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This paper addresses the question of whether higher levels of education contribute to greater tolerance of homosexuals. Using survey data for Ireland and exploiting a major reform to education, the abolition of fees for secondary schools in 1968, it is shown that increases in education causes individuals to be significantly more tolerant of homosexuals. Ignoring the endogeneity of education leads to much lower estimates of the effect of education. Replicating the model with data for the United Kingdom generates very similar results.

Keywords: education, homophobia, tolerance, social returns

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Introduction

There is a vast literature in economics measuring the returns to education, for example Card (1999). In most cases, these estimates refer to the effects in people's earnings and are typically based on the well known Mincer human capital model. Other outcomes that have been studied include labour market status, for example Ashenfelter & Ham (1979) and Nickell (1979) consider the effect on unemployment, and the effect on health has been considered by Lleras-Muney (2005) or Silles (2008) amongst others.

These returns are all private: how an individual benefits from their own education. However, it seems to be commonly believed that education is socially desirable in that it also conveys benefits to society beyond the private benefit to the individual i.e. that education generates substantial positive externalities. These benefits could arise in many ways, for example human capital spill-overs where one person's productivity increases someone else's but in a way that is not rewarded in the labour market for example Moretti (2004). Aside from such economic externalities, a number of studies have recently addressed the question of whether there are civic returns to education: that education somehow leads individuals to be better citizens, for example Helliwell & Putnam (2007) consider whether education is associated with higher levels of social engagement. Understanding whether such externalities exist is not just an academic curiosum: if all the benefits to education are private then the case for public subsidy of education is a good deal weaker particularly if credit constraints are not important as some studies suggest (e.g. Carneiro & Heckman 2002).

It is not difficult to find a correlation between individual outcomes, whether it is their health, earnings or other behaviours. But it could very easily not reflect the effect of education because of unobserved heterogeneity. That is, education may simply be correlated with some factor (for example preferences) which are correlated with the outcome of interest.

A number of papers on the civic returns to education have tackled this problem directly. Dee (2004) and Milligan, Moretti & Oreopoulos (2004) investigate if education increases the probability of individuals voting. The former paper also considers individuals' support for free speech. Since education and political behaviour could be correlated for many reasons, to identify a causal relationship it is necessary to find factors that exogenously vary education. The latter two papers use a variety of such instrumental variables including the proximity to college, child labour laws and changes in the minimum school leaving age. In both studies, significant positive civic returns to education are estimated.

Recent work in this vein however has produced quite different results. Siedler (2010), who uses reforms to school leaving age in Germany, finds however that correcting to endogeneity leads in general to small, statistically insignificant effects on voting and related measures of citizenship such as membership of a political party or interest in politics. Berinski & Lenz (2010) use the Vietnam military draft as an instrument with US data and also find little evidence of an effect on voter turnout. Pelkonen (2010), another instrumental variable study, finds weak evidence of education on voter turnout in Norway.

Gibson (2001) does not use instrumental variable estimation but instead uses sibling difference methods to remove unobserved family specific heterogeneity. This leads to the somewhat surprising result that education leads to individuals to a lower supply of voluntary labour, a negative external effect if anything.

This paper contributes to the literature on external effects of education by examining whether exogenous changes in educational levels cause changes in people's attitudes. The outcome is individual's stated tolerance of homosexuals based on their response (from a list of 5 possibilities) to the statement "Gays and lesbians should be free to live life as they wish"[1].

This outcome differ from those discussed above since, arguably, there is no "right answer": individuals are entitled to whatever opinions they hold on these matters however disagreeable somebody else might find them. Nonetheless, it seems uncontroversial to say that tolerance towards others is widely regarded as socially desirable and that it is therefore useful to know what factors cause this to be higher or lower. This paper is positive rather than normative, its purpose is to reveal whether education causes this attitude to be more or less likely without advocating any particular stance.

It seems very plausible that education has an effect on people's attitudes to other people. Offe and Fuchs (2002) suggest that school " …is the first non-familial context in an individual's life that trains…moral and cognitive capacities favouring cooperation". Furthermore, they argue, schools serve as institutional environments that favour informal associability amongst peers and fellow members. Ellison (1992) provides evidence that educated people are, in a particular sense, "nicer" but given the somewhat subjective nature of the measurement (an assessment by the interviewer) it is not clear how much one can generalize from this. Uslaner (1999) measures the effect of education on a set of eight indicators of moral behaviour but it is not statistically significant in any of them. Bettinger & Slonim (2006) use an experiment (the randomized receipt of education vouchers) to show how education can increase young people's altruism towards charitable organizations – although not their peers.

While there is extensive research on homophobia (or homo-negativity) it is difficult to find studies which directly address the role of education. Štulhofer & Rimac (2009) find a negative association between education and homo-negativity although it is not well determined. There are a number of studies that consider whether tolerance in other domains (particularly political views) is associated with education, for example Jackman (1978), Weil (1985), Bobo & Lacari (1989) and Golebiowska (1995). An important feature of this work in general is that it does not deal with the potential endogeneity of education or the possible correlation with unobserved heterogeneity. So it is difficult to ascribe a causal interpretation to any observed association between the outcome and education.

Other research on people's attitudes has looked at, *inter alia*, attitudes towards immigration. For example O'Rourke & Sinnott (2006) model how individuals view immigration on the basis of factor proportions trade theory and find that, as predicted,

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[1] The five possible responses are: agree strongly, agree, neither agree nor disagree, disagree, disagree strongly.

higher skilled individuals are less opposed to immigration than low skilled individuals. Thus trade theory and the endowments of individuals provide a simple economic rationale for being opposed to immigration. By contrast, it is difficult to think of any rational explanation for an individual responding negatively to the statement above about homosexuals. Hence it seems plausible to regard such responses as purely representing variations in tolerance.

The paper uses a natural experiment, the abolition of tuition costs for secondary schooling in Ireland in 1968, to generate instrumental variables. Use is also made of a second natural experiment, the raising of the minimum school leaving age, to generate an instrumental variable.

Ireland is a relatively homogenous country culturally and socially. In US terms, it has about the same population as South Carolina and is somewhat smaller geographically. About 95% of the population describe themselves as Roman Catholic. Homosexuality was decriminalized in 1993. An act of parliament recognizing civil partnerships between same-sex couples was passed in 2011.

Data and methods

The data used is the European Social Survey (ESS), waves 1 to 4 for Ireland. The waves are collected at 2 year intervals, starting in 2002 with close to 2000 observations in each year. The data is a random sample of the residential (non-institutional) population aged 15 or over[2]. Data for Ireland was chosen because of the identification strategy used.

Since the outcome of interest is an ordered response, an ordered probit is used to model this. To take account of the endogeneity of education, a bivariate model consisting of a simultaneous ordered probit and linear regression is estimated by

[2] Further information is at http://www.europeansocialsurvey.org

Maximum Likelihood. This is a generalization of the familiar “IV probit” estimator of Ameniya (1978) and Newey (1987). Specifically, for the simple ordered probit estimator:

$$\Pr(\boldsymbol{outcome}_j) = \Pr(\boldsymbol{\kappa}_{j-1} < \boldsymbol{\beta S} + \boldsymbol{\gamma X} + \boldsymbol{u} \leq \boldsymbol{\kappa}_j \ ) \qquad j = 1...5 \qquad (1)$$

Where S is years of education, X is a vector of controls (to include a constant), ***u*** is assumed to be distributed normally, κ_0 and κ_5 are assumed to be $-\infty$, $+\infty$ respectively. In the bivariate model, a second equation models years of education:

$$\boldsymbol{S} = \boldsymbol{\alpha Z} \ + \boldsymbol{\lambda X} \ + \boldsymbol{v} \qquad (2)$$

Z is a vector of one or more instruments. The model is estimated by Limited Information Maximum Likelihood (LIML) assuming that the disturbance terms ***u*** and ***v*** are distributed bivariate normal[3]. This amounts to a Seemingly Unrelated Regression model in which the equations are treated as independent but with jointly normal disturbance terms. This “SUR” approach to LIML estimation has wide applicability. For the binary dependent variable models, the two step estimator of Newey (1987) is used.

I use a small set of controls: a quadratic in age, sex, father’s education in levels, whether the individual lives in a city and whether they were born in another country. The age at which individuals move to the country is not recorded so it is possible they were not affected by the reform. As a check, I also report results excluding those born abroad. It is possible to find other variables which are statistically significant in the model. The ESS contains information on many other attitudes (towards religion, towards emigrants, the government for example) as well as characteristics such as trade union membership. However it does not make sense to “explain” one attitude in terms of another not least because these are likely to be endogenous also so they are not included.

[3] Roodman’s (2008) Maximum Likelihood Stata program for conditional mixed process recursive models is used.

Identification

Identification in the model is principally based on an educational reform in Ireland in 1968 and follows Denny & Harmon (2000). In this reform, education in secondary (high) school was made free for all school-age youths. Prior to the reform the fee-paying aspect to secondary education was potentially a major hurdle for families. Annual fees per pupil at the time of introduction of the policy were approximately two weeks wages for the average manual worker with no organised schemes nationally or at school level for the waiver of the fee[4]. Taking into account the large family sizes that prevailed at that time, secondary school fees could be of the order of one-sixth of total household income.

The reform, the announcement of which had not been anticipated, had a significant effect on the participation rate in education. Archer (1998) notes that contemporaneous accounts of the policy change recorded a 33% increase in the number of children participating in a school transport scheme in the two years post-reform. In an econometric analysis of time series data on participation rates Tussing (1978) found increases in enrolment for age groups that were significantly in excess of the trend in participation rates, which had been increasing in the period prior to the policy announcement. The long run increase in school participation due to the reform, abstracting from other trends in the data, is estimated by Tussing to be about 20%.

What is important for present purposes is not the aggregate change in education but the distribution effect. Prior to the reform those that received secondary (and by implication third level) education came from a wealthier socio-economic background. Thus the elimination of fees for secondary schooling had a differential effect, with larger increases in participation for those from less well off backgrounds. To generate instrumental variables from this reform I create a step dummy (labelled “No fees”) equal to 0 if the year of birth was before 1952 (as these would not have been affected), equal

[4] No national merit award or other scholarship scheme existed. Some individual schools may have granted partial or full remission of fees in particular cases but these were exceptions.

to 1 if year of birth was between 1952 and 1955 inclusive (as these would have been in secondary school and hence possibly affected) and equal to 2 if born after 1955 as these would started secondary school after the reform. Small changes in the specification of this variable (i.e. by one year) do not produce very different results. This step variable will be used to generate two instruments. To take account of the distributional effect of the reform, this variable is interacted with dummy variables representing the education of the father since one expects the reform to have a smaller effect for those from a higher SES.

A second educational reform that can be used to generate an instrumental variable is the raising of the minimum school leaving age in Ireland (from 14 to 15) in 1972, see Murtin & Viarengo (2008). Such reforms, often referred to as “Rosla”, have been widely used in the schooling returns literature to generate instrumental variables. The effects of this reform have not been widely studied in Ireland however and it is unclear that it had a significant effect, unlike the equivalent reforms in the United Kingdom where Harmon & Walker (1995) and others have shown it caused a significant effect on years of education. Less reliance is made of this instrument partly as the sample size available here may make it difficult to get a precise effect of such a one-off reform.

To ensure that the instruments are not picking up trends, I use a subset of the available data: individuals born between 1942 and 1965 inclusive, a 10 year window on either side of the “No fees” reform which will also include those affected by the reform to the minimum school leaving age. I will also experiment with a 5 year window.

Results

The basic results, treating education as exogenous, are shown in Table 2. The first column considers two possible measures of educational attainment, years of full time education completed and a set of three dummy variables for highest level completed. Unlike the human capital/Mincer model, it is not obvious that years of education will be what drives people's attitudes and "sheepskin effects" seem plausible in this context. However, it transpires that the linear (years) measure of education dominates the education levels and one cannot reject the hypothesis that the three coefficients on the latter are jointly zero (p=.7151). Nonetheless, the positive coefficients are consistent with the hypothesis education is generally associated with individuals being more tolerant of homosexuals.

The second column removes these dummy variables and is our "base case". To get a sense of what the coefficient on years of education means, the associated marginal effects are shown in the first column of Table 5. One year of education decreases the probability of an individual being in any of the four less tolerant categories and increases the probability of the first (most tolerant) category by a corresponding amount, about 1.27 percentage points. Since from Table 1(a), one can see that 22.74% are in this category the proportionate effect is not huge. The third model uses the full sample and does not change the coefficient of interest by much.

These results make no allowance for the potential endogeneity of education. Clearly education could be correlated with attitudes for many reasons, notably unobserved heterogeneity. Table 3 reports the estimates of the bivariate model which controls for this using the identification strategy discussed above. The lower panel shows the education equation. Considering column 1, one can see that the reform increased years of education and this increased with "exposure" to the reform i.e. there was a smaller effect for those in secondary education at the time relative to those who had yet to start secondary education. This model also includes as instruments the interaction of the reform variable (treated as continuous for simplicity) with paternal education.

To see how these work, one needs to consider the direct effects of paternal education. These coefficients show a clear positive socio-economic gradient: an individual whose father has tertiary education could expect to have about 3.8 more years of education than if their father has the lowest level of education i.e. has not completed lower secondary education. The direct effects refer to the situation before the abolition of school fees. The negative coefficients on the interactions mean that this socio-economic gradient is lower after the reform, or equivalently, that the reform had a bigger effect for those from a lower SES. The effect is not quite monotonic since the coefficient on the last interaction is small and not well determined. In general, these results are as expected and are very close to those reported in Denny & Harmon (2000) using a different dataset.

The implications for the coefficient of interest are striking, it increases to 0.176 (from .0362). The marginal effects for this model are given in the second column of Table 5. As one would expect, they are qualitatively similar to the simple probit models but much larger in magnitude: an extra year of education increases the probability of an individual being in the most tolerant category by about 5 percentage points and hence reduces the probability of being in the lower four categories by the same amount. A test for the statistical significance of the correlation between two error terms (ρ_{12}) can be thought of as a test for exogeneity and is easily rejected. The second column in this table adds as an instrument, the raising of the school leaving age in 1972 which increases the education coefficient to 0.194. The final column drops the interaction terms as instruments and this reduces the education coefficient slightly.

The finding the controlling for the endogeneity of education leads to larger effects is quite in estimates of earnings effects from human capital models. Some of this could be explained by an attenuation bias caused by measurement error although it seems implausible that it could be a major factor. Another explanation is that the coefficient is a Local Average Treatment Effect – the return to the sub-group who have been induced to change by the instrument. For non-linear models, it is unclear that either of these explanations is of much assistance. Both Dee (2004) and Milligan *et al*. (2004) find that estimated civic returns are higher with IV than OLS although the differences are generally smaller than found here.

A number of additional models are presented in Table 4 which include only details of the parameters of interest but for which the control variables are the same as in Tables 2 and 3. As a robustness test, I considered a smaller “window” around the event which generates the instrumental variable. Taking the first model in Table 3 and shrinking the window from 10 years on either side to 5 reduces the sample size to 1578. The results are shown in the first column. The coefficient on years of education is slightly higher (0.193) and is still well determined. A second robustness test arises from the fact that not all individuals may have been in the country at the time of the reform: it is known if they were born abroad but not whether they had completed their education when they arrived in Ireland. Taking the first model in Table 3 and excluding those born abroad reduces the sample size to 2450. The results are shown the second column. The coefficient on years of education is slightly higher (0.209) and is well determined.

Finally, since the marginal effects in Table 5 shows that the effect of an additional year of education shifts individuals from the lower four categories into the top one (i.e. into the “agree strongly” category), I generated an alternative binary dependent variable equal to 1 for that top category and 0 otherwise. The estimation of the models differs from the previous ones in that it uses the two-step Minimum Chi-square model of Newey (1987). The reason for this is that a test for over-identification (due to Lee (1992)) is available. The results of this specification are shown in column 3 of Table 4. As before, a year of education has a well determined positive effect on the outcome: the education coefficient is 0.224 (t= 2.97)[5]. If this model is estimated by simple probit, with no control for endogeneity, the coefficient is 0.048 (t=5.36). One can reject exogeneity of education and the test for over-identification is passed easily. Comparing this model with those in Table 3, it seems there is not much to be gained from using an ordered probit to examine the full set of responses over a simple binary choice specification.

As a further test, I replicated the last model (i.e. Table 4, column 3) using data from the United Kingdom from the European Social Survey. The specification is identical other than the instruments. As instruments I use the two increases in the minimum school leaving age which have been used by several authors including Harmon & Walker (1995), Silles (2008) and Milligan *et al*. (2004). The results are shown in the Appendix and are very similar to the results for Ireland: treating education as endogenous increases its coefficient by a multiple of around five although the parameter is less well determined.

To summarize, in naïve models which assume education to be exogenous, it can be seen that an additional year of education increases the probability of an individual

[5] Using instead the MLE approach of Ameniya (1978) leads to very similar results: the parameter of interest is 0.206 (t=4.34).

being tolerant towards homosexuals. The effect is small however. A bivariate model controlling for the endogeneity of education generates much larger effects increasing the coefficients and marginal effects four- or five-fold. The results appear to be robust to choice of instruments.

Conclusion

This paper contributes to research on civic returns to education by asking whether a higher level of education contributes to individuals being more tolerant towards homosexuals. Using survey data for Ireland and exploiting an educational reform to generate independent variation in education, it is shown that education does indeed reduce individuals' homophobia. Ignoring the endogeneity of education leads to much lower estimates of the effect. Replicating the model for the United Kingdom produces very similar results. Simple correlations between measures of attitudes and education are unlikely to be very informative.

In aggregate, any impact that education has on homophobia may not be considered important since only a small proportion of the population are homosexual, although the true proportion is necessarily difficult to discern and clearly for the population concerned, discrimination and intolerance is distressing. Nonetheless, this paper adds to the body of evidence that education changes not just people's labour market prospects but also how they think and interact in society.

Table 1: Descriptive statistics

(a) "Gays and lesbians should be free to live life as they wish"

	Restricted sample	%	Full sample	%
Agree strongly	609	22.74	1600	24.81
Agree	1618	60.42	3615	56.06
Neither agree nor disagree	282	10.53	768	11.91
Disagree	133	4.97	359	5.57
Disagree strongly	36	1.34	106	1.64
	2678	**100**	**6448**	**100**

(b) Descriptive statistics for restrictive sample

	Mean	Std deviation
Years of education	12.9469	3.308
Lower secondary education	.233	.423
Upper secondary education	.240	.427
Tertiary education	.332	.471
Age	52.069	7.230
Female	.554	.497
Father : lower secondary education	.114	.318
" : upper secondary education	.104	.306
" : tertiary education	.082	.274
City	.293	.455
Foreign	.085	.279
"No fees"=1	.166	.373
"No fees"=2	.420	.494
"No fees" x Father : lower secondary education	.129	.477
"No fees" x Father : upper secondary education	.127	.475
"No fees" x Father : tertiary education	.099	.415
ROSLA	.380	.486

N=2678

Table 2: ordered probit model of tolerance variable

	(1) Restricted sample	(2) Restricted sample	(3) Full sample
Years education	0.0362***	0.0424***	0.0455***
	(3.52)	(5.87)	(9.68)
Lower secondary education	-0.00580		
	(0.08)		
Upper secondary education	0.00368		
	(0.05)		
Tertiary education	0.0716		
	(0.75)		
Age	-0.00349	-0.00305	0.00336
	(0.08)	(0.07)	(0.74)
$Age^2/1000$	-0.0317	-0.0365	-0.139**
	(0.08)	(0.09)	(3.10)
Female	0.162***	0.162***	0.178***
	(3.71)	(3.72)	(6.42)
Father lower secondary education	-0.00832	-0.00440	0.0147
	(0.12)	(0.06)	(0.35)
Father upper secondary education	-0.00583	0.00511	0.0368
	(0.08)	(0.07)	(0.80)
Father tertiary education	-0.130	-0.116	0.0432
	(1.50)	(1.35)	(0.87)
City	0.0299	0.0320	-0.00669
	(0.62)	(0.66)	(0.22)
Foreign born	-0.0115	-0.00857	-0.138**
	(0.14)	(0.11)	(2.97)
N	2678	2678	6448

Absolute *t* statistics in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 3: instrumental variable ordered probits

	(1)	(2)	(3)
Years education	0.176***	0.194***	0.187**
	(3.94)	(4.60)	(3.29)
Age	-0.0002	0.0004	0.0000
	(0.00)	(0.01)	(0.00)
Age^2/1000	0.0111	0.0174	0.0156
	(0.03)	(0.04)	(0.04)
Female	0.0992*	0.0878	0.0928
	(1.97)	(1.73)	(1.67)
Father lower secondary education	-0.202*	-0.229*	-0.217*
	(2.09)	(2.43)	(1.97)
Father upper secondary education	-0.371*	-0.425**	-0.402*
	(2.51)	(2.97)	(2.23)
Father tertiary education	-0.601***	-0.667***	-0.639**
	(3.30)	(3.83)	(2.87)
City	0.0244	0.0227	0.0234
	(0.51)	(0.48)	(0.49)
Foreign born	-0.0831	-0.0938	-0.0893
	(1.01)	(1.15)	(1.06)
Years education			
"No fees"=1	1.030***	1.020***	0.883***
	(4.78)	(4.68)	(3.95)
"No fees"=2	1.616***	1.209***	0.948**
	(6.05)	(3.53)	(2.75)
Father lower secondary education x "No fees"	-0.608**	-0.595**	
	(3.24)	(3.21)	
Father upper secondary education x "No fees"	-0.698***	-0.703***	
	(3.57)	(3.66)	

Father tertiary education x "No fees"	-0.244	-0.240	
	(1.07)	(1.08)	
Raising of school leaving age (ROSLA)		0.544	0.472
		(1.83)	(1.55)
Age	-0.0242	0.0400	0.0596
	(0.20)	(0.31)	(0.46)
$Age^2/1000$	0.407	-0.145	-0.334
	(0.35)	(0.12)	(0.27)
Female	0.341^{**}	0.344^{**}	0.339^{**}
	(2.94)	(2.96)	(2.91)
Father lower secondary education	2.098^{***}	2.078^{***}	1.446^{***}
	(7.54)	(7.53)	(7.69)
Father upper secondary education	3.538^{***}	3.529^{***}	2.724^{***}
	(11.73)	(11.81)	(13.95)
Father tertiary education	3.808^{***}	3.795^{***}	3.558^{***}
	(11.08)	(11.21)	(16.20)
City	0.0473	0.0509	0.0501
	(0.37)	(0.39)	(0.39)
Foreign born	0.547^{*}	0.538^{*}	0.507^{*}
	(2.57)	(2.53)	(2.38)
constant	11.26^{***}	9.403^{**}	9.057^{**}
	(3.43)	(2.75)	(2.64)
Ln σ_2			
	1.091^{***}	1.091^{***}	1.094^{***}
	(79.83)	(79.81)	(80.05)
ρ_{12}			
	-0.475^{**}	-0.416^{**}	-0.449^{*}
	(3.51)	(2.96)	(2.48)
N	2678	2678	2678

Absolute *t* statistics in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. $\sigma_1 = 1$ (normalized).

Table 4: Additional models

	(1) +/-5 yr window	(2) No foreign	(3) Binary probit
Years education	0.193***	0.209***	0.224**
	(4.15)	(4.85)	(2.97)
Years education			
"No fees"=1	0.973***	0.951***	1.119***
	(4.01)	(4.30)	(5.30)
"No fees"=2	1.642***	1.544***	1.592**
	(5.15)	(5.65)	(5.85)
Father lower secondary education x "No fees"	-0.650**	-0.617**	-0.614
	(2.46)	(3.15)	(3.08)
Father upper secondary education x "No fees"	-0.928***	-0.576***	-0.646
	(3.59)	(2.95)	(3.08)
Father tertiary education x "No fees"	0.166	0.166	-0.318
	(0.53)	(0.69)	(1.31)
Ln σ_2	1.090***	1.08***	
	(61.24)	(75.56)	
ρ_{12}	-0.468**	-0.501**	
	(3.15)	(3.63)	
Wald test of exogeneity $\chi^2(1)$			6.33
P-value			0.012
Over-identification $\chi^2(4)$			3.185
P-value			0.527
N	1578	2450	2678

Absolute *t* statistics in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. σ_1 =1 (normalized). All models contain the same covariates as in Tables 2 and 3, details available on request. The dependent variable in column 3 is a binary variable equal to 1 for the first ("agree strongly") category and 0 otherwise.

Table 5: Marginal effect of one year of education

"Gays and lesbians should be free to live life as they wish"	(1) Table 2, col. 2	(2) Table 3, col. 1
Agree strongly	.0127	.0507
	(5.62)	(5.18)
Agree	-.0023	-.0045
	(3.67)	(2.69)
Neither agree nor disagree	-.0053	-.0175
	(5.45)	(9.44)
Disagree	-.0037	-.0167
	(5.18)	(3.84)
Disagree strongly	-.0015	-.0122
	(4.50)	(1.63)

Absolute t ratios in parentheses. Coefficients show the effect of an increase in one year of education on the probability of each of the five outcomes occurring. Note that .05 corresponds to 5 percentage points. Column totals may not add to zero due to rounding.

Appendix: replication for the United Kingdom

The table below replicates the model shown in Table 4, column 3 where the dependent variable is a binary variable equal to one for the top category (the “agree strongly” response) and is zero otherwise using the United Kingdom data from the same data source, the ESS. The covariates (a quadratic in age, sex, paternal education, living in a city and foreign born) are the same. The instruments are the raising of the minimum school leaving age (from 14 to 15) generated by legislation in 1944 (England and Wales) and 1945 (Scotland) – this is “Rosla 1” and a further increase (from 15 to 16) in 1973 (England, Wales and Northern Ireland) and 1976 (Scotland) – this is “Rosla 2”.

Table A1

	(1)	(2)	(3)	(4)
	+/- 10 yr window		+/-5 year window	
	Probit	IV probit	Probit	IV probit
Years education	0.0416***	0.242*	0.0418***	0.184
	(6.73)	(2.13)	(6.11)	(1.84)
Years education				
Rosla 1		0.512**		0.586**
		(2.91)		(3.06)
Rosla 2		0.641**		0.542**
		(3.46)		(2.58)
Wald test of exogeneity $\chi^2(1)$		3.75		2.72
p-value		0.0527		0.0991
Over-identification $\chi^2(1)$		0.141		1.768
p-value		0.7074		0.1836
N	5261		4252	

Absolute *t* statistics in parentheses. Model specification is identical to Table 4, column 3. Coefficients on the other variables are available on request.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

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