.0166 |
|  | (0.0825) | (0.0535) | (0.0222) | (0.0152) |
| Female | 0.9186\*\*\* | 0.5060\*\*\* | 0.0822 | 0.0802\* |
|  | (0.2068) | (0.1456) | (0.0518) | (0.0467) |
| Education (years) | -0.0208 | -0.0190 | 0.0129 | 0.0241\*\* |
|  | (0.0474) | (0.0156) | (0.0080) | (0.0108) |
| Age | 0.0199\*\*(0.0082) |
|  |
| Age (squared) | -0.0002\*\*\*(0.0001) |
|  |
| No children | 0.7141\*\*\* | -0.2619\*\* | -0.2102\*\*\* | -0.1634\*\* |
|  | (0.2288) | (0.1264) | (0.0724) | (0.0701) |
| Part-time worker | 0.0243(0.0656) |
|  |
| Less than part-time | 0.0996(0.0652) |
|  |
| Unemployed | 0.1064(0.0996) |
|  |
| In education | 0.0459 | 1.6828\*\* | 0.4049\* | 0.2007 |
|  | (1.0441) | (0.6564) | (0.2305) | (0.1946) |
| Retired | 0.1311(0.1030) |
|  |
| Country-level variables |
| Trade openness (2000) | 0.0279\*\*\* | 0.0199\*\*\* | 0.0078\* | 0.0051\*\* |
| (0.0101) | (0.0062) | (0.0041) | (0.0026) |
| Inequality (2000) | 0.0909\*\*\*(0.0320) |
| Public education spending (2000) | -0.2338(0.1776) |
| Constant | 0.6544 | 0.4726 | -1.6867 | -3.6209\* |
|  | (2.2752) | (1.8899) | (2.0598) | (1.9977) |
| Model fit |
| N | 17,394 |
| Log pseudolikelihood | -19550.523 |
| Pseudo R2 | 0.0298 |

\* *p*<0.1; \*\* *p*<0.05; \*\*\* *p*<0.01. Standard errors in parentheses. Reference category for labor market status dummies is full-time employment. For those variables, where the ordered logit assumption held, only one coefficient is shown for all categories (e.g., “age”) and can be interpreted as an ordered logit coefficient. For those variables, which violated the assumption, different coefficients for each answer category are shown (e.g., “income”). For interpretation of partial proportional odds models see description in text above and Williams (2006).

**Interpretation of Table D (Online-Appendix):**

A well-known and common weakness of ordered logit models is, that this assumption is often violated. This was also the case for our sample. The two standard solutions to this problem are to either ignore this shortcoming and still report the results, or to abstain from using ordered logit models by using multinomial or generalized ordered logit models instead.

As a more convincing solution, we used a partial proportional odds model.[[4]](#footnote-4) This model relaxes the proportional odds assumption for those variables, which violate it, but keeps it for those variables, which do not violate the assumption. Put simply, a partial proportional odds model is an ordered logit model for those variables, which meet the proportional odds assumption, and a series of logit models for those variables, which do *not* meet the assumption. The main advantage of the model is that it uses more of the available information as (in contrast to logit) all answer categories are used and (in contrast to multinomial models) the fact that the answer categories are ordered is taken into account. The interpretation of the coefficients is slightly different than in logit models: The coefficients of those variables, for which the ordered logit assumption holds, can be interpreted as standard ordered logit coefficients. When the assumption is violated, the model estimates a series of logit models. For these, *positive* coefficients indicate that the variable increases the likelihood that a respondent is in a *higher* answer category than the current. *Negative* coefficients, vice versa, mean that the respondent is likely to be *in the current or a lower* category on the answer scale.

Results of the partial proportional odds model, using the same variable specification as model 3 in Table 1, are presented in Table D.[[5]](#footnote-5) Again, the effect of trade openness remains robust: the coefficients of the openness variable are positive and significant for all answer categories, which indicates that trade openness has an almost linear positive effect on education spending preferences. This justifies our decision to dichotomize the dependent variable to facilitate interpretation. To be precise, the effect decreases slightly in size over the answer categories, indicating that trade openness decreases opposition to more spending more drastically than it increases support. In any case, however, this indicates that trade openness has the theorized effect and the results are not driven by our decision to dichotomize the dependent variable.

**Additional references (for online appendix)**

Armingeon, K., D. Weisstanner, S. Engler, P. Potolidis, and M. Gerber (2012), *Comparative Political Data Set I 1960-2010*, Bern: Institute of Political Science, University of Bern.

Williams, R. (2006), ‘Generalized Ordered Logit/Partial Proportional Odds Models for Ordinal Dependent Variables’, *The Stata Journal*, 6, 1, 58-82.

1. Many correlations (Spearman resp. Pearson) are below 0.4, most below 0.1. The highest correlations were obviously detected between age and age-squared, and between age and the retirement-dummy. We tested these separately and the effects remained robust. [↑](#footnote-ref-1)
2. In the French case, this might (partly) be due to the fact that a different occupation-coding scheme is applied. For Canada and Finland, however, no obvious explanation for the skew is apparent and the index remains questionable when applied to ISSP-data. [↑](#footnote-ref-2)
3. To be precise, the effect decreases slightly in size over the answer categories, indicating that trade openness decreases opposition to more spending more drastically than it increases support. [↑](#footnote-ref-3)
4. We estimate these using Stata’s “gologit2“-ado with the “autofit“-option (Williams 2006). [↑](#footnote-ref-4)
5. The model reveals why we do not find an effect of gender on education policy preferences. Gender indeed does have an effect, but it is not linear across all answer categories (indicated by the decreasing sizes and changing significance levels of the effect): Women are significantly less likely than men to oppose education spending. For the third answer category, which discriminates between “more spending” vs. “the same or less” (just as our main analyses above have done), however, this effect disappears. In other words, women are less likely than men to take extreme positions on public education spending. Women seem to hold more moderate positions than men. In a similar vein, the model shows that income indeed does have an effect, as soon as we investigate the answers across answer categories: the higher a respondent’s reported income, the less likely that she will favor “much less” spending. In other words: especially poorer respondents seem to tick the option “much less education spending”. We abstain from discussing this surprising finding in-depth, but strongly encourage future research to address this relationship in more detail. [↑](#footnote-ref-5)