

Inequality and Union Membership: The Influence of Relative Earnings and Inequality Attitudes

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Abstract

Using surveys from the International Social Survey Programme covering the period 1985–2002 for seven European countries (West and East Germany, Sweden, Norway, Italy, the Netherlands and Great Britain), we examine the effect of relative earnings on union membership and show that union density is higher among workers in the intermediate earnings group than among low or high earners. Next, we examine the association of inequality attitudes with union membership and demonstrate that union membership is not only motivated by instrumental considerations related to relative earnings, but also by normative concerns about inequality. We interpret our findings suggesting that rising earnings inequality is in itself a source of union decline.

1. Introduction

Trade union membership and density is in decline in many countries (OECD 2004; Visser 2006). Its causes are not fully known. Cross-national empirical studies of unionization trends have mostly analysed aggregate statistics, distinguishing between structural, cyclical and institutional causes (Checchi and Visser 2005; Ebbinghaus and Visser 1999; Scruggs and Lange 2002; Western 1997). Although some researchers have made use of individual or micro data to explain differences in unionization across countries (Bryson and Gomez 2005; Bryson *et al.* 2005; Schnabel and Wagner 2003, 2005), a combination of longitudinal and cross-sectional research based on micro data is rare (exceptions being Blanchflower 1996, 2006; Schnabel and Wagner 2005), which has limited the scope of research on international variation in unionization

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trends. In this article, we aim to widen the scope by examining two groups of individual characteristics that have often been neglected: relative earnings and social attitudes towards inequality.

We start by analysing the relationship between relative earnings and union membership. Our hypothesis is that both low and high earners are, for different reasons, less likely to join the union or retain membership. Our starting point is that unions generally favour wage compression (see Visser and Checchi 2009, for an overview). This is true, in particular, for unions that try to organize workers with low *and* high earnings or skills in the same sectoral organization, as is the case in most of the European countries analysed here (Ebbinghaus and Visser 2000). In order to attract low earners, such sectoral unions must propose egalitarian policies in which wages are disconnected from the relative skills and market power of particular groups of workers, and they must rely, in part, on the stronger market power of those with higher skills to achieve improvements for those with less market power (Moene and Wallerstein 1997; Pontusson 2001; Streeck 2005). However, in order to be organizationally viable and prevent higher-earning member groups from seceding, the union must soften its *ex ante* egalitarian stance in central negotiations, for instance, by allowing *ex post* wage drift (Holden 1987).

From this we deduct our hypothesis that the union is likely to disappoint lower-earning workers to the same extent as it tries to remain attractive for higher-earning workers. Whereas the explanation why high earners are less attracted to a union is straightforward, especially if the union tries to compress wages, the reason why low earners are underrepresented in the union is more complex and related to disappointment with the lack of egalitarian policies and the perception that the union does not halt the rise in inequality.

In addition to estimating the effect of relative earnings on union membership, we analyse the influence of social attitudes towards inequality. These attitudes are likely to differ across individuals and may have changed over time. Generally, trade unions expound a philosophy of equality and advertise their actions as contributing to more fairness in the treatment of workers and more equality in opportunities and rewards (Visser and Checchi 2009). Our hypothesis is that workers who subscribe to the view that society is too unequal, that greater equality is desirable and collective action is needed to reduce inequality, will be more attracted to unions than workers who do not hold these views. Such normative views may also be held by workers located in the upper tail of the earnings distribution who themselves would lose from wage-compressing union policies. These views may weaken or cancel instrumental reasons for joining or leaving the union and thus weaken the hypothesized association between relative earnings and union membership. The relative strength of and interaction between these instrumental and normative concerns based on relative earnings and social attitudes, respectively, are of prime interest for our analysis.

If our hypothesis that workers at the lower and upper tails of the earnings distribution are underrepresented in the union proves correct, and if it is true that in most advanced economies the tails of the earnings distribution have

become larger and wider apart (Alderson and Nielsen 2002; EC 2005; Gottschalk and Joyce 1998; OECD 1996; Rueda and Pontusson 2000; Visser 2004), the combination of these two observations could in itself be an additional explanation for the observed decline in unionization. Such a development may be tempered or exacerbated by a change in attitudes or a change in the association between attitudes, relative earnings and union membership.

We are aware that we can only offer a partial solution to the puzzle of the current trend of union decline in advanced economies. Most researchers have started from the other side and treated union decline as one of the causes of rising earnings inequality (Alderson and Nielsen 2002; DiNardo *et al.* 1996; Freeman 1980; Rueda and Pontusson 2000; Wallerstein 1999). Comparing the effect of unions on wage dispersion in 1984–1985, 1991–1992 and 2001 in the USA, Canada and the UK, Card *et al.* (2003) demonstrate that unions have become less effective in containing rising earnings inequality within and across skill groups among male wage earners. They relate this to the sharp decline in male union density. Comparing data for 1998 and 1978, Metcalf *et al.* (2001) show that British unions have become less successful in narrowing the pay gap between male and female, manual and non-manual, immigrant and native workers. We do not quarrel with these findings. The novelty of our research is that we provide empirical evidence consistent with a potential reverse link (to be interpreted as causal in some cases, but not in others), where increased inequality affects union membership rates by lowering the incentives for becoming union members at both tails of the earnings distribution. Our analysis thus contributes to an understanding of the current trend of union decline as a dynamic process that potentially feeds on itself. Seen in this perspective, our contribution is in line with the explanation originally proposed by Hirsch (1982), where union membership and earnings inequality are jointly determined, with density negatively affecting earnings dispersion and dispersion exhibiting ambiguous effects on union membership.

The article proceeds as follows. In Section 2, we first present a simple model relating union membership to individual and industry characteristics, followed by a discussion of union wage policies in relation to relative earnings and social attitudes. In Section 3, we describe our dataset, derived from the International Social Survey Programme (ISSP), and set out our empirical strategy. In Section 4, we present and discuss our results. The article ends with conclusions and a discussion of the findings.

2. Trade unions and inequality

Determinants of Union Membership

Following the framework originally proposed by Farber and Saks (1980), we consider the individual decision to obtain union membership as deriving from a comparison between the expected utility of being a union member and the expected utility of not being a union member.¹ If the worker is risk or

inequality averse, she will take into account not only the expected wage, but also the expected distribution under unionization. Assuming that wages on average are higher and less dispersed under unionization, we expect the propensity of union membership to rise with the size of the (expected) union wage premium and the extent of wage dispersion without unions. Other factors are the strength of risk or inequality aversion of workers and the costs of membership. From these predictions, it would be straightforward to derive a union status equation, if information about workers' expectations about the wage distribution under alternative regimes were available. Unfortunately, this information is rarely available, and we are forced to rely on observable worker characteristics, to be interpreted as proxies for some of these determinants or motivations.² In this article, we focus on the individual's relative income position and her preferences or attitudes concerning inequality. High-earning workers are less likely to be union members because they may be less risk averse (if risk aversion declines with individual wealth). Furthermore, they may exhibit greater acceptance of inequality (for reasons of cognitive dissonance), and they may face higher absolute costs of union membership (if dues are somewhat proportional to monthly wages, as is often the case in European trade unions).

Also, low-wage earners can be less likely to become union members, due to *ex post* disillusion. While *ex ante* each worker would expect a positive wage premium from union membership, some workers may end up in occupations where the premium is low or non-existent. This is particularly true for marginal occupations, where their bargaining power is limited and the 'union premium' is squeezed by the competitive/decentralized working of the labour market and by the pressure of the informal/illegal sector (janitors, cleaning, retail, hotel and restaurants).³ A worker ending up in these occupations will experience a lower (than average) wage and no wage premium (Ebbinghaus and Visser 1999). Moreover, in these sectors, collective agreements tend to be declared binding by public law in many European countries irrespective of union membership, and statutory minimum wages are especially relevant in these sectors (EC 2004; Hassel 2008).

One reason why such disillusion may take place results from the compromises which broad sectoral unions of the European type must strike within their own ranks if they want to keep their higher- and middle-earning members with the strongest bargaining capacity on board. In addition, they must strike a compromise with employers who usually favour larger wage differentials. Thus, if it is true that low-earning workers expect *ex ante* improvements in their earnings position from participating in the union, the chances of *ex post* disappointment are large.⁴ In fact, during the 1970s and 1980s, many European trade unions did try to defend the interests of lower-paid workers in the context of centralized wage policies, but their policies were blunted by higher wage drift for skilled workers, with the complicity of employers, and by the rise of independent unions for higher-earning staff, or secessions from existing ones (Ebbinghaus and Visser 2000; Holden 1987;

Visser 1990). This led to a general correction of union policies away from radical attempts to change existing earnings distributions, in part by moderating highly centralized wage and incomes policies (EC 2004; Pontusson 2001; Visser 2004).

Thus, unless their risk- or inequality-aversion is very strong, or they are normatively motivated by a strong dislike of inequality, low-earning workers may have limited incentives for becoming union members. These arguments are reinforced when we consider that low-earning workers face stronger liquidity constraints and have stronger incentives for saving the money otherwise spent on union membership dues.

By jointly considering the previous arguments, we expect an asymmetric relationship between relative earnings positions and the propensity of joining the union. We operationalize this model by taking the median earnings as a reference point, expecting a decline in union membership the further away we move from the median. The (relative to the median) earnings position does not rule out more traditional explanations based on personal (age, gender, education, occupation) and industry (concentration, capital intensity, employment composition) characteristics.⁵ For these reasons, in addition to relative earnings (distinguishing between being a high earner or low earner), we include in the present analysis age, gender, marital status, full-time or part-time hours, occupational status (supervisory position), type and level of education, and sector (working for a public or private firm) among the potential determinants of union membership.

Unfortunately, the survey data we use do not allow us to check for the extensions proposed by Booth (1985) and Naylor (1990) and take into account norms of membership established and sanctioned by reference groups like parents, friends or colleagues.⁶ We partially consider this aspect by including the worker's attitude towards inequality, as measured by opinions expressed in surveys. In this way, we are able to identify various individual features that are positively correlated with union member status without necessarily being correlated among themselves. For example, high-earning workers have a lower propensity of union membership because they share an anti-egalitarian ideology; they hold managerial responsibilities within the firm; or they have a lower demand for insurance offered by the unions (or any combination of the three motives). If the first two explanations are properly controlled for, the finding of a significant correlation with relative earnings suggests a genuine new effect. The same holds for the effect on union membership of an earnings position below the median, once we control for working hours and educational attainment.

Social Attitudes

There are three reasons to include attitudes to inequality in a study of unionization trends. First, it is substantively interesting to assess their effects on union membership. Second, by including them in our models, we can

examine to what extent well-known variations in membership across sub-groups are explained by differences in attitudes. Third, attitudes serve as a control variable for our main explanation of membership decline related to relative earnings. Since our key hypothesis is derived from assumptions about how the decision to become a union member is influenced by the advantage or disadvantage of egalitarian union wage policies, we need to control for attitudinal factors to ensure that the variation in the propensity of union membership across earnings positions is not influenced by normative rather than instrumental reasons. This is especially important, since research has shown that attitudes towards inequality are in fact correlated with financial resources (Svallfors 1993, 1997).

Attitudes towards inequality relevant for union membership can be articulated in (1) a general view about the extent of inequality in the society; (2) specific opinions concerning the acceptable gap between actual and legitimate earnings; and (3) a view about collective action needed to combat inequality. The first is the most general and simply comprises the affirmation that *'inequality in society is too large'*. The second type is more specific and assumes that individuals, when they evaluate the extent of inequality in a particular society, have specific ideas about the perceived *actual* and *legitimate* earnings of particular occupations. Relating the estimates of respondents of what they believe are 'acceptable' or 'legitimate' earnings, we can construct a preference for more or less egalitarian outcomes or 'tolerance of inequality' attitude (e.g. Jasso 1999; Osberg and Smeeding 2006). The third type of attitude that may influence workers' motivations to join a union is more specifically concerned with the perceived need for collective action in order to reduce inequality. If egalitarian attitudes motivate people to join the union, this may coincide with the conviction that collective action is necessary. We call this the 'collective action' attitude. It is measured by asking people if they agree with the statement that *'inequality in society exists because people do not join together'*.

With regard to the relative importance of these three types of attitudes, there are two scenarios. First, it might be argued that more specific attitudes have a stronger influence than general inequality attitudes, the latter being vague, less contextualized and less committing. Research on intentions has shown that the more these can be put in concrete terms, the stronger their predictive power for subsequent behaviour (Ajzen and Fishbein 1980). This would imply that the 'collective action' attitude should have a larger effect than the other two and that the 'general' inequality attitude should have the weakest effect. However, people may think about other forms of collective action, for instance, political action and taxation, or they may criticize the unions for being ineffective while they score high on the 'collective action' attitude. In that case, a more general attitude towards inequality should have a stronger association with union membership than the more specific attitude on collective action. Given this uncertainty, we will use all three indicators to analyse the relevance of attitudes.

3. Data and operationalization

Data

Using the cross-national surveys of the ISSP, we analyse the data from 1985 to 2002 for seven countries: West Germany, East Germany, Sweden, Norway, Italy, the Netherlands and Great Britain. We have selected these countries according to two criteria. On the one hand, we take into account the presence of data on membership, relative earnings and social attitudes; on the other hand, we consider some institutional variations. The seven countries give a good spread of different employment and industrial relations arrangements (Ebbinghaus and Visser 1997; Esping-Andersen 1990; Gallie 2007), from 'inclusive' and co-ordinated (Sweden, Norway), to 'dual', status-based (Germany,⁷ the Netherlands, Italy) to market-based, uncoordinated (UK) ones. This variation allows us to distinguish between countries with different collective bargaining *institutions*, varying from the relatively centralized wage bargaining systems in Scandinavia to the highly decentralized one in Britain (see Table 1).

These institutional differences are likely to be reflected, first, in different union policies (more egalitarian in inclusive and co-ordinated systems) and in their impact on membership, and, second, in different attitudes to inequality. The adverse effects on higher-earning members should be lower where union or bargaining coverage is high, since it will be difficult to set up rival unions. Conversely, where there is a statutory minimum wage, that is, a publicly guaranteed protection mechanism that may be perceived as an alternative for the union (in our sample this applies only to the Netherlands and, since 1999, to the UK), lower-paid members may have fewer instrumental reasons to join the union.

Our overall sample consists of 55,109 individuals, unevenly distributed across countries: West Germany (1985–2002, 18 surveys, 12,094

TABLE 1
Cross-National Variations in Institutional Characteristics

<i>Employment regime</i>	<i>Inclusive</i>		<i>Dual</i>			<i>Liberal</i>
	<i>Sweden</i>	<i>Norway</i>	<i>Germany</i>	<i>The Netherlands</i>	<i>Italy</i>	<i>UK</i>
Union density	80.1	54.4	24.6 25.5 (W) 19.0 (E)	22.6	34.7	29.6
Bargaining coverage	91.0	72.0	63.0 68.0 (W) 58.0 (E)	86.0	60.0	36.3
Union centralization	0.541	0.530	0.518	0.643	0.388	0.302
Rival unions	No	Yes	No	Yes	Yes	Yes
Statutory minimum wage	No	No	No	No	No	Yes (1999–)

Source: ICTWSS Database Visser (2009), Schnabel and Wagner (2003), and Ellguth and Kohaut (2005) for East and West Germany; all data for 2000–2001.

observations), East Germany (1991–2002, 12 surveys, 4,094 observations), Sweden (1994–2002, 8 surveys, 5,194 observations), Norway (1989–2002, 14 surveys, 11,615 observations), Italy (1985–1998, 12 surveys, 4,628 observations), the Netherlands (1987–2002, 11 surveys, 7,901 observations) and Great Britain (1985–2002, 18 surveys, 9,583 observations). Within each country/year sample, we have included in our analysis wage and salary earners in employment, excluding the unemployed, the self-employed and those outside the labour force. If we exclude one survey (Britain in 1989), the ISSP data on membership are in line with the evidence obtained from administrative data and national surveys (Ebbinghaus and Visser 2000; Visser 2006, 2009).

The attitudes regarding inequality are part of the ISSP surveys of 1987, 1992 and 1999. Unfortunately, not all three years are available for all countries. Trends across three survey years could be analysed for West Germany and Britain. For Norway and East Germany, two years are available (1992 and 1999). For three countries, only one survey year was available: the Netherlands (1987), Italy (1992) and Sweden (1999).

Relative Earnings Positions

We describe the relative earnings position of each individual by the relative distance from the survey median income (computed within each country/year sample), normalized by the median income. Positions above or below the median are kept distinct in order to distinguish different attitudes of people at both tails of the earnings distribution. If we define y_{ict} as the earnings of individual i in country c and survey year t , and \hat{y}_{ct} as the median earnings for the same country/year, our measures are given by

$$R_{ict}^{above} = \frac{y_{ict} - \hat{y}_{ct}}{\hat{y}_{ct}}; \quad R_{ict}^{below} = 0 \quad \text{if } y_{ict} > \hat{y}_{ct}. \quad (1)$$

$$R_{ict}^{above} = 0; \quad R_{ict}^{below} = \frac{\hat{y}_{ct} - y_{ict}}{\hat{y}_{ct}} \quad \text{if } y_{ict} < \hat{y}_{ct}. \quad (2)$$

A higher score indicates a larger distance between individual earnings and the median earnings in the country/year under study.

Social Attitudes

We consider three different attitudes towards inequality. The ‘general inequality’ attitude is based on the survey item ‘*Earnings inequality in my country is much too high*’. A higher value indicates agreement. The ‘tolerance of inequality’ attitude is based on work by Osberg and Smeeding (2006). For five occupations (doctor, chair of a large company, skilled worker, unskilled worker and minister), individual respondents estimate the actual earnings that they believe these occupations pay as well as the earnings that these

occupations in their opinion *should* pay.⁸ Next, for each respondent, an ordinary least squares (OLS) regression coefficient is computed (on $N = 5$ occupations), regressing the *should-earn* answer (representing the dependent variable Y) onto the *do-earn* answer (as our predictor X). A higher slope coefficient expresses a greater acceptance or tolerance of inequality. Finally, the ‘collective action’ attitude is based on the survey item ‘*Inequality exists because people do not join together*’. A higher score indicates agreement.

For all three types of attitude, we took the standardized z -value across all country/year combinations.⁹ The correlation between the ‘general inequality’ and the ‘inequality tolerance’ attitude is -0.336 ; between ‘inequality tolerance’ and the ‘inequality needs collective action’ attitude -0.104 ; and between the ‘general inequality’ and the ‘inequality needs collective action’ attitude 0.276 .

Additional Controls

In addition to relative earnings and social attitudes, we control for the usual demographics (gender, age, marital status), for education (measured in four categories: lower secondary or less, corresponding to the OECD definition of ‘dropout’; upper secondary; beyond secondary; completed university degree) and some job characteristics (working hours, public/private, supervising someone else).

4. Results

Earnings Position and Union Membership

Our first analysis consists of an empirical estimation of the effects of various observable characteristics, including earnings position, on union membership. We do this for all available surveys from 1985 to 2002.¹⁰ Figure 1 depicts the kernel densities of earnings by union status, showing the differences between unions and non-union members.¹¹ Clearly, in some countries (notably Sweden and Norway), the earnings distribution of union members is more compressed. For the remaining countries, the density distribution exhibits a larger mass around the modal values, with some indication of a union premium shifting the union member distribution to the right; in this case, membership is less likely among low-wage earners and more likely among those with higher earnings. Britain seems the most telling example of such a shift.

In order to explore the issue further, we estimate a probit model predicting union membership. The coefficients are marginal effects, indicating the difference in the probability if the independent variable undergoes a unitary change (see Table 2). We observe that women are less likely to be in the union than men. However, in the two Scandinavian countries, this effect is smaller, if at all significant, testifying to the ‘inclusive’ character of Scandinavian

FIGURE 1
Kernel Density of (Log)Earnings, by Union Membership and Countries.

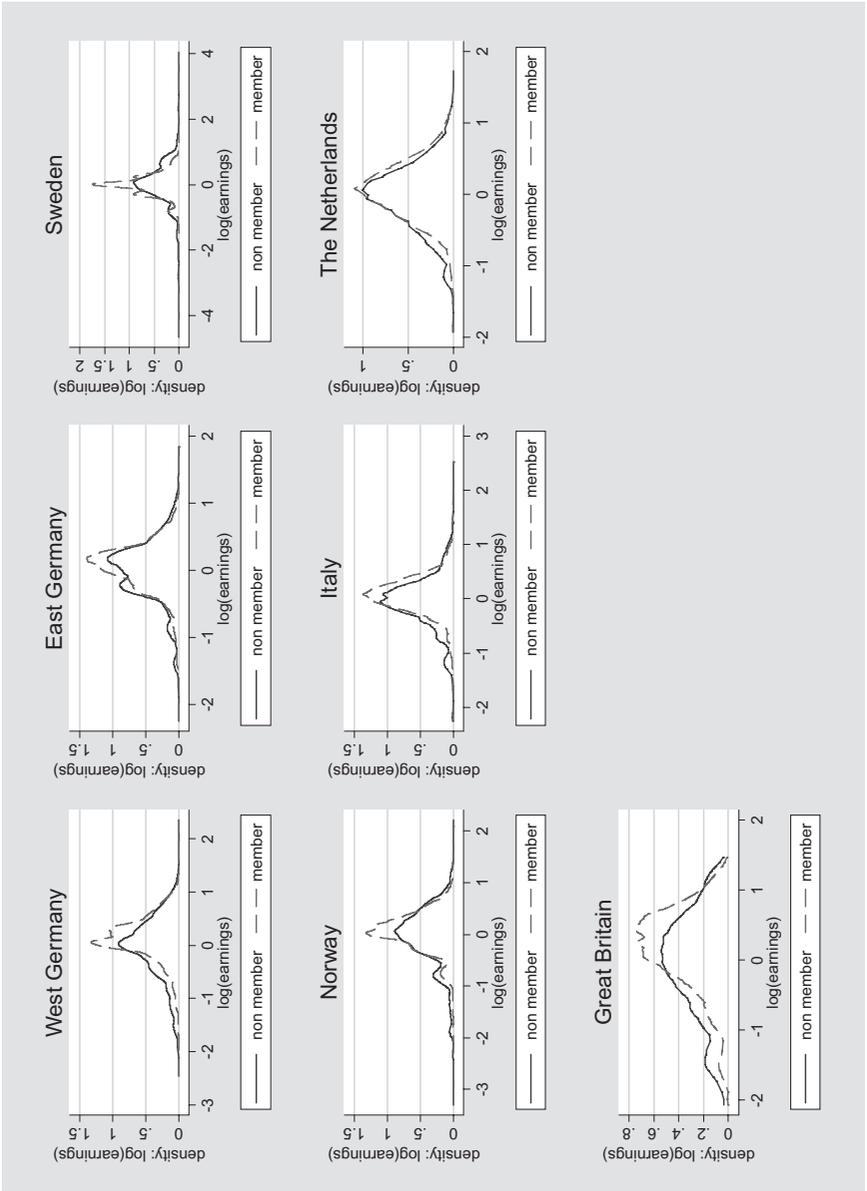


TABLE 2
 Probit Regression of Union Membership (Marginal Effects) — ISSP 1985–2002

	West Germany	East Germany	Sweden	Norway	Italy	The Netherlands	Britain
Relative distance from the median earnings (above)	-0.064 [0.014]***	-0.113 [0.030]***	-0.012 [0.006]*	-0.085 [0.015]***	-0.027 [0.018]	-0.042 [0.015]***	-0.045 [0.012]***
Relative distance from the median earnings (below)	-0.251 [0.030]***	-0.278 [0.054]***	-0.206 [0.034]***	-0.347 [0.034]***	-0.120 [0.055]**	-0.076 [0.031]**	-0.392 [0.030]***
Female	-0.135 [0.011]***	-0.036 [0.020]*	0.013 [0.012]	-0.031 [0.013]**	-0.091 [0.021]***	-0.105 [0.015]***	-0.078 [0.013]***
Age	0.010 [0.002]***	0.021 [0.005]***	0.019 [0.002]***	0.030 [0.003]***	0.025 [0.005]***	0.022 [0.004]**	0.023 [0.003]***
Age ²	-0.00009 [0.000]**	-0.00021 [0.000]***	-0.0002 [0.000]***	-0.00029 [0.000]***	-0.00026 [0.000]**	-0.00018 [0.000]**	-0.00025 [0.000]***
Married	0.009 [0.011]	-0.031 [0.021]	0.012 [0.010]	0.023 [0.011]*	0.076 [0.022]***	-0.010 [0.013]	0.018 [0.012]
Working less than full-time	-0.094 [0.016]***	0.073 [0.029]**	0.012 [0.014]	-0.067 [0.019]***	-0.109 [0.028]***	-0.032 [0.015]**	-0.041 [0.017]**
Working for government/public owned	0.126 [0.011]***	0.085 [0.020]***	0.134 [0.010]***	0.392 [0.009]***	0.185 [0.025]***	0.148 [0.013]	0.409 [0.012]***
Supervising someone	-0.038 [0.012]***	0.010 [0.020]	-0.058 [0.009]***	-0.032 [0.011]***	0.071 [0.021]**	0.000 [0.015]**	-0.057 [0.012]***
Upper secondary completed	-0.052 [0.017]***	-0.065 [0.036]*	-0.058 [0.015]***	0.016 [0.013]	-0.027 [0.021]**	0.001 [0.015]	-0.019 [0.016]
Beyond secondary/incomplete college	-0.075 [0.020]***	0.018 [0.038]	-0.079 [0.022]***	-0.017 [0.017]	0.015 [0.034]	0.034 [0.022]	-0.004 [0.015]
University degree completed	-0.135 [0.016]***	-0.069 [0.035]*	-0.075 [0.019]***	-0.030 [0.020]	-0.119 [0.037]***	-0.078 [0.018]**	-0.124 [0.021]***
Female*university degree completed	0.187 [0.040]***	0.057 [0.058]	0.034 [0.018]*	0.105 [0.021]***	0.080 [0.063]	0.171 [0.031]**	0.168 [0.031]***
Observations	8831	3141	4899	10040	2988	7066	9262
Pseudo R ²	0.08	0.09	0.12	0.18	0.09	0.06	0.16
Log likelihood	-4953.58	-1860.49	-1748.72	-5367.79	-1832.99	-4253.14	-5276.91
Test (coeff. below median > coeff. above median) (<i>p</i> -value)	1.00	0.99	1.00	1.00	0.97	0.94	1.00

Robust standard errors in brackets — * significant at 10%; ** significant at 5%; *** significant at 1% — year dummies included. Missing values are replaced with sample averages in order to retain sample size.
 ISSP, International Social Survey Programme.

unions. In all countries, the probability of union membership exhibits an inverted *U*-shaped age profile, with peaks ranging between the age of 47 (Sweden, Italy and Britain) and 61 (the Netherlands), with Germany and Norway in between. As a consequence, young workers (say below the age of 30) are less likely to be union members than older workers (Blanchflower 1996, 2006). In Italy, and to a lesser extent in Norway, married workers are more likely to be unionized than those not married. Supervisors are less likely to be in the union, except in East Germany and the Netherlands. There is a very strong association of public sector employment with union membership in all countries, particularly so in Norway and Britain.

The association with education varies across countries. The general pattern is that people with at most lower secondary level (the excluded case, corresponding to high school dropout) have the highest probability of union membership, whereas membership of workers holding a university degree is less likely. West Germany is the only country where the probability of membership reduces monotonically for each additional step on the educational ladder. There is also a rather large education effect in Sweden, although the probability of union membership does not significantly vary between university graduates and workers with some post-secondary education. In Italy, the Netherlands and Britain, male university graduates have a lower probability of membership than workers who dropped out of school (i.e. with lower secondary education or less). Among women, university graduates do not have a lower likelihood of membership than women with lower levels of schooling (as can be seen from the positive interaction terms relative to the negative main effects of university degree). This may be related to the larger proportion of highly educated women employed in unionized public service professions like teaching and nursing.

With regard to the impact of relative earnings, the data do not contradict our theoretical expectation that trade unions mainly attract workers from the intermediate earnings groups. The further one's personal income is from the median, the lower the estimated likelihood of membership.¹² This holds in all countries, except for 'being above the median' in Italy, where the coefficients are only significant at the 20 per cent confidence interval. Our finding is in accordance with Addison *et al.* (2003: 16), who claim that in the UK 'the distribution of union density has become less pro-poor over time, shifting for example from the less educated to the better educated.' For Germany, however, Goerke and Pannenberg (2004) report a negative correlation between membership and earnings position.

An inspection of Table 2 reveals that in all countries, the negative effect on the probability of union membership of earnings *below* the median is stronger than that of the earnings *above* the median (as witnessed by the one-side test, reported at the bottom of Table 2). This is also confirmed by rerunning the probit estimations for each survey or country/year combination. The number of statistically significant coefficients (at the 10 per cent level) is generally higher for relative earnings below the median than for those above. When we analyse each survey/year separately, we do not find a systematic change or

trend in the association between relative earnings (distance above or below the median) and the probability of union membership. This suggests that it is the change in the earnings *distribution*, that is, the contraction of the middle, rather than an increased disaffection of lower- or higher-earning workers, which is relevant for the recent decline in union membership.

This interpretation can be accepted only if the position in the earnings distribution can be considered as (strongly) exogenous with respect to union membership decisions. Empirical research on this is rare and only done in the context of voting for membership in the USA (Farber and Saks 1980; Hirsch 1982). However, with union contracts covering most firms and making no distinction between union and non-union member, it is not a priori clear why union members should earn more than non-members.¹³ But even if unions were able to shift the entire earnings distribution to the right for their members only (which, however, is not evident in the raw data — see Figure 1), our results — showing a lower propensity of lower- and higher-paid workers to be union members — still begs for an explanation.

With the data available, we cannot exclude the potential endogeneity of the individual income position due to unobservables (like ability, political orientation, religion and so on). In such a case, our results would be reflecting spurious correlation between the likelihood of union membership and relative earnings. However, we can check the robustness of our results by resorting to instrumental variable (IV) estimation. A good instrument should exhibit correlation with the independent variable (relative earnings) without affecting union membership status. Examples of shifters of relative earnings are educational attainment, full-time employment or employment sector, but these variables are known to be associated with union membership and have already been introduced as covariates in our analysis. Unfortunately, another candidate, parental background, in this dataset is available only for West Germany. Other potential candidates are variables related to family conditions (family size and/or being married, which we have already seen being uncorrelated with union membership in all countries but Italy), on the argument that establishing a (large) family puts pressure on workers to earn more. However, we need at least two instruments since we have two relative earnings positions that are to be instrumented. For this reason, we have extended the set of instruments to the upper and lower family income quartiles. Under the assumption of assortative mating (people tend to marry with others in the same income position), we expect that this variable affects relative earnings without itself changing the likelihood of union membership. The first stage *F*-tests are satisfying, suggesting that the selected instruments are effective in shifting the relative income position. They are good instruments (i.e. uncorrelated with our main dependent variable) in five cases out of seven (West and East Germany, Sweden, Norway, Italy — upper part of Table 3). By replacing family size with marital status, we obtain satisfactory results in the Hansen overidentification test for the remaining two countries (the Netherlands and Britain — lower part of Table 3). Examining the relevant coefficients, we find that for four countries (Sweden, Norway, the Netherlands and

TABLE 3
 Linear Probability Regression of Union Membership (Marginal Effects) — ISSP 1985–2002 Instrumental Variables — Alternative Specifications

	<i>IV: family size, 1st and 4th family income quartile</i>						
	<i>West Germany</i>	<i>East Germany</i>	<i>Sweden</i>	<i>Norway</i>	<i>Italy</i>	<i>The Netherlands</i>	<i>Britain</i>
% distance from the median earnings (above)	-0.053 [0.032]	-0.1 [0.116]	-0.361 [0.069]***	-0.218 [0.044]***	0.01 [0.074]	-0.088 [0.033]***	-0.08 [0.035]**
% distance from the median earnings (below)	-0.162 [0.158]	-0.44 [0.279]	-0.802 [0.484]*	-0.469 [0.149]***	-0.177 [0.151]	-0.117 [0.076]	-0.333 [0.113]***
Observations	8841	3150	4909	10046	3720	3808	9265
Adjusted R ²	0.08	0.1	-0.91	0.21	0.11	0.06	0.19
RMSE	0.44	0.45	0.48	0.43	0.46	0.45	0.44
F-test on the first stage (above the median)	55.83	97.75	60.86	337.51	152.24	508.94	616.82
F-test on the first stage (below the median)	49.48	66.96	40.15	169.2	194.46	596.93	333.88
Hansen J	0.22	3.01	0.01	0.84	1.06	10.36	9.82
Hansen J p-value	0.64	0.08	0.92	0.84	0.30	0.00	0.00
Log likelihood	-5235.93	-1978.39	-3371.6	-5723.39	-2408.82	-2391.56	-5562.86

	<i>IV: married, 1st and 4th family income quartile</i>						
% distance from the median earnings (above)	-0.073 [0.067]	-0.1 [0.115]	-0.361 [0.067]***	-0.192 [0.046]***	0.022 [0.069]	-0.071 [0.025]***	-0.127 [0.039]***
% distance from the median earnings (below)	-0.387 [0.494]	-0.357 [0.274]	-0.831 [0.465]*	-0.346 [0.186]*	-0.117 [0.158]	-0.092 [0.048]*	-0.593 [0.143]***
Observations	8847	3150	4942	10987	4078	7121	9267
Adjusted R ²	0.07	0.1	-0.91	0.21	0.1	0.06	0.18
RMSE	0.44	0.45	0.48	0.43	0.47	0.46	0.44
F-test on the first stage (above the median)	191.94	110.98	62.26	325.21	62.41	1112.63	653.43
F-test on the first stage (below the median)	46.77	71.05	39.87	97.54	173.19	1508.92	240.07
Hansen J	0.91	3.12	0.31	4.73	14.05	0.01	2.18
Hansen J p-value	0.34	0.08	0.58	0.03	0.00	0.91	0.14
Log likelihood	-5277.62	-1973.98	-3390.91	-6253.07	-2665.32	-4519.65	-5625.52

Standard errors in brackets — * significant at 10%; ** significant at 5%; *** significant at 1% — year dummies included. ISSP, International Social Survey Programme; RMSE, root mean squared error. Covariates as in Table 2.

Britain) IV estimates of the relative income position retain their signs and statistical significance, with the magnitude of the coefficient being higher than the corresponding OLS coefficients. The IV estimates also retain the feature that being below the median exhibits a stronger effect on union membership than being above the median.

The validity of the results hinges strictly on the quality of the instruments, which is admittedly suboptimal, since their validity is based on the assumption of assortative mating with regard to income. It is therefore not surprising that specification tests reject some instruments for some countries though not for others. It is, however, reassuring that, irrespective of the type of instrument adopted, in four of the seven countries (Sweden, Norway, the Netherlands and the UK), we cannot reject the hypothesis that the causal arrow runs from relative income position to union membership.¹⁴ When we use marriage status as an IV, we can only reject the null hypothesis of no causal impact at a confidence level of 10 per cent for three of the other four countries. In the case of Italy and Germany, we find no evidence of any impact of the relative income position on the probability of union membership when income position is exogenously shifted by family income position. Whether this is related to specific union policies or institutional features of collective bargaining is impossible to say with the available data.

IV estimates do not shed additional light on the reasons why workers whose earnings are further away from the median tend to be less represented in the union. The attitudes expressed in the questionnaires, used in the ISSP surveys and discussed below, concern the general perception of inequality, not the assessment of the individual or family position in the earnings distribution. Only in-depth interviews would be able to discriminate among alternative motivations favouring entering or quitting union membership. Had the ISSP been a panel study following the same individuals over time and with data on individuals entering and exiting union membership (Elias 1996; Visser 2002), we could provide a stronger test. At present, using repeated cross sections, we cannot exclude that worker ability (yielding higher earnings or favouring individualistic behaviour) and risk aversion (leading workers to accept lower wages in exchange for longer contract duration, as well as raising the demand for union protection — see Goerke and Pannenberg (2007)) — may also account for part of our results.¹⁵

The Impact of Inequality Attitudes on Union Membership

When we consider aggregate data in (average) inequality attitudes, we observe that the lowest preference for levelling earnings relative to the perceived situation (i.e. a higher regression slope of the Osberg–Smeeding index) is found in Norway (1992 and 1999), in the Netherlands (1987) and in West Germany (1999).¹⁶ In addition, we notice that, particularly in the 1990s, there was an increasing tolerance towards inequality in all countries where we can compare different years. Similarly, support for the statement that ‘earnings inequality is too high’ declined slightly in the 1990s in both East and West

Germany and the UK (after a rise between 1987 and 1992), but only in the case of West Germany (1999) do we detect a statistically significant trend. Finally, we find no significant trend in the collective action (for achieving lower inequality) attitude.

We examine the correlation of social attitudes with union membership by estimating a probit regression analysis on the pooled sample. The results are reported in Table 4, showing four model specifications. Model 0 uses the same variables as in Table 2, but now with the fewer datasets that include the attitude items, and with a pooled analysis for all countries including country dummies. Model 1 adds the three attitudinal variables. Model 2 adds the multiplicative interactions between relative earnings and social attitudes, and model 3 replaces these interactions with interactions between survey year and attitudes.

In model 0, we find support for most of the findings reported earlier for the full dataset with all years. Most importantly for our purpose, people with earnings that are further away from the median (above and particularly below) are less likely to be found in the unions. In model 1, we see that attitudinal factors have the expected association with union membership. People who accept greater inequality have a lower probability of membership. Individuals who find inequality too high in their society have a higher likelihood of union membership. This effect is more than twice as large as the effect of inequality tolerance. However, those who think that inequality exists because people do not join together have no higher chance of membership than people who disagree with this statement.

The next question is how attitudinal factors modify the effects of other variables on union membership. The coefficients of model 1 are very similar to those of model 0. The gender difference in membership cannot be explained by gender differences in inequality attitudes (coefficient not shown). This is unsurprising, as women are generally more critical towards the existing earnings distribution than men (van de Werfhorst and de Graaf 2004). Neither the effect of schooling nor the difference between public and private sector workers can be explained by differentials in inequality attitudes. Thus, attitudinal factors give us no additional *interpretation* of some well-known variations in union membership. The exception is supervisory status. Its effect becomes smaller and drops below significance once attitudes are controlled for. Thus, the lower propensity of supervisors to membership may well be related to their lesser concern about inequality and greater tolerance for it.

With regard to the role of relative earnings, we see that the association between union membership and the distance of one's earnings above the median is only partly related to differences in inequality attitudes. The coefficient of this variable decreases from -0.101 to -0.077 . Thus, only to a limited extent is it true that high-earners tend to distance themselves from the unions because of their greater tolerance of and lesser concern about inequality, as expressed in these attitudes. Most of the main effect remains, which supports the view that their non-membership is more likely based on instrumental rather than explicitly normative considerations.

TABLE 4
 Probit Regression of Union Membership Including Attitudinal Variables (Marginal Effects)
 — ISSP 1987, 1992, 1999 — All Countries

	<i>Model 0</i>	<i>Model 1</i>	<i>Model 2</i>	<i>Model 3</i>
Relative distance from the median earnings (above)	-0.101*** [0.024]	-0.0772*** [0.024]	-0.0560** [0.027]	-0.0785*** [0.024]
Relative distance from the median earnings (below)	-0.237*** [0.045]	-0.244*** [0.046]	-0.225*** [0.047]	-0.246*** [0.046]
Inequality tolerance		-0.0207** [0.008]	-0.0218* [0.011]	-0.0580*** [0.017]
General inequality attitude		0.0421*** [0.009]	0.0391*** [0.013]	0.0278 [0.017]
Collective action inequality attitude		0.0134 [0.008]	0.00894 [0.012]	0.0115 [0.015]
Distance earnings (below) × tolerance for inequality attitude			0.0366 [0.046]	
Distance earnings (above) × tolerance for inequality attitude			-0.0126 [0.025]	
Distance earnings (below) × general inequality attitude			-0.0199 [0.049]	
Distance earnings (above) × general inequality attitude			0.0157 [0.023]	
Distance earnings (below) × collective action inequality attitude			0.00606 [0.043]	
Distance earnings (above) × collective action inequality attitude			0.017 [0.024]	
1992 × tolerance for inequality attitude				0.0475** [0.023]
1999 × tolerance for inequality attitude				0.0593** [0.023]
1992 × general inequality attitude				0.00705 [0.022]
1999 × general inequality attitude				0.038 [0.025]
1992 × collective action inequality attitude				0.0141 [0.020]
1999 × collective action inequality attitude				-0.0209 [0.023]
Year 1992	-0.022 [0.025]	-0.0379 [0.026]	-0.0379 [0.026]	-0.0336 [0.026]
Year 1999	-0.162*** [0.027]	-0.167*** [0.027]	-0.168*** [0.027]	-0.169*** [0.027]
Observations	4455	4455	4455	4455
Pseudo R^2	0.167	0.174	0.174	0.175
Log likelihood	-2566	-2545	-2542	-2539

Robust standard errors in brackets — * significant at 10%; ** significant at 5%; *** significant at 1% — country dummies included. Controls for gender, age, age², marital status, part-time, public employee, supervisor, educational attainment (three dummies) and gender interacted with university degree included but not reported.

Missing values are replaced with sample averages in order not to reduce sample size. ISSP, International Social Survey Programme.

In model 2, we see that there is no statistically significant interaction between earnings position and inequality attitudes. Thus, the association between inequality attitudes and union membership is independent of one's own relative earnings position. Model 3 shows that the effect of the tolerance of inequality attitude is only significant in 1987 (the excluded case) and reduced to about zero in later years. The coefficient of the general inequality attitude seems to have slightly increased between 1987 and 1999, with a positive albeit insignificant interaction effect of 0.038. Interestingly, according to our analysis, actual preferences to levelling earnings inequalities appear to have been a reason for union membership only in the 1980s; in later years, such preferences no longer seem to have any effect. Yet the general view that inequality is too high has had a more stable effect over the years. A possible explanation for these findings is that since the 1990s, people no longer regard trade unions as organizations that can successfully manipulate the earnings distribution across society. Still, a more general uneasiness about inequality may move workers towards union and covary with union membership.¹⁷

5. Conclusions

In this article we have studied the connection between union membership and inequality. First, we examined the relation between union membership and one's place in the earnings distribution, finding that trade unions are mostly representing workers in the middle. One implication of this finding is that, if earnings inequality increases, union membership tends to decrease as a result, because more and more workers find themselves further away from the median and perceive union action in this area as ineffective or contrary to their interests. This fits Streeck's recent observation on unions and labour markets, that 'the numbers are rising of those who have enough market power to do without collective organization, as well as of those who have too little market power to be capable of it' (Streeck 2005: 278). We do not find support for the hypothesis, however, that the association between relative earnings position and union membership itself changed during the period under study.

Our second line of investigation is concerned with the association between inequality attitudes and union membership. We show that union membership is also affected by normative concerns. People are partly motivated to become trade union members because they feel that economic inequality in society is too high and the distance between perceived actual and legitimate earnings should be reduced. We also show that well-known variations in union membership (e.g. by gender, skill level, supervisory status or sector of employment) are not explained by differences in attitudes across these subgroups. Consequently, by controlling for people's attitudes, we can be more confident that the association between relative earnings and union membership is not itself the result of different attitudes.

We can go back to the literature on union membership decline with additional insight. Our results suggest that workers are in the union either because they hold egalitarian opinions and/or they are not excessively anxious about the effects of egalitarian policies on themselves, as they are not that far from the median earnings position that such policies seem to reflect and benefit most. People belonging to the top tail of earnings distribution may self-interestedly 'vote with their feet' by leaving the union or they never joined a union in the first place. They may see the 'insurance' offer by the unions against unforeseen mishaps in the future as too costly or irrelevant for them or they may accept a higher risk for themselves together with a higher level of inequality in the world around them. Conversely, lower-paid workers may leave the unions, or never become members, when the union's policies defending the bottom tier in the labour market and closing the gap in relative earnings with the middle are perceived as ineffective.

Taken at face value, our results imply that any exogenously given increase in earnings inequality, for instance, as the consequence of skill-biased technological change or of increased globalization, has negative implications for trade unions. This is not to say that the observed increase in earnings inequality *has caused* the decline observed in union membership. As suggested in the introduction, the connection that we investigate in this article may be only one element in a chain. Union decline, along with changing technology and institutions, for instance, decentralization of collective bargaining, may be among the causes of increased earnings dispersion (Card *et al.* 2003; Hibbs and Locking 1996; Iversen 1999; Mosher 2002; Visser and Checchi 2009). Our article directs the attention to the *self-reinforcing* properties of this process, of increasing wage inequality weakening the incentives for union membership and making it more difficult for unions to create bridging coalitions that are attractive for both tails in the earnings distribution. Whether this is part of a more general process of increased heterogeneity, which erodes the basis for membership-based collective action organizations (like trade unions), is a matter we cannot decide on the basis of our findings, but it would certainly be an important issue for future research on unions and industrial relations.

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Notes

1. In the US context, workers who vote in union recognition elections would have to compare the expected pay-off of taking a union job (with higher wage but lower probability of continuation) and the expected pay-off of taking a non-union job. See Farber and Saks (1980).
2. In most of the cases, the interpretation of these proxies is observationally equivalent to many other models, the most traditional ones being framed in a demand and supply framework. The demand for union services or protection is a function of relative benefits deriving from membership, the costs of joining and social attitudes favouring union services over alternative provisions. On the supply side, one has the value or price of union services and the costs of providing them. Because of data deficiencies and difficulties in modelling union goals and behaviour, the estimation of structural demand and supply functions is usually replaced by a union status function in a reduced-form equation. See, for example, Blanchflower (2006), Hirsch and Addison (1986), and Schnabel (2003).
3. In the North American literature, the standard assumption is that low-paid workers should express a greater desire for union membership in union recognition elections, since there is more for them to gain (Farber and Saks 1980). Our empirical evidence, based on some European countries, contrasts with this assumption.
4. This effect should show in lower union retention rates among low-earning workers. The few studies on retention rates known to the authors (Klandermans and Visser 1995) do actually confirm this hypothesis, but also suggest additional causes such as more frequent loss of jobs, longer spells of unemployment and so on.
5. In an international and longitudinal comparison, institutional characteristics like a change in the law or differences in the conditions for union organizing or collective bargaining have been considered as well (Ebbinghaus and Visser 1999; Freeman and Pelletier 1990; Western 1997).
6. For an empirical analysis along that line, see Blanden and Machin (2003) on British data, Goerke and Pannenberg (2004) on German data and Visser (2002) on Dutch data.
7. We have decided to treat East and West Germany as two separate cases, although they have been politically unified since 1990 and the same industrial relations institutions apply, producing a number of convergent trends (Schnabel and Wagner 2003). However, with regard to concerns about actual and legitimate inequality, we expect differences as part of the legacy of communist labour relations (see Blanchflower and Freeman 1997).
8. The three ISSP surveys that include these questions vary with regard to the occupations that are listed. In 1987 and 1992, these questions were asked about 11 (partly different) occupations, in 1999, about 10 occupations. We chose to restrict our measurement to those five occupations that were included in all three surveys. However, if we use all occupations for each survey, allowing variation across surveys in the occupations that are judged in terms of perceived and legitimate

pay, the results of our regression analyses are very similar to the ones we present here (available from the authors). This is exemplified in high correlations between the employed measure and the measure that uses all available occupations, with a value of $r = 0.98$ in 1987, 0.79 in 1992, 0.87 in 1999, and across all three surveys 0.88. Given that we are interested in changing effects across time of inequality tolerance on union membership, we chose to employ a measure that is based on the same set of occupations in each survey year.

9. We have normalized the distribution for two reasons. First, it makes it easier to compare the effects of the different attitude measures, and second, because we display marginal effects that can now be easily interpreted as the effect for one standard deviation change. Standardization within country/year combinations largely yielded the same results.
10. In order to keep the maximum sample size, when missing values were encountered for some variable (typically the case of 'supervising someone' and 'beyond secondary/uncompleted college'), we have replaced them with country averages. Results are unaffected by this change, but standard errors for other variables are smaller.
11. In order to neutralize the influence of inflation, earnings have been regressed on year dummies, and what are reported are the residuals from those regressions.
12. Since we control for age and age squared (confirming the existence of an inverted U-shaped profile), the earnings effect is not a life cycle effect but must be explained by other factors.
13. In fact, it has been difficult to find a significant union wage premium in continental Europe (Teulings and Hartog 1998) and nearly all empirical research on union relative wage effects is limited to the USA, Canada and the UK (Blanchflower and Bryson 2003; Booth 1995). Yet, according to our dataset, union members tend to earn more, as can be detected by running an OLS regression of (log of) earnings on union status and standard covariates. This may be due to either self-selection into union membership due to unobservables (which should be different from the available covariates in our estimations) or unions offering benefits to their members, like access to training, promotion opportunities, additional employment protection or longer tenure, all of which are factors that are associated with higher earnings.
14. Although when using family size as instrument, support is only found for the relative income position below the median.
15. Our results do not exclude the existence of a reverse causal link, with the decline in union bargaining power having reduced their capability to compress earnings, thus shifting workers' earnings away from the median. However, this alternative story should be tested on groups of workers, for whom it is possible to measure their membership and associated bargaining power. If one would possess a panel of matched employer-employees data (a repeated version of the data used by Farber and Saks (1980), for example), one could, in principle, distinguish between firms that are exposed to exogenous increases in inequality (by sectors, looking at their exposure to international trade and/or to technological competition) and check whether workers in 'treated' firms are less likely to be union members. Similarly, one may think of geographical/temporal variations in the possibility of union membership (due to legislation changes), and check whether the earnings distribution in the 'treated' firms would be more/less unequal than in the 'untreated' ones. Unfortunately, this type of data is not available for most of the countries in our sample, and we are forced to confine ourselves to the analysis of

- a representative sample of individual data, where only relative wage can be correlated to personal union status.
16. This contrasts with Svallfors' (1997) findings, but it should be noted that his measure of the range of legitimate earnings did not take into account perceived *actual* pay levels. So, although Svallfors' analysis showed that the Norwegian population would prefer a lower dispersion of earnings across different occupations than respondents in Germany, Australia and the USA, this does not translate in a stronger preference for levelling if we take account of the perceived actual pay of occupations in Norway. This can obviously only be the case if the (perceived) actual earnings dispersion in Norway is smaller than elsewhere, which according to our data is indeed the case.
 17. As a robustness check, we have replicated the analysis of Table 4 by country (available from the authors). This analysis reveals that the effect of the inequality tolerance attitude, in 1987, was significant only in Britain and the Netherlands. The association of the general inequality attitude with union membership is more widespread, although no significant effects were found in the Netherlands, East Germany and Great Britain. Interaction models, furthermore, show that there is no systematic variation between earnings groups in the association between attitudes and union membership, which confirms the results of the pooled analysis.

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