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Laszlo Goerke, Yue Huang

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Abstract

Using panel data from 1985 to 2019, we provide the first comprehensive investigation of the relationship between trade union membership and job satisfaction in Germany. Cross-sectional analyses reveal a negative correlation, while fixed effects estimates indicate an insignificant relationship. This is also true if we incorporate information on collective bargaining coverage or the existence of works councils in subsamples for which this data is available. To address the endogeneity of union membership, we generate information on the union density individuals faced in their industry and region. This time-variant IV suggests no causal impact of individual union membership on job satisfaction. Finally, using different estimation models, we investigate whether the effects vary by gender, age, birth year, and employment status.

JEL-Codes: I310, J280, J510.

Keywords: exit-voice framework, German socio-economic panel, instrumental variable, job satisfaction, sorting, trade union membership.

Laszlo Goerke
IAAEU (Institute for Labour Law and Industrial Relations in the European Union
Trier University, Campus II
Germany – 54286 Trier
goerke@iaaeu.de

Yue Huang
IAAEU (Institute for Labour Law and Industrial Relations in the European Union
Trier University, Campus II
Germany – 54286 Trier
huang@iaaeu.de

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1 Introduction

Job satisfaction is often viewed as an indicator of the utility associated with work. Combining this interpretation with the assumption that individuals become trade union members if the associated benefits exceed the resulting costs, we can expect trade union members to exhibit higher job satisfaction than individuals who choose not to become union members. Early empirical studies, such as by Freeman (1978) and Borjas (1979), do not observe the predicted positive correlation and have sparked an intense debate, generating a multitude of investigations. Most subsequent cross-sectional analyses report a negative relationship. If employing panel data or accounting for the endogeneity of union membership, relatively little evidence for a negative impact of union membership remains. However, the theoretical expectation of a positive correlation rarely finds empirical support.

The vast majority of studies, for example, considered in the meta-analysis by Laroche (2016) and the surveys by Hammer and Avgar (2005), Artz and Heywood (2021), and Goerke (2021), are based on data from the UK and the US. Given the specificity of their industrial relations system, we cannot easily generalise the findings. Meanwhile, the relationship between union membership and job satisfaction for other economies has not been well documented. Moreover, analyses that can truly identify causal effects are rare.

In this paper, we fill part of the research lacunae implicitly described above. As our first contribution, we comprehensively analyse the linkage between trade union membership and job satisfaction for Germany, using data from the German Socio-Economic Panel from 1985 to 2019. Germany is an interesting subject of research because trade unions play an important role. They have almost eight million members. Net union density is around 16% and has declined substantially over the last decades (OECD and AIAS, 2021). Moreover, the linkage between collective bargaining coverage and union membership is much weaker in Germany than in, for example, the US, indicating different incentives to become a trade union member. In addition, co-determination at the firm level via works councils plays a vital role. Therefore, we provide additional insights on the function of bargaining coverage and co-determination for the relationship between union membership and job satisfaction. As our second contribution, we present an innovative approach using a time-variant instrumental variable (IV) to account for the possible endogeneity of union membership. In particular, we rely on the concurrent strength of unions at the sectoral and regional level. Third, we conduct a comprehensive subgroup analysis and investigate whether the relationship between union membership and job satisfaction varies by gender, age, birth year, and employment status.

We document a negative correlation between trade union membership and job satisfaction in Germany in Ordinary Least Squares (OLS) models. We do not discern a statistically significant relationship in Fixed Effects (FE) specifications, unless we separate between individuals who join the trade union and those who leave it. Adding information on collective bargaining coverage or co-determination at the workplace for a subset of years clarifies that the above-mentioned correlations are not determined by such labour market institutions. Specifications based on time-variant exogenous variations in a trade union membership suggest the absence of any causal effect of union membership on job satisfaction. Finally, we find stronger correlations in OLS models for female, older, and part-time employees. However, estimates from FE and IV models are mostly statistically insignificant. Therefore, subgroup analyses yield qualitatively similar results as for the whole sample.

The further paper proceeds as follows: Section 2 describes the institutional background, clarifies what effects of union membership on job satisfaction may be expected in Germany, and summarises the extant empirical evidence. In Section 3, we introduce the data and describe the empirical model. Section 4 presents the findings. Section 5 concludes the paper. The appendix provides additional information and documents the results from a number of robustness checks, which we briefly report in the main text.

2 Background

2.1 Institutional Setting

In Germany, trade union membership is not tied to a job or employment, but results from an individual's decision. A member is entitled to financial support, for example, in case of a strike, to legal advice and support, inter alia, if there is a dispute with the employer, to extensive informational offers, and to a host of financial advantages, such as the possibility to obtain some insurances at reduced prices. The membership fee usually amounts to one percent of the wages.

Union membership is distinct from coverage by a collective bargaining contract. Such a contract determines a lower bound for wages and often regulates a variety of working conditions. Further, it may include agreements on fringe benefits in excess of legal minimum levels. Currently, about 40% of the private sector and most public sector employees work in establishments that are covered by sectoral collective bargaining agreements. About 8% of employees are paid according to collective agreements negotiated at the firm level (Ellguth and Kohaut, 2021). The resulting bargaining coverage substantially exceeds union density of about 16% (OECD and AIAS, 2021). Collective agreements bind the firms covered by it and trade union members. Since trade unions cannot oblige employees to become members, the share of union members among all employees is usually well below 100% even in firms covered by collective bargaining. Moreover, employers of firms covered by collective agreements generally apply its content to all employees, irrespective of their membership status. These features may explain why there is no evidence of a union membership wage premium in Germany, once observable characteristics are accounted for (Schmidt and Zimmermann, 1991; Fitzenberger et al., 1999; Goerke and Pannenberg, 2004).[1]

While collective bargaining is often argued to constitute the first important pillar of industrial relations in Germany, co-determination at firm level via works councils is then regarded as the second decisive element. Works councils can only be established in private-sector firms with at least 5 employees and require a vote by the workforce about their members. Since such a vote is not compulsory, councils exist in less than 10% of all eligible establishments. As this is the case mostly in large firms, about 40% of all private-sector employees are employed in firms in which there is a works council. Collective bargaining and works councils often co-exist. However, there is a substantial fraction of firms in which only one of the two pillars of industrial relations can be observed (Ellguth and Kohaut, 2021).

In sum, it is of great essence to discover how trade unions work in Germany and affect employees' well-being. Given the specificities of the German industrial relations system, the insights on the linkage between trade union membership and job satisfaction obtained for Anglo-Saxon countries cannot be simply generalised. Moreover, as union membership is not decided upon at the firm level, alternative empirical approaches than for, e.g., the US are required to cater for the endogeneity of

[1] In a recent paper, Hirsch et al. (2022) show that slightly less than 10% of employees working in firms that are covered by a collective contract are not paid according to it. This is particularly true for men, managers and employees at both ends of the wage distribution. The data does not enable Hirsch et al. (2022) to investigate a correlation with union membership.

union status.

2.2 Expectations

From a rational choice perspective, an individual can be expected to become a member of a trade union if the expected private benefits exceed the respective costs. If, additionally, the net benefits of membership are associated with the job and not, for example, located in the private domain, the job-related utility of union members should be higher than that of comparable non-members. Since trade unions also provide benefits outside the job, and job-related utility can spill over to overall utility, union members can also be conjectured to exhibit a higher level of overall utility. While it is difficult, if not impossible, to measure utility directly, we often have information about the satisfaction individuals express with their life. Assuming a monotone relationship between utility and satisfaction, the latter may represent a suitable empirical proxy for the former (Hamermesh, 1977, 2001; Clark and Oswald, 1996; Frey and Stutzer, 2002). If this line of arguments extends to utility from the job, trade union members can be expected to exhibit higher job satisfaction than otherwise equivalent non-members. A priori, this line of reasoning is applicable to all societies in which union membership results from an act of individual choice and, therefore, to Germany as well.

The expectation of a positive association between union membership and job satisfaction has been refined in light of the exit-voice framework going back to Hirschman (1970, 1974) and applied to trade unions by Freeman and Medoff (1984). It posits that employees can respond to unsatisfactory working conditions by either voicing discontent and attempting to improve them or, alternatively, by leaving the employer. Since better working conditions constitute a public good and unions 'collect' their members' voices, the exit-voice approach predicts that union members have greater incentives to improve their working conditions than non-members do. The above reasoning assumes that trade unions offer a bundle of goods that employees can purchase by becoming members. However, the extent of improvements in working conditions depends substantially on the strength of employees' voices. If a trade union emphasises adverse working conditions and "manufactures discontent" (Hammer and Avgar, 2005, p. 243; see also Borjas, 1979; Heywood et al., 2002; Freeman and Medoff, 1984, pp. 139f), workers' voices are more likely to be heard. Hence, the exit-voice perspective can be consistent with a greater degree of stated job dissatisfaction among union members than non-members because lower satisfaction is an input for the production of voice activities. In Germany, the exit-voice-framework may be less applicable with regard to trade unions than in Anglo-Saxon countries because works councils can act on behalf of employees and help to improve working conditions. However, works councils are by no means universal. Therefore, this institution does not preclude the existence of an exit-voice effect on job satisfaction via trade union membership.

While there does not appear to be a membership wage premium in Germany, there is some evidence that trade union members obtain higher fringe benefits and are better protected against job loss (Goerke and Pannenberg, 2011; Goerke et al., 2015). Moreover, unions emphasise the provision of free legal advice and support in case of job-related conflict to members. This suggests that individual employees can benefit from union membership. In sum, we expect that trade union members on average exhibit higher job satisfaction than comparable non-members in Germany.

2.3 Previous Findings

In contrast to the theoretical expectation of a positive relationship, empirically "(t)here is a well-established negative correlation between union membership and job satisfaction" (Bryson and White, 2016a, p. 898). The findings referred to in the above quote are often based on cross-sectional estimates, employing data from Anglo-American countries. While the British Household Panel Survey (BHPS; and its successor) allows for a differentiation between the impact of collective bargaining and an individual's union membership (see, for example, Bryson et al., 2010; Bryson and White, 2016a; Green and Heywood, 2015), analyses for the US usually do not make this distinction (see Heywood and Wei, 2006; Artz, 2012; Artz et al., 2021).

Empirical investigations for other countries generate a somewhat different picture. While positive correlations are rare (see Donegani and McKay, 2012; van der Meer, 2019; Blanchflower et al., 2022), more often insignificant effects are observed (García-Serrano, 2009, 2011; Hipp and Givan, 2015; Hauret and Williams, 2017; Lightman and Kevins, 2019). Furthermore, a possible change in the direction of the correlation from negative to positive, as Artz et al. (2021) and Blanchflower et al. (2022) find in the last two decades and for later birth cohorts in the US and the UK, has not been diagnosed for other countries.[2]

[2] A partial exception is Blanchflower et al. (2022) who use three waves of the European Social Surveys and report a positive correlation for a sample of 38 countries. However, they do not provide country-specific findings.

It is noteworthy that cross-sectional analyses for Germany either represent a by-product of an investigation with a different focus or only provide indicative evidence. Jirjahn and Tsertsvadze (2006) consider the correlation between job satisfaction and works councils, using SOEP data for 2001. They also include information on union membership and observe a positive relationship with job satisfaction for the entire sample and some subgroups. Moreover, Hipp and Givan (2015) use the 2005 wave of the International Social Survey Program (ISSP). There is no discernible difference in overall job satisfaction between union members and non-members in an OLS regression.

Most cross-sectional analyses face the problem that trade unions cannot regulate the entire set of factors, which affect job satisfaction. Therefore, employers may respond to, for example, wage increases by reducing other components of labour costs. If such compensatory responses affect working conditions negatively, they are likely to reduce job satisfaction. Consequently, empirical analyses, which omit control variables capturing such adverse compensation effects, may indicate a negative correlation (Hersch and Stone, 1990; Renaud, 2002; García-Serrano, 2009). In Germany, this reasoning is unlikely to be relevant because there is no evidence of a union membership wage premium and wages and working conditions tend to be the same for all workers in a specific firm, as establishments covered by a collective bargaining agreement generally apply it to the vast majority of staff. If compensatory responses occur, they are therefore unlikely to be restricted to union members. In our empirical analysis, we can evaluate the relevance of this argument using a subsample of observations for which information about collective bargaining coverage is available.

A further explanation why the linkage between union membership and job satisfaction may be inadequately identified in cross-sectional analyses is that individuals with certain characteristics, which affect stated levels of job satisfaction, are more likely to be a union member than others who do not have these characteristics. Accordingly, in some studies for the UK and the US, which take time-invariant individual characteristics into account, the negative correlations reported in cross-sectional estimations cannot be observed anymore in FE specifications (see, inter alia, Pouliakas, 2010; Artz, 2010; Kosteas, 2011).[3] However, various analyses report negative correlations between union membership and job satisfaction also in FE models (see Heywood et al. (2002), Pouliakas and Theodossiou (2009), Green and Heywood (2015) for members covered by collective bargaining

[3]In line with these findings, Artz (2012) documents that an insignificant cross-sectional estimate of the union dummy for males turns positive in a FE model. However, Heywood and Wei (2006) report the reverse.

agreements, and Artz (2012) for women).

Since identification in FE models relies on individuals who change their union membership status, the results can only be generalised if those who join or leave a trade union are representative of the population. However, switches in union status are endogenous, so that such representativeness is at least questionable (Bryson and White, 2016b). We will take up this issue and investigate whether individuals who change their union status are comparable to those who do not.

To cater for the possibility of reverse causality, sorting in accordance with time-varying personal characteristics, and also the endogeneity of entry into and exit out of a trade union, instrumental variable approaches have been pursued. Previous studies use a variety of time-invariant instruments, such as geographical and industry information (Borjas, 1979), attitudes towards unions in combination with industry and occupational dummies (Bender and Sloane, 1998), assessments of the industrial relations climate (Bryson et al., 2004; Laroche, 2017), state-level right-to-work laws in the US (Gius, 2013), and information on establishment age and single-plant firms (Bryson et al., 2010).[4] The estimated coefficients for the union membership dummy are often not significantly different from zero (Bender and Sloane, 1998; Bryson et al., 2004; Laroche, 2017), suggesting that the estimated correlation may be far from causal.[5] In this paper, we can utilise a time-variant instrumental variable, which allows us to exploit the panel nature of the data also in the analysis of a possible causal relationship.

In sum: Previous contributions provide a mixed picture of the relationship between trade union membership and job satisfaction. The sign of the correlation appears to vary across countries and industrial relations systems, possibly over time, as well as between those who switch their union status and those who do not. Causal analyses do not yield a clear-cut image, either. As argued above, an analysis of Germany can resolve some of the ambiguities because of the features of its industrial relations system and the exceptional quality of our data, which we will explain next.

[4] Miller (1990) and Pfeffer and Davis-Blake (1990) also report findings from specifications with IVs, though without precisely defining the instruments employed.

[5] Borjas (1979) and Bryson et al. (2010) report negative estimates, whereas Pfeffer and Davis-Blake (1990) find positive effects also in an IV estimation.

3 Data and Methodology

For the empirical analysis, we use the German Socio-Economic Panel (SOEP, version 36), a high-quality and representative longitudinal dataset of private households in Germany (see Goebel et al. (2019) for more information). We consider employed individuals aged 16 to 65 and exclude self-employed respondents. The overall observation period ranges from 1985 to 2019. We focus on those SOEP waves, which contain a question on union membership.

3.1 Definition of Variables

The outcome variable is based on the following question: "*How satisfied are you today with your job?*" Feasible answers range from '0' (*completely dissatisfied*) to '10' (*completely satisfied*). The main explanatory variable takes the value of one if the respondent is a union member, and zero otherwise. The respective question was asked in the years 1985 and 1989 in West Germany, 1990 in East Germany, and 1993, 1998, 2001, 2003, 2007, 2011, 2015 and 2019 to respondents from both parts of the then united country.

Being a trade union member could be correlated with socio-demographic and labour market characteristics. Individuals with specific traits or features may also show different levels of satisfaction with the job. The first group of control variables to account for such effects includes demographic characteristics, such as age (level and quadratic term), gender, being married, years of schooling (level and quadratic term), and the number of children in the household. Second, we consider two sets of labour market characteristics. The first set of variables is relatively exogenous and includes being a blue-collar worker, being a civil servant, working in a public sector, four firm-size dummy variables, and a group of dummy variables for the 16 industrial sectors (NACE, see Table B-1 in the appendix). The second set of controls includes the logarithm of the monthly wage, the logarithm of actual working hours, years of tenure (level and quadratic term), and a dummy variable for having a new job (tenure of less than one year). We finally incorporate dummy variables for the survey years and the federal states of residence.

The four covariates of the second group of labour market variables may be affected by the consequences of union membership and also have an impact on job satisfaction, i.e., they may constitute bad controls. For instance, changing jobs is positively correlated with job satisfaction

(Chadi and Hetschko, 2021). Union members are less likely to change jobs and have, on average, longer tenure (see Table 1). In this case, having a new job could be interpreted as a bad control. However, it is also possible that employees decide to join the union after they move to a new job, facing new working conditions, indicating that having a new job is a relevant covariate. Similar lines of reasoning could apply to the wage, despite the lack of evidence concerning a union membership wage premium in Germany, working hours, and tenure. To ascertain the relevance of such concerns, we show results of specifications with and without the second set of labour market characteristics in Table 2 and Table 4.

Excluding respondents with missing values for the outcome variable or covariates, the final estimation sample consists of 81,169 observations from 36,422 individuals. In this sample, 27,230 individuals (54,761 observations) stated never to have been a union member, while 5,483 individuals (11,010 observations) were always union members. Moreover, 3,709 individuals (15,398 observations) reported changes in the membership status. In particular, 1,126 individuals joined, 1,757 left, and 826 reported both an entry and exit.

3.2 Empirical Strategy

Since the outcome is an ordered variable, the obvious empirical approaches to estimate the correlation between job satisfaction and union membership are an ordered probit model, possibly combined with individual random effects and Mundlak-type corrections, or an ordered logit model with or without taking into account individual fixed effects. However, for a comparable question in the SOEP about life satisfaction, Ferrer-i-Carbonell and Frijters (2004) show that OLS models yield results that are equivalent to those for non-linear models. This virtual equivalence is also exploited in many analyses of job satisfaction. Since linear models are more straightforward to interpret, we subsequently focus on them and demonstrate for our main specifications that results are basically unaffected.

3.2.1 Correlation Analysis

We initially report OLS estimates. In order to control for time-invariant observable and unobservable factors that might be correlated with union membership and job satisfaction, we also employ

FE regressions. We estimate the following model:

$$JS_{it} = \beta_0 + \beta_1 TU_{it} + \boldsymbol{X'_{it}}\ \boldsymbol{\beta_2} + \lambda_i + \lambda_t + \varepsilon_{it}, \tag{1}$$

where JS_{it} is the job satisfaction of individual i interviewed in year t and TU_{it} is a dummy variable that indicates trade union membership status. $\boldsymbol{X_{it}}$ is a vector of individual-level socio-demographic and labour market characteristics, as discussed in Section 3.1. λ_i indicates individual fixed effects, while λ_t represents a set of year dummy variables. ε_{it} is the error term. We cluster standard errors at the individual level, unless stated otherwise.

Measures of job satisfaction are, by construction, ordinal in nature, and the mean ranking of ordinal variables is difficult to identify (Oparina and Srisuma, 2022). Bond and Lang (2019) argue that the ranking of the means can be identified only if the distribution of well-being states for one group of individuals first-order stochastically dominates that of the other. However, the actual distribution of states is not observable. This concern can be mitigated if we apply FE models because an individual's satisfaction distribution should be approximately identical over time and uncorrelated with their union membership status.

3.2.2 Causal Analysis

Joining a trade union is not random. The possible existence of reverse causality and omitted variables may bias the estimated coefficient on trade union membership in the correlation analysis. Therefore, in this section, we propose a strategy for estimating the causal impact of union membership.

In the US, for example, collective bargaining and union membership are often closely aligned, in contrast to Germany. Comparing certification elections with close outcomes, it can then be argued that union membership is exogenous from the perspective of an individual employee because this person can hardly affect the election outcome.[6] In Germany, a person can decide to become a member, or abstain from joining a trade union individually. Consequently, being a union member is not exogenous and can be correlated with a respondent's unobservable characteristics. Such

[6] See DiNardo and Lee (2004) for the basic idea and an application to the analysis of collective bargaining on firm outcomes. The approach has been employed to investigate a variety of questions, such as by Sojourner et al. (2015) or Knepper (2020). Frandsen (2021) indicates that the exogeneity assumption may not be warranted, since election outcomes can be disputed.

factors may also be relevant for their job satisfaction. Moreover, it is possible that deteriorating work conditions lower job satisfaction, providing employees with an incentive to join a union, i.e., reverse causality can arise. Since there is no such institution as certification elections, we cannot apply a regression discontinuity design. Instead, we employ a time-variant instrumental variable (IV) approach and estimate Two-Stage Least Squares (2SLS) and IV Fixed Effects (IVFE) models to determine the causal effect of trade union membership on job satisfaction.

We hypothesise that an individual is more likely to join a union if more employees in the same industry, region, and year are members, i.e., the higher the industry-specific, regional, contemporaneous union density is.[7] Therefore, our identification strategy exploits the exogenous variation in the individual-level union membership, induced by the change in the union density at the aggregate level. Potential individual-specific confounders can be accounted for by person fixed effects. More specifically, we calculate the share of *other* employees who are union members in the same region in which the individual lives and in the same industry for each wave containing information of union membership. We apply a broad definition of sectors to ensure a sufficient number of observations in a region-industry-year unit.[8] Moreover, we focus on the German spatial planning region (ROR, *Raumordnungsregion*). On average, each such region has a population of somewhat less than one million inhabitants, of which about one hundred are SOEP interviewees.[9]

We estimate the following regressions with the 2SLS and IVFE models, respectively.

$$JS_{it} = \beta_0 + \beta_1 \widehat{TU}_{it} + \boldsymbol{X'_{it}}\ \boldsymbol{\beta_2} + \lambda_i + \lambda_t + \tau_{2it}, \tag{2}$$

$$TU_{it} = \alpha_0 + \alpha_1 Density_{it} + \boldsymbol{X'_{it}}\ \boldsymbol{\alpha_2} + \lambda_i + \lambda_t + \tau_{1it}. \tag{3}$$

$Density_{it}$ is the union density of other employees in year t computed for individual i. The first assumption for a valid IV requires that union density is associated with an individual's union membership status, as it is the case (see Table 4). The second assumption requires that the IV is exogenous. That is, the union density at an aggregate level is not correlated with time-variant

[7]See Hadziabdic and Baccaro (2020), who study the impact of union membership on individual political attitudes. Berton et al. (2021) employ the average NACE-two-digit-sector-level incidence of the presence of union representatives at the workplace to instrument the current likelihood of having a formal union representation.

[8]The indicator includes eight groups of industrial sectors (see Table B-1 in the appendix).

[9]There was a ROR reform in 1996 in which a few regions were combined to larger RORs. We re-code the region information for individuals surveyed before 1996 by applying the ROR classification after the reform.

uncontrolled variables at the individual level based on covariates. Moreover, trade union density does not directly impact on an individual's job satisfaction or influence it via other channels. While we obviously cannot prove the exogeneity assumption and our IV is not perfect from nature, we provide a set of robustness analyses to show that our IV is *arguably* exogenous.

First, there might be a concern that individuals' time-invariant characteristics, such as preferences for working in a certain sector, could be correlated with sector-specific union densities and decisions to join a union. Since the IV varies across individuals and years, we are able to control for person fixed effects. However, there might be other factors influencing the contemporary individual-level union membership and the aggregate-level union density. Therefore, the IV might be more exogenous if we employ the lagged union density of other individuals in the same region and industry as IV, which will be shown in one robustness check.

A second concern is that individuals may change their place of residence or industry during the observation period based on own preferences, seeking for higher or lower union densities. Such preferences may have a direct impact on job satisfaction. These employees may be a specific group. If preferences are time-invariant, they can be accounted for by person fixed effects. Since we cannot be sure that such invariance exists, as an additional robustness check, we investigate the impact of union membership on job satisfaction for the more homogeneous group of non-movers (i.e., individuals who changed neither their residence nor industry). For this sample, we are further able to cluster standard errors at the ROR-industry level, at which the IV varies. This method allows us to take into account the correlation between individuals working in the same region and industry.

Third, there might be a concern that union density influences job satisfaction independently of individual union membership, either through other channels or directly. For instance, a high trade union density may improve working conditions, therefore resulting in higher job satisfaction. Such situation may lead to an overestimated coefficient on union membership in the second stage. Furthermore, some social custom models of trade union membership propose that individuals are, ceteris paribus, more likely to join the union, the higher union density is. This effect comes about because the payoffs from conforming to the social custom rise, and a higher density can enhance collective bargaining power. Both effects imply that the utility level of union members rises.[10]

[10] See, for example, Naylor and Cripps (1993), Naylor and Raaum (1993), and Booth and Chatterji (1993). We are grateful to an anonymous referee for pointing out this possible channel by which union density can affect job satisfaction.

However, it is not obvious how the utility change translates into an alteration of job satisfaction because the social custom component relates to the non-job-related part of utility. In any of the above-mentioned situations, the estimated coefficient on union membership in the second stage may be biased due to the violation of the exclusion restriction. Since this restriction is not directly testable, we apply the strategy in Conley et al. (2012), allowing that our IV could impact job satisfaction partly through other channels, and show that our IV is *plausibly exogenous*.

Finally, another remaining bias could be that some unobservables in the first stage may be correlated with both union density and the individual decision to join a union. With a positive bias, the coefficient on the IV in the first stage will be overestimated and the coefficient on union membership in the second stage may be underestimated.

4 Results

In this section, we first present descriptive evidence. Second, we report the results from OLS and FE estimations. Third, we consider the impact of collective wage agreements and works councils. Afterwards, we introduce results from the causal analysis. Finally, we investigate whether the relationship between union membership and job satisfaction differs by gender, age, birth cohort and employment status.

4.1 Descriptive Evidence

Table 1 indicates the summary statistics of the main variables.[11] On average, 22.76% of individuals (observations) are trade union members. This number is moderately lower than the figure provided by the OECD and AIAS (2021) for a slightly earlier time period.

Figure 1 shows the development of trade union membership over time. It is evident from Panel (a) that the share of union members among all employees has decreased for males and females, though males show a slight increase in the share from 2015 to 2019. On average, 16.97% of female observations are union members, while this share amounts to 27.34% for male observations. Panel (b) clarifies that the share of females among all union members has increased substantially over the last decades (see also OECD and AIAS (2021)).

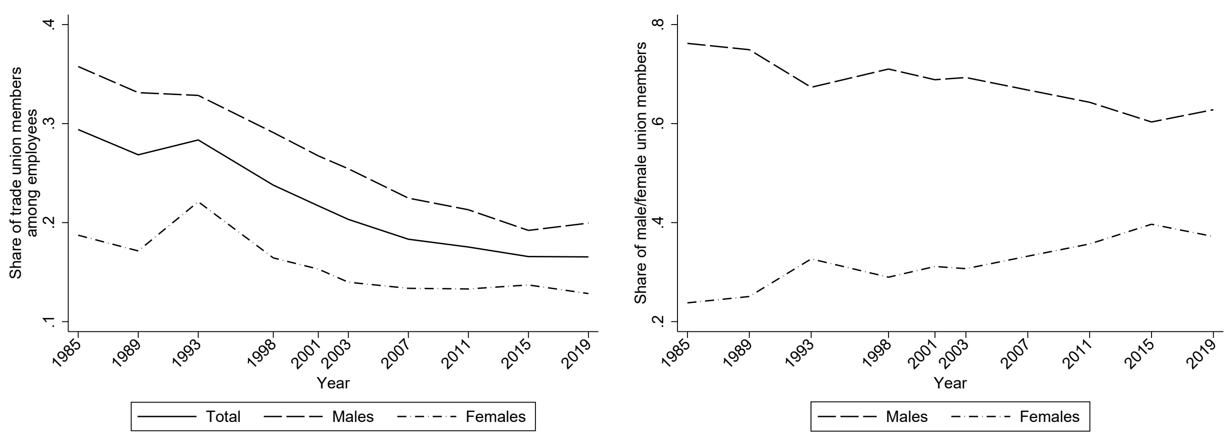
[11]The summary statistics of all variables in the baseline specification are shown in Table A-1 in the appendix.

Table 1: Summary statistics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	All observations		Union member		Non-union member		(3) − (5)	
	Mean	S. D.	Mean	S. D.	Mean	S. D.	Difference	p-value
job satisfaction	7.1067	2.0296	6.9947	2.1119	7.1397	2.0035	-0.1450	0.000
life satisfaction	7.2156	1.6553	7.0811	1.7369	7.2553	1.6283	-0.1741	0.000
TU member	0.2276	0.4193						
female	0.4420	0.4966	0.3296	0.4701	0.4751	0.4994	-0.1455	0.000
age	41.0022	11.7737	42.8717	11.2564	40.4514	11.8659	2.4203	0.000
tenure	10.7077	10.1118	14.6491	10.9412	9.5464	9.5488	5.1027	0.000
public sector	0.2793	0.4487	0.3612	0.4804	0.2552	0.4360	0.1061	0.000
civil servant	0.0742	0.2620	0.1136	0.3173	0.0625	0.2421	0.0511	0.000
blue-collar worker	0.3180	0.4657	0.4368	0.4960	0.2831	0.4505	0.1537	0.000
having a new job	0.1302	0.3366	0.0667	0.2496	0.1490	0.3560	-0.0822	0.003
full-time employment	0.7657	0.4236	0.8444	0.3625	0.7425	0.4373	0.1019	0.000
firm size								
< 20	0.2150	0.4108	0.0668	0.2497	0.2586	0.4379	-0.1918	0.000
$[20, 200)$	0.2828	0.4504	0.2309	0.4214	0.2981	0.4574	-0.0672	0.000
$[200, 2000)$	0.2358	0.4245	0.2844	0.4511	0.2215	0.4152	0.0629	0.000
≥ 2000	0.2664	0.4421	0.4179	0.4932	0.2218	0.4155	0.1961	0.000
Observations	81,169		17,656		63,513			

Notes: This table shows summary statistics of the estimation sample using weights. In columns (1)–(2), we show statistics for the whole sample, in columns (3)–(4) for trade union members, and in columns (5)–(6) for non-union members. Column (7) shows the difference in means between column (3) and (5), and columns (8) depicts the p-value. Individuals working more than 36 hours per week are defined as full-time employees.

Figure 1: Trade union membership



(a) Trade union density

(b) Share of male and female trade union members

Data source: SOEP, estimation sample, using weights, own calculation.
Notes: Panel (a) shows the share of trade union members for all individuals, males, and females, respectively. Panel (b) shows the share of male as well as female trade union members. Data for 1990, available for East Germany only, are not used.

Overall, respondents show a level of job satisfaction of 7.1. Union members report a significantly lower level (see columns (7)–(8) of Table 1). Moreover, the difference is much larger in magnitude for females than males (results not depicted). Table 1 also indicates that trade union members have higher tenure than non-members and are more likely to be blue-collar workers and to work full time, in large firms and in the public sector.

4.2 OLS and FE Results

Table 2 provides OLS estimates. In column (1), we depict the estimated coefficients for a specification in which individual characteristics, federal state FE, and year FE are accounted for. In column (2) we add information on relatively exogenous labour market characteristics. Subsequently, we include the second group of labour market factors, namely information on monthly wages, working hours, tenure, and acquiring a new job. We show the latter results in column (3), which is also our baseline specification.[12] All three columns depict a statistically significant and quantitatively comparable negative relationship between trade union membership and job satisfaction.

Incorporating the second group of labour market factors in column (3) does not change the estimate qualitatively in comparison to column (2), since the size of the coefficient in absolute terms decreases only marginally. This suggests that our findings hold even if bad controls would exist. The absolute value of the estimated coefficient on union membership depicted in column (3) is two times larger than the coefficient on the number of children and resembles the coefficient on education (see column (1) of Table A-2). Therefore, the job satisfaction differential associated with union membership is of economic relevance. Furthermore, its magnitude is in line, for example, with estimates for the UK (Bryson et al., 2004; Bryson and White, 2016a).

However, we have to keep in mind that the coefficient in OLS models could be biased. There might be some uncontrolled factors influencing both union membership and job satisfaction. To test the relevance of unobservables, we apply the strategy in Oster (2019) and show results in Table

[12]The full estimation results of the baseline specification are shown in Table A-2 in the appendix. Findings from ordered probit models are presented in Table A-3.

Table 2: Trade union membership and job satisfaction

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS			FE		
TU member	-0.1160***	-0.1218***	-0.1160***	-0.0405	-0.0473	-0.0198
	(0.0227)	(0.0233)	(0.0233)	(0.0339)	(0.0339)	(0.0337)
Observations	81,169					
Number of respondents	36,422					
Control variables						
State & year FE	X	X	X	X	X	X
Demographic characteristics	X	X	X	X	X	X
Labour market factors (1)		X	X		X	X
Labour market factors (2)			X			X

Notes: This table shows the relationship between trade union membership and job satisfaction. The observation period is 1985–2019. Demographic characteristics include age (level & quadratic term), being female, being married, years of schooling (level & quadratic term), and the number of children in the household. As the first set of labour market factors, we include being a blue-collar worker, working in the public sector, being a civil servant, and a set of dummy variables for the firm size and industrial sectors. We also control for the logarithm of the monthly wage, the logarithm of actual working hours, years of tenure (level & quadratic term), and having a new job, denoted as the second set of labour market factors. Federal state fixed effects and year fixed effects are accounted for in all specifications. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

A-4.[13] We conclude that the selection on the unobservables does not appear to be of great relevance for the OLS estimates.

Adding person fixed effects to the model can account for the individual-specific and time-invariant factors. Results from estimating FE specifications are presented in columns (4)–(6) of Table 2, again successively adding groups of control variables. The estimated coefficients are insignificant in all specifications. This systematic difference between OLS and FE results could have occurred because individuals with time-invariant unobservable characteristics that are negatively associated with job satisfaction may sort into union membership.

We can attempt to evaluate this hypothesis by investigating the linkage between union membership and life satisfaction. If, for example, more pessimistic individuals are more likely to become members of a union, we can expect to observe a negative correlation between membership and

[13] We depict in column (1) of Table A-4 the coefficient from the baseline model, i.e., column (3) of Table 2, if all covariates are included in the regression model. The identified set is shown in column (2). The lower bound is estimated if the proportional degree of selection on unobservables to the selection on observables is zero ($\delta = 0$), i.e., the coefficient depicted in column (1). The upper bound is estimated if we set the selection on unobservables equal to the selection on observables ($\delta = 1$), and the R-squared R_{max} is equal to the 1.3 times of the R^2 in column (1). The identified set suggests that the correlation between union membership and job satisfaction is negative. Furthermore, column (3) shows the value of δ which would produce the estimated coefficient on union membership equal to zero, given the value of $R_{max} = 1.3(R^2)$. The result suggests that the selection on the unobservables would have to be 8.27 times higher than the selection on the observables for the coefficient on union membership to become insignificant.

life satisfaction in OLS specifications as well, which vanishes or becomes less pronounced in FE estimates. This is indeed what we observe (see Table A-5 in the Appendix).[14]

Before concluding from the comparison of OLS and FE estimates that individuals with particular time-invariant features sort into union membership, another aspect deserves closer scrutiny. While all observations in the sample contribute to the identification of the correlation between union membership and job satisfaction in OLS estimates, in the FE model only individuals changing the membership status help to identify the relevant coefficient. However, membership-status switchers and non-switchers (i.e., never-members and always-members) may differ in personal characteristics. In consequence, the findings from the FE model may not be generalisable.[15]

The descriptive statistics suggest that switchers and non-switchers indeed differ from each other in various individual characteristics (see Table A-6 in the Appendix). Thus, we re-estimate the specifications in column (3) and (6) of Table 2 using an OLS model for non-switchers and OLS and FE specifications for switchers. Results presented in columns (1)–(3) of Table A-7 suggest that non-switchers show a significant and negative correlation, while for switchers we find significant correlations between job satisfaction and union membership neither in OLS nor in FE models. The insignificant results for individuals who switch their union status may have come about because those who join a trade union experience a change in job satisfaction, which lies in opposition to the variation for employees who leave the union. To scrutinise this conjecture, we differentiate the group of switchers further. In particular, we generate a dummy variable, *only entry*, which equals one if an individual only reported an entry in the union during the observation period, and zero otherwise. The OLS estimates indicate that the correlation between union membership and job satisfaction is negative and statistically significant for only-entry individuals, while we do not find such a pattern in the FE model (see columns (4)–(5) of Table A-7 in the Appendix). This finding is compatible with our above hypothesis that unsatisfied or pessimistic employees join, whereas we discern no such sorting effect for individuals who leave the trade union.

The findings from OLS and FE models will not be informative concerning the effect of union membership on job satisfaction if reverse causality exists, i.e., if individuals join a union because

[14]Note that our finding of a negative correlation between union membership and life satisfaction contrasts with results for other countries for which positive linkages have been reported (Radcliff, 2005; Keane et al., 2012; Flavin and Shufeldt, 2016; Artz et al., 2021; Blanchflower et al., 2022).

[15]We are very grateful to an anonymous referee for suggesting this fruitful line of enquiry.

they are unsatisfied with their job. In order to mitigate this concern, we regress job satisfaction in year $t+1$, i.e., the leading value of JS, on union membership in year t. Since we need at least two consecutive observations, the sample shrinks by about 20% without qualitative impact on the baseline specifications (as defined by columns (3) and (6) in Table 2). If we use job satisfaction from $t+1$ instead of t, the estimated coefficients neither in the OLS nor the FE model change qualitatively (the OLS results are even quantitatively virtually the same; see Table A-8 in the Appendix). Therefore, our main findings generally hold.

In sum, the OLS specification suggests a negative correlation between an individual's trade union membership and job satisfaction. The FE models and the differentiation between those who join the trade union and those who leave it indicate that this may be due to the sorting of individuals who decide to become a union member.

4.3 Institutions

So far, our analysis has neglected the role of labour market institutions, which are specific to Germany, namely collective bargaining agreements, which apply to all covered employees and works councils. These institutions will be considered next. Currently about 40% of private sector employees are subject to collective bargaining agreements and about the same percentage work in establishments with co-determination via works councils. While the two institutions often overlap, a substantial fraction of employees is covered by only one of them (Ellguth and Kohaut, 2021).

4.3.1 Collective Wage Agreements

Collectively negotiated wages are generally paid to union members and non-members alike in covered firms. However, being a union member could be correlated with working in a covered firm. Therefore, it is feasible that some of the benefits and costs associated with union activities may result from collective bargaining but not from union membership. To investigate this possibility, we utilise the fact that the SOEP provides information on collective wage agreements since 2014.

From 2014 to 2017, the relevant question was "*Are you paid according to a collectively agreed wage agreement?*". There are several possible answers: (1) "*Yes, a legally binding company wage agreement*"; (2) "*Yes, paid according to a collective wage agreement that is not legally binding for this sector/company*"; (3) "*Yes, a legally binding collective wage agreement*"; (4) "*No, my job *is exempt from the collective wage agreement in place where I work*"; and (5) "*No, there is no collective wage agreement*". In 2018 and 2019, the questionnaires asked whether the respondent is "*paid according to a collectively agreed wage agreement*". A positive response to this question corresponds to answers (1)–(3) of the earlier (2014–2017) question. Accordingly, we generate a dummy variable, referred to as collectively determined wage (CDW). It is equal to one if any of the first three answers was given in the years 2014–2017 or the response was 'Yes' in 2018 or 2019, and zero otherwise.[16] Since the SOEP contains information on union membership in 2015 and 2019, we utilise the information from these two years when discussing CDW. In our data, the correlation coefficient between union membership and CDW is about 0.22.

TABLE 3: Trade union membership and job satisfaction:
Collective wage agreement & works council

	(1)	(2)	(3)	(4)	(5)	(6)
		OLS			FE	
Panel (A): Collective wage agreement – collective determined wage (CDW)						
TU member	-0.1093***	-0.1409***	-0.3866***	0.1520	0.1383	-0.0433
	(0.0422)	(0.0425)	(0.0918)	(0.1116)	(0.1116)	(0.1927)
CDW		0.1822***	0.1491***		0.1684**	0.1388*
		(0.0321)	(0.0334)		(0.0717)	(0.0755)
TU member × CDW			0.3112***			0.2255
			(0.1013)			(0.1954)
Observations			21,347			
Number of respondents			15,899			
Panel (B): Works council						
TU member	-0.1506***	-0.1617***	-0.3868***	-0.0095	-0.0242	0.2350
	(0.0433)	(0.0439)	(0.1245)	(0.0893)	(0.0897)	(0.2133)
WC in firm$_{imp}$		0.0674	0.0462		0.1744*	0.2033**
		(0.0410)	(0.0413)		(0.0967)	(0.0996)
TU member × WC in firm$_{imp}$			0.2576*			-0.2917
			(0.1317)			(0.2167)
Observations			20,329			
Number of respondents			12,928			

Notes: This table shows whether the existence of collective wage agreements or of a works council influences the relationship between trade union membership and job satisfaction. We apply the indicator of collective wage determination (CDW) in Panel (A) and a dummy variable for being covered by works council in Panel (B). The observation years in Panel (A) are 2015 and 2019, and the observation period in Panel (B) is from 2001 to 2019. OLS estimates are depicted in columns (1)–(3) and FE estimates are shown in columns (4)–(6). The covariates are the same as in column (3) of Table 2. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

[16] We can also utilise information from 2014 to 2017 and focus on collective bargaining coverage (CBC). The dummy variable CBC takes the value of one if respondents choose either of the answers (1), (3) or (4), and zero otherwise. The OLS results for CBC (not depicted) are comparable to those reported for CDW and available upon request.

We depict results from OLS specifications in columns (1)–(3) and FE estimates in columns (4)–(6) in Panel (A) of Table 3, respectively. In each case, we first estimate the baseline specification with the new samples (columns (1) and (4)), which yield results that are comparable to those for the main sample (see Table 2). These findings do not qualitatively change if collective wage determination (CDW) is accounted for (see columns (2) and (5)). The OLS coefficient on union membership becomes even more negative in terms of magnitude. The results also emphasise the positive correlation between collective bargaining and job satisfaction. In columns (3) and (6), we include the interaction terms between union membership and the dummy variable for collective agreements. The OLS estimates suggest that the association between union membership and job satisfaction is statistically negative for uncovered employees, but not significantly different from zero for covered employees. The FE specifications constantly show an insignificant coefficient on union membership.

Our analysis, furthermore, indicates that job satisfaction is highest for respondents who do not belong to a trade union and whose wages are determined by collective agreements. It is somewhat lower for union members paid in line with a collective wage agreement, even lower for uncovered non-members, and lowest for uncovered union members. This ranking, first, implies that free riders on the benefits of collective bargaining are more satisfied with their job than union members are. Second, there is a positive coverage effect on job satisfaction, which cannot be due to higher wages since the wage measure is included as a control variable. Third, the results for Germany differ from those for the UK, which suggest a negative association between coverage and job satisfaction (see Bryson and White (2016a), Green and Heywood (2015), and also Powdthavee (2011)).

In sum, although the negative correlation in column (1) of Table 3 is driven by individuals uncovered by collective wage agreements, the seeming-paradox that a negative correlation is found in Table 2 is not due to the lack of information about coverage. The coefficient remains negative if the coverage is accounted for. Moreover, the findings in Panel (A) of Table 3 provide no support for the hypothesis that collective bargaining results in a deterioration of working conditions, which, in turn, lead to lower job satisfaction.

4.3.2 Works Councils

The existence of a works council is correlated with job satisfaction. However, the direction of this relationship is ambiguous (see Jirjahn and Tsertsvadze (2006), Grund and Schmitt (2011), and Bellmann et al. (2019)). In addition, employees in firms with a works council are more likely to be trade union members than those in establishments without one (Goerke and Pannenberg, 2021). Lastly, the effect of union membership on job satisfaction may be different for employees in works council firms if such institutions alter the scope for satisfaction-enhancing union activities. Therefore, not controlling for the existence of a works council may bias our results.

We use two questions in the SOEP to generate the variable on whether there is a works council in the firm in which the respondent works. The first question, included in the waves in 2001, 2006, 2011, 2016 and 2019, is "*Does an employees' council exist at your place of work?*".[17] The second question, asked in the years 2001, 2003, 2006, 2007, 2011, 2015 and 2019, asks if the respondent is a member of the council. In case of a positive response, we can be sure that there is a works council at the place of work. To enhance the number of observations, we impute works council information for intermittent years. In particular, we do so if the respondent reported the same answer to the question concerning a works council in two neighbouring survey years, holding the firm/job constant conditional on tenure.[18] We restrict the sample to the private sector and those respondents who work in a firm with more than five employees because a works council can only be elected in establishments above this size threshold. In the new estimating sample, the correlation coefficient between union membership and the existence of a works council is 0.27. Furthermore, the works council variable is weakly correlated with job satisfaction, and the raw correlation coefficient is 0.02.

Panel (B) of Table 3 shows the results of the baseline specification using the new estimation

[17]The German questionnaire distinguishes between a works council (in the private sector) and a staff or personnel council (in the public sector). The English translation of the questionnaire provided by the SOEP subsumes them under the label of employees' council, which we do not use here in order not to mix up the terminology.

[18]Taking the years 2001 to 2006 as an example, we impute the response to the question on the existence of a works council from 2001 to years 2002–2005 if the answers in 2001 and 2006 are identical and individuals' tenure has increased monotonically during this period. Moreover, if the information on being a member of a council is available in 2003, we first do an imputation for the year 2002 with the information from 2001 and 2003, and do another imputation for years 2004–2005 using the information from 2003 and 2006, also taking the increase in tenure into consideration. Applying this method to all years available, we can have raw and imputed information on works councils in each year from 2001 to 2019. Considering the question on trade union membership, the final sample for estimation consists of individuals interviewed in the years 2001, 2003, 2007, 2011, 2015 and 2019.

sample in columns (1) and (4), for a specification in which we account for the existence of a works council in columns (2) and (5), and for a specification including the interaction term between union membership and works council in columns (3) and (6). Columns (1) and (4) show qualitatively the same results as in Table 2. Columns (2) and (5) suggest that including the existence of a works council in the model does not change the findings qualitatively. Furthermore, the OLS estimates in column (3) show that the lower job satisfaction of union members may be more pronounced in firms without a works council. However, the FE estimates depicted in column (6) do not corroborate this conjecture, at least for those individuals who change their union membership status. Moreover, column (3) does not indicate a direct job satisfaction effect of a works council, in line with results by Bellmann et al. (2019), while the FE estimate in column (6) hints at a positive correlation.

In summary, controlling for the existence of a works council does not change the main results qualitatively: There is a negative correlation between union membership and job satisfaction in OLS specifications, which can no longer be observed once we incorporate person fixed effects. Therefore, in the German context, the correlation between union membership and job satisfaction will not be strongly biased or qualitatively changed if collective bargaining agreements or works councils cannot be accounted for due to, e.g., data availability.

4.4 2SLS and IVFE Results

In this section, we depict and discuss the results from the IV analysis. Similar to the correlation analysis, we present results with and without person fixed effects, i.e., 2SLS estimates and IVFE estimates, and excluding and including the second set of labour market characteristics.[19]

All of the specifications in Table 4 show a statistically significant and positive impact of union density on an individual's membership in the first stage. The size of the coefficient gets smaller if person fixed effects are accounted for. The F statistics are well above 10, indicating that our IV is not weak. In the second stage, we find no significant impact of the union membership on individuals' job satisfaction. Therefore, these findings provide no evidence in support of the conjecture that trade union membership in Germany causally affects job satisfaction.

As discussed in Section 3.2.2, we cannot entirely rule out the possibility that the IV is correlated

[19] Compared to Table 2, the number of observations depicted in Table 4 decreases because we dropped individuals who only responded once in the observation period.

Table 4: Trade union membership and job satisfaction:
Instrumental variable analysis

	(1)	(2)	(3)	(4)
	2SLS		IVFE	
Second stage:				
TU member	-0.3718	-0.3200	-0.9668	-0.7504
	(0.2362)	(0.2425)	(0.6932)	(0.6915)
First stage:				
TU density of others	0.2943***	0.2863***	0.1113***	0.1106***
	(0.0161)	(0.0159)	(0.0128)	(0.0128)
F statistic	334.53	323.25	75.66	74.96
Mean of outcome	7.1112			
Mean of IV	0.2130			
S. D. of IV	0.1767			
Observations	63,672			
Number of respondents	19,416			
State & FE	X	X	X	X
Demographic characteristics	X	X	X	X
Labour market factors (1)	X	X	X	X
Labour market factors (2)		X		X

Notes: This table shows results of IV-estimations using the trade union density in an individual's region, industry and survey year as the instrumental variable. The observation period is 1985–2019. 2SLS estimates are depicted in columns (1)–(2) and IVFE estimates are shown in columns (3)–(4). The covariates are the same as in the specifications (2) and (3) of Table 2, except that the industry dummy variables are defined in a broader way. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

with the error term, has a direct impact on job satisfaction, or affects it through other channels than an individual's union membership. If that were the case, our conclusion mentioned above that there is no causal impact might be premature. To mitigate such concerns, we undertake a set of robustness checks (see Table A-9 in the Appendix for details of the results).

First, the IV utilises regional, sectoral, and temporal variations in union membership. We have accounted for industry and year fixed effects in the baseline model. However, it is possible that some regions always have a high or low union density. Moreover, individuals from specific regions (ROR) may have different attitudes towards working conditions and, therefore, distinct levels of job satisfaction. Accordingly, we add ROR fixed effects to the model. The estimates do not change qualitatively, and the ROR-level effect does not affect our results (compare columns (1) and (2) in Table A-9 in the Appendix).

Second, we apply a broader definition of sectors. This enables us to avoid a situation in which one industry-ROR-year unit has few individuals, such that the corresponding union density may be driven by particular employees. Using a restricted sample in which we solely consider individuals from industry-ROR-year units with at least 10 employees does not change our findings (see column (3) in Table A-9 in the Appendix).

Third, as mentioned in Section 3.2.2, individuals may change their place of residence or the sector in which they work. Focusing on employees who never change their residence and sector, i.e., non-movers, allows us to observe a more homogeneous group of people than in our base sample (see Table 4). Moreover, we can cluster the standard errors at the ROR-industry level. The 2SLS estimate does not change much compared to the baseline result, while the IVFE estimate decreases in magnitude, but remaining statistically insignificant (compare columns (1) and (4) in Table A-9 in the Appendix).

Fourth, we employ the lagged union density of other employees as an IV. This IV is more exogenous than the one in the baseline model. However, its use requires that the respondents stay in the survey for several years. This restriction reduces the sample size by half. The first-stage estimates are comparable to those of the baseline specification and the second stage continues to yield insignificant estimates (see column (5) in Table A-9 in the Appendix).

Fifth, union density is calculated using SOEP data at the ROR-industry-year level. It would certainly be preferable to obtain this information from a different source. However, we are not aware of union density information for Germany at such a disaggregated level based on an alternative data set. Therefore, as a further robustness check, we aggregate union density differently. First, we calculate it at the state-industry-year instead of the ROR-industry-year level. This method allows us to take into consideration the possibility that the union density of a specific industry in one ROR region affects the individual decision to join a union in another ROR region within the same federal state. Second, we calculate union density at the ROR-year level, neglecting industry information. This approach caters for the possibility that the union density of one industry influences employees from another sector. In both cases we use current and lagged union density as IV. Irrespective of the specification, our main findings hold (see columns (6)–(9) in Table A-9 in the Appendix).

Finally, the IV may affect job satisfaction independently of individual union membership. To show the robustness of our results, we follow the strategy in Conley et al. (2012) and assume that the IV has a direct effect, γ, on job satisfaction in the second stage. If the IV is valid, γ must be zero, i.e., the exclusion restriction holds. Then, following Nybom (2017), we define λ as the share of the reduced-form effect of the IV on job satisfaction that is independent of the IV, and λ ranges between zero and one.[20] Figure A-1 plots the 2SLS and IVFE estimates for different values of λ.[21] These figures show that if the plausibility of the exclusion restriction goes down, the point estimates become less negative, but are never statistically significant.

In sum, our conclusion that there is no evidence of a causal impact of trade union membership on job satisfaction in Germany is robust to alternative specifications of the IV approach.

4.5 Effect Heterogeneity

When comparing findings from OLS and FE specifications, we observe differences between individuals who do not change their union status and those who do, as well as between respondents who join a trade union and who leave it. Such differences in the correlation between union membership and job satisfaction could also arise across groups of employees. It may be the case, for example, that trade unions better represent the interests of male members, given their quantitative dominance in Germany. This could result in a gender-specific job satisfaction differential of union membership. A similar line of argument may apply to working time, if trade unions primarily act on behalf of full-time employees. Finally, Blanchflower and Bryson (2020), Blanchflower et al. (2022), and Artz et al. (2021) argue that the relationship under investigation may be different for younger individuals and may have changed in recent years. Therefore, in this section, we present results from according subgroup analyses in Table 5, based on OLS, FE, 2SLS, and IVFE estimates.

Gender

Panel (A) in Table 5 documents a significant and negative correlation between union membership and job satisfaction in the OLS model for males and females (columns (1) and (5)). The size of the coefficient is larger for females. In all other estimates, we observe no significant effects. Consequently, there are no clear-cut gender difference in the relationship between union membership

[20] $\lambda = 0$ correspondences to the notation $\gamma = 0$ in Conley et al. (2012).

[21] We generate the figures using the STATA command *plausexog* by Clarke and Matta (2018) with the local to zero approach. The point estimates for $\lambda = 0$ present the corresponding 2SLS and IVFE estimates shown in column (2) and (4) of Table 4, respectively. The grey solid areas depict the 90% confidence intervals if we assume that the variance of γ is zero, while the grey dashed lines are the 90% confidence intervals if we apply the variance of the reduced-form coefficient.

TABLE 5: Trade union membership and job satisfaction:
Heterogeneous effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	FE	2SLS	IVFE	OLS	FE	2SLS	IVFE
Panel (A): Gender	**Males**				**Females**			
TU member	-0.0853***	0.0120	-0.2931	-0.2064	-0.1734***	-0.0678	-0.3551	-1.8434
	(0.0295)	(0.0424)	(0.2824)	(0.7975)	(0.0382)	(0.0552)	(0.5625)	(1.3860)
F statistic			211.78	46.56			84.00	26.23
Observations	42,767		33,988		38,402		29,684	
Number of respondents	18,589		10,115		17,833		9,301	
Panel (B): Age	**Older than 42**				**42 or younger**			
TU member	-0.1790***	-0.0976*	-0.0796	0.4853	-0.0390	0.0368	-0.5734*	-1.1515
	(0.0318)	(0.0544)	(0.3965)	(1.4238)	(0.0311)	(0.0501)	(0.3162)	(1.1873)
F statistic			119.76	24.34			213.23	24.82
Observations	38,990		29,624		42,179		29,698	
Number of respondents	19,509		10,321		22,601		10,392	
Panel (C): Birth year	**Born before 1964**				**Born in/after 1964**			
TU member	-0.1713***	-0.0526	-0.3325	-0.7165	-0.0157	0.0324	-0.2639	-1.2609
	(0.0311)	(0.0439)	(0.3241)	(0.7629)	(0.0348)	(0.0523)	(0.4001)	(1.5181)
F statistic			168.67	60.04			134.97	16.51
Observations	40,006		33,900		41,163		29,772	
Number of respondents	15,344		9,549		21,078		9,867	
Panel (D): Employment	**Full-time**				**Part-time**			
TU member	-0.1040***	-0.0062	-0.1903	-0.0677	-0.1782***	-0.1693*	-1.1730	-8.0815
	(0.0256)	(0.0383)	(0.2633)	(0.7451)	(0.0516)	(0.0934)	(0.8927)	(5.6485)
F statistic			260.22	58.62			35.48	3.63
Observations	59,937		45,735		21,232		13,207	
Number of respondents	28,024		14,199		12,739		4,794	

Notes: This table shows heterogeneous relationships between trade union membership and job satisfaction by gender, age, birth year and employment status. The observation period is 1985–2019. The covariates are the same as in column (3) of Table 2. The number of observations decreases in specifications using 2SLS and IVFE estimations because we dropped individuals who only responded once in the corresponding subgroup. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

and job satisfaction in Germany. This result is similar to findings for the UK and the US (Bender and Sloane, 1998; Bender et al., 2005; Artz, 2012; Bryson and White, 2016a; Powdthavee, 2011).

Age and Birth Cohort

In our estimation sample, the median age is 42. Using this cut-off, Panel (B) in Table 5 shows results for the two age groups, i.e., individuals older or younger than 42. Additionally, we classify individuals into two groups by using the sample median of the birth year, i.e., individuals born before 1964 and in/after 1964, and depict results in Panel (C). OLS estimates are significant and negative for older individuals, while the estimates are statistically insignificant for younger people. FE models yield mostly insignificant results. Similarly, if we take the endogeneity of union membership into consideration, the estimates are mostly insignificant, too.

Accordingly, our results suggest that it is less likely to observe a negative correlation between union membership and job satisfaction for younger people, but the association does not turn positive in Germany. These findings are to some extent consistent with those reported by Artz et al. (2021), Blanchflower and Bryson (2020), and Blanchflower et al. (2022) who document a negative relationship between union membership and job satisfaction for older individuals in the US and the UK, and a positive one in later years without accounting for the endeogeneity of the union membership.

Employment Status: Full-time vs. Part-time

Finally, we distinguish according to an individual's employment status. We refer to individuals working more than 36 hours per week as full-time employees and, for simplicity, to those reporting fewer working hours as part-time employees. It is evident from Table 1 that the probability of having a full-time job is significantly higher for union members than non-members. We estimate the baseline specification separately for the two groups of individuals. Results for full-time employees are shown in columns (1)–(4) of Panel (D), and for individuals working part time in columns (5)–(8).[22]

Both full-time and part-time employees show negative correlations in OLS models, but the size

[22] We also use 32 hours as a cut-off. Additionally, we apply the self-reported information on full-time employment to group individuals. Our findings are robust to the use of these alternative classifications. Results are available upon request.

of the coefficient is much larger for part-time workers. The FE estimate for full-time employees is close to zero and insignificant, while it is negative and borders on significance for part-time employees. If we apply IV regression models, we do not find any significant effects.[23]

We can summarise the heterogeneity analysis as follows: OLS specifications indicate a negative correlation between union membership and job satisfaction for most subgroups. Once time-invariant unobservable factors are accounted for, such a correlation can no longer be observed, with the possible exception of part-time and older employees. Importantly, the IV analyses suggest the absence of a causal impact for all subgroups, corroborating our findings for the entire sample.

5 Summary

Researchers have devoted much effort to explaining the negative relationship between trade union membership and job satisfaction. They have offered several explanations for the apparent paradox and suggested various strategies to estimate the 'true' causal impact. In this paper, we provide the first comprehensive analysis of the association between union membership and job satisfaction in Germany. Using SOEP data from 1985 to 2019, we find a negative and significant relationship between union membership and job satisfaction in OLS models and no significant association in FE models. Using the leading values of job satisfaction in OLS and FE specification, instead of contemporaneous information, provides similar estimates, indicating that reverse causality may not be an issue. Taking into account information on collective bargaining coverage or the existence of works councils in subsample analyses does not qualitatively change the main findings. Our results also indicate that the association between union membership and job satisfaction tends to be more negative for employees who are not covered by the two institutions than for those who are. To better account for time-variant confounders, we use union density in the same industry, region, and year as an IV. We do not find a significant impact of union membership on job satisfaction. Therefore, we do not obtain evidence suggesting that trade union membership causally affects job satisfaction in Germany.

The difference between OLS and FE specifications and the insignificance of IV estimates indicate that the negative association between union membership and job satisfaction observed in the

[23] Note that the F statistic in column (8) is smaller than 10. It is intuitive because part-time employees are less likely to join the union and the sample size is much smaller.

correlation analysis may be due to sorting effects. Our somewhat preliminary analysis of individuals who become union members and of those who leave the trade union also suggests that sorting effects may be relevant. In future research, it will, therefore, be desirable to better understand such sorting effects.

Moreover, IV approaches have usually focused on time-invariant instruments, while we provide one which changes over time. In the absence of an exogenously-caused variation in union membership, this IV, to our mind, deserves closer scrutiny since its use can help to progress towards a causal interpretation in industrial relations settings in which institutions such as certification elections do not exist.

On a broader scale, the IV approach, though not perfect, can be considered and potentially employed for the analysis of other consequences of union membership in Germany and further countries. Meanwhile, necessary robustness checks should be provided in order to produce convincing results. While we interpret our findings as demonstrating the absence of a causal impact of union membership on job satisfaction, we cannot solve the puzzle as to why many individuals voluntarily become union members, and incur the costs of membership fees, without a discernible positive impact.

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Job Satisfaction and Trade Union Membership in Germany

Laszlo Goerke

IAAEU - Trier University*, IZA, Bonn, CESifo, München and GLO

goerke@iaaeu.de

Yue Huang

IAAEU - Trier University* and GLO

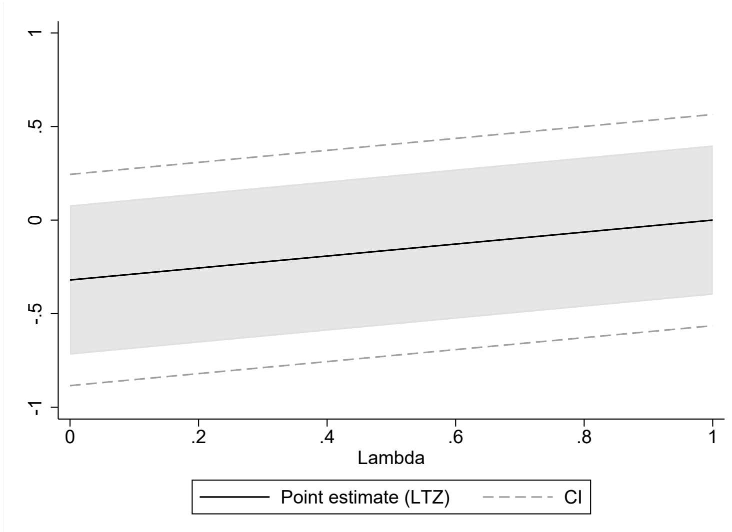
huang@iaaeu.de

Supplementary Material

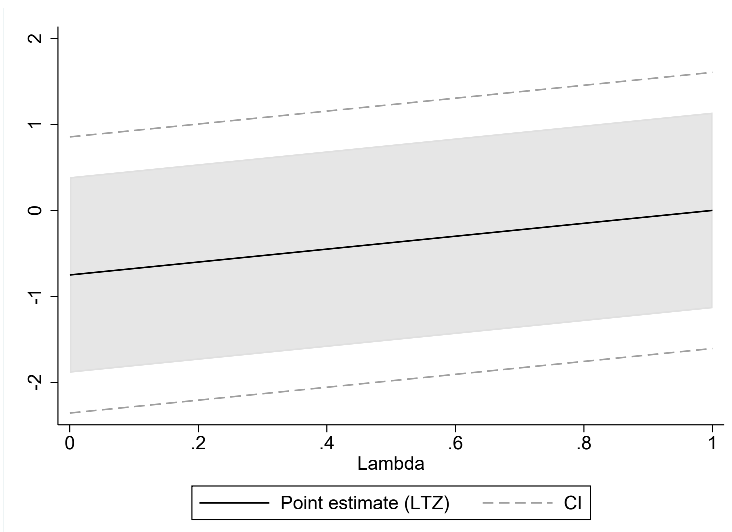
*Institute for Labour Law and Industrial Relations in the European Union, Trier University, Campus II, D-54286 Trier, Germany

Appendix A Figures and Tables Cited in the Main Text

FIGURE A-1: Instrumental variable analysis: Plausibly exogenous



(a) 2SLS



(b) IVFE

Notes: We use the Stata command *plausexog* to calculate the point estimate and confidence interval, and generate the figures. Lambda represents the share of the reduced-form effect of the IV (TU density of others) on the outcome (job satisfaction) that is independent of the individual union membership. The black line is the point estimate of union membership for different values of lambda. The Grey solid areas depict the 90% confidence intervals if we assume that the variance of the effect of the IV on job satisfaction independent of the individual union membership is zero, while the grey dashed lines are the 90% confidence intervals if we apply the variance of the reduced-form coefficient. The figures show that the effect of TU on JS is statistically insignificant at different values of lambda.

Table A-1: Summary statistics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	All observations				Trade union member				Not trade union member			
	Mean	S. D.	Min	Max	Mean	S. D.	Min	Max	Mean	S. D.	Min	Max
job satisfaction	7.1067	2.0296	0	10	6.9947	2.1119	0	10	7.1397	2.0035	0	10
life satisfaction	7.2156	1.6553	0	10	7.0811	1.7369	0	10	7.2553	1.6283	0	10
TU member	0.2276	0.4193	0	1								
age	41.0022	11.7737	16	65	42.8717	11.2564	17	65	40.4514	11.8659	16	65
female	0.4420	0.4966	0	1	0.3296	0.4701	0	1	0.4751	0.4994	0	1
married	0.5699	0.4951	0	1	0.6344	0.4816	0	1	0.5509	0.4974	0	1
children number	0.5718	0.8836	0	10	0.5663	0.8743	0	10	0.5734	0.8863	0	8
years of education	12.1201	2.5741	7	18	11.7860	2.3432	7	18	12.2185	2.6302	7	18
ln(monthly wages)	7.4465	0.8136	0	11.5129	7.5394	0.6644	0	10.1807	7.4191	0.8506	0	11.5129
ln(working hours)	3.5903	0.4206	0	4.3820	3.6603	0.3039	0	4.3820	3.5697	0.4472	0	4.3820
public sector	0.2793	0.4487	0	1	0.3612	0.4804	0	1	0.2552	0.4360	0	1
civil servant	0.0742	0.2620	0	1	0.1136	0.3173	0	1	0.0625	0.2421	0	1
blue-collar worker	0.3180	0.4657	0	1	0.4368	0.4960	0	1	0.2831	0.4505	0	1
tenure	10.7077	10.1118	0	50.9170	14.6491	10.9412	0	49.2500	9.5464	9.5488	0	50.9170
having a new job	0.1302	0.3366	0	1	0.0667	0.2496	0	1	0.1490	0.3560	0	1
firm size												
< 20	0.2150	0.4108	0	1	0.0668	0.2497	0	1	0.2586	0.4379	0	1
$[20, 200)$	0.2828	0.4504	0	1	0.2309	0.4214	0	1	0.2981	0.4574	0	1
$[200, 2000)$	0.2358	0.4245	0	1	0.2844	0.4511	0	1	0.2215	0.4152	0	1
≥ 2000	0.2664	0.4421	0	1	0.4179	0.4932	0	1	0.2218	0.4155	0	1
NACE categories												
1	0.0132	0.1141	0	1	0.0092	0.0954	0	1	0.0144	0.1190	0	1
2	0.0044	0.0664	0	1	0.0133	0.1144	0	1	0.0018	0.0427	0	1
3	0.2814	0.4497	0	1	0.3884	0.4874	0	1	0.2499	0.4330	0	1
4	0.0131	0.1139	0	1	0.0214	0.1446	0	1	0.0107	0.1030	0	1
5	0.0656	0.2476	0	1	0.0486	0.2150	0	1	0.0706	0.2562	0	1
6	0.1198	0.3247	0	1	0.0658	0.2479	0	1	0.1357	0.3424	0	1
7	0.0239	0.1527	0	1	0.0069	0.0831	0	1	0.0289	0.1675	0	1
8	0.0604	0.2382	0	1	0.1036	0.3048	0	1	0.0477	0.2130	0	1
9	0.0404	0.1970	0	1	0.0231	0.1502	0	1	0.0456	0.2085	0	1
10	0.0688	0.2530	0	1	0.0224	0.1478	0	1	0.0824	0.2750	0	1
11	0.0932	0.2908	0	1	0.1278	0.3339	0	1	0.0831	0.2760	0	1
12	0.0627	0.2424	0	1	0.0624	0.2420	0	1	0.0627	0.2425	0	1
13	0.1141	0.3180	0	1	0.0777	0.2677	0	1	0.1248	0.3305	0	1
14	0.0348	0.1832	0	1	0.0270	0.1621	0	1	0.0371	0.1890	0	1
15	0.0031	0.0552	0	1	0.0006	0.0238	0	1	0.0038	0.0615	0	1
16	0.0010	0.0324	0	1	0.0019	0.0433	0	1	0.0008	0.0283	0	1
Observations	81,169				17,656				63,513			

Notes: This table shows the summary statistics for the estimation sample of the baseline specification.

Table A-2: Trade union membership and job satisfaction (full table)

	(1)	(2)
	OLS	FE
TU member	-0.1160***	-0.0198
	(0.0233)	(0.0337)
age	-0.0429***	-0.0048
	(0.0055)	(0.0092)
age^2	0.0004***	-0.0002**
	(0.0001)	(0.0001)
female	-0.0710***	
	(0.0215)	
married	0.0755***	0.0072
	(0.0211)	(0.0334)
education	-0.1124***	-0.2541**
	(0.0317)	(0.1027)
education^2	0.0031***	0.0078**
	(0.0012)	(0.0038)
# of children	0.0416***	0.0167
	(0.0092)	(0.0137)
ln(monthly wages)	0.2281***	0.2084***
	(0.0167)	(0.0264)
ln(working hours)	-0.2183***	-0.0714*
	(0.0263)	(0.0369)
public sector	0.0917***	0.0695*
	(0.0283)	(0.0415)
civil servant	0.0585	0.0721
	(0.0437)	(0.1095)
blue-collar worker	-0.3308***	-0.2607***
	(0.0232)	(0.0375)
tenure	-0.0083***	-0.0468***
	(0.0031)	(0.0041)
tenure^2	0.0001	0.0005***
	(0.0001)	(0.0001)
new job	0.1890***	0.1848***
	(0.0256)	(0.0334)
firm size (reference: <20)		
$[20, 200)$	-0.1083***	0.0208
	(0.0238)	(0.0349)
$[200, 2000)$	-0.1102***	0.1011**
	(0.0269)	(0.0400)
≥ 2000	-0.0522*	0.1637***
	(0.0272)	(0.0426)

Notes: This table continues on the next page.

TABLE A-2: Trade union membership and job satisfaction (full table continued)

	(1)	(2)
	OLS	FE
NACE categories (reference: group 1)		
group 2	-0.0363	0.2976
	(0.1297)	(0.2224)
group 3	-0.0497	0.1127
	(0.0799)	(0.1248)
group 4	0.0296	0.2838*
	(0.1052)	(0.1591)
group 5	-0.0691	0.1225
	(0.0841)	(0.1311)
group 6	-0.1898**	0.0992
	(0.0813)	(0.1269)
group 7	-0.1179	0.2313
	(0.0924)	(0.1622)
group 8	-0.1533*	0.1606
	(0.0856)	(0.1380)
group 9	-0.1413	0.1725
	(0.0900)	(0.1561)
group 10	-0.1032	0.1347
	(0.0831)	(0.1287)
group 11	-0.0019	0.2864**
	(0.0864)	(0.1341)
group 12	0.1584*	0.4216***
	(0.0866)	(0.1408)
group 13	0.0144	0.3564***
	(0.0825)	(0.1358)
group 14	0.0799	0.3200**
	(0.0890)	(0.1361)
group 15	-0.3325**	-0.2529
	(0.1613)	(0.2491)
group 16	0.4231	0.7021*
	(0.2806)	(0.4070)
Observations	81,169	
Number of respondents	36,422	
State & year fixed effects	X	X

Notes: This table shows the relationship between trade union membership and job satisfaction. The observation period is 1985-2019. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A-3: Trade union membership and job satisfaction: Ordered probit model

	(1)
Panel (A): Coefficients	
TU member	-0.0556***
	(0.0120)
Panel (B): Marginal effects	
satisfaction level	
0	0.0010***
	(0.0002)
1	0.0008***
	(0.0002)
2	0.0019***
	(0.0004)
3	0.0030***
	(0.0007)
4	0.0029***
	(0.0006)
5	0.0060***
	(0.0013)
6	0.0037***
	(0.0008)
7	0.0026***
	(0.0005)
8	-0.0051***
	(0.0012)
9	-0.0078***
	(0.0017)
10	-0.0090***
	(0.0019)
Observations	81,169
A full set of controls	X

Notes: This table shows ordered probit regression results of the relationship between trade union membership and job satisfaction. The observation period is 1985–2019. Covariates are the same as in column (3) of Table 2. Individual fixed effects are not controlled for. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A-4: Trade union membership and job satisfaction: Selection on unobservables

	(1)	(2)	(3)
	Coefficient from baseline model	Identified set	δ for $\beta = 0$ given R_{max}
TU member	-0.1160***	[-0.1160, -0.1110]	8.37

Notes: This table shows the validation results for the analysis of the correlation between trade union membership and job satisfaction. The coefficient in column (1) is the same as the OLS estimate in column (3) of Table 2 where all covariates are included in the regression model. The identified set in column (2) is bounded below by the estimate if the proportional degree of selection on unobservables to the selection on observables is 0 ($\delta = 0$), i.e., the coefficient shown in column (1). The upper bound is calculated based on the situation that the selection on unobservables is equal to the selection on observables ($\delta = 1$) and $R_{max} = 1.3(R^2)$, where R^2 is from column (1). Column (3) shows the value of δ which would produce $\beta = 0$ given the value of R_{max}. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A-5: Trade union membership and life satisfaction

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS			FE		
TU member	-0.0647***	-0.0823***	-0.1082***	-0.0501*	-0.0588**	-0.0603**
	(0.0193)	(0.0197)	(0.0197)	(0.0257)	(0.0257)	(0.0257)
Observations	81,169					
Number of respondents	36,422					
Control variables						
State & year FE	X	X	X	X	X	X
Demographic characteristics	X	X	X	X	X	X
Labour market factors (1)		X	X		X	X
Labour market factors (2)			X			X

Notes: This table shows the relationship between trade union membership and life satisfaction. The observation period is 1985–2019. The description of covariates can be found in Table 2. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

TABLE A-6: Differences in characteristics between switchers and non-switchers

	Switchers − Non-switchers	
	Coeff.	S.E.
age	-0.084	(0.138)
female	-0.071***	(0.006)
married	0.040***	(0.006)
educ	-0.328***	(0.031)
# of children	0.076***	(0.010)
ln(monthly wage)	0.043***	(0.009)
ln(working hours)	0.081***	(0.004)
public sector	0.074***	(0.006)
civil servant	0.026***	(0.004)
blue-collar worker	0.105***	(0.006)
tenure	0.815***	(0.124)
new job	-0.030***	(0.004)
firm size		
< 20	-0.097***	(0.004)
$[20, 200)$	0.013**	(0.006)
$[200, 2000)$	0.034***	(0.005)
≥ 2000	0.050***	(0.006)

Notes: This table represents the difference in respondents' characteristics between union status switchers and non-switchers. To show the difference, we regress the variable of characteristics on the dummy variable for being a switcher using OLS models and applying weighting factors. Robust standard errors in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A-7: Trade union membership and job satisfaction: Switchers and non-switchers

	(1)	(2)	(3)	(4)	(5)
	Non-switchers	Switchers			
	OLS	OLS	FE	OLS	FE
TU member	-0.1344***	-0.0158	-0.0156	0.0526	0.0207
	(0.0290)	(0.0365)	(0.0347)	(0.0450)	(0.0432)
Only entry				0.0689	
				(0.0653)	
TU member × Only entry				-0.2215**	-0.1160
				(0.0862)	(0.0848)
Observations	65,771	15,398			
Number of respondents	32,713	3,709			
A full set of controls	X	X	X	X	X

Notes: This table shows the relationship between trade union membership and job satisfaction for union status non-switchers (column (1)) and switchers (columns (2)–(5)). For switchers, we also distinguish the association between individuals only reporting an entry in the union and those reporting an exit from the union or multiple status changes. The observation period is 1985–2019. The covariates are the same as in column (3) of Table 2. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A-8: Trade union membership and job satisfaction:
Satisfaction measured in t and $t+1$

	(1)	(2)	(3)	(4)
	OLS		FE	
	t	$t+1$	t	$t+1$
TU member	-0.1169***	-0.1159***	-0.0351	0.0058
	(0.0263)	(0.0268)	(0.0383)	(0.0380)
Observations	58,798			
Number of respondents	26,693			
A full set of controls	X	X	X	X

Notes: This table shows regression results of the relationship between trade union membership and job satisfaction. Job satisfaction and union membership are measured in the same year in columns (1) and (3), while the satisfaction indicator is measured one year later in columns (2) and (4). The observation period is 1985–2019. Covariates are the same as in columns (3) and (6) of Table 2. Robust standard errors (clustered at individual level) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

TABLE A-9: Effect of trade union membership on job satisfaction: Robustness checks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Baseline	ROR FE	≥ 10 persons	Non-movers	Lagged IV	State-industry-year IV		ROR-year IV	
Panel (A): 2SLS									
Second stage:									
TU member	-0.3200	-0.5193	-0.1353	-0.2590	-0.4154	-0.2992	-0.1685	0.0013	-0.0978
	(0.2425)	(0.3327)	(0.2859)	(0.3228)	(0.2884)	(0.3115)	(0.3531)	(0.2638)	(0.3451)
First stage:									
TU density of others	0.2863***	0.2124***	0.3552***	0.3654***					
	(0.0159)	(0.0157)	(0.0237)	(0.0331)					
Lagged TU density of others					0.2726***				
					(0.0183)				
TU density of others (state-industry-year)						0.4374***			
						(0.0311)			
Lagged TU density of others (state-industry-year)							0.3186***		
							(0.0261)		
TU density of others (ROR-year)								0.5748***	
								(0.0355)	
Lagged TU density of others (ROR-year)									0.4105***
									(0.0341)
F statistic	323.25	183.03	224.76	122.16	221.98	197.68	148.75	262.55	145.35
Observations	63,672	63,672	50,613	36,206	31,413	64,273	31,832	64,199	31,696
Number of respondents	19,416	19,416	16,032	12,010	9,385	19,534	9,475	19,524	9,445
A full set of controls	X	X	X	X	X	X	X	X	X

Notes: This table continues on the next page.

TABLE A-9: Effect of trade union membership on job satisfaction:
Robustness checks (continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Baseline	ROR FE	$\geq$ 10 persons	Non-movers	Lagged IV	State-industry-year IV		ROR-year IV	
Panel (B): IVFE									
Second stage:									
TU member	-0.7504	-0.7846	-0.8494	-0.0763	-1.0473	-0.7018	0.0200	-0.6334	-0.7312
	(0.6915)	(0.7010)	(0.8584)	(0.7523)	(0.9626)	(0.5817)	(0.6242)	(0.5420)	(0.6797)
First stage:									
TU density of others	0.1106***	0.1097***	0.1361***	0.1453***					
	(0.0128)	(0.0128)	(0.0196)	(0.0203)					
Lagged TU density of others					0.0885***				
					(0.0150)				
TU density of others (state)						0.2656***			
						(0.0264)			
Lagged TU density of others (state)							0.2013***		
							(0.0242)		
TU density of others (ROR-year)								0.3346***	
								(0.0326)	
Lagged TU density of others (ROR-year)									0.2316***
									(0.0314)
F statistic	74.96	73.71	48.32	51.37	34.59	101.54	69.18	105.19	54.57
Observations	63,672	63,672	50,613	36,206	31,413	64,273	31,832	64,199	31,696
Number of respondents	19,416	19,416	16,032	12,010	9,385	19,534	9,475	19,524	9,445
A full set of controls	X	X	X	X	X	X	X	X	X

Notes: This table shows robustness checks for the causal analysis of the effect of trade union membership on job satisfaction. The observation period is 1985–2019. Column (1) shows the baseline results, e.g., columns (2) and (4) of Table 4. We additionally account for ROR fixed effects in column (2). In column (3), we construct the IV by considering that there are at least 10 individuals in the ROR-industry-year unit. In column (4), we only consider individuals whose ROR-industry classification did not change during the observation period. Column (5) reports results of specifications with lagged union density of others as an IV. The number of observations decreases in column (5) because we have no information on the lagged union density for individuals interviewed in year 1985 and some individuals who did not take part in the survey in specific years and whose information could not be matched. In columns (6) and (7), we aggregate the union density using the information on federal states in stead of RORs. In columns (8) and (9), we aggregate the union density using the information on RORs and survey years. All other control variables are the same as in the baseline specification, except that industrial sectors are classified in a broader way. Robust standard errors (clustered at the ROR-industry level in column (4) and at the individual level in columns (1)–(3) and (5)–(9)) in parentheses: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Appendix B Statistical Classification of Economic Activities in the European Community

Table B-1: Industry classifications

(1)	(2)	(3)	(4)	(5)	(6)
Baseline	Broader	NACE Rev. 1.1		NACE Rev. 2	
Classification	Classification	Section	Description	Section	Description
1	I	A B	Agriculture, hunting and forestry Fishing	A	Agriculture, forestry and fishing
2	I	C	Mining and quarrying	B	Mining and quarrying
3	II	D	Manufacturing	C	Manufacturing
4	III	E	Electricity, gas and water supply	D E	Electricity, gas, steam and air conditioning supply Water supply, sewerage, waste management and remediation activities
5	III	F	Construction	F	Construction
6	IV	G	Wholesale and retail trade: repair of motor vehicles, motorcycles and personal and household goods	G	Wholesale and retail trade; repair of motor vehicles and motorcycles
7	V	H	Hotels and restaurants	I	Accommodation and food service activities
8	V	I	Transport, storage and communications	H J	Transportation and storage Information and communication
9	VI	J	Financial intermediation	K	Financial and insurance activities
10	VI	K	Real estate, renting and business activities	L M N	Real estate activities Professional, scientific and technical activities Administrative and support service activities
11	VII	L	Public administration and defence; compulsory social security	O	Public administration and defence; compulsory social security
12	VII	M	Education	P	Education
13	VII	N	Health and social work	Q	Human health and social work activities
14	VIII	O	Other community, social and personal services activities	R S	Arts, entertainment and recreation Other service activities
15	VIII	P	Activities of private households as employers and undifferentiated production activities of private households	T	Activities of households as employers; undifferentiated goods- and services-producing activities of households for own use
16	VIII	Q	Extraterritorial organisations and bodies	U	Activities of extraterritorial organisations and bodies

Notes: This table shows industry classifications. NACE Rev. 1.1 is depicted in columns (3)–(4) and Rev. 2 in columns (5)–(6). We apply NACE Rev. 1.1 to observations interviewed before 2018 and NACE Rev. 2 to observations interviewed in 2018 and 2019. Column (1) shows the industry classification applied in the baseline specification. Column (2) shows the broader classification that is used to generate the instrumental variable, i.e., union density.
Data source: NACE Rev. 2 – Statistical classification of economic activities in the European Community. Eurostat Methodologies and Working Papers. ISSN 1977-0375.