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Evidence from Norway

Doctoral thesis

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Preface

This thesis consists of three independent essays addressing different topics related to the impact of unions and collective agreements on labor market outcomes within the Norwegian labor market. The essay in Chapter 1 is joint work with Professor Ragnar Nymoen at the Department of Economics at the University of Oslo. Chapter 2 is authored by me alone. The essay in Chapter 3 is written in cooperation with Fredrik Bakkemo Kostøl at the Department of Economics at the Norwegian University of Science and Technology.

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Introduction

This thesis addresses the question of what unions do within the context of the Norwegian labor market. Specifically, I examine how the presence of unions alters wage inequality, productivity, and low pay risk at the local (i.e., establishment/firm) level in the Norwegian private sector in the period 2000–2018. All analyses are conducted on matched employer-employee data sets, based on micro data made available by Statistics Norway.

The objectives and impact of unions in the labor market have been the subject of extensive research for decades. The seminal work of Freeman & Medoff (1984) is a much-cited reference in the literature, drawing a map of the core functions of unions from both the theoretical and empirical perspectives. In theory, unions may be portrayed as having two faces: the monopoly face and the exit voice/institutional response face. The former addresses the monopoly power attained by unionized workers through collective bargaining, enabling them to raise wages. In this regard, unions distort labor market efficiency by adding a union premium to the competitive market wage. The exit voice/institutional response face refers to the collective voice unions provide workers with. This voice may contribute to alter employment relations in the workplace, and thereby potentially lead to favorable outcomes for both employers and firms. Empirically, most of the research on unions in the labor market is focused on the Anglo-Saxon countries. The findings from these parts of the world do not automatically translate to the Nordic and Norwegian settings due to differences in institutional contexts. The implications of union presence in an economy are shaped by many institutional factors, including differences in legislation, the degree of centralization and coordination in wage bargaining, extension of collective agreements and regulations of membership in employee and employer organizations. For these reasons, empirical research from different countries and time periods, based on several data sources and updated methodological approaches, is necessary to understand the functions of unions in the labor market.

The potential effect of unions on *wages* is one of the oldest questions investigated in labor economics. Early on, Adam Smith (1776) argued that unions had an impact on wages, while Milton Friedman in 1950 believed that the influence of unions was exaggerated. The first wave of empirical studies conducted towards the end of the 20th century revealed a positive partial correlation between wage levels and unions, mainly in the Unites States and Canada (Freeman & Medoff 1984, Lewis 1986). Subsequent studies from other parts of the world have revealed that the *union wage premium* is in general smaller in countries where sectoral or national bargaining predominates due to the extension of negotiated wage outcomes to uncovered parts of the economy (Bryson 2010, 2014). Another general finding is that the impact of unions varies substantially across the wage distribution. In particular, unions tend to have a larger effect on wages in the lower part of the wage distribution (Card 1996, 2001; DiNardo & Lemieux 1996; Firpo et al. 2009; Farber et al. 2021). Unions have also been shown to compress wages across countries and over time (Rueda & Pontusson 2000; Pontusson 2013; Vlandas 2018). The presence of unions and collective agreements is further associated with reduced low-pay risk within countries, even for non-union members (see Benassi & Vlandas 2021 for Germany, Jordfald et al. 2021 for Norway, Schmitt 2008 for the US). Most of the studies of the equalizing impact of unions on wages are concentrated on aggregate levels, that is, across or within countries, sectors, or industries. Less attention has been paid to how union strength within the workplace affects wage levels in different groups of wage earners. The essays in Chapter 1 and Chapter 2 both contribute in this regard by providing evidence of the role of union density in shaping the within-establishment wage distribution.

In contrast to the union-wage literature, the potential effect of unions on *productivity* is assumed to be more indirect. Union presence may have an impact on productivity through several channels, such as firm-level investments in different forms of capital or employee behavior as measured by voluntary turnover, job satisfaction, and organizational commitment (Doucouliagos et al. 2017). For instance, Freeman (1976) and Freeman & Medoff (1984) claim that by providing workers with a means of expressing discontent through a collective voice, unions can reduce turnover and improve morale, motivation, job satisfaction, and cooperation, thereby enhancing productivity. The additional information provided by a collective voice can moreover enable firms to choose a better mix of working conditions, workplace rules, and wage levels (Laroche 2020). Empirically, the evidence on how unions affect productivity has been inconclusive. A recent meta-analysis by Doucouliagos et al. (2017) indicates that the overall association between unions and productivity is close to zero, but also that the strength of such a relationship may vary significantly across countries and industries. Institutional differences can be one factor contributing to variation in the implications of unions (Blanchflower & Freeman 1992). The institutions that enable and constrain union efforts vary greatly between countries. In Norway, collective agreements constitute an important organizational institution through which unions may alter employment relations at the firm level. The essay in Chapter 3 contributes to existing knowledge regarding the relationship between unions and productivity and, in particular, how this relationship is affected by the quality of industrial relations as measured by the presence of a collective agreement.

While unions are an important part of many developed economies, they represent very different institutions depending on the economic and political setting (Bryson 2007). Unlike in the US and the UK, union bargaining remains of great significance in many European countries, Norway included. The relationship between employers and employees in Norway is organized through an interaction between legislation and collective agreements, with the latter being of relatively high importance compared to other countries. Together with the other Nordic countries, Norway is one of the countries in the world with the highest density of union members (ILO 2021). Approximately half of all employees are members of a union, a share that has been relatively stable for the last two decades. Strength characterizes the employers' associations as well, with an organization rate of 73 percent. The parties, together with the state, have a long tradition of cooperation at the national level. Importantly however, the strong ties between the parties have their counterparts within companies. The management and the union representatives have a responsibility to implement the national accords and the results of collective bargaining at the company level, and they participate in productivity enhancement, restructuring, and organizational development (Løken et al. 2013). This feature of the Norwegian model implies that the workplace level is central to understanding the mechanisms underlying the relationship between unions and labor market outcomes.

The regularities uncovered in the union wage premium literature are primarily based on observational studies using cross-sectional data (Bryson 2007, Barth et al. 2020). In recent years, researchers have been increasingly concerned with identifying *causal* effects of unions on labor market outcomes. One methodological challenge in this regard is the potential endogenous determination of union presence. Unionization is not randomly distributed across either firms or individuals, and factors affecting wages, productivity, or other outcomes may also be determinants of union presence in a firm or of individual union membership. The increased availability of longitudinal matched employer-employee data is likely to reduce bias in estimating union effects (Bryson 2007). Following workers (workplaces) over time allows for netting out time-invariant factors associated with the worker (workplace) that may influence unionization and the outcome variable of interest. Still, unobservable differences between workers and firms may change over time and thus remain a source of endogeneity. Strategies to tackle this selection on unobservables include the use of instrumental variables and, in some rare cases, regression discontinuity designs (DiNardo & Lee 2004, Lee & Mas 2012). Although establishing a causal relationship between the presence of unions and wages or productivity is an appealing aim, it is not easy. Reliable sources of exogenous variation in unionization are necessary to produce unbiased 2SLS estimates, but they are hard to come by. Nevertheless, assessing results from different methodological approaches together may increase our understanding of how unions alter different outcomes.

The thesis includes three essays. The different research questions have motivated different methodological choices and empirical strategies. The dynamic framework applied in the first essay allowed us to evaluate possible short- and long-term consequences of changes in unionization, as well as potential feedback. The two subsequent essays are concerned with handling the endogenous nature of union density to estimate effects that may be given causal interpretations.

In the first essay, we model the empirical relationship between union density and wage inequality within a sample of relatively large and "long-lived" establishments in the period 2000–2018. A particular focus of the study is the dynamic relationship between the two variables. In general, the evolution of union membership and wage inequality may be thought of as gradual processes rather than an instantaneous change. The long time span covered by the data in our sample allows us to elaborate on the strength, direction, and interdependence of the relationship between union density and wage inequality. We apply dynamic panel data models that are estimated for different wage inequality measures. The results show a negative relationship between union membership and wage inequality which is robust with respect to the choice of measurement and estimation method. Although the models have no direct causality interpretation for the contemporaneous regression coefficient, they provide interpretable results about the dynamic relationships. In particular, the evidence is based on changes in union membership within establishments and measures potential effects of increasing membership on within-establishment inequality conditional on existing inequality and existing membership (lagged values). We find that the strength of the relationship increases with the permanency of the shift in union membership. Furthermore, the recursive interpretation of the system is confirmed empirically by the result of a structural break test. The findings suggests that union membership is a more important explanatory variable for the lower part of the wage distribution than for the upper part. Furthermore, the significant negative relationship between union density and wage inequality in our sample is conditional on the presence of a collective agreement.

The second essay examines the impact of workplace-level bargaining power on individual low-pay probability. One of the core objectives of unions is to raise the wages of the lowest paid workers. In the Nordic countries, these objectives are referred to as solidaristic wage policies. While previous literature has shown that strong unions are associated with lower wage inequality in their environments, particularly in the lowest part of the wage distribution, less is known about the relationship between union bargaining strength and individual low-pay risk within establishments. Due to the Norwegian system of two-tiered wage bargaining, the local level represents an important arena for evaluating the influence of union presence. By exploiting changes in tax deductions for union members in Norway as a source of exogenous variation, a negative effect of increased union density on low-pay risk is identified within job spells. The time period the data covers is characterized by a huge increase in immigration to Norway. A second question raised in the study is whether the potential reduction of low-pay probability attributable to the bargaining power of the union is heterogeneous among immigrants and natives. The results suggest that the effect of local bargaining power on individual low-pay probability was larger among immigrants than among natives. One interpretation of this finding is that immigrants are worse off, in the sense that they hold less bargaining power than natives in the first place and therefore derive greater benefit from the solidaristic wage policy unions employ.

The third essay seeks to expand the current knowledge of what unions do to firm-level productivity in Norway and, in particular, how this relationship is affected by the quality of industrial relations as measured by the presence of a collective agreement. We estimate a range of model specifications and apply several different estimators in order to identify the effect of unionization on productivity. Specifically, we introduce a new source of exogenous variation in union membership by utilizing information on intergenerational transmission of union preferences. Our results show that the qualitative interpretation of what unions do to total factor productivity depends on whether a firm is covered by a collective agreement. In the absence of an agreement, increases in union density among the workers in a firm are estimated to reduce productivity. However, the implementation of a collective agreement is estimated to moderate this negative impact. Moreover, when evaluated at average union density, the implementation of a collective agreement is estimated to increase productivity in most model specifications. In the Norwegian context, collective agreements formally acknowledge the importance of workers' voices and their contributions to productivity growth by establishing a system of collaboration, communication, and participation. Collective agreements thus represent an institutionalization of a particular way of managing industrial relations. In the absence of this institution, union activity may be more poorly organized and less predictable. Similarly, it may be difficult to utilize the productivity-enhancing potential of collective agreements in the absence of union activity. Our results indicate that a sufficiently high union density and a collective agreement combined have a positive impact on firm-level productivity. However, care should be taken in interpreting our results as the possible endogenous decision to enter or exit a collective agreement remains a caveat in the study.

The three essays build up to a common conclusion: the presence of unions and collective action organized by unions leave traces in the labor market. Where unions are relatively strong, the findings suggest that they contribute to labor market outcomes that diverge from those in areas where unions have a weaker position. However, the fact that the entire Norwegian economy to some extent is affected by a common institutional context represents an important caveat when it comes to generalizing the results to other countries and settings. In more decentralized wage negotiation systems, it makes more sense to draw a clear distinction between the unionized and non-unionized parts of the economy. Still, as is apparent from the studies included in the thesis, the extent to which unions and collective agreements have coverage in the labor market in Norway varies greatly. Exploiting this variation provides interesting knowledge about the role of unions in shaping labor market outcomes.

Working with this project has provided some insights regarding the empirical operationalization of *union presence*. A main takeaway is that different measures of union presence may have very different implications depending on the labor market outcome of interest. The term "union presence" is by no means unambiguous or self-explanatory. The range of meanings attributed to union presence includes individual representation, bargaining strength (within firms, industries, or countries), formal recognition in the workplace (e.g., majority vote), the presence and/or coverage of collective agreements, employee voice through elected representatives, and different forms of codetermination and participation in management decisions. All these interpretations may be meaningful but should be discussed in each specific setting. Furthermore, the dimensions of unionization often interact in a way that is difficult to disentangle for the econometrician. Hence, econometric studies should be followed by qualitative research to get a better understanding of the mechanisms at play.

The title of the thesis raises an important and ambitious question. The findings presented here by no means fully answer the question of what unions do. Nevertheless, I believe that the findings do provide some new insights. As unions are in decline throughout the world, inequality is on the rise. National statutory minimum wages are currently being introduced across large parts of the EU. In many countries (e.g., the US), unions are commonly viewed solely as labor cartels, limiting job creation and destroying the economy. The evidence laid forth in this thesis suggests that unions may be important regulators of wage inequality, as well as contributing to enhance productivity. Consequently, trying to understand what unions could do as part of the web of society should be of interest to those making decisions to achieve political goals and social change. Although the results are not directly transferable to other contexts, some of the hypothesized mechanisms may be.

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Chapter 1

Wage inequality and union membership at the establishment level: An econometric study using Norwegian data

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OXFORD

Wage inequality and union membership at the establishment level: An econometric study using Norwegian data

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Abstract

We model the empirical relationship between the within-establishment union membership rate and wage inequality in Norway. The data set is a panel of 2,285 private-sector establishments observed in the period 2000–18. The statistical model represents joint feedback between wage inequality and union membership. Dynamic panel data models are estimated for different wage inequality measures, with gini as the reference measure. The results show a negative relationship between union membership and wage inequality, which is robust with respect to different inequality measures and estimation method. The strength of the relationship increases with the permanency of the shift in union membership. We find evidence that union membership is a more important explanatory variable for the lower part of the wage distribution than for the upper part. Furthermore, the impact of union density on wage inequality is conditional on the presence of a collective agreement.

JEL classifications: C22, C23, C26, C51, E02, E11, E24

1. Introduction

The ways in which wages depend on union presence is one of the longest studied topics in labour economics. There are several, somewhat ambiguous, ways in which unions can affect the wage distribution, with large variations depending on sample characteristics, time periods, macroeconomic conditions, and institutional context.

Since the work of Freeman (1980, 1982), it has become common to assume that unions reduce wage inequality through standardization of union members' wages. Even though the empirical evidence is mixed, the ability to compress wages is often referred to as one of the core functions of unions. According to Checchi *et al.* (2010), 'unions expound a philosophy of equality and advertise their actions as contributing to more fairness in opportunities and reward'. This may be particularly true in Norway, where the trade unions are known to share egalitarian values (Moene and Wallerstein, 2003, 2006; Dølvik and Visser, 2009). A highlighted feature in the so-called Norwegian and Nordic models of working life has been the ability of unions to reduce the need for government redistribution through a kind of pre-distribution negotiated directly by employers and workers. In turn, this pre-distribution also tends to equalize financial outcomes, creating less of a gap between the higher and lower earners in the economy (see e.g. Agell and Lommerud, 1993; Ahlquist, 2017).

Most existing studies consider how unions affect wage inequality across workplaces (e.g. comparing unionized and non-unionized workers). There is less econometrically based knowledge about how unions alter the within-component of wage inequality. An interesting question is therefore whether the presence of collective bargaining and unions' wage policies is detectable within establishments. An empirical study based on Norwegian data might be of particular interest, due to the two-tiered system of wage negotiations. In this article, we study the relationship between union membership and wage inequality within Norwegian private-sector establishments. We do so by utilizing a panel of 2,285 establishments observed in the period 2000–18. The panel has matched employer–employee data set, containing individual wage data merged with population-wide administrative register data. A particular feature of our data set is that all the establishments are present in the data for a relatively long time span. Union membership and wage inequality typically adjust gradually to changes in underlying institutional and individual determinants, and therefore require a dynamic modelling framework, and consequently sufficient within variation in the data.

Our approach is to combine conditional model equations for wage inequality and marginal model equations for the union membership rate. Though simple, this modelling strategy allows us to test for the existence of a relationship, both as a contemporaneous phenomenon and through feedback. We make use of standard panel data estimation methods, namely the Least Square Dummy Variable (LSDV) estimator and the Arellano and Bond General Method of Moments (GMM) estimator (Arellano and Bond, 1991). In an attempt to elucidate the direction of the relationship between wage inequality and union membership, we exploit the long time dimension in our data set to split the sample into two, and test for parameter constancy and invariance across 'regimes'.

The dynamic panel data models are estimated for different wage inequality measures, with gini as the reference measure. The results show a negative relationship between union membership and wage inequality which is robust with respect to the choice of measurement and estimation method. In particular, the evidence is based on changes of union membership within firms and measures potential effects of increasing membership on within-firm inequality conditional on existing inequality and existing membership (lagged values). We find that the strength of the relationship increases with the permanency of the shift in union membership. Furthermore, union membership seems to be a more important explanatory factor for the lower part of the wage distribution than for the upper part. Finally, the results show that the presence of a collective agreement is an important conditioning factor for our empirical results.

The article is organized as follows: In Section 2, we review a selection of important existing studies that have relevance for our research purpose. In Section 3, we discuss the Norwegian institutions that are most pertinent to our study. Section 4 presents the data set, and Section 5 gives a brief account of the econometric modelling framework. Section 6 presents the empirical results. Section 7 contains a summary and a brief discussion of the implications concerning the results for the role unions can play in a process aimed at both limiting government intervention and keeping inequality low.

2. Existing studies

How unions affect the wage distribution is likely to depend on several factors, such as collective agreement coverage, potential extension mechanisms, the structure and pattern of wage bargaining, the degree of bargaining coordination, and union density. Theoretically, the influence of unions on the dispersion of wages is therefore ambiguous. Empirically, a considerable number of studies have investigated the effects of unions and of labour market institutions on wage dispersion. Several studies have concluded that a high proportion of workers being members of a union is associated with lower wage inequality in their environments, that is across countries, industries, and establishments (e.g. Freeman, 1980, 1982; DiNardo *et al.*, 1996; DiNardo and Lemieux, 1997; Kahn, 1998, 2000, Jaumotte, 2003, Frandsen, 2012; Card *et al.*, 2020). The empirical literature has also showed that differences in the rate of de-unionization are correlated with differences in the growth of inequality (Card *et al.*, 2004; Dustmann *et al.*, 2014; Biewen and Seckler, 2019).

Two commonly used measures of union influence are union membership and collective agreement coverage. There are large cross-country differences in both the levels of these measures and the gap between the two. The gap is for instance larger in European countries than in the USA, Canada, and the UK. Differences between countries are of importance when evaluating the impact of unions on the wage structure (Visser and Checchi, 2009). In Canada and the USA, union representation and collective bargaining are regulated by the legal framework known as the 'Wagner Act' model. Within this framework, workers who meet the statutory definition of an employee have the right to union representation and collective bargaining. A Labour Relation Board works as an administrator for the procedures defining appropriate bargaining units and for certifying bargaining representatives (Card *et al.*, 2004). If a group of workers choose to be represented by a union, usually by majority vote, the union becomes the only bargaining representative of all employees in the particular bargaining unit, irrespective of union membership.

In contrast to the highly decentralized firm-by-firm bargaining in Anglo-Saxon countries, centralized bargaining between unions and groups of employers is the norm in many European countries. In some of these countries, collective agreements set legally binding minimum pay levels for all employers in an industry. In such cases, there may not be a clear relationship between union membership and collective agreement coverage. However, in countries such as Norway, industry-wide contracts are not necessarily binding for all employers, even if many employers have traditionally adhered to the wage provisions in the agreements. The implication is that collective agreement coverage is a relevant potential determinant of wage differences across establishments in Norway. However, within the workplace, the coverage of a collective agreement is extended to the non-union workers as well as union members in occupations covered by the agreement. Consequently, the potential impact of a collective agreement on wages does not discriminate between unionized and non-unionized workers within an establishment. This does not rule out that changes in the union membership rate have an impact on wage levels and wage inequality in the workplace. Whereas the presence of a collective agreement is closer to measuring the effectiveness of unions in providing and defending minimum standards of wages and employment protection, firm-level union density can be considered an indicator of potential union bargaining pressure (Visser, 2003). Empirical studies show that establishment-specific factors, such as union density, have an impact on individual wages in Norway, see Barth *et al.* (2000), Balsvik and Sæthre (2014), Bryson *et al.* (2020).

Even though they represent different dimensions of union presence, collective agreements and union density are more or less a function of one another in many European countries. This is also true in the case of Norway. Specifically, the union membership rate needs to exceed a certain level before the employees can demand a collective agreement at a workplace (usually somewhere between 10% and 50%, depending on the provisions of the particular union). When the purpose is to explain the role of unions in shaping the establishment-level wage distribution, it is therefore important to represent both dimensions. A few studies consider both elements when evaluating how unions affect wages at the level of the workplace. For example, Fitzenberger *et al.* (2013) find that in Germany, union density reinforces the effect of collective agreements when wage bargaining occurs on the local level, and tends to reduce the wage dispersion.

Most studies on union wage effects in Norway have examined the effect of union membership on wage levels. The estimates indicate a union wage premium of around 7% (Barth *et al.*, 2000; Balsvik and Sæthre, 2014). Barth *et al.* (2020) exploit tax-induced exogenous variance in the price of union membership to identify the causal effect of changes in firm union membership on firm productivity and wages over the period 2001–12. They find that both productivity and wage levels increase with union membership. The few Norwegian studies addressing how wage differentials are related to union presence indicate that unions contribute to a more compressed wage distribution. Barth *et al.* (2012) find that the introduction of performance-related pay increase wage inequality in non-union firms, but not in firms with high union density. Christensen (2019) investigates how collective agreements influence wage levels and wage dispersion in Norway from 1997 to 2012. Her results suggest that collective agreements decrease wage inequality by compressing the wage distribution at both ends.

The majority of studies assessing how unions alter the wage distribution, evaluate inter-firm wage inequality. In other words, they examine if the wages of unionized workers are more compressed than those of non-union workers (see Dell'Aringa and Lucifora, 1994; Hibbs Jr and Locking, 1996; Palenzuela and Jimeno, 1996; Checchi and Pagani, 2005; Dahl *et al.*, 2013). Recently, more attention has been drawn to the within-component of wage inequality. As within-workplace wage inequality constitutes a substantial part of the total increase in wage inequality in several countries (Fournier and Koske, 2013; ILO, 2016), Norway included, it is of relevance to further investigate if union presence has a role in shaping the wage distribution within firms. Some studies address the subject. Addison *et al.* (2014) show a modest widening of within-establishment wage dispersion for establishments that abandon sector-level collective bargaining in Germany. Cirillo *et al.* (2019) find that firm-level bargaining have heterogeneous effects across countries and time.

In most countries within the Organisation for Economic Co-operation and Development (OECD), there has been a development towards a more decentralized wage formation system over the last decades (Calmfors *et al.*, 2001). In the wake of this movement away from centralized bargaining, several studies have tried to uncover how the level of centralization shapes the wage distribution. In Denmark, Dahl *et al.* (2013) found that wages in Denmark were more dispersed under firm-level bargaining compared to more centralized wage-setting systems, caused by a higher wage premium for workers at the top of the wage distribution. Contrary to the Danish results, Andréasson (2014) found that decentralized and two-tiered bargaining in Sweden compressed the wage structure by awarding relatively higher wage premiums to low-wage earners in particular in decentralized regimes. The partly conflicting results for two Scandinavian countries, which from an onlooker's perspective appear to have quite similar social institutions, illustrate the need for country-specific studies in the field.

Another issue that has been analysed in the literature is whether the relationship between the union membership rate and wage inequality is interdependent. Herzer (2016) finds evidence of a two-way relationship between unionization and income inequality in a sample of 20 countries. Specifically, the results indicate that an increase in unionization on average reduces income inequality, but also that higher inequality leads to *lower* unionization rates. The findings are in line with those of Checchi *et al.* (2010), who show that the further an individual's earnings are from the median, the lower the estimated likelihood of their being a union member. The authors' interpretation is that trade unions primarily attract workers from the intermediate-earnings group. An implication of this finding may be that a secular increase in wage inequality leads to reduced union membership, because 'more and more workers find themselves further away from the median and perceive union action in this area as ineffective or contrary to their interests' (p. 101). However, another possible mechanism is that increasing inequality might cause workers to unionize because they feel that they are treated unfairly. Union members have been known to be more likely than other individuals to support redistribution (Finseraas, 2009).

We hope to contribute to the literature in several ways. First, we provide evidence on the role of union density in shaping the within-establishment wage distribution in Norway. We do this by exploiting a matched employer–employee panel of Norwegian establishments in operation over a relatively long span of time, enabling us to empirically investigate both short-run and more long-run relationships. We thus acknowledge that both wage inequality and union density change slowly over time, and consequently apply a dynamic modelling framework.

Secondly, we elaborate on the strength, direction, and interdependence of the within-establishment relationship between union density and wage inequality in Norway. We apply six separate measures of wage inequality in order to characterize how the wage distribution is shaped by the presence of strong unions. Furthermore, we draw on theory on super-exogeneity and invariance to infer about the direction of the relationship.

Finally, studies from different countries have a role to play in the understanding of how unions may operate in modern economies. The somewhat disparate (hard to reconcile) findings about the role of unions that operate in conjunction with different national institutions have proven the importance of this point. Norway may be of particular interest because of the two-tiered bargaining system, where the implications of local negotiations for wage inequality are likely to depend on the presence of both a collective agreement and union density. In this study, we specifically examine how the impact of union density on wage inequality depends on whether the establishment is part in a collective agreement.

3. Institutional framework for labour market regulation

The Norwegian system of labour market regulation has developed over a long period, going back to the industrialization of the Norwegian economy at the start of the 20th century. The system is a mixed one. Collective bargaining exists side by side with individualistic wage contracts, also within industries. A machinery for interest dispute resolution was established quite early. The 'peace obligation' in disputes of rights (in practice everything that is regulated by collective agreements) goes back to the Basic agreement from 1935. There has been a relatively low threshold for the use of compulsory arbitration.

The system of pattern wage bargaining is an important part of the wage formation at the national level. The Technical Calculation Committee was established in 1967 by a tripartite agreement and is vested with elaborating a common understanding about recent wage developments and about the forecast for cost of living, as well as other parameters of relevance for the upcoming agreement revisions, see Longva (1994). The state mediator has had a strong position, and the period of validity of agreements has become coordinated (2 years). At the establishment level, unions negotiate wage adjustment for their members each year, and the wider set of issues every second year. One defining trait of the Norwegian system is the limited reach of a collective agreement (Evju, 2014a). An agreement is only binding for the establishment that has negotiated it with the union. However, once an agreement is put into place, it applies to all employees belonging to the current category of profession or job description: union members or not. This application follows from the *principle of invariability*, which has developed over the last 100 years and which is based on case law.[1,2] There are, however, other benefits to joining a union than pure wage considerations. The benefits include representation in grievance procedures related to disputes over unfair or arbitrary treatment. There is also a (partial) tax deduction for the union fee.

Table 2 shows the collective agreement coverage rate in Norway. As can be expected, these rates are consistently higher than the worker organization rates (Table 1). However, in comparison with other western countries, the Norwegian bargaining coverage is not particularly high. The reason is that there are formal extension mechanisms in many countries.

There is a distinction between formal bargaining coverage, as measured in Table 2, and the effective bargaining coverage that results when employers without membership in a confederation choose to offer their workers compensation in line with the relevant collective agreement. It is a custom to assume that voluntary extension (adoption of a wage norm) has been a reality in Norway and to point at the historically long periods of near full employment after World War II as an underlying factor. It could have been rational for unorganized establishments to pay the going wage, as a way of avoiding cost-increasing wage bidding rounds.

However, the system of labour market regulation is not static. A relatively new element is The General Application Act (of Collective Agreements) of June 1993. Although it was far from a semi-automatic extension mechanism, and considering that it targeted social dumping, the act was contested by organizations on both sides of the bargain at the time. Its use has increased after 2007 and 2009, see Evju (2014a,b), possibly as a response to practical problems of maintaining collective bargaining as a main regulating mechanism in industries with many European Union (EU) labour immigrants.

Another dimension of the Norwegian private-sector bargaining system, significant to our analysis, is that it is two-tiered. In practice, a large part of the total wage regulation in any given year may be determined at the local level, a phenomenon known as wage drift, see Holden (1989), Moene *et al.* (1993). Local negotiations (collective and individual) has

1 Parts of the principle are also established in legislation (the Labour Dispute Act §6). The purpose of the principle is to ensure that wage differentials do not undermine the significance of collective agreements. Hence, what it implies for employees in covered workplaces in most cases is a binding wage floor.

2 Not all agreements have wage rate provisions, but most of the agreements without wage rates comprise occupations with relatively high wage levels.

Table 1. Organization densities in Norway in selected years

Year	Unionization rate (%)	Employer organization (%)
1948	50	
1972	51	
1990	57	50
2005	50	60
2013	49	65
2015	49	69
2018	49	71

Source: Stokke *et al.* (2013) and Nergaard (2018).

Table 2. Collective agreement coverage in Norway in selected years

Year	Private sector (%)	Production of goods (%)	Service (%)
1998	63	71	58
2004	60	63	58
2005	59	64	56
2008	59	65	55
2013	58	62	56
2017	52	56	51

Source: Nergaard (2018, Table 2.5).

played an increasingly important role in the Norwegian wage formation during the 1990s and 2000s (Dølvik *et al.*, 2018). This feature of the Norwegian wage formation system indicates that union membership (bargaining strength) may be one of the factors influencing establishment-level wage distributions. The observed wage will, however, always be the outcome of both collective bargaining and employers' unilateral choices.

One can speculate about the possibility of maintaining a system like Norway's, in which the confederate organizations play a major role, without establishment-level negotiations, at least as a supplement. A completely centralized collective agreement would also need to be implemented in the wage scale of the individual workplace. The central agreements determine only a base wage or a norm. The individual worker's actual wage compensation will be partly determined at the establishment level and it is easy to imagine that it can become influenced by both establishment-specific factors and by the local unions negotiating strength and preferences about low-pay 'profile'.

4. The data set

We make use of a matched employer–employee data set drawn from the administrative registers of Statistics Norway. Our primary data source for the period 2000–14 is Statistics Norway's wage statistics. For the remaining years, 2015–18, our data are collected from the 'a-ordning', a coordinated service used by employers to report information about income and employees to the Norwegian Labor and Welfare Administration, Statistics Norway, and the Norwegian Tax Administration. In the a-ordning, all establishments in the private sector are included. Before 2015, wage statistics were only collected for a sample of private-sector establishments.[3] However, all employees in the sampled establishments are included.

The individual wage data are primarily reported as monthly earnings.[4] In order to compare full-time and part-time workers, we have calculated an hourly wage based on the monthly wage and reported contractually agreed working hours. To minimize the impact of outliers on the calculation of the wage inequality measures, we have set a lower limit of 70 Norwegian Krone (NOK) and an upper limit of NOK 2000 (adjusted for inflation with the Consumer Price Index, base year 2015) on hourly wage.[5] The hourly wage was used to compute the wage inequality measures for each establishment. Our main wage inequality variable is the gini coefficient, which is commonly used to measure inequality within populations. The gini coefficient, sometimes referred to as the gini index or gini ratio, is a measure of statistical dispersion, derived from the Lorenz curve of cumulative income distribution (Gini, 1921). A gini coefficient of 0 corresponds to a 45-degree straight Lorenz curve and indicates perfect equality: i.e. everybody earns the same. A gini coefficient of 1 means that one individual has all the earnings. The gini is independent of the size of the population and it uses information from the entire wage distribution (Trapeznikova, 2019). One drawback of the gini is that it puts more weight on the observations in the middle of the distribution. Another weakness is that two establishments with the same gini may still have quite different wage distributions, and thus it does not provide much information about the type of wage inequality in each workplace.

To make sure our results are robust, and to examine which segments of the wage distribution that are most affected by the presence of unions, we include five other inequality measures in addition to the gini: The standard deviation of log wage (*sdl*), the coefficient of variation (*cv*), and three different relative wage-level measures. The sdl and the cv are alternatives to the gini as single-valued measurements of the entire wage distribution.

Unlike the measures of wage dispersion (gini, sdl, and cv), percentile ratios focus on specific segments of the wage distribution. We consider three such ratios: $p90/p10$, $p90/p50$, and $p50/p10$. In particular, the 'interdecile ratio' ($p90/p10$) shows the income level of individuals at the top of the income distribution (top 10%) relative to the income level of those at the bottom of the distribution (bottom 10%).

A focus variable in the study is the establishment-level union membership rate. Our data set contains information on whether a union membership fee is paid by each individual and reported to the tax authorities. Based on these payments, we calculate union density as the ratio of paying union members relative to the number of employees in each workplace.

Whether an establishment participates in a collective agreement or not is derived from membership data from the mutual arrangement for private sector collectively agreed pension scheme ('Fellesordningen for AFP'), in which all establishments who are members are also part in a collective agreement.

3 The selection method applied by Statistics Norway was based on stratified random, systematic cluster selection, where the stratification was made by enterprise size (number of employees) in each industry, with complete counting in the largest companies, and cut-off in the smallest. https://www.ssb.no/omssb/tjenester-og-verktoy/data-til-forskning/lonn/data_lonn.

4 Monthly earnings include basic monthly salary, variable additional allowances, and bonuses. Overtime pay is not included: //www.ssb.no/en/arbeid-og-lonn/wage-terms#Monthly_earnings

5 Around 5% of the observations are excluded due to this restriction. The results are, however, very robust to less strict trimming, see Table B3 in Supplementary Appendix.

Our data further include a rich set of individual/job characteristics and establishment/industry characteristics, see Table A1 in Supplementary Appendix.

We have made some restrictions to our sample. In order to apply a dynamic modelling framework, it is important that the establishments are present for a sufficient number of time periods. Therefore, to be included in the data set, an establishment cannot have more than 2 years of missing wage observations during the 19-year time period.[6] Furthermore, establishments are required to have been existing and in operation for at least 12 of the 19 years in our data sample period. This leaves us with a sample average T of about 16.5 years. To calculate a representative measure of wage inequality in each workplace and to reduce the influence of extreme values, we have left out establishments with less than 25 employees. Our final sample consists of 37,656 observations from 2,285 establishments during a 19-year period.[7] The establishments included in the analysis are representative of a wide scope of industries (see Table A4 in Supplementary Appendix).

5. The modelling framework

As noted above, the unionization rate at the national level changed moderately over the second half of the last century, and it has been relatively stable so far in the new millennium. Stability and gradual adjustment (i.e. dynamic) rather than instantaneous changes also seem to characterize the union membership rate at the establishment level in our data set. In the time domain, wage distributions have the same characteristics: although the gini can change considerably as a result of changes in the labour market and in wage setting institutions, the length of adjustment periods is usually longer time spans.

Hence, we use a dynamic modelling framework, exploiting the long time dimension of our data. In unrestricted form, the modelling framework treats *in* (inequality measure) and *um* (union membership rate) as endogenous variables. We present results for several operational definitions of inequality, but within the same statistical framework.

Let $\boldsymbol{y}_{it}$ denote the vector with in_{it} and um_{it} where i is the cross-section index (establishment) and the time index t (years). Following custom, we define $t = 1, \ldots, T$, $i = 1, \ldots.., n$. We let $\boldsymbol{x}_{it}$ denote a vector with non-modelled variables while $\boldsymbol{\varepsilon}_{it}$ denotes the vector with the error-terms. Their joint statistical distribution is conditional on $\boldsymbol{y}_{it-1}$ and $\boldsymbol{x}_{it}$. In order to save notation, and without loss of generality, we do not specify any lags of the x-variables, but lagged terms will be used in the empirical models.

A main decision to make in empirical modelling of an evolving system is the order of dynamics. Under-specification will typically make it impossible to maintain a model assumption about non-autocorrelated residuals, which is important for validity of the statistical model. However, the issue is more pressing with quarterly and monthly data than with annual data, and in the following we mainly use first-order dynamics as specified by

$$\boldsymbol{y}_{it} = \boldsymbol{\Phi}\boldsymbol{y}_{it-1} + \Gamma_i \boldsymbol{x}_{it} + \boldsymbol{\varepsilon}_{it}. \tag{1}$$

6 Due to the sampling method applied before 2015, not all establishments were included in the data set every year between 2000 and 2014, even if they were in operation.

7 Approximately 20% of the establishments are part of firms with more than one workplace. We have conducted the estimations in a sample excluding these establishments, and the results remain robust.

In the terminology used to characterize panel data model equations, (1) is specified with homogeneous parameters in the Φ matrix, while coefficients can vary between units in the Γ_i-matrix. The simplest example of such heterogeneity is that there are $n-1$ coefficients for the constant term (1_{it}) contained in $\boldsymbol{x}_{it}$.

We can now specify the requirements for 'no relationship' between wage inequality and union membership in the model (1). It is that the off-diagonal elements of ϕ are zero and that the covariance between the error terms in $\boldsymbol{\varepsilon}_{it}$ is zero.

A practical way of testing these hypotheses is to make use of conditional modelling. To simplify notation, assume that $\boldsymbol{x}_{it}$ only contains two elements: the constant term and a single random variable x_{it}. The system (1) can be re-expressed as the conditional model equation of in_t given um_t and the marginal model equation for um_t:

$$in_{it} = \beta_{11,1}in_{it-1} + \beta_{12,0}um_{it} + \beta_{12,1}um_{it-1} + \beta_{1x}x_{it} + \alpha_{1i} + \epsilon_{1it} \tag{2}$$

$$um_{it} = \phi_{21,1}in_{it-1} + \phi_{22,1}um_{it-1} + \phi_{2x}x_{it} + \gamma_{2i} + \varepsilon_{2it} \tag{3}$$

where the coefficient $\beta_{12,0}$ is the regression coefficient and the other coefficients in (2) are also parameters of the conditional expectation of in_{it} given um_{it}. α_{1i} and γ_{2i} denotes establishment fixed effects. ϵ_{1it} and ε_{2it} are error terms, assumed to be normally distributed and I.I.D.

Independence (no relationship between wage inequality and union membership as defined above) implies the following restrictions on (2)–(3):

1. $\beta_{12,0} = 0$,
2. $\beta_{12,1} = 0$,
3. $\phi_{21,1} = 0$.

If restrictions 1 and 2 can be rejected but restriction 3 cannot, changes in union membership affect wage inequality both contemporaneously and dynamically, while there is no effect from inequality back on union membership. However, if the third restriction also can be rejected, the relationship goes both ways. Specifically, if there is an autonomous increase in inequality in year t, the expected change in union membership the following year is measured by $\phi_{21,1}$.

Moreover, there are other second and third round effects. Hence, when we present the empirical results, the focus will not only be on the coefficients that capture the short-run relationships, but also on the long-term effects that are implied by the steady-state solution of the system.

Below, we estimate (2) and (3) using standard panel data methods, namely the LSDV estimator (the within estimator) and the GMM estimator for dynamic panel data models.

Ever since Hurwicz (1950), it has been known that the OLS estimator of an autoregressive model contains a finite sample bias, which, however, is small in magnitude unless the degree of persistence is high (close to unit root non-stationarity). The small sample bias problem carries over to the LSDV estimator applied to dynamic panel models in that it contains a Hurwicz-type bias even when N is very large, see Nickell (1988), Judson and Owen (1999).[8] Therefore, we also estimate our model by the use of a GMM estimator which instruments the pre-determined variables. The basic idea of the Arellano and Bond

8 The problem is more serious for the random-effects model, all least squares estimators will contain a bias that remains even when T is very large and approaches infinity.

estimator (Arellano and Bond, 1991; Arellano and Bover, 1995; Blundell and Bond, 1998), AB for short, is to obtain GMM instruments by utilizing the orthogonality conditions that exist between the lagged values of the dependent variable and the disturbances. This method removes the above-mentioned bias asymptotically (i.e. when N is infinite).

As with all methods of moments estimators of conditional models, there is a trade-off between theoretical large sample consistency and larger estimated coefficient standard errors (N is, after all, a finite number). Unknown finite sample bias due to weak instruments is generic. It has been pointed out that liberal use of GMM instruments, with some of them relatively weak, can bias the GMM estimators (Newey and Windmeijer, 2009). It is thus of importance to examine the robustness of the estimates with respect to the number of GMM instruments applied.

6. Results

In this section we present results for estimation of (2) and (3), using the LSDV (within) estimator and the AB (GMM) estimator for dynamic panel data models. Fixed effects estimation enables us to control for unobserved time-invariant individual heterogeneity between establishments, when assessing the relationship between the union membership rate and wage inequality. We also examine robustness and possible heterogeneity across different inequality measures, investigate the direction of the relationship through the exploitation of a structural break, and finally explore the relevance of collective agreements.

6.1 Main results

Our initial set of results is displayed in Table 3. We first note that the estimated coefficients that test the null of no-relationship in the conditional model equation in column (1) are significantly different from zero, both individually and jointly (the F-statistic is 14.25 which is significant at an arbitrary low level of significance). The sum of the coefficients is negative with a t-value of –4. Hence, in the model, a permanent change in union membership leads to short- and long-term reductions in wage inequality.

In column (2), we see that the lagged gini-coefficient is positive and statistically significant at the 5% level. Taken at face value, larger wage inequality in one year predicts a certain increase in union membership the next year.

The two-way dependency implies that the estimated long-term effect on inequality of an autonomous increase in membership is found from the steady-state solution of the system (1) and (2). The estimate turns out to be a reduction in gini by –0.005 for an increase in um by 10 percentage points. This may appear as numerically insignificant. However, a 0.005 change is in fact relatively large in our data set, given that the mean gini in the sample is 0.14, cf. Table A1 in Supplementary Appendix.

If the two-way dependency between um and in is ignored (i.e. look at column (1) in isolation), the estimated effect of the same change in um is smaller in magnitude (–0.004). Hence, the two-way dependency increases the estimated long-run coefficient of an autonomous change in union membership on wage inequality.

The AB estimation results in columns (3) and (4) in Table 3 show that the results are robust.[9] The coefficient of um_t in (3) and of um_{t-1} in (4) are significant at the 1% level.

9 As mentioned, a high number of GMM instrument may bias the GMM estimator. In our case, an average T of around 16 gives approximately 195 GMM instruments, which seems to be a high

Table 3. LSDV and AB estimation results for the parameters in (2) and (3)

	(1) LSDV: *gini*	(2) LSDV: *um*	(3) AB: *gini*	(4) AB: *um*
um_t	–0.0206***		–0.0241***	
	(–4.69)		(–3.96)	
$gini_{t-1}$	0.378***	0.0349*	0.351***	0.116***
	(39.93)	(2.18)	(18.30)	(4.52)
um_{t-1}	0.00772*	0.597***	0.00634	0.524***
	(2.06)	(56.38)	(1.22)	(18.92)
Sum of *um*-coefficients	–0.0128**		–0.0178***	
	(–3.99)		(–2.71)	
R^2	0.328	0.489		
N	32,951	32,951	29,464	29,464
Establishments	2,275	2,275	2,268	2,268
Avg. obs.	14.48	14.48	12.99	12.99
IS-test, order 1	176.84***	86.58***		
IS-test, order 2	357.19***	219.67***		
Arellano–Bond test for AR(1)			–20.314***	–17.191***
Arellano–Bond test for AR(2)			3.6184***	3.9085***

Note: All models include establishment-level educational shares, establishment-level age shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories, and time dummies. For models (1) and (2), full estimation table is available in the Supplementary Appendix. The IS-test refers to the Inoue–Solon test for serial correlation. Robust standard errors clustered at establishment level, *t*-statistics in parenthesis.
*$p < 0.05$;
**$p < 0.01$;
***$p < 0.001$.
Source: Authors' calculations.

Hence the AB-results support joint dependency between gini and the union membership rate. The point estimates are practically the same as in column (1). One difference is that the lagged membership variable is insignificant in the inequality equation (5), which entails a more negative short-run relationship between membership and the gini. However, the estimated long-run effect changes only a little. Using the AB estimator (GMM), the estimated long-run effect on gini of a change in *um* of 0.1 turns out to be –0.007, as opposed to –0.005 for the LSDV estimated model.

In contrast to the findings of Herzer (2016) and Checchi *et al.* (2010), our results show that an increase in gini gives rise to an *increase* in the union membership rate within establishments. The mechanism suggested by the authors, namely that union members end their trade union membership/do not become members if inequality increases, is therefore not supported by our findings. One possible explanation as to why this may be is that local wage inequality stimulates mobilization and recruitment in the workplace. Furthermore, discontent about local wage inequality may be more easily directed against the local union in decentralized bargaining regimes than in the Norwegian system, where a part of the

number. We have therefore re-estimated the model with a limited number of lags, but this does not notably change the estimates.

wage formation is centralized, and also coordinated through pattern bargaining. The threshold for ending the union membership following an increase in the establishment-level wage inequality may be higher in Norway than in countries with more decentralized bargaining regimes. Hence, increasing wage differentials in the workplace may induce discontent among employees, but in a way that motivates them to join the union as opposed to leaving it.

6.2 Different wage inequality measures

In this section, we look at alternative operational measures of wage inequality, using the same estimation method (LSDV). The results are shown in Table 4.

We see that the regression coefficients $\beta_{12,0}$ of *um* are estimated with negative signs that are statistically significant for all six measures of wage inequality. Note that the results in column (4) are for the gini and are therefore identical to the results of the previous section. The other key coefficients of the system are also robust across the different measurements of wage inequality. In particular, we note that the coefficient of um_{it} is larger in magnitude than the coefficient of um_{it-1} in the conditional equations. Hence, there are no examples of changed signs between the short-run and long-run relationships. They are negative across all the six models with different measures of wage inequality.

The results in Table 4 (bottom part of the table) confirm that also the feedback effect from inequality to um_t is robust across the different measurements of inequality. If we imagine a permanent (autonomous) increase in union membership rate by 10 percentage points, we see that it is associated with a reduction in $p90/p10$ by 0.0195 in the first year (column (1) in the table). The estimated long-run effect of the hypothetical change, which takes into account the two-way feedback mechanisms, is, however, larger in magnitude: –0.046. An interpretation of the negative relationship can involve at least three mechanisms: (i) a higher membership rate reduces the difference between high-wage earners and middle wage earners within the establishment since there is a tendency that the middle percentiles are more saturated with union members than the upper percentiles; (ii) that mechanism is strengthened by the institutional arrangement that a collective agreement implies: equal pay for identical work for union and non-union workers; and/or (iii) wage policies by the union, aimed at delivering a notable wage premium at the lower end of the distribution.

It is interesting therefore that the hypothetical change in membership gives different results when we measure it by $p90$ relative to the median (column (2)) compared to what we obtain when we consider the median relative to $p10$ (column (3)). For $p90/p50$, the short- and long-term coefficients are $(-0.05; -0.08)$ while they are $(-0.08; -0.13)$ for $p50/p10$. Hence, there is an indication that the lower end of the wage distribution is more influenced by changes in the union membership rate than the upper-half.

The fact that the upper part of the distribution is affected by a hypothetical autonomous change in union membership is consistent with mechanisms (i) and (ii). However, that the lower half of the distribution appears to be even more affected, indicates the mechanism (iii) plays a significant role as well. The statistical significance of the focus parameters extends to the estimates in columns (5) and (6) of Table 4, for the standard deviation of log hourly wage (*sdl*) and the coefficient of variation (*cv*).

Table 4. LSDV for the parameters in (2) and (3), for different operational measures of wage inequality

	(1) $p90/p10$	(2) $p90/p50$	(3) $p50/p10$	(4) *gini*	(5) *sdl*	(6) *cv*
um_t	–0.195***	–0.0517*	–0.0861***	–0.0206***	–0.0469***	–0.0513***
	(–4.40)	(–2.27)	(–4.22)	(–4.69)	(–6.85)	(–5.57)
$(p90/p10)_{t-1}$	0.282***					
	(10.41)					
$(p90/p50)_{t-1}$		0.351***				
		(18.53)				
$(p50/p10)_{t-1}$			0.235***			
			(15.98)			
$gini_{t-1}$				0.378***		
				(39.93)		
sdl_{t-1}					0.336***	
					(37.52)	
cv_{t-1}						0.426***
						(19.08)
um_{t-1}	0.0629	0.00498	0.0393*	0.00772*	0.0199***	0.0267**
	(1.61)	(0.25)	(2.20)	(2.06)	(3.35)	(3.25)
R^2	0.265	0.194	0.250	0.328	0.333	0.282
N	32,951	32,951	32,951	32,951	32,951	32,951
Establishments	2,275	2,275	2,275	2,275	2,275	2,275
Avg. obs.	14.48	14.48	14.48	14.48	14.46	14.46

	(1) *um*	(2) *um*	(3) *um*	(4) *um*	(5) *um*	(6) *um*
$(p90/p10)_{t-1}$	0.00361**					
	(2.69)					
$(p90/p50)_{t-1}$		0.00557*				
		(2.24)				
$(p50/p10)_{t-1}$			0.00527			
			(1.70)			
$gini_{t-1}$				0.0349*		
				(2.18)		
sdl_{t-1}					0.0207*	
					(2.31)	
cv_{t-1}						0.0104*
						(2.04)
um_{t-1}	0.597***	0.597***	0.597***	0.597***	0.608***	0.608***
	(56.54)	(56.43)	(56.53)	(56.38)	(62.01)	(61.90)
R^2	0.437	0.442	0.441	0.489	0.437	0.437
N	32,950	32,951	32,950	32,951	32,924	32,924
Establishments	2,275	2,275	2,275	2,275	2,275	2,275
Avg. obs.	14.48	14.48	14.48	14.48	14.47	14.47

Notes: All models include establishment-level educational shares, establishment-level age shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories, and time dummies. Robust standard errors clustered at establishment level, *t*-statistics in parenthesis.

*$p < 0.05$;

**$p < 0.01$;

***$p < 0.001$.

Source: Authors' calculations.

6.3 Reversed regression and invariance

The joint distribution of in_{it} and um_{it} in (1) can alternatively be put on model form with a conditional equation of um_{it} given in_{it} and a marginal model equation for in_{it}. This is the case of reversed or inverted regression.

It is known from the theory of super-exogeneity and invariance that structural breaks in the joint distribution of two variables may represent information that can aid the discrimination between the two regression directions, see (Nymoen, 2019, Ch. 1.8, 8.5). The basic idea is that a conditional model that has relatively constant parameters with respect to structural breaks elsewhere in the system represents structure, and can be used to estimate effects of changes (i.e. policy analysis), (Engle and Hendry, 1993). In practice, feasible tests of this form of parameter constancy, known as invariance, are done with respect to structural breaks in the marginal model.

Heuristically, testing for invariance can be done with reference to the non-invertibility of stable conditional models under regime shift. In our case, letting σ^2_{in} and σ^2_{um} denote the variances of the two error terms of the reduced form system (1), we have the following relationship between the two regression coefficients:

$$\beta_{12,0} = \beta'_{21,0}\frac{\sigma^2_{in}}{\sigma^2_{um}}. \tag{4}$$

$\beta_{12,0}$ is the partial regression coefficient of um_t in the conditional model of in_t (i.e. as in (2)), while $\beta'_{21,0}$ is the coefficient of in_t in the inverted regression. Hence if there is a structural break in for example σ^2_{in} and $\beta_{12,0}$ is constant, $\beta'_{21,0}$ cannot be constant. And vice versa: if $\beta'_{21,0}$ is stable, $\beta_{12,0}$ cannot be invariant to the structural break.

This argument demonstrates that if the marginal models exhibit enough change, at least one of the 'directions of regression' can be ruled out on non-constancy grounds. Hence if only one of the regression directions provides evidence of stability and invariance, we have empirical support for the hypothesis that the relationship also runs in that direction.

We have estimated the two conditional model equations with data from two subsamples: 2000–7 (regime 1) and 2008–18 (regime 2).[10] The sample-split is relevant for testing for a structural break since the first sub-sample was a period of relative stability in labour market institutions (regime 1), while in the second sub-sample the potential for disruption that followed after EU labour market enlargement began to be noticeable (regime 2), see e.g. Evju (2014a) and the references therein. The financial crisis also placed new strains on industrial relations, although it did not develop into the same job crisis in Norway as it did internationally.

We look at the reference case, where wage inequality is measured by the *gini*. The results are summarized in Table 5. We see that both coefficients are reduced numerically (they are closer to zero) in regime 2 compared to regime 1. However, there is a higher degree of stability in the estimated $\beta_{12,0}$ than in the estimated coefficient of the inverted model. There is also a notable difference in how the associated confidence intervals change between the two regimes. The two intervals for $\beta_{12,0}$ overlap a great deal, whereas there is no overlap between the two confidence intervals for $\beta'_{21,0}$.

Although informal, the outcome of the tests supports the interpretation that the conditional model of the gini is more invariant with respect to the regimes shift(s) between the

10 The conditional models are estimated by LSDV (within) estimation in the reported results. We have also applied the AB-estimator, and the results are robust.

Table 5. Split sample: coefficients and confidence intervals from conditional models of gini and union membership in two subsamples (LSDV estimation)

Time period	*um* → *gini*	95% CI	*gini* → *um*	95% CI
2000–7	−0.0326 (−5.82)	[−0.04360, −0.02163]	−0.2859 (−6.68)	[−0.38136, −0.19061]
2008–18	−0.0141 (−2.22)	[−0.02658, −0.00165]	−0.0640 (−1.84)	[−0.12039, −0.00777]

Notes: All models include establishment-level educational shares, establishment-level age shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories, and time dummies. Robust standard errors clustered at establishment level, *t*-statistics reported in parenthesis.
Source: Authors' calculations.

samples than what can be said of the alternative (inverted) regression. The results are in line with the hypothesis that a change in union membership will have a change in wage inequality as a consequence.

6.4 The relevance of collective agreements

As discussed in Sections 2 and 3, the importance of establishment-level union density is likely to depend on the presence of a collective agreement. In particular, collective agreements may act as a formal recognition of the unions' right to bargain over wages and to get their wage policies implemented in the workplace. Union membership in uncovered establishments may thus be motivated by other reasons than the preference for redistribution or a relatively higher wage floor. As mentioned, there are other benefits to joining a union than pure wage considerations, such as access to help solving problems that arise in the workplace. Furthermore, not all trade unions have a stated preference for redistribution. In particular, some of the trade unions organizing white collar workers (e.g. MBA candidates and lawyers) in Norway have more individually oriented wage policies. Hence, if the workplace is not covered by a collective agreement, a high union membership rate is less likely to be reflected in smaller wage differentials.

In the following, we estimate our model separately for covered and uncovered establishments.[11] In addition to the gini, we also show results for $p90/p50$ and $p50/p10$, keeping in mind that the 'union wage bite' may assert itself to a greater extent in the bottom share of the wage distribution, in line with the estimates in Table 4. The results from the estimations are shown in Table 6.

The results support the interpretation that the impact of union membership on wage inequality is conditioned by the presence of a collective agreement (ca). The estimated *um*-coefficients are not statistically significant for the no-ca models, and the magnitudes of the dependency coefficients are also smaller than in the with-ca models. Specifically, when inequality is measured by the gini, the with-ca results (1a columns) are very close to the full sample results in Table 4 (column (4)), while the no-ca model coefficients (1b columns) give no statistical support for a relationship between the membership variable and gini.

11 Another way of assessing the impact of collective agreements would be to include the collective-agreement dummy in the within estimations. However, this approach would require more within-variation in coverage-status than what is observed in our data.

Table 6. LSDV for the parameters in (2) and (3), for establishments with collective agreement (ca) and without collective agreements (no ca)

	(1a) *gini*, ca	(1b) *gini*, no ca	(2a) $p90/p50$, ca	(2b) $p90/p50$, no ca	(3a) $p50/p10$, ca	(3b) $p50/p10$, no ca
um	−0.0257***	−0.00843	−0.0793**	−0.0267	−0.0921***	−0.0519
	(−5.16)	(−0.98)	(−2.98)	(−0.60)	(−3.83)	(−1.32)
$gini_{t-1}$	0.360***	0.375***				
	(33.03)	(18.69)				
$p90/p50_{t-1}$			0.315***	0.369***		
			(22.15)	(8.03)		
$p50/p10_{t-1}$					0.239***	0.178***
					(16.65)	(6.30)
um_{t-1}	0.00580	0.0109	−0.0180	0.0652	0.0439*	−0.00264
	(1.36)	(1.56)	(−0.78)	(1.49)	(2.11)	(−0.08)
R^2	0.550	0.586	0.426	0.473	0.342	0.339
N	24,815	8,756	24,815	8,756	24,814	8,756
Establishments	1,825	844	1,825	844	1,825	844
Avg. obs.	13.60	10.37	13.60	10.37	13.60	10.37

	(1a) *um*, ca	(1 b) *um*, no ca	(2a) *um*, ca	(2 b) *um*, no ca	(3a) *um*, ca	(3 b) *um*, no ca
$gini_{t-1}$	0.0196	0.0161				
	(0.98)	(0.67)				
$p9050_{t-1}$			0.00162	0.00885*		
			(0.49)	(2.35)		
$p50/10_{t-1}$					0.00286	0.00189
					(0.72)	(0.44)
um_{t-1}	0.553***	0.547***	0.553***	0.548***	0.553***	0.547***
	(44.79)	(17.55)	(44.93)	(17.50)	(45.11)	(17.55)
R^2	0.835	0.940	0.835	0.940	0.835	0.940
N	24,815	8,756	24,815	8,756	24,814	8,756
Establishments	1,825	844	1,825	844	1,825	844
Avg. obs.	13.60	10.37	13.60	10.37	13.60	10.37

Notes: All models include establishment-level educational shares, establishment-level age shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories, and time dummies. Robust standard errors clustered at establishment level, t-statistics reported in parenthesis.

*$p < 0.05$;

**$p < 0.01$;

***$p < 0.001$.

Source: Authors' calculations.

The results for $p50/p10$ (3b columns) show coefficients for the no-ca model that are of the same magnitude as in the model of establishments with a collective agreement (3a columns). The pair of short- and long-term coefficients are estimated to be $(-0.05; -0.15)$ while the corresponding pair from the collective agreement model is $(-0.09; -0.14)$. However, only the second pair is based on coefficients that are individually statistically significant.

We note that the collective agreement coverage rate in our sample is higher than for the private sector in total (1,825 of the 2,285 establishments in our sample are covered). As is seen from Table A4 in the Supplementary Appendix, a large share of the establishments in our sample are placed within the manufacturing sector, where coverage tends to be higher than average. Further, both establishment size and period of existence are positively correlated with collective agreement coverage in Norway. All the establishments in our sample are both relatively large and long-lived.

With these remarks in mind, care must be taken in the generalization of our results. There are good reasons to assume that our sample is more representative of workplaces with a higher likelihood of being 'unionized' (i.e. having a collective agreement in place and/or a significant proportion of the employees unionized) than the Norwegian private sector as a whole. However, the results may be indicative of the joint role collective agreements and union density can play in relation to wage outcomes within establishments more generally.

7. Summary and concluding remarks

In this study, we have modelled the empirical relationship between the union membership rate and wage inequality in Norwegian private-sector establishments. We have analysed a panel of 2,285 establishments in the period 2000–18. The econometric framework treats wage inequality and union membership as two endogenous variables determined in a system, allowing us to empirically investigate various aspects of interdependence.

We have used standard panel estimation methods in order to quantify the models. Our estimation strategies are complementary and elucidate different aspects of the empirical relationship between wage inequality union membership. The operational definition of wage inequality used as a reference has been the gini coefficient. However, all of the models have also been estimated using alternative inequality measures.

We have estimated model equations of wage inequality conditional on union membership, and systematically completed the models with the marginal model equations for union membership. Although these models have no direct causality interpretation for the contemporaneous regression coefficient, they provide interpretable results for the dynamic relationships. The results suggest that higher union membership moderately reduces within-firm inequality. The estimation is based on changes in union membership within firms and the results represent potential consequences of increasing membership on within-firm inequality.

The magnitude of the estimated reduction in wage inequality is numerically significant, although not huge. The choice of operational definition for wage inequality plays a role. For example, it appears that the redistributive impact of unions may be stronger in the lower part of the wage distribution than in the upper part. This is an interesting aspect to note, which supports the idea that strong unions provide a form of protection against relatively low wages. We have used linear functional forms in our estimations, and an interesting aim in future work could be to test whether tipping points can be estimated, applying relevant functional forms.

Our empirical model captures that union membership may instantaneously increase union bargaining power while it is unlikely that workers directly join unions if inequality is very high in the same period. By making use of the relatively long time dimension of our data set, we have introduced the idea (from the econometric exogeneity literature) that the degree of constancy of conditional models can support a specific direction of a relationship. The result of this test, which utilizes a sample split associated with regime shifts, supports the interpretation that the within-year relationship direction runs deepest from union membership to wage inequality. In other words, the recursive interpretation of the system is confirmed empirically by the result of the structural break test.

Finally, we have examined how the impact of union density on within-establishment wage inequality depends on the presence of collective agreements. The results show that the wage compressing impact of union density on wage inequality is conditioned on the presence of a collective agreement.

In summary, our results represent empirical evidence that unions exert a negative impact on wage inequality within establishments. A highlighted feature in the so-called Norwegian and Nordic models has been the ability of unions to reduce the need for government redistribution through a kind of pre-distribution negotiated directly by employers and workers. In turn, this pre-distribution also tends to equalize financial outcomes, creating less of a gap between the higher and lower earners. Our findings support the view that unions contribute to lower inequality through the compression of within-establishment wage distributions in the modern Norwegian economy. A wider implication of our results is therefore that a decline in union membership could be a concern for policy makers who want to keep wage inequality low without increasing government intervention and regulation.

Supplementary Material

Supplementary material is available online at the OUP website. The supplementary material comprises an Online Appendix, a note on data availability, program and replication, as well as replication files. The main data used in the analysis are not provided, as restrictions apply to the availability of the data which were used under license for this study. Researchers affiliated with an approved research institution or a public authority can apply to data from Statistics Norway (https://www.ssb.no/en/data-til-forskning/utlan-av-data-til-forskere).

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Appendices

Wage inequality and union membership at the establishment-level. An econometric study using Norwegian data.

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This document contains data description and additional estimation results for Svarstad and Nymoen (2021).

A Data description

Figure A1 is a bin scatter, plotting the distribution of establishment level gini coefficients across union densities in our sample.

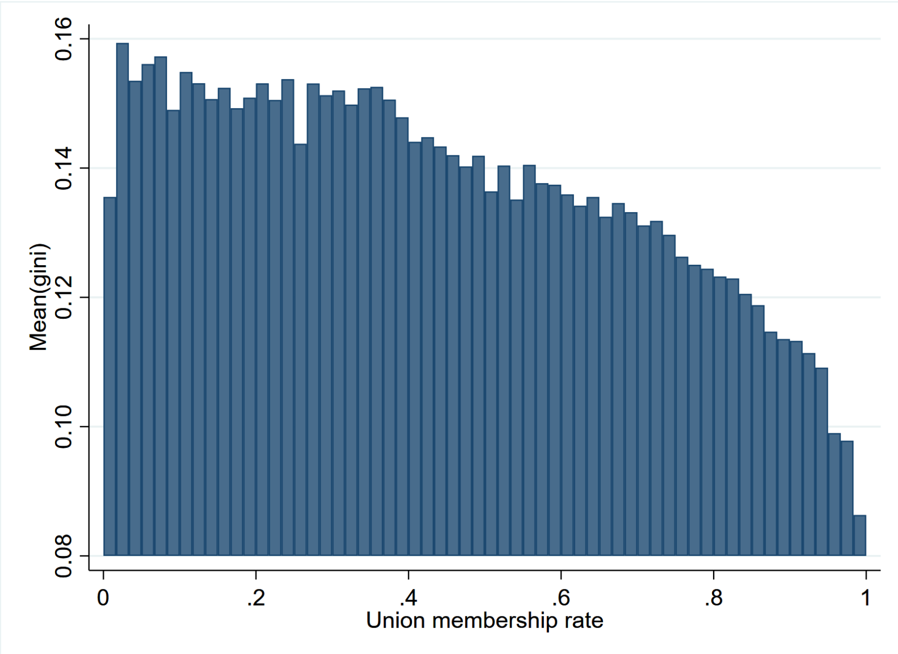

Figure A1: Distribution of mean gini coefficients over union membership rate. Bin scatter. Time period: 2000-2018. N=37 656. Source: Authors' calculations.

A.1 Between and within variation

The subject of the study has been the relationship between wage inequality and union density *within* establishments. Accordingly, within estimators were applied to all the model equations in the article. However, there is a good deal of variation between workplaces in our data as well. This is shown in Table A6 where we have decomposed the variation in our six inequality measures, as well as union density, into a between and within component. In Appendix B (Table B4), we show robustness of our main results with respect to estimation of random effects models, and hence utilization of between variation.

Table A1: Descriptive statistics.

	Mean	SD
gini	0.14	0.051
sdl	0.25	0.085
cv	0.29	0.12
p90/p10	1.83	0.51
p90/p50	1.43	0.26
p50/p10	1.27	0.21
Union membership rate	0.513	0.278
Part of collective agreement (dummy)	0.73	0.44
Immigrant share	0.13	0.16
Share of women	0.31	0.24
Share with primary school education*	0.19	0.13
Share with high school education	0.51	0.19
Share with high education, lower degree	0.227	0.15
Share with high education, higher degree	0.073	0.14
Share of employees aged 20-24	0.095	0.12
Share of employees aged 25-29	0.12	0.085
Share of employees aged 30-34	0.13	0.069
Share of employees aged 35-39	0.13	0.064
Share of employees aged 40-44	0.13	0.062
Share of employees aged 45-49	0.12	0.063
Share of employees aged 50-54	0.11	0.064
Share of employees aged 55-59	0.092	0.065
Share of employees aged 60-66	0.073	0.062
High skill occupation share**	0.27	0.32
Medium skill occupation share	0.21	0.29
Low skill occupation share	0.28	0.35
25-50 employees (dummy)	0.35	0.48
51-75 employees (dummy)	0.20	0.40
76-100 employees(dummy)	0.12	0.32
100 or more employees (dummy)	0.33	0.47
Share of part-time workers	0.13	0.21

*2-digit NUS2000 codes, translatable to ISCED97. **1-digit ISCO-08. Source: Authors' calculations.

Table A2: Yearly mean establishment-level gini coefficient.

Year	Mean establishment gini	Observations
2000	0.1286	1164
2001	0.1278	1367
2002	0.1296	1601
2003	0.1290	1767
2004	0.1311	1884
2005	0.1321	2040
2006	0.1354	2099
2007	0.1351	2112
2008	0.1365	2135
2009	0.1349	2195
2010	0.1363	2183
2011	0.1377	2149
2012	0.1374	2132
2013	0.1379	2078
2014	0.1364	2005
2015	0.1592	2230
2016	0.1533	2211
2017	0.1502	2191
2018	0.1261	2113
N		37 656

Source: Authors' calculations.

Table A3: Yearly mean establishment-level union membership rate.

Year	Mean establishment union membership rate	Observations
2000	0.5178	1164
2001	0.5193	1367
2002	0.4882	1601
2003	0.5083	1767
2004	0.5131	1884
2005	0.5170	2040
2006	0.5177	2099
2007	0.5153	2112
2008	0.5153	2135
2009	0.5143	2195
2010	0.5188	2183
2011	0.5139	2149
2012	0.5163	2132
2013	0.5059	2078
2014	0.5056	2005
2015	0.5114	2230
2016	0.5203	2211
2017	0.5188	2191
2018	0.5086	2113
N		37 656

Source: Authors' calculations

Table A4: Industry distribution. Standard Industrial Classification 2007 (SIC 2007).

Agriculture, forestry and fishing	119
Mining and quarrying	1074
Manufacturing	8429
Electricity, gas, steam and air conditioning supply	461
Water supply; sewerage, waste management and remediation activities	296
Construction	4816
Wholesale and retail trade; repair of motor vehicles and motorcycles	7389
Transportation and storage	3283
Accommodation and food service activities	2625
Information and communication	1893
Financial and insurance activities	19
Real estate activities	144
Professional, scientific and technical activities	2095
Administrative and support service activities	3309
Education	94
Human health and social work activities	891
Arts, entertainment and recreation	593
Other service activities	326
N	37 656

Source: Authors' calculations

Table A5: Mean establishment level union density and collective agreement coverage rate, by industry.

Industry	Mean UD	Mean CA
Agriculture, forestry and fishing	0.374	0.697
Mining and quarrying	0.665	0.801
Manufacturing	0.673	0.949
Electricity, gas, steam and air conditioning supply	0.794	0.466
Water supply; sewerage, waste management and remediation activities	0.429	0.781
Construction	0.559	0.846
Wholesale and retail trade; repair of motor vehicles and motorcycles	0.337	0.738
Transportation and storage	0.639	0.598
Accommodation and food service activities	0.329	0.790
Information and communication	0.562	0.621
Financial and insurance activities	0.578	0.789
Real estate activities	0.387	0.632
Professional, scientific and technical activities	0.427	0.306
Administrative and support service activities	0.400	0.560
Education	0.420	0.148
Human health and social work activities	0.570	0.451
Arts, entertainment and recreation	0.477	0.524
Other service activities	0.475	0.680

*2-digit NUS2000 codes, translatable to ISCED97. **1-digit ISCO-08. Source: Authors' calculations

Table A6: Decomposition of variation in wage inequality and union density in the sample.

Variable		Mean	SD	Min	Max	Observations	
gini	overall	0.137	0.051	0.000	0.602	*N*	37 656
	between		0.038	0.025	0.390	Establishments	2285
	within		0.033	-0.085	0.463	Avg. Obs.	16.48
sdl	overall	0.255	0.085	0.000	1.174	*N*	37 614
	between		0.062	0.050	0.626	Establishments	2285
	within		0.059	-0.077	1.035	Avg. Obs.	16.46
cov	overall	0.292	0.119	0.000	2.446	*N*	37 614
	between		0.083	0.050	1.033	Establishments	2285
	within		0.085	-0.362	2.111	Avg. Obs.	16.46
$p90/p10$	overall	1.833	0.505	1.000	20.567	*N*	37 654
	between		0.351	1.146	5.269	Establishments	2285
	within		0.364	-1.200	18.998	Avg. Obs.	16.48
$p90/p50$	overall	1.433	0.257	1.000	7.077	*N*	37 656
	between		0.184	1.067	3.459	Establishments	2285
	within		0.180	-0.327	5.052	Avg. Obs.	16.48
$p50/p10$	overall	1.272	0.209	1.000	6.700	*N*	37 654
	between		0.130	1.021	2.259	Establishments	2285
	within		0.164	0.492	6.223	Avg. Obs.	16.48
um	overall	0.513	0.277	0	1	*N*	37 640
	between		0.263	0.012	0.99	Establishments	2285
	within		0.089	-0.319	1.315	Avg. Obs.	16.48

Notes: Calculated using - *xtsum* - in Stata. Source: Authors' calculations.

B Additional estimation results and robustness

B.1 Data break in 2015

There is a break in the data in 2015 when the "a-ordning" was put into place. The a-ordning is a coordinated service used by employers to report information about income and employees to the Norwegian Labour and Welfare Administration (NAV), Statistics Norway and the Norwegian Tax Administration. The new service involved the inclusion in the wage statistics of the entire population of wage earners in the private sector, as opposed to the representative sampling method previously used.[1] Because we follow the same establishments in our analyses, the changes to the sample size are of minor importance. However, the new reporting system also involved a new system for establishments to report working hours. The transition is likely to have lead to measurement errors in the working time variable, which has consequences for how the hourly wage evolves after 2014. It is partly for this reason that we leave out of the analysis hourly wage observations lower than NOK 70 and higher than NOK 2000 (CPI-adjusted, with base year 2015). Even with this restriction however, we are not able to completely avoid the discontinuity in the data.

The drop in the estimated $\beta_{12,0}$ in Table 5 in section 6.3 may in part be explained by the break in the data series. A closer inspection of the yearly mean establishment gini (see Table A2) reveals a significant rise in 2015, which is likely to drive the observed reduction in the *um* coefficient. We have re-estimated the model for the period 2000-2014 (before the data break, see Table B1). The results are comparable to those in Table 3 in the main article. This strengthens the credibility of the results of the structural break test in Table 5.

[1]Until 2014, Statistics Norway only collected wage statistics for a sample of private-sector establishments. The selection method applied by Statistics Norway was based on stratified random, systematic cluster selection, where the stratification was made by enterprise size (number of employees) in each industry, with complete counting in the largest companies, and cut-off in the smallest. `https://www.ssb.no/omssb/tjenester-og-verktoy/data-til-forskning/lonn/data_lonn`

Table B1: LSDV estimation of the parameters in (2) and (3) for the time period 2000-2014.

	(1) *gini*	(2) *um*
um	-0.0355***	
	(-8.22)	
$gini_{t-1}$	0.378***	0.0332
	(30.05)	(1.47)
um_{t-1}	0.0109**	0.528***
	(2.94)	(42.67)
R^2	0.210	0.367
N	25 157	25 157
Establishments	2285	2285
Avg. obs.	11.01	11.01

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: All models include establishment-level educational shares, age bin shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories and time dummies. Robust standard errors clustered at establishment level, t-statistics in parenthesis. Source: Authors' calculations.

B.2 Sample restrictions

To minimize the impact of outliers on the calculation of the wage inequality measures in the study, we have set a lower limit of NOK 70 and an upper limit of NOK 2000 (CPI-adjusted with base year 2015) on hourly wage. 5 percent of the individual wage observations were excluded following this restriction. Table B3 shows estimation results for the *um*-coefficient in (2) with alternative trimming strategies. Model (1) is the reference model, corresponding to Table 3 in the main article. Model (2) Winzorices 1% of hourly wages, while model (3) is estimated on a sample without trimming.

The sample used in the study have further been 'curtailed' by exclusion of workplaces that have been present less that less than 12 years, have more than two gaps, or have less than 25 employees. Since this not random it entails that the sample is not statistically representative for the whole population of private sector establishments. Also, since there was an inflow and outflow of establishments during the period, the results cannot be extrapolated to all private sector workplaces in Norway in that period. Table B5 show estimation results for the parameters in equation (2) and (3) without the mentioned restrictions. Specifically, the estimations are done on all private sector establishments present in the data set (model 1), and on establishments with at least 25 employees (model 2).

B.3 Higher order dynamics

The residual diagnostics show signs of significant auto-correlation. This is apparent in Table 3, where the Inoue-Solon test (denoted IS) for serially-correlated residuals in the within model is highly significant for both first- and second-order serial correlation. Similarly, the AR(1) and AR(2) tests indicate the same conclusion when applying the Arellano and Bond, AB, estimator. Although auto-correlation and heteroscedasticity consistent standard errors were used in the results reported above, it is still possible that the results change markedly when additional lags are included in the model equations.

We have therefore re-estimated the conditional model using three lags of *gini* in the empirical version of (2). This removes residual auto-correlation entirely. The set of results show that the estimate of the focus coefficient, um_{it}, is robust to extension of the dynamics. When the AB-estimator is used, the coefficient changes from -0.0241 with one lag, to -0.0223 with two lags, and finally to -0.0250 when three lags of the gini are included in the model. The rather small change in the um_{it} coefficient illustrates that the results for our

Table B2: LSDV estimation results for the parameters in (2) and (3). Full estimation table corresponding to columns (1) and (2) in Table 3 in main article.

	(1)	(2)
	$gini$	um_t
um_t	-0.0206***	
	(-4.69)	
$gini_{t-1}$	0.378***	0.0349*
	(39.93)	(2.18)
um_{t-1}	0.00772*	0.597***
	(2.06)	(56.38)
Share with high school education	-0.0240***	-0.0123
	(-3.91)	(-0.75)
Share with high education, lower degree	0.0252***	-0.133***
	(3.74)	(-7.05)
Share with high education, higher degree	-0.0176	-0.0712*
	(-1.25)	(-2.22)
Share of employees aged 25-29	-0.00429	0.155***
	(-0.56)	(8.85)
Share of employees aged 30-34	-0.00557	0.163***
	(-0.72)	(9.62)
Share of employees aged 35-39	-0.0162*	0.168***
	(-2.22)	(8.22)
Share of employees aged 40-44	-0.0118	0.198***
	(-1.60)	(10.87)
Share of employees aged 45-49	-0.000804	0.200***
	(-0.12)	(11.25)
Share of employees aged 50-54	-0.000270	0.202***
	(-0.04)	(8.72)
Share of employees aged 55-59	0.00264	0.224***
	(0.33)	(9.78)
Share of employees aged 60-66	0.0104	0.225***
	(1.25)	(10.67)
Medium skill occupation share	0.0127***	-0.0297***
	(6.69)	(-7.14)
High skill occupation share	0.0460***	-0.0467***
	(25.19)	(-11.05)
51-75 employees (dummy)	0.00124	0.00509*
	(1.50)	(2.54)
76-100 employees(dummy)	0.00198	0.0118***
	(1.87)	(4.42)
100 or more employees (dummy)	0.00444***	0.0166***
	(3.61)	(4.95)
Share of part-time workers	-0.0323***	0.000886
	(-12.82)	(0.19)
Share of women	0.0131*	-0.0147
	(2.46)	(-0.90)
Immigrant share	-0.0103	-0.0356*
	(-1.87)	(-2.48)
2002	-0.00166	-0.0170***
	(-1.38)	(-4.82)
2003	-0.00226*	-0.000609
	(-1.98)	(-0.20)
2004	-0.00224*	0.000880
	(-2.02)	(0.29)
2005	-0.00173	-0.00563
	(-1.54)	(-1.84)
2006	0.00151	0.00291
	(1.40)	(0.95)
2007	-0.000599	-0.00425
	(-0.54)	(-1.33)
2008	0.000426	0.00431
	(0.37)	(1.33)
2009	-0.00218	0.00110
	(-1.79)	(0.33)
2010	0.0000943	0.00796*
	(0.07)	(2.33)
2011	0.000799	-0.000352
	(0.61)	(-0.10)
2012	-0.000140	0.00366
	(-0.10)	(1.00)
2013	-0.000494	-0.00337
	(-0.35)	(-0.90)
2014	-0.00201	0.00388
	(-1.34)	(0.97)
2015	0.0358***	-0.0173***
	(18.29)	(-3.88)
2016	0.0218***	-0.0147**
	(11.63)	(-3.17)
2017	0.0207***	-0.0176***
	(11.26)	(-3.75)
2018	-0.00238	-0.0320***
	(-1.24)	(-6.86)
R^2	0.328	0.489
N	32 951	32 951
Establishments	2275	2275
Avg. obs.	14.48	14.48

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: Reference categories: share with primary school education, share of employees aged 20-24, low skill occupation share, size dummy 25-50 employees. 2001 year dummy omitted due to collinearity. Robust standard errors clustered at establishment level, t-statistics in parenthesis. Source: Authors' calculations.

Table B3: Estimation of the *um*-parameter in (2) for different levels of hourly wage trimming in the calculations of the gini coefficient.

	(1) *gini*, reference model	(2) *gini*, 1% Winsorization	(3) *gini*, no trimming
um	-0.0206***	-0.0264***	-0.0280***
	(-4.69)	(-5.30)	(-6.90)
R^2	0.335	0.272	0.291
N	32 951	32 951	32 951

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: All models include establishment-level educational shares, age bin shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories and time dummies. Robust standard errors clustered at establishment level, t-statistics in parenthesis. Source: Authors' calculations.

Table B4: Random effects estimation results for the coefficients in (2) and (3).

	(1) *gini*	(2) *um*
um	-0.0319***	
	(-7.37)	
$gini_{t-1}$	0.673***	-0.0150
	(78.89)	(-1.11)
um_{t-1}	0.0174***	0.926***
	(4.12)	(295.62)
R^2(between)	0.954	0.995
R^2 (within)	0.296	0.425
R^2 (overall)	0.650	0.934
N	32 951	32 951
Establishments	2275	2275
Avg. obs.	14.48	14.48

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: All models include establishment-level educational shares, age bin shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories and time dummies. Robust standard errors clustered at establishment level, t-statistics in parenthesis. Source: Authors' calculations.

Table B5: LSDV estimation of the parameters in (2) and (3). Model (1) includes all private sector establishments. Model (2) is restricted to establishments with at least 25 employees.

	(1) Private sector		(2) Private sector, >24 employees	
	gini	*um*	*gini*	*um*
um	-0.00509***		-0.0295***	
	(-7.17)		(-11.37)	
$gini_{t-1}$	0.00735***	-0.00953***	0.172***	0.00486
	(5.80)	(-4.01)	(53.23)	(1.16)
um_{t-1}	-0.00284***	0.277***	0.0103***	0.472***
	(-4.12)	(223.29)	(4.40)	(182.81)
R^2	0.089	0.663	0.374	0.842
N	752 004	752 328	106 916	106 921
Establishments	186 614	186 746	19 155	19 156
Avg. obs.	4.030	4.029	5.582	5.582

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: All models include establishment-level educational shares, age bin shares, occupational shares, share of women, part-time share, immigrant share, establishment size categories and time dummies. Robust standard errors clustered at establishment level, t-statistics in parenthesis. Source: Authors' calculations.

focus parameters are robust despite the residual auto-correlation picked up by the IS test in Table 3.

Chapter 2

Do unions show solidarity in wage bargaining? Evidence from Norway

Elin Svarstad

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Do unions show solidarity in wage bargaining?*

Evidence from Norway

Elin Svarstad[1 2 †]

Abstract

One of the core objectives of unions is to raise the wages of the lowest paid. Utilizing a panel of individual matched employee-employer data covering the Norwegian private sector in the period 2000-2014, I investigate how workplace union density is related to individual low-pay risk. By exploiting changes in tax deductions for union members in Norway as a source of exogenous variation, a negative effect of increased union density on low-pay risk is identified within jobs. The results further suggest that the effect of local bargaining power on individual low-pay probability wais larger among immigrants than among natives.

JEL Classification: C23, C26, J31, J50, J51, J52

Keywords: unions, low pay, wage differentials, panel data, instrumental variables

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1. Introduction

“Upgrading pay for the lowest paid groups in the trade must be an objective.” (Riksavtalen 2022, p. 8).

The citation above is retrieved from one of the largest collective agreements within the hospitality sector in Norway and represents a core mission statement internalized by many Norwegian trade unions. Throughout their existence, unions have been known for opposing inequality, and perhaps mainly through raising the wages of those at the bottom of the wage distribution. In Norway this aim, referred to as ‘solidaristic wage policy’ is commonly thought of as one of the core features of the *Nordic model of working life*.

In many countries, a growing share of low paid employees is an important driver of increasing inequality. The extent to which the low pay segment has expanded over time, however, varies substantially across the most advanced economies (McKnight et al. 2016). These developments have shed new light on factors both contributing to and counteracting the prevalence of low paid jobs. While megatrends such as skill-biased technological change, globalization and immigration have been shown to explain several dimensions of increases in inequality,[1] they explain less of why the development in inequality has evolved at different speed across countries. To gain more insight into these differences, attention has been turned to institutional factors, and in particular to the dimensions of the institutional framework that are distinctive to specific countries (Doucouliagos 2017). There are large variations in both regulatory frameworks and the strength of collective institutions across countries, with correspondingly different implications for labor market outcomes.

A distinction is usually made between centralized and decentralized wage negotiation systems. The Nordic countries represent an intermediate case, referred to as *centralized decentralization*: wage bargaining is coordinated at the central level, but industry-wide negotiations are usually supplemented by local negotiations at the establishment level. This aspect of the bargaining system implies that workplace-level factors are likely to influence the final wage outcome. Furthermore, most of the collective agreements in the predominantly low-wage industries covering blue collar occupations in Norway are so-called minimum wage agreements, establishing an absolute wage floor with the intention that increments additional to the established wage rates should be negotiated. The actual wage levels are therefore likely to depend on the local bargaining strength of the union, as well as the objectives it brings to the negotiations. Unions in Norway, and those affiliated with the Norwegian Confederation of Trade Unions (LO) in particular, are known to promote solidaristic wage policies, in the sense that they have an explicit agenda to raise the wages of the lowest paid. At the central level, these policies are reflected in the demands and priorities in the industry-wide negotiations. It is less clear how unions use their bargaining strength to achieve their goals at workplace level[2]. One purpose of this study is to reduce this knowledge gap, by investigating the effect of local union bargaining power on individual low-pay risk.

The time period the data cover is characterized by a huge increase of immigration to Norway. Following the EU enlargements in 2004 and 2007, Norway emerged as one of the countries in Western Europe with the highest relative rates of immigrants from new EU member states in Central and Eastern Europe (Friberg 2016b). The inflow of new migrants represented a massive shock to parts of the Norwegian labor market. Despite substantial language barriers, employers seeking manual labor gained access to a cheap and flexible pool of workers from countries with wage levels well below Norway's. The supply shock quickly exerted downward pressure on wages, and the term 'social dumping'[3] was frequently used in the public debate. Given the vulnerable position of many immigrants in low-wage occupations, a second question raised in this study is whether the gain from having a solidaristic union in the workplace, in terms of reduced low-pay risk, is more pronounced among immigrants than among natives.

Utilizing a high-quality matched employer-employee dataset covering the entire Norwegian private sector in the period 2000 to 2014, I raise two questions. First, I ask whether local bargaining power, as measured by workplace level union density, has an effect on the individual's probability of being low paid. As such, the study explicitly tests a stated ambition among Norwegian trade unions to raise the wages of those at the bottom of the wage distribution at establishment level. By exploiting exogenous variation in public subsidization of union membership in a 2SLS regression, the estimated effect may be given a causal interpretation. Second, I investigate whether the potential reduction of low-pay probability attributable to the bargaining power of the union is heterogeneous among immigrants and natives, respectively. Possible heterogeneity may provide policymakers with valuable knowledge about how institutions in the labor market handle immigration and the consequences at workplace level.

The remainder of the paper is organized as follows: Section 2 reviews related literature on how unions alter labor market outcomes in terms of wages and low pay. Section 3 is a description of the Norwegian wage bargaining system, union wage policies and the implications for the extent of low pay. In Section 4, I describe the data, define low pay as the term is used in the study, and present some descriptive statistics. In Section 5, I outline the empirical methodology and discuss identification, while Section 6 documents the results. Section 7 provides a discussion and some concluding remarks.

2. Related Literature

What unions do has been the subject of extensive research for decades. The review by Freeman & Medoff (1984) acts as the leading reference in the literature, drawing a map of the core functions of unions from both a theoretical and an empirical perspective. The question of what unions *do* differs from the related question of what unions *want*, i.e., what are their 'objectives'. The short answer to the latter question is that unions are formed to ensure fair wages, benefits, and better working conditions for their members. The former question concerns the ability of unions to apply their priorities in practice and is more difficult to answer. Unions have the potential to influence a long list of labor market outcomes, including wage levels and wage inequality, productivity, profits, investments, and technological change, to mention some. On the one hand, unions negotiate with employers over various aspects of the employment contract, and thus use their bargaining power to pursue their goals at the given bargaining level (i.e., workplace, industry, nationally). On the other hand, they represent a political force through their role as large collective organizations and may be able to gain benefits not as easily achieved through bargaining, through the political process. It is the former role of unions that is of interest in this study, i.e., their ability to implement their wage policy through bargaining at the micro level (i.e., workplace).

Theoretically, whether unions achieve their wage goals depends on their ability to limit the supply of labor if their wage demands are not met. Workers represented by unions thus have the potential to extract rents from employers and receive payment above perfect competition market wages. The observed wage outcome of union bargaining is assumed to depend on several factors, such as the relative bargaining power of the two parties, which in turn is a function of potential conflict outcome, the price elasticity of demand for labor, the capital-labor ratio and worker support for the union (Oswald 1985).

Although the objectives of unions vary across countries and environments, a common goal seems to be some form of wage standardization, which often translates into a wage-compressing effect. Specifically, union wage policies are often guided by 'a fair day's pay for a fair day's work,' implying that wages are attached to jobs rather than to individuals' attributes (Bryson 2014). The Nordic union movement has a long history of solidaristic wage policies, in the sense that low-wage workers should receive larger percentage increases than high-wage workers. However, 'taking wages out of the competition' is a standard formulation among North American unions as well (Rosenfeld 2014, p. 70). Nevertheless, it is not obvious why unions should support redistributive wage policies. According to Freeman & Medoff (1984), there are three reasons for why unions prefer equality. The first is that they want to promote the interests of the median member, who in most cases earns less than the mean wage earner. The other explanations are that single-rate wage agreements reduce managerial discretion, and that greater wage equality increases 'worker solidarity and organizational unity' (p. 80). Although aspects of these explanations seem reasonable, there are many potential objections as well.

The shortcomings of the economic rationality approach have shifted the focus towards more sociological explanations based on norms (Elster 1989; Swenson 1989). One strand of the empirical literature is focused on relating norms, values, and attitudes to union membership. Across OECD countries, union membership is shown to be associated with support for redistribution (see, e.g., Finseraas 2009; Checchi et al. 2010). In a sample of twenty-one European countries over the period 2002–14, Mosiman & Pontusson (2017) showed that union membership is associated with support for redistribution among low-wage workers and even more so among high-wage workers. The authors interpret the solidarity effect of union membership as the internalization of distributive norms, and perhaps beliefs promoted by unions about the relationship between inequality and economic growth. The findings imply that unions and their members may have material interests that lead them to support redistributive policies even if such policies do not conform to short-term income maximization for the median union member.

Regardless of the theoretical motivation, a large strand of empirical literature has established a relationship between unions and labor market institutions, and the wage structure, i.e., on wage differentials across industries, firms, skills, gender, age, migratory background etc.[4] Research from the US has shown that de-unionization has been an important factor in explaining the rise in wage inequality, mainly through the diverging impact on wages in the lower and middle part of the wage distribution (Card 1996, 2001; DiNardo & Lemieux 1996; Firpo et al. 2009; Farber et al. 2021). Studies using data for OECD countries suggest that unions compress wage differentials across countries and over time (Rueda & Pontusson 2000; Pontusson 2013; Vlandas 2018). The presence of unions and collective agreements is also shown to be associated with reduced low-pay risk within countries, even for non-union members (see Benassi & Vlandas 2021 for Germany, Jordfald et al. 2021 for Norway, Schmitt 2008 for the US). For Norway, Reite (2020) found significant heterogeneity in returns for both union membership and union density across the wage distribution. Specifically, manipulating the distribution of union density yielded positive wage effects at the median wage level and below, and negative effects above the 77th percentile. The results suggest that unions reduce wage level dispersion in the left tail of the wage distribution.

The vast majority of the studies on the equalizing impact of unions on wages are concentrated on aggregate levels, that is within countries, sectors or industries. Less attention has been directed towards how union strength within the workplace affects individual wage levels in different groups of wage earners. Studies from Norway focused on detecting average union wage premiums find little or no wage advantage associated with individual union membership, but detect substantial wage rises in workplaces with a higher union density (Barth et al. 2000; Balsvik & Sæthre 2014; Bryson et al. 2020). Analyzing intra-establishment wage inequality, Svarstad & Nymoen (2022) show that increases in workplace level union density can contribute to a more compressed wage structure in successive years in a sample of private sector workplaces over a 19-year period. The relationship was especially pronounced in the lowest part of the wage distribution. These results indicate that we should expect unions to have an impact on low-pay risk at local level.

The second question I raise in the study is whether the effect of union strength in the workplace on low-pay risk may be different for immigrants and for natives. Research on immigrants in the labor market is largely focused on the effects of immigration on wages, the human capital provided by immigrants and their assimilation into society. Less attention has been devoted to the part played by unions in altering wage responses to immigration, in particular at workplace level. It is not self-evident how unions view and react to immigration. They may oppose it because it poses a threat to the native labor force. The influx of labor into the labor market may undermine union power, since the majority of migrants are non-unionized (Avci & McDonald 2000). However, once the migrants have been admitted, unions have a strong interest in policies concerning their rights, to prevent immigration from causing a deterioration in wages or working conditions (Menz 2010; Boräng et al. 2020). In Norway, LO initially endorsed transitional arrangements that imposed restrictions on access to the labor market for individual jobseekers from the new EU member states but made it clear that they welcomed migrants provided that they worked under the same conditions as natives (Hardy et al. 2012).

Studies examining the impact of unions on wages have yielded mixed results. In a study of 18 countries, including Norway, Boräng et al. (2020) show that since the 1980s, countries with strong unions have extended more social and economic rights to migrants relative to those extended to citizens, than countries with weak unions. In the US, Schmitt (2010) finds that immigrants that are union members earn significantly more that non-union members and are more likely to have a retirement plan. For Ireland, Turner et al. (2014) report that unionized Irish nationals are more likely to earn more than the median hourly wage than unionized immigrants, implying that unionized nationals enjoy greater benefits from membership than unionized immigrant workers. However, unionized immigrants were found to be almost twice as likely as non-unionized immigrants to earn above the median hourly pay. Finseraas et al. (2020) showed that the increase in labor supply in Norway due to the EU enlargement had negative effects on the earnings and employment prospects of native workers facing tougher labor market competition, but no evidence that the increase in immigrant labor had any effects on natives' tendency to unionize. They do not, however, consider immigrant wages. In general, immigrants in Norway have a lower tendency to unionize than natives (Nergaard & Ødegård 2022). Although this may be partly due to attitudes or cultural differences, Cools et al. (2020) show that immigrants in Norway are subject to sorting in the labor market, because they tend to be employed in firms and industries with lower levels of unionization.

3. Institutional Context: Unions, Wage Bargaining and Low Pay in Norway

The relationship between different dimensions of union presence and labor market outcomes varies across institutional contexts. It is therefore essential to discuss the implications of union presence in the context of how the labor market is organized in a particular country, sector or industry. Norway is one of the few countries in the the OECD without a national legal minimum wage. Wages and other working conditions are instead negotiated between the social partners at industry level. Bargaining takes place at both industry and establishment level, although central coordination plays an important role in ensuring sound macroeconomic outcomes. In an international context, Norway, as well as the other Nordic countries, has compressed wage distribution and a high minimum wage rate (Eurostat 2016).

Coordination is a key feature of Norwegian wage formation, at both central and local level. The so-called 'front-runner model' is one of the main coordinating institutions in the Norwegian labor market. The premise of the model is that '*wage growth must be adjusted to a level which over time is capable of sustaining the competitiveness of import and export competing industries'* (Nymoen 2017, p. 13). In practice, this is done by letting the exposed industries bargain first and establish a wage norm based on what is considered a sustainable wage level relative to competing countries. By setting the premise for wage development in the rest of the economy, the norm ensures that wages in the sheltered industries neither exceed nor lag behind the industries competing internationally. The front-runner model has been an essential contribution to keeping wage inequality low across different parts of the labor market. It ensures that the groups that possess the lowest bargaining power (the lowest paid) benefit the most, as they receive the wage growth obtained by groups with greater market power.

Collective agreements also play a pivotal role in the prevention of low pay, by introducing binding industry-specific minimum wage rates. In order for these wage floors to 'bite', a certain level of coverage is necessary, as only establishments that are bound by collective agreements are obliged to adhere to wage rates and adjustments. Collective agreement coverage in the private sector is approximately 52 per cent (Nergaard 2022), although the effect of collective agreements in Norway has been strengthened through a system of general application (Eldring & Alsos 2012). The unionization rate in the Norwegian private sector is around 38 percent but it varies a great deal across industries, ranging from just over 70 per cent in electricity, gas, steam, and air-conditioning supply to under 20 percent in accommodation and food service activities (Nergaard 2022).[5] The corresponding organization rate among employers is 73 percent.

As noted, the wage bargaining system in Norway is 'two-tiered': the central negotiations are usually supplemented by local wage negotiations conducted at establishment level. How wage growth is distributed centrally and locally varies across industries. In parts of the private sector, as much as 60 percent of the annual wage growth among blue-collar workers is negotiated at workplace level (NOU 2013:13). Most blue-collar workers in the private sector are covered by so-called minimum wage agreements. These agreements establish an absolute wage floor, which the employer cannot deviate from. Furthermore, the agreements stipulate the negotiation of increments in addition to the minimum wage rates, often based on criteria related to productivity and the financial situation of the establishment (Stokke 2012; Alsos & Nergaard 2021). The final wage outcome thus depends on the result of local wage negotiations between employer and union at each workplace. The local negotiations are not subject to sanctions such as strike or lockout, although some agreements contain provisions allowing unionized workers to lower their productivity during the negotiations to put pressure on their employers. Furthermore, as noted by Moene et al. (1993), peace clauses do not mean that the employees are powerless: 'Workers may engage in work-to-rule actions where they follow work instructions in a pedantic way, decline to work overtime, and generally refuse to co-operate with the firm' (p. 102).

Wage compression was adopted in the 1950s as a goal for the trade union movement in both Norway and Sweden under the name 'solidaristic wage policy'. Solidaristic wage policies are often associated with a wage compressing outcome of centralized wage negotiations (Moene & Wallerstein 1995).[6] At central level, the solidaristic wage policy of Norwegian unions is reflected in the inclusion of income guarantee provisions in collective agreements, ensuring that wage increments benefit workers in low-wage industries. This is done by demanding nominal rather than percentage increases, and making sure agreements without local wage formation and industries with low average wages receive a higher increase than others (Alsos & Nergaard 2022). However, the values and norms underlying the union wage policies are likely to have an impact on their priorities at every level where bargaining occurs. LO and its affiliated unions have long traditions of promoting equality and fairness by working against low pay. For example, many unions provide guidance in the form of written directions on how to conduct local negotiations for the employee elected representatives. The Norwegian Food and Allied Workers Union (NNN), which organizes workers in the food industry, states the following about local negotiations: '*Traditionally, NNN's pay policy has been based on the smallest possible pay differences between employees, as this strengthens both cohesion and common solutions.*' (NNN 2022). Furthermore, the demand for nominal supplements as opposed to percentage increases in order to raise the lowest wages is common not only at central level but in local negotiations as well.

4. Data and descriptive statistics

4.1 Data

The primary data sources used in this study are the Norwegian Employer-Employee Register (AA register) and the Register of End of the Year Certificate (LTO register). The AA register is a basic data register of employment in Norway and contains all jobs in the Norwegian labor market that have more than four contracted hours per week and that last for at least one week. It contains detailed information about establishments and employees. Because employers are legally obligated to report all changes in the stock of employees, the coverage is close to complete. Information about earnings is collected from the LTO register. Educational statistics are attached, as well as occupation, country of origin, gender, year of birth and several establishment characteristics. Variables such as industry and sector are obtained from the Register of Legal Entities and Statistics Norway's Business and Enterprise

Register (VoF). Personal attributes are obtained from the Central Population Register (DSF). Each individual, workplace and firm has its own unique identifying number, thus allowing the units to be tracked over time.

The dataset is constructed as an individual-year-panel. In cases where an employee has jobs in more than one establishment a specific year, the job with the highest number of days in the calendar year is kept as the most representative job. The sample is restricted to employees in the private sector, working at least 20 hours each week. The restriction is imposed to ensure a certain level of attachment to the labor market, as well as a wage measure less sensitive to measurement errors. Because union density is the preferred indicator of union bargaining strength, most estimations are conducted on a sample of workplaces with at least ten employees. The final sample consists of 2,017,393 individuals within 61,152 establishments, encompassing 3,357,995 unique job spells. The total number of observations in the dataset amounts to 11,830,262.

Earnings is measured as total payments, including base salary, bonus payment, and overtime payments.[7] The hourly wage is constructed from the tax data based on job-specific annual earnings, job spell duration and contracted weekly working hours.

Individual union membership is obtained from data on union membership fees, which are reported to the tax authorities by the unions. Union density is calculated as the yearly leave out mean of workers members of a union within an establishment. Whether an establishment participates in a collective agreement or not is obtained from membership data from the private sector collectively agreed pension scheme ('Fellesordningen for AFP'), whereby all workplaces that are members are also parties to a collective agreement.

4.2 Definition of Low Pay

There is no generally accepted limit for what qualifies as low paid work across countries. There seems to be agreement that low wages should be defined as wages below a threshold designating a socially acceptable remuneration, but it remains difficult to determine what 'socially acceptable' translates into. These difficulties have led researchers to adopt different thresholds, expressed as a proportion of the median or average wage of all workers. Such relative measures have the advantage that they are easy to compare across countries. A relative measure also captures "*a sense of the degree of social and economic inclusion among a country's workforce that is sensitive to societal notions of relative deprivation or relative disadvantage"* (Grimshaw 2011, p. 4). The OECD defines low pay as less than two-thirds of median earnings, and this definition seems to have gained acceptance in research and statistics.

The low pay threshold used in this study is relative in nature, but highly country specific. In the following, low pay will refer to an hourly wage level of less than 85 percent of the mean for manufacturing workers. The manufacturing worker is an important point of reference in the

Norwegian context, as a representative of the exposed sector in the front-runner model. Furthermore, the wage level of the manufacturing worker is located close to the middle (median) of the Norwegian wage distribution, making it a convenient measure for monitoring the extent to which the wage distribution remains compressed over time. The definition is applied by the Technical Reporting Committee on Income Settlements (Teknisk Beregningsutvalg, TBU),[8] and is a frequently used point of reference for the income guarantee provisions in collective agreements. 85 percent of the mean annual wage of manufacturing workers amounted to NOK 445 740 (approximately USD 43 950) in 2021 (NOU 2022:4), which is higher than most other definitions of low pay. The rationale behind the choice of definition constitutes the purpose of this study. In order to examine whether the local bargaining power of unions affects the individual propensity to be low paid, the threshold applied should reflect what the unions themselves define as low pay.

4.3 Sample Statistics

Table 1 reports the annual low-pay limits according to the definition of less than 85 per cent of the mean wage of manufacturing workers, as well as the share of workers paid below the threshold in the estimation sample.

Table 1 Annual/hourly low-pay thresholds (nominal) and share of low paid employees. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees. N=11 830 262.

Year	Low-pay limit, annual wage* (NOK)	Low-pay limit, hourly wage (NOK)	Share below low-pay limit (percent)
2000	215 800	111	21
2001	226 400	116	20
2002	237 700	122	20
2003	246 600	126	20
2004	252 600	130	19
2005	260 600	134	18
2006	270 100	139	17
2007	284 600	146	18
2008	301 200	154	19
2009	312 300	160	21
2010	321 800	165	22
2011	333 000	171	23
2012	345 400	177	23
2013	356 800	183	23
2014	366 400	188	24

*Defined as a less than 85 percent of the mean hourly wages of manufacturing workers. Source: Annual reports, TBU.

As is apparent from the table, there has been an overall increase in the share of low paid employees in the sample during the 15 years from 2000 to 2014, despite the somewhat diverging trends in the first and second halves of the period.

Because low pay is measured as a binary state, it is of interest to examine the extent of changes in individual low-pay status. Most model specifications in the following rely solely on within variation, implying that the estimated effect of union density on low pay is exclusively based on variation whereby the individual actually changes to or from low-pay status from one year to the next. Tables 2 and 3 explore this issue further, by showing transitions in low-pay status in the estimation sample. Although the majority of individuals in the sample remain in one pay category (low paid or otherwise) for the entire period, there are a good number of switches as well. Note that individuals who are only observed in one year are excluded from the matrix.[9]

Table 2 Transition probabilities at individual level. Low-pay status. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.

Pay status year *t-1*	Pay status year *t* (above/below Low-pay threshold) Above	Below	Total
Above	93.07	6.93	100
Below	30.38	69.62	100
Total	78.2	21.8	100

Note: 'Low paid' is defined as a wage level of less than 85 percent of the mean hourly wage of manufacturing workers.

The rows in Table 2 reflect the initial values, and the columns reflect the final values. Each year, some 93 percent of the higher paid individuals in the data remained higher paid the following year.:The remaining 7 percent became low paid. While higher paid individuals only had a 7 percent chance of becoming low paid each year, low paid individuals had a 30 percent chance of rising (or returning) to a pay level above the threshold. Table 3 is the corresponding transition matrix within a particular job, i.e., for the same individual in the same workplace. The transition probabilities are not that different from those in Table 2.

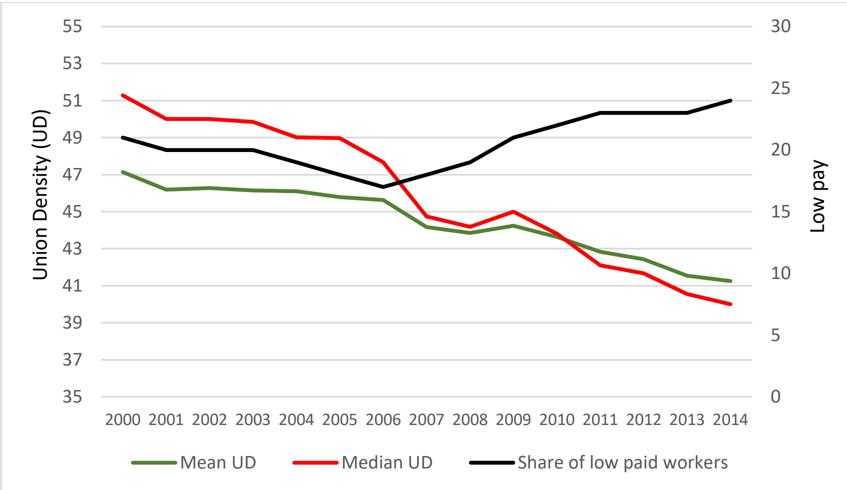
Table 3 Transition probabilities within jobs. Low-pay status. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.

Pay status year *t-1*	Pay status year *t* (above/below Low-pay threshold) Above	Below	Total
Above	94.2	5.8	100
Below	26.41	73.59	100
Total	79.52	20.48	100

Note: 'Low paid' is defined as a wage level of less than 85 percent of the mean hourly wage of manufacturing workers.

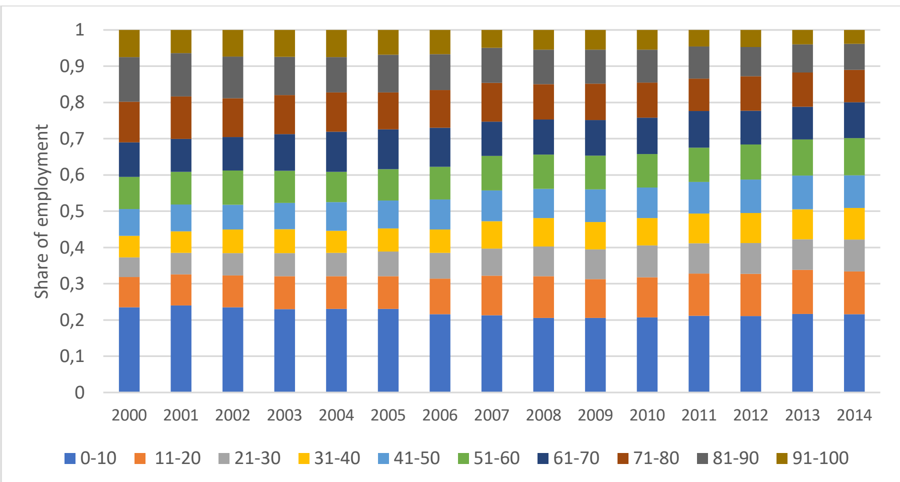
The following study is concentrated on the establishment level. Accordingly, Figure 2 shows the development in mean union density as well as the mean share of low paid workers within an establishment. The figure illustrates that on average the workplace union density in the sample fell by just above 5 percentage points during the 15-year period. If we consider the median, the decline was more extensive, from around 50 percent in 2000 to 40 percent in 2014. The share of low paid workers is included in the figure to illustrate the opposing trends in the two variables. While union density displays a negative trend, the share of low paid workers moves in the opposite direction.

Figure 1 Workplace level union density (UD) and share of low paid workers. Weighted by the number of employees. Private sector fulltime employees in workplaces with more than 9 employees. N=11,830,262.



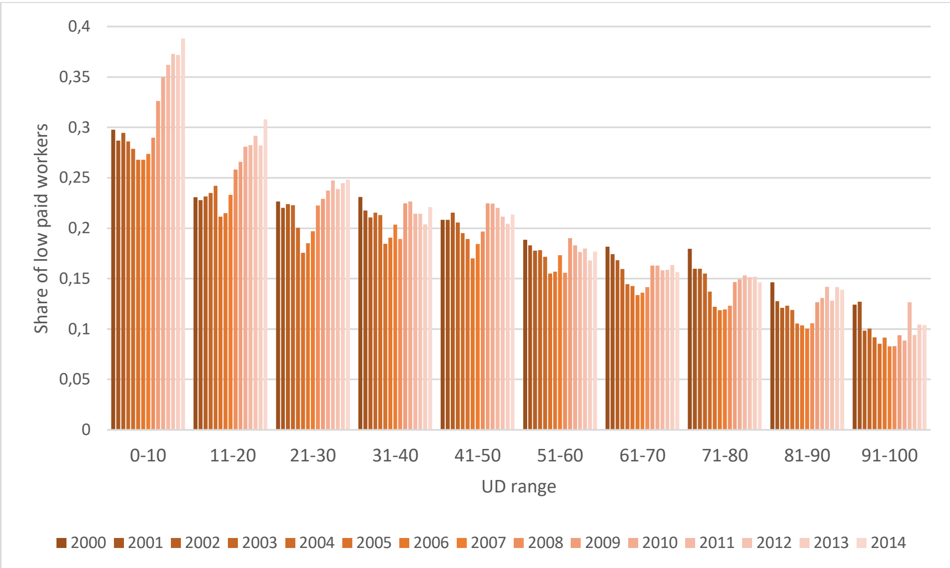
The larger drop in the median than in the mean implies a fall in the share of employees in workplaces with higher levels of union density. This is strongly supported by Figure 2, which shows the share of employees in the sample working in establishments with different levels of unionization.

Figure 2 Share of wage earners by establishment level union density levels. Private sector fulltime employees in workplaces with more than 9 employees. N=11,830,262.



As is apparent from the figure, there has been a shift towards a larger share of employees working in establishments with lower levels of union density. This trend has been accompanied by a steeper negative relationship between establishment level union density and share of low paid employees in the second part of the period, illustrated in Figure 3.

Figure 3 Mean share of low paid workers by establishment level union density ranges. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees. N=11,830,262.



5. Empirical Approach

In order to evaluate the effect of union density on the probability of being low paid, I estimate several specifications of the following linear probability model:

$$LP_{\mathrm{ij}t} = \alpha_i + \gamma UD_{jt} + \boldsymbol{X_{it}\delta} + \lambda_t + \varepsilon_{ijt},$$

where $LP_{\mathrm{ij}t}$ is a binary variable taking the value 1 if individual *i* in establishment *j* is low paid in year *t* (i.e., hourly wage less than or equal to 85 percent of the mean for manufacturing workers), and 0 otherwise. α_i denotes time-invariant individual fixed effects, while λ_t represents time-specific effects reflecting shocks, events, and changes of economic environment common to all individuals. The primary variable of interest is workplace union density (UD_{jt}), calculated as the mean share of workers within an establishment that are members of a union, excluding the value of individual *i*. The reason for leaving out the individual's own value is the concern that individual membership status in itself may be the driver of switches in low-pay status. Workplace union density is a continuous variable measured in percent.[10] $\boldsymbol{X_{it}}$ represents the vector of control variables reflecting demographic and occupational characteristics such as gender, age, immigration status, educational attainment level, occupational category, and 1-digit SIC industry codes of the current workplace. Finally, ε_{ijt} is an error term, assumed to be normally distributed and i.i.d.

Identifying the true effect of workplace unionization on the individual's probability of being low paid is challenging. My strategy involves a stepwise exploration of the relationship between the two variables by means of several functional forms, estimators, and sample restrictions. As a starting point, and to provide a benchmark for subsequent estimations, I run an ordinary least squares (OLS) regression. A drawback of the OLS estimator is that it may provide biased estimates in the presence of unobserved heterogeneity across individuals. Employees maintain a range of capabilities not captured by the data which may or may not contribute to low pay. These capabilities may also vary systematically with the parts of the labor market where low paid employees typically work, which in turn may be a predictor of the level of unionization in the workplace. Failure to control for unobserved variables that are correlated with both low pay and union density may lead to omitted variable bias. I therefore estimate the model equation using a within estimator, allowing for individual fixed effects. Because estimating fixed effect coefficients soaks up all the between-individual variation, both observed and unobserved, the variation left in the data is less likely to be attributed to unobserved differences in capabilities among employees. Utilizing within-individual variation only, i.e., considering how individual changes in low-pay status are associated with changes in union density across time, reduces the threat of omitted variable bias.

A threat to the identification strategy remains, however, if changes in unionization are correlated with job switches. Establishments differ in their ability and willingness to pay employees above or below the low-pay threshold, both within industries and within occupations. If by changing her job (i.e., establishment), an employee goes from being low paid to earning above the threshold, while simultaneously moving from an establishment with a low unionization rate to a highly unionized workplace, the estimated coefficient does not capture changes in local bargaining power. I therefore move on to estimating a model that includes job fixed effects (i.e., a combination of individual and workplace), relying exclusively on changes in low-pay status associated with job variation in unionization across time.

Even when the same individual is considered within the same establishment, there may still be omitted variables affecting both union density and individual earnings, thereby causing the estimated coefficient on union density to be biased. Changes in the demand and/or supply of workers are examples of such variables. The dramatic increase in the supply of immigrant workers following the EU expansions in 2004 and 2007 is an illustrative case. Most immigrants entered industries already prone to low pay, such as construction, industrial cleaning and the hospitality sector, providing additional downward pressure on wages.[11] As immigrant workers are in general less likely than natives to become union members, shifts in the labor supply may have overestimated the negative relationship between the level of unionization in the workplace and the probability of being low paid. On the other hand, increased relative demand for high-skilled labor due to technological changes may have slowed down wage growth among existing lower skilled employees, while simultaneously *raising* union density, as higher skilled workers are more likely to be union members. This would appear as a positive relationship between low pay and union strength in the workplace.

As the final step in my analysis, in an attempt to bypass any remaining endogeneity, I utilize changes in tax deductions for union membership fees in Norway as a source of exogenous variation in unionization, to estimate the effect of workplace union density on low pay.

6. Results

In this section, I present the results of the empirical analysis. Table 4 displays estimation results based on linear probability models, capturing the relationship between establishment level union density and the individual probability of being low paid in the period 2000-2014.

Table 4 Linear probability models of the impact of workplace level union density on individual probability to be low paid. Private sector fulltime employees. 2000-2014.

Model *Estimator*	**1a** *OLS*	**1b** *Within*	**1c** *Within*	**1d** *Within*	**1e** *Within*	**1f** *Within*
Union density (UD)	-0.00139***	-0.00108***	-0.00107***	-0.0000808*	-0.00130*	0.00000550
	(-30.85)	(-37.98)	(-27.55)	(-2.29)	(-2.54)	(0.10)
Year dummies	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓
Ind. fixed effects		✓	✓			
Job. fixed effects				✓	✓	✓
Min no. of employees	10	10	20	10	20	10
Group of workers	All	All	All	All	All	Covered establishments
R^2	0.237	0.603	0.609	0.688	0.682	0.668
N	11816315	11413055	8904843	10584720	8305183	5989629

*Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than the mean hourly wage of 85 percent of manufacturing workers. Union density is measured in percent. Model 1a contains the following controls: gender, immigration status, age, age squared, occupation (1-digit ISCO 08), educational attainment level (1-digit ISCED 2011) and industry of current occupation (1-digit SIC 2007). Models 1b and 1c include controls for educational attainment level and industry of current occupation (1-digit SIC 2007), while Models 1d and 1e control for educational attainment level. Robust standard errors clustered at establishment level, t-statistics in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$*

The first column (Model 1a) shows results from a pooled model estimated by means of OLS, including a set of control variables. The estimated coefficient on union density (UD) is negative and statistically significant at the 0.1 percent level, implying that a 10-percentage point increase in the establishment level union density is associated with a reduction of about 1.4 percent in the individual low-pay probability.

Because the OLS estimate may be partly driven by unobserved individual heterogeneity, Model 1b includes individual fixed effects, thus controlling for time-invariant average differences across individuals. Allowing for individual fixed effects causes the estimated UD-coefficient to drop from 0.00139 to 0.00108, indicating that the OLS -estimates overestimate the importance of unionization on the probability of being low paid. The results from Model 1b suggest that a 10-percentage point increase in workplace union density reduces the probability of being low paid by just over 1 percent. The result is robust to restricting the sample to larger establishments (Model 1c).

We may worry that the changes in union density associated with changes in low-pay status captured by the UD-coefficient in Model 1b may be correlated with individual job changes. The results in column 4 are from a model estimated with *job* fixed effects, hence only exploiting variation originating from individuals in specific establishments[12] across time. The estimation of this specification completely alters the results. The estimated UD-coefficient now has an absolute value close to zero. This pattern indicates that when mobility across workplaces is considered, there is no systematic relationship between local bargaining power and the individual probability to be low paid. The conclusion remains unchanged when only workplaces exceeding 19 employees are considered (Model 1e).

A non-existent relationship between union density and low-pay risk within jobs may be surprising, given the redistribution preferences exhibited by Norwegian unions. A possible explanation may be the potential absence of collective agreements. The right to bargain over wages at the workplace is established in the local agreement entered into by the particular establishment. Furthermore, as highlighted in Barth et al. (2000), wage formation in the uncovered sector differs from that in the covered sector. To explore this possibility, Model 1f is estimated on a sample consisting only of covered establishments. This restriction does not, however, do much to change the results. The estimated coefficient is insignificant and approximately equal to zero. Table A13 in the Appendix reports separate results for selected industries in the private sector. Apart from administrative and support service activities, the results of none of the included industries differ from the results in the previous section.

6.1 Endogenous Unionization

The above results suggest that there is virtually no relationship between low-pay status and workplace unionization within jobs. Yet, there may be reason to doubt the causal interpretation of this result. Even when considering the same individual in the same establishment, union density in the workplace may be endogenously determined by the individual's propensity to be low paid. Because union density is calculated excluding the individual's own potential contribution to local bargaining power, the biggest concern is related to unobservable variables simultaneously affecting both the union membership decisions of colleagues (i.e. union density) and the individual's low-pay probability. As mentioned, one such factor could be shifts in the demand for or supply of certain kinds of labor. If, for example, there is an inflow into the workplace of high paid workers with a relatively high propensity to unionize, their wage levels are likely to negatively affect the relative wage growth of existing workers. At the same time, union density would increase within the workplace, generating a positive correlation between low pay and local bargaining power. Correspondingly, increased access to low paid workers with an inherently lower probability of unionizing would lower union density, while possibly putting downward pressure on wages. This would negatively bias the estimated coefficient of union density.

Another noteworthy shortcoming of the fixed-effects approach to uncovering union wage effects, as pointed out by Vella & Verbeek (1998), is that this approach only eliminates the endogeneity operating through the individual (job-)specific effects (p. 171). Any time-varying endogeneity continues to contaminate the estimates. To counter the potential remaining endogeneity, I instrument for workplace union density with changes in tax subsidies for union membership during the period of analysis.

6.1.1 Public Subsidization of Union Membership

In Norway, employees who pay union fees are entitled to a tax deduction. The deduction is, however, limited upward by a cap. During the 15-year period of the analysis, the size of the cap was increased several times as a result of political priorities by the left-wing government in power for the majority of the years the data cover. These changes in deductions of taxable income led to a significant change in the net price of union membership. Under the assumption that union membership is an ordinary good, price reductions are followed by an increase in the individual demand for unionization. Empirically, this assumption is supported by Barth et al. (2020a), who found strong support for a positive relationship between the subsidy rate and the individual propensity to unionize. As the workforce of an establishment constitutes the sum of employees, the sum of demand changes following the policy adjustments is likely to have an impact on union density within workplaces. Given that the price changes have no impact on individual low-pay status, the subsidy is eligible as an instrument for union density. This identification strategy to tackle the endogeneity of union density was first applied by Barth et al. (2020b), and later by Donini et al. (2021).

The instrument is constructed by utilizing data on actual individual payments of union membership fees. As changes in tax rules for union members affect incentives to unionize, also among individuals who are not union members, hypothetical unions based on 3-digit occupational codes and 2-digit industry codes are constructed, in line with Barth et al. (2020b). For each existing union member, I calculate the average membership fee for each hypothetical union each year, excluding the individual’s own contribution to the mean. The tax subsidy is then calculated as the product of the marginal income tax (28 percent) and the minimum of the average fee and the cap on tax deductions.

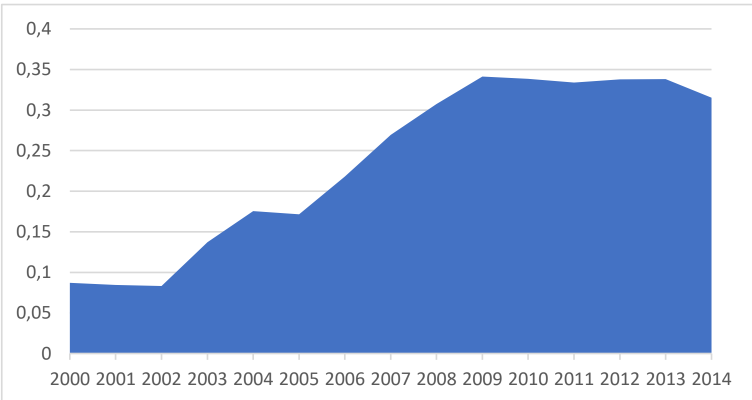
That is, $s = 0.28 \times \min(\overline{fee},\ cap)$. The subsidy is measured relative to the net union membership fee, such that

$$S_{ratio_t} = \frac{s_t}{\bar{f}_0 - s_t}$$

where s_t is the subsidy amount in year *t*. The average union membership fee in the workplace is fixed at the first year of observation, $(\bar{f}_0)$, to avoid potential endogeneity arising from price responses from the unions following increases in the subsidy, as well as adaptation of the occupational composition of workplaces by employers. Because the net union fee may be influenced by low-pay status, I also include the inverse of the historical net union fee as a control variable in all the regressions.

Figure 4 illustrates how the subsidy ratio evolved in the period 2000-2014.

Figure 4 The subsidy ratio. 2000-2014.



Note: The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee.

In order for the instrument to be valid, the subsidy ratio should not be correlated with individual low-pay status through channels other than union density (instrument exogeneity assumption). Because some trade unions relate their membership fees to earning levels, it is not self-evident that the exogeneity assumption holds. However, there are large variations in how Norwegian unions calculate the fee: some are fixed, some progressive, some limited upward by caps. As the synthetic membership fees are calculated on the basis of all members of the data set, representing different membership fee schemes and excluding the individual's own contribution to the mean, I argue that the subsidy ratio is indeed a valid instrument for union density in this case.

6.1.2 2SLS estimates

Table 5 shows the results of 2SLS estimations, using the subsidy ratio as an instrument for workplace union density.

Table 5 Estimated effect of union density on individual probability to be low paid. 2SLS. Private sector fulltime employees. 2000-2014.

Model	**2a**	**2b**	**2c**	**2d**
Union density	-0.00336***	-0.00297***	-0.00723***	-0.01011***
	(-18.41)	(-14.01)	(-4.84)	(-4.70)
Year dummies	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Ind. fixed effects	✓	✓		
Job fixed effects			✓	✓
Min no. of employees	10	20	10	20
Group of workers	All	All	All	All
First stage:				
Subsidy ratio	30.70***	37.43***	20.21***	21.44***
	(7.26)	(8.88)	(6.07)	(5.14)
Weak instrument test:				
Cragg–Donald F:	110419.6	117278.7	11036.7	13960.2
Kleibergen–Paap F:	412.4	396.7	59.27	37.18
N	11412278	8904560	10584290	8304974

Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than the mean hourly wage of 85 percent of manufacturing workers. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Models 2a and 2b control for educational attainment level (1-digit ISCED 2011) and industry of current occupation (1-digit SIC 2007), while models 2c and 2d control for educational attainment level. The inverse of the historical net union membership fee is included in all models. Robust standard errors clustered at establishment level, t statistics in parentheses.
** $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$*

As is apparent from the first stage estimation in Table 5, the subsidy ratio is indeed significantly correlated with union density. Specifically, an increase in the subsidy ratio of 10 percentage points is estimated to increase union density by somewhere in the range of 2-3.7 percentage points. Instrumenting union density by the subsidy ratio has the effect of enhancing the negative relationship between low-pay risk and union density in the models, including individual fixed effects, and making it reappear in the models with job fixed effects. The 2SLS results for Model 2a, which includes individual fixed effects, suggest that a 10-percentage point increase in union density reduces the risk of being low paid by approximately 3.4 percent.[13] The estimated effect is similar when the sample is restricted to larger establishments (Model 2b). The inclusion of job fixed effects (Model 2c) results in a larger estimated coefficient, implying that the individual low pay probability is reduced by 7.2 percent following an increase in union density of 10 percentage points within the same job- spell. Overall, the results indicate that the estimated models in the previous section underestimated the effect of union density on the individual propensity to be low paid.

As the tradition for conducting local negotiations may in practice vary across the private sector, the Appendix includes separate 2SLS regressions for five main industries (A13). While the direction of the results is in general accordance with the estimated effect from Table 4, the effect varies a great deal between industries. This is primarily explained by the first stage, i.e., that the subsidy ratio affects membership differently in different parts of the labor market. It is important to note that the instrument recovers the local average treatment effects (LATE), rather than an average treatment on the treated effect (ATT). Consequently, some caution must be shown in interpreting the results. For example, Barth et al. (2022) show that tax subsidies tend to stimulate union membership more in segments of the labor market where density is low in the first place. However, immigrants and low-wage workers are, in general, shown to be among those with the highest elasticity of union membership with respect to the subsidy.

6.2 Natives and Immigrants

A second area of interest in this study is whether union bargaining strength affects the propensity to be low paid to the same extent among natives and immigrants. Most immigrants entering Norway in the period of the analysis were low-skilled workers arriving from low-income countries with a lot to gain from leaving their country of origin. By way of illustration, in 2007 average hourly wages in Norway were 50 percent higher than in Sweden, almost eight times higher than in Poland and almost fifteen times higher than in Romania (Friberg et al. 2012). Consequently, one might expect that willingness to work for low wages would be greater on average among immigrants than among natives. The potential gains from a solidaristic union in the workplace would thus be correspondingly larger in the former group. Table 6 shows different specifications of the model equation in Section 2, estimated separately for natives and immigrants.

Table 6 Estimated effect of union density on individual probability to be low paid. 2SLS. Natives and immigrants. Private sector fulltime employees. 2000-2014.

Model *Estimator*	**3a** *Within*	**3b** *Within*	***t-stat diff.***	**3c** *Within*	**3d** *Within*	***t-stat diff.***	**3e** *2SLS*	**3f** *2SLS*	***t-stat diff.***
Union density	-0.00105*** (-34.03)	-0.00166*** (-27.42)	-8.97	-0.0000775* (-2.20)	-0.0000952 (-0.90)	-0.16	-0.00987*** (-5.17)	-0.02143** (-3.20)	-1.95
Year dummies	✓	✓		✓	✓		✓	✓	
Controls	✓	✓		✓	✓		✓	✓	
Ind. fixed effects	✓	✓							
Job. Fixed effects				✓	✓		✓	✓	
Min empl.	10	10		10	10		10	10	
Group	Natives	Immigrants		Natives	Immigrants		Natives	Immigrants	
First stage									
Subsidy ratio							20.39*** (5.96)	16.64*** (4.40)	
Weak instrument test:									
Cragg–Donald F:							12857.5	515.7	
Kleibergen–Paap F:							35.54	14.31	
R^2	0.576	0.632		0.677	0.703				
N	10059552	1353503		9365655	1219157		9365273	1219017	

Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than the mean hourly wage of 85 percent of manufacturing workers. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Models 3a and 3b contain the following controls: educational attainment level (1-digit ISCED 2011) and industry of current occupation (1-digit SIC 2007). Models 3b-3f include controls for educational attainment level. The inverse of the historical net union membership fee is included in all models. The t-stat diff. refers to a test for equality of the union density coefficients in the estimated models between natives and immigrants. Robust standard errors clustered at establishment level, t-statistics in parentheses.
** p < 0.05, ** p < 0.01, *** p < 0.001*

The pattern in Table 6, when moving from models including individual fixed effects (3a and 3b) to job fixed effects (3c and 3d), and then finally instrumenting union density with the exogenous changes in the subsidy ratio (3e and 3f), is similar to that shown in Table 4. However, the results of the separate estimations reveal a difference between natives and immigrants wth respect to how local bargaining power affects the probability of being low paid. Indeed, the results from Models 3e and 3f indicate that the reduction in low-pay probability resulting from an increase in union density of 10 percentage points is more than twice as large among immigrants as among natives (21 vs. 10 percent). While the estimated effect of union density for natives is similar to that in the whole sample (Model 2c in Table 4), the estimated effect seems to differ substantially for immigrants. It should be noted, however, that the 2SLS estimate is somewhat imprecisely estimated in the immigrant sample. A t-test reveals that the differences between the estimated coefficients in the separate samples of natives and immigrants are statistically significant at the one percent level for models 3a/3b, not statistically significant for models 3c/3d, and only marginally for models 3e/3f.

An interesting observation is that the inclusion of job fixed effects in models 3c/3d, i.e., controlling for where people work, completely eliminates the difference between the groups as it appears in models 3a/3b. This suggests that there is selection into workplaces, i.e., that natives and immigrants are disproportionally distributed in establishments with strong unions. However, when the exogenous variation in the tax subsidy is applied as an instrument, the difference reappears.

What might explain why immigrants benefit more than natives in terms of reduced low-pay risk when unions grow stronger in the workplace? One possible interpretation is that immigrants possess lower bargaining power in the first place, and thereby have relatively more to gain from the presence of a strong union in the workplace. Most immigrants entering the Norwegian labor market following the EU enlargements in 2004 and 2007 had few outside options. The majority of them came from poor countries in Eastern Europe and were willing to work for what qualifies as low wages in a Norwegian context.[14] The same cannot be said to the same extent about natives, who were protected by the Norwegian social security network, as well as having a comparative advantage regarding job mobility within the Norwegian labor market. Provided that unions work to promote the conditions of those who need it most, it may not be that surprising that immigrants benefit more from the presence of a strong union.

Of course, migrants entering Norway during the years 2000-2012 were different along several dimensions, including skills level and bargaining power. The results from estimated models of the four largest subsamples of migrants broken down by origin are reported in the Appendix (A15). Immigrants from the EU/EEA appear to be the group driving the result in Table 6, supporting the hypothesis that it is groups with relatively low bargaining strength that primarily profit from union strength in the workplace.

7. Concluding Remarks

Unions have been known to compress wage inequality in their environments. In the Nordic countries, these objectives are referred to as solidaristic wage policies, and primarily involve raising the wages of the lowest paid. At the macro level, unions work to achieve their solidaristic wage objectives by including income guarantee provisions in the collective agreements, demanding nominal rather than percentage increases, and making sure agreements without local wage formation or industries with low average wages receive higher increases than others. However, a significant part of wage formation in the Norwegian private sector happens at local level, i.e., within the workplace. While previous literature has shown that strong unions are associated with lower wage inequality in their environment, particularly in the lowest part of the wage distribution, less is known about the relationship between union bargaining strength and individual low-pay risk within establishments.

Utilizing a panel of individual matched employee-employer data covering the Norwegian private sector in the period 2000-2014, this study has examined the relationship between local bargaining power, as measured by workplace level union density, and the individual propensity to be low paid. The results show that increases in union density have a significant negative effect on individual low-pay risk within job spells. Specifically, an increase in union density of 10 percentage points is estimated to reduce low-pay risk by 7.2 percent. The findings strongly suggest that the objective of Norwegian unions to raise the wages of the lowest paid has been achieved at local level in the sample and period analyzed in the study. Although the results appear robust, the estimated effect varies across industries.

A second finding of the study is that immigrants have comparatively more to gain from strong unions in the workplace than natives. One interpretation of this finding is that immigrants are worse off, in the sense that they hold less bargaining power than natives in the first place and therefore derive greater benefit from the solidaristic wage policy unions exhibit. This hypothesis is supported by the heterogeneity found across different groups of migrants. While the estimated effect is stronger than the overall effect in the subsample consisting of immigrants from EU/EEA, it is practically absent among migrants from European countries outside the EU and Africa and Asia. EU/EEA migrants were in a particularly vulnerable position in the years following the EU enlargements and may therefore have derived a greater advantage from stronger unions in the workplace.

The results of the study imply that unions may have been important regulators of low pay at the local level in Norway during the period of the analysis. This is important knowledge in the context of the ongoing debate about a statutory minimum wage across the EU. Both unions and employer organizations in the Nordic countries have opposed this suggestion, as the principle that wages are the responsibility of the social partners stands strong in these countries. The principle entails that the social partners, particularly the trade unions, have assumed a responsibility to ensure an acceptable wage floor. However, there are threats to this strategy. Most importantly, bargaining strength requires a sufficient union density level. Although high in some parts of the labor market, the level of unionization is low in many private sector industries in Norway. The evidence presented above shows that union strength has an impact on low-pay risk in a sample where workplace level union density averages around 45-50 precent, indicating that the impact of unions on low pay is not conditioned on very high levels of union density in the workplace. However, in Norway, as in Sweden and Denmark, the greatest decline in union density in recent years has occurred in typical low-wage industries (Alsos & Nergaard 2022). This trend should perhaps be the greatest worry in countries which believe that the issue of ensuring a sufficiently high wage floor should be resolved between the unions and the employers' organizations.

Notes

[1] See e.g. Acemoglu & Autor (2011), Helpman (2018).
[2] Throughout the paper, the terms 'workplace' and 'establishment' are used synonymously, both referring to the lowest functional unit at a single, physical location that produces or distributes goods or performs services.
[3] The Norwegian government defines social dumping as follows: Social dumping is deemed to be present both if foreign employees are subject to breaches of health, safety and working environment regulations and if they are paid wages that are unacceptably low. https://www.regjeringen.no/en/topics/labour/the-working-environment-and-safety/innsikt/social-dumping/id9381/
[4] See Card (2020) for a review of the literature.
[5] The Norwegian union membership rate is low compared to the other Nordic countries, where trade unions have traditionally administered the unemployment benefit funds and thus have had better recruitment opportunities.
[6] Rehn and Meidner (1953) argued that wage equalization across workplaces and industries should be a goal, because it would stimulate economic development via faster conversion to new and more efficient technology, and thus higher productivity. The idea was that equal pay for equal work and a compressed wage distribution would raise wages in low-productivity establishments, while restraining wages in high-productivity workplaces. This arrangement implies indirect subsidizing of highly productive establishments and indirect taxation of those with low productivity. The result is faster replacement of low productivity in favor of new and more modern workplaces, through a process similar to creative destruction (Schumpeter 1942).
[7] Because overtime is included in the wage measure, some robustness checks are done to make sure the results are not driven by overtime payments. In the Appendix (A5), the hourly wage is compared to that of individuals included in another register source ('Lønnsstatistikken', the Earnings statistics), which is a representative sample from the same time period. The Earnings statistics are regarded as more accurate, as they are collected for the purpose of wage negotiations. The calculated hourly wages from the two sources are similar for the individuals included in both samples.
[8] TBU was established in 1967 and plays a central role in ensuring that the social partners and the authorities have a shared understanding of the statistical material underlying the wage negotiations. The committee submits annual reports that form the basis for wage negotiations, including the share of low-paid fulltime wage earners.
[9] Table A4 in the Appendix displays the number of distinct individuals changing low-pay status each year in the period of analysis.
[10] To explore possible non-linearities in the relationship between union density and the probability of being low paid, results based on a less restrictive version of the model are reported in Table A7 in the Appendix. The model includes a set of dummies representing different bands of unionization. The point estimates of the dummy variables are monotonically increasing, and the linear specification seems to be a good approximation.
[11] Within the construction industry, wage growth was shown to be lower in trades with rising immigrant employment shares during the period 1998-2005 (Bratsberg & Raaum 2012).
[12] 'Job' is defined as individual within the same establishment. However, the results are robust to a more restrictive definition of 'job', namely 'individual within occupation within establishment'. The fact that there is practically no difference between the two operational definitions indicates that changes in occupation are of little importance for changes in low-pay status.
[13] As a sensitivity analysis, I have re-estimated the models using alternative low-pay thresholds; see Tables A8 and A9 in Appendix. The estimated effects of union density reported there are very similar to those reported in this section.
[14] It should be noted that many of the immigrants entering Norway to work following the EU enlargements often had typical, precarious employment, and a significant share were posted workers (Friberg et al. 2016a). Some of the most vulnerable immigrant groups are thus not included in the sample consisting of fulltime employees in mid-size and large establishments.

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Appendix

A1 Summary statistics on key variables. 2000-2014. Private sector full time employees in workplaces with more than 9 employees.

Variable		Mean	Std. dev.	No. of obs.
Low paid (0/1)	overall	0.21	0.44	11 830 262
	between		0.42	2 017 393
	within		0.28	5.86
Union density	overall	0.44	0.31	11 830 262
	between		0.28	2 017 393
	within		0.13	5.86
Female (0/1)	overall	0.32	0.47	11 830 262
	between		0.48	2 017 393
	within		0.00	5.86
Immigrant (0/1)	overall	0.13	0.33	11 830 262
	between		0.41	2 017 393
	within		0.00	5.86
Age	overall	40.29	11.99	11 830 262
	between		12.50	2 017 393
	within		3.47	5.86
Education (bins)	overall	4.30	1.81	11 830 262
	between		2.09	2 017 393
	within		0.32	5.86
Collective agreement (0/1)	overall	0.55	0.50	11 830 262
	between		0.45	2 017 393
	within		0.24	5.86

Note: Calculated using -xtsum- in Stata 17.

A2 Annual mean of key variables. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.

Year	Low paid (0/1)	Union density (0/1)	Female (0/1)	Immigrant (0/1)	Age	Education (bins)	Collective agreement
2000	0.21	0.47	0.32	0.07	39.16	3.98	0.53
2001	0.20	0.46	0.32	0.07	39.29	4.02	0.58
2002	0.20	0.46	0.32	0.08	39.62	4.06	0.57
2003	0.20	0.46	0.32	0.08	39.98	4.09	0.57
2004	0.19	0.46	0.32	0.08	40.30	4.13	0.57
2005	0.18	0.46	0.31	0.08	40.59	4.16	0.57
2006	0.17	0.46	0.31	0.09	40.46	4.21	0.57
2007	0.18	0.44	0.31	0.11	40.27	4.26	0.57
2008	0.19	0.44	0.32	0.13	40.27	4.31	0.57
2009	0.21	0.44	0.32	0.14	40.76	4.36	0.57
2010	0.22	0.44	0.32	0.15	40.61	4.41	0.53
2011	0.23	0.43	0.32	0.17	40.59	4.46	0.53
2012	0.23	0.42	0.31	0.18	40.63	4.53	0.53
2013	0.23	0.42	0.31	0.20	40.67	4.60	0.53
2014	0.24	0.41	0.31	0.21	40.78	4.66	0.52
Total	0.21	0.44	0.32	0.13	40.29	4.30	0.55

A3 Number of years the individuals were observed in the dataset. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.

Years observed	Freq.	Percent	Cum.
1	403 517	3.41	3.41
2	514 166	4.35	7.76
3	583 986	4.94	12.69
4	629 056	5.32	18.01
5	703 640	5.95	23.96
6	654 792	5.53	29.49
7	722 449	6.11	35.6
8	741 232	6.27	41.87
9	770 130	6.51	48.38
10	727 890	6.15	54.53
11	693 649	5.86	60.39
12	747 324	6.32	66.71
13	783 211	6.62	73.33
14	1 019 760	8.62	81.95
15	2 135 460	18.05	100
Total	11 830 262	100	

A4 Number of distinct individuals changing to/from low-pay status, sorted by years observed in dataset. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.

Years observed in dataset	Distinct individuals changing low-pay status
1	0
2	56010
3	69607
4	70910
5	68703
6	59310
7	56946
8	51180
9	47700
10	40701
11	35101
12	33018
13	30299
14	31860
15	48668
Total	700013

A5 Comparison of mean and median hourly wage levels (nominal) in the earnings statistics and the LTO register.

	Mean		Median	
	ES	*LTO*	*ES*	*LTO*
2000	159	163	139	149
2001	166	173	145	157
2002	175	181	153	165
2003	180	187	158	171
2004	187	194	164	177
2005	195	205	170	186
2006	204	216	176	194
2007	215	229	187	204
2008	229	241	198	215
2009	236	238	204	210
2010	244	245	211	215
2011	254	254	219	222
2012	263	262	227	230
2013	275	273	236	239
2014	283	279	244	245

A6 Distribution of employees in workplaces with different ranges of union density. 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.

Union density range	Percent
0-10	21.97
11-21	10.39
21-30	7.49
31-40	7.28
41-50	8.27
51-60	9.31
61-70	9.85
71-80	10.12
81-90	9.6
91-100	5.74
Total	100

A7 Alternative specification of Model 1b. Union density specified as rate (1) vs. bins (2). Linear probability models of the impact of workplace level union density on individual propensity to be low paid. Private sector fulltime employees in workplaces with more than 9 employees. 2000-2014.

	(1)	(2)
Union density	-0.108***	
	(-37.98)	
Union density dummies		
Ref. (< 0.2)		-
[0.2-0.4)		-0.0302***
		(-18.64)
[0.4-0.6)		-0.0515***
		(-25.65)
[0.6-0.8)		-0.0657***
		(-29.03)
>0.8		-0.0788***
		(-29.01)
R^2	0.603	0.610
N	11426644	11426644

*Note: (1) corresponds to Model 1b in the main paper. The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than the mean hourly wage of 85 percent of manufacturing workers. Both models contain the following controls: educational attainment level (1-digit ISCED 2011) and industry of current occupation (1-digit SIC 2007). Robust standard errors clustered at establishment level, t-statistics in parentheses. * p < 0.05, ** p < 0.01, *** p < 0.001*

A8 Estimated effect of union density on individual propensity to be low paid. 2SLS. Higher threshold: 85 percent of mean wage for manufacturing worker plus 5 percent. Private sector fulltime employees. 2000-2014.

	(1)	(2)	(3)	(4)
Union density	-0.00427***	-0.00369***	-0.0138***	-0.0141***
	(-22.61)	(-16.76)	(-5.21)	(-4.47)
Year dummies	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Ind. fixed effects	✓	✓		
Job fixed effects			✓	✓
Min empl.	10	20	10	20
Group of workers	All	All	All	All
First stage:				
Subsidy ratio	30.70***	37.43***	20.21***	21.44***
	(7.26)	(8.88)	(6.07)	(5.14)
Weak instrument test:				
Cragg–Donald F:	111385.0	117999.4	13833.6	15.427.6
Kleibergen–Paap F:	416.1	398.9	36.83	26.38
N	11412278	8904560	10584290	8304974

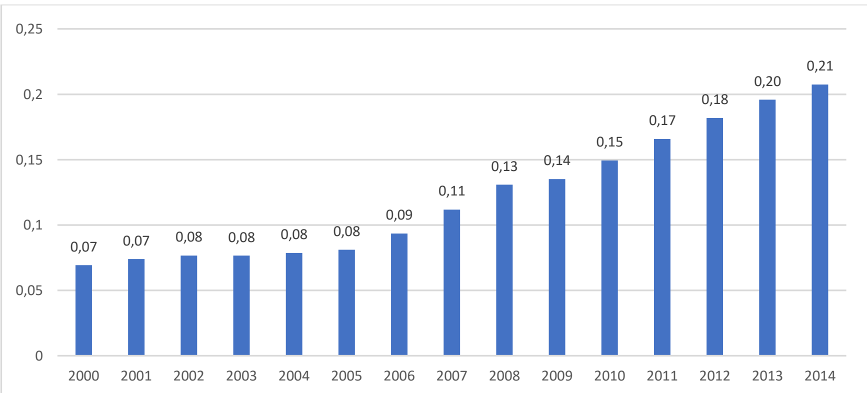
*Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than the mean hourly wage of 85 percent of manufacturing workers plus 5 percent. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Models 2a and 2b control for educational attainment level (1-digit ISCED 2011) and industry of current occupation (1-digit SIC 2007), while models 2c and 2d control for educational attainment level. The inverse of the historical net union membership fee is included in all models. Robust standard errors clustered at establishment level, t-statistics in parentheses. * p < 0.05, ** p < 0.01, *** p < 0.001*

A9 Estimated effect of union density on individual propensity to be low paid. 2SLS. Lower threshold: 85 percent of the mean for manufacturing workers minus 5 percent. Private sector fulltime employees. 2000-2014.

	(1)	(2)	(3)	(4)
Union density	-0.00433***	-0.00376***	-0.0141***	-0.0134***
	(-22.67)	(-16.91)	(-5.42)	(-4.56)
Year dummies	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Ind. fixed effects	✓	✓		
Job fixed effects			✓	✓
Min no. of employees	10	20	10	20
Group of workers	All	All	All	All
First stage:				
Subsidy ratio	30.70***	37.43***	20.21***	21.44***
	(7.26)	(8.88)	(6.07)	(5.14)
Weak instrument test:				
Cragg–Donald F:	111385.0	117999.4	13833.6	15427.6
Kleibergen–Paap F:	416.1	398.9	36.83	26.38
N	11412278	8904560	10584290	8304974

*Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than 85 percent of the mean hourly wage of manufacturing workers minus 5 percent. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Models 2a and 2b control for educational attainment level (1-digit ISCED 2011) and industry of current occupation (1-digit SIC 2007), while models 2c and 2d control for educational attainment level. The inverse of the historical net union membership fee is included in all models. Robust standard errors clustered at establishment level, t-statistics in parentheses. * p < 0.05, ** p < 0.01, *** p < 0.001*

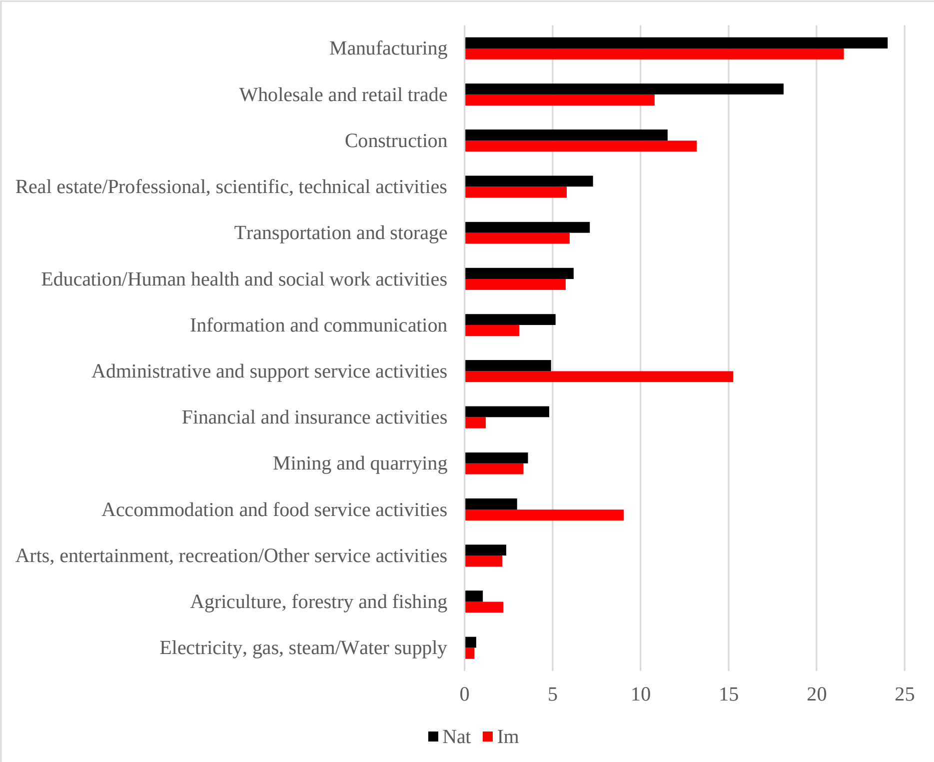
A10 Share of immigrants. Private sector fulltime employees. 2000-2014.



A11 Annual share of low pay and mean workplace level union density among immigrants (Im) and natives (Nat). Private sector fulltime employees. 2000-2014.

	Low-pay share		**Mean UD**	
	Im	Nat	Im	Nat
2000	0.31	0.20	39	48
2001	0.31	0.19	39	47
2002	0.31	0.19	39	47
2003	0.30	0.19	39	47
2004	0.30	0.18	39	47
2005	0.28	0.17	38	46
2006	0.27	0.16	38	46
2007	0.30	0.17	36	45
2008	0.31	0.17	37	45
2009	0.36	0.19	37	45
2010	0.39	0.19	35	45
2011	0.39	0.20	34	45
2012	0.39	0.19	33	45
2013	0.38	0.19	33	44
2014	0.41	0.20	32	44
Total	0.36	0.19	35	46

A12 Distribution of employees by industry. Percentage shares of natives and immigrants in main industries (ISCED 2011). 2000-2014. Private sector fulltime employees in workplaces with more than 9 employees.



A13 Linear probability models of the impact of workplace level union density on individual propensity for being low paid. Selected private sector industries. Fulltime employees. 2000-2014.

	(1)	(2)	(3)	(4)	(5)
Industry	*Manufacturing*	*Construction*	*Wholesale and retail*	*Accommodation and food service activities*	*Administrative and support service activities[a]*
Union density	-0.000827 (-0.08)	0.00535 (0.68)	0.00262 (0.42)	-0.00903 (-0.58)	-0.0559** (-2.65)
Year dummies	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓
Job fixed effects	✓	✓	✓	✓	✓
Min no. of employees	10	10	10	10	10
R^2	0.655	0.640	0.677	0.672	0.685
N	2623379	1249871	1823329	353408	321996

*Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than 85 percent of the mean hourly wage of manufacturing workers. Union density is measured in percent. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Controls contain educational attainment level (1-digit ISCED 2011). [a]Employment activities (Nace 78) is excluded from the sample. Robust standard errors clustered at establishment level, t-statistics in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$*

A14 Estimated effect of union density on individual propensity to be low paid. 2SLS-estimates. Selected industries in the private sector. Full time employees. 2000-2014.

	(1)	(2)	(3)	(4)	(5)
Industry	*Manufacturing*	*Construction*	*Retail*	*Accommodation and food service activities*	*Administrative and support service activities*[a]
Union density	-0.0161***	0.0315	-0.0202***	-0.0176	-0.0349
	(-3.86)	(1.27)	(-3.83)	(-1.10)	(-0.79)
First stage:					
Subsidy ratio	40.83***	27.58	20.00***	20.59	10.73
	(5.10)	(1.36)	(4.34)	(1.57)	(0.89)
Year dummies	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓
Job fixed effects	✓	✓	✓	✓	✓
Min no. of employees	10	10	10	10	10
Weak instrument test:					
Cragg–Donald F:	10867.5	252.9	584.69	163.2	79.60
Kleibergen–Paap F:	26.05	1.839	18.86	2.454	0.785
N	2623379	1249871	1823329	353408	321996

*Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than 85 percent of manufacturing workers' mean hourly wage. Union density is measured in percent. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Controls contain educational attainment level (1-digit ISCED 2011). [a]Employment activities (Nace 78) is excluded from the sample. The inverse of the historical net union membership fee is included in all models. Robust standard errors clustered at establishment level, t-statistics in parentheses * p < 0.05, ** p < 0.01, *** p < 0.001*

A15 Estimated effect of union density on individual propensity to be low paid. 2SLS estimates. Selected countries of origin. Fulltime employees. 2000-2014.

	(1)	(2)	(3)	(4)
Country of origin	*EU/EEA*	*European countries outside the EU*	*Africa*	*Asia*
Union density	-0.0315**	0.0162	0.0130	-0.00304
	(-2.79)	(1.47)	(0.98)	(-0.57)
First stage:				
Subsidy ratio	15.01**	19.34	17.85*	21.37***
	(3.28)	(2.81)	(2.09)	(4.00)
Year dummies	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Job fixed effects	✓	✓	✓	✓
Min no. of employees	10	10	10	10
Weak instrument test:				
Cragg–Donald F:	274.8	78.04	47.02	237.5
Kleibergen–Paap F:	10.78	7.882	4.362	16.00
N	596065	93546	63120	248078

*Note: The dependent variable is a binary variable equal to 1 if the individual is low paid and 0 otherwise. Low pay refers to a pay level less than 85 percent of the mean hourly wage of manufacturing workers. Union density is measured in percent. The subsidy ratio is calculated as the marginal tax rate (28 per cent) multiplied with the minimum of actual membership payments and the maximum deductible amount, measured relative to the net membership fee. Union density is measured in percent. Controls consist of educational attainment level (1-digit ISCED 2011). The inverse of the historical net union membership fee is included in all models. Robust standard errors clustered at establishment level, t-statistics in parentheses. * p < 0.05, ** p < 0.01, *** p < 0.001*

Chapter 3

Unions, collective agreements and productivity: A firm-level analysis using Norwegian matched employer–employee panel data

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ORIGINAL ARTICLE

Unions, collective agreements and productivity: A firm-level analysis using Norwegian matched employer–employee panel data

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Abstract

What is the role of collective agreements in explaining how unions affect firm-level productivity? Using matched employer–employee panel data for the Norwegian labour market, comprising almost 21 million individual-year observations in the period 2002–2018, we find that the presence of a collective agreement in a firm is associated with higher productivity. Without a collective agreement, higher union density is estimated to reduce productivity. However, if a collective agreement is implemented in the firm, not only is the estimated negative effect reduced—in some cases it becomes positive. This result remains significant, numerically and statistically, across several model specifications and different estimation methods. In particular, we provide a new source of exogenous variation in union memberships by utilizing information on intergenerational transmission of union preferences. Besides regulating terms and conditions for wage formation and working hours, collective agreements have a profound impact on how firms organize and formally recognize the voice of workers. In this regard, our finding supports the conclusion of Free-

man and Medoff that the quality of institutional systems is crucial to understand what unions do to productivity.

1 | INTRODUCTION

What unions do to productivity, as well as for other aspects of corporate performance, has been the subject of extensive research for decades. In the seminal works by Freeman and Medoff (1979, 1984), unions are portrayed with two faces: the *monopoly face* and the *exit voice/institutional response face*. While the former refers to the monopoly power attained by unionized workers through collective bargaining, the latter covers the various mechanisms through which unions may alter industrial relations. As these effects generally work in opposite directions, the effect of unions on productivity is theoretically ambiguous. The question of how unions affect productivity is therefore a question that must be answered empirically. However, despite the vast body of empirical literature, the evidence is mixed and inconclusive, reflecting various strengths of the two faces of unionism in different contexts (Doucouliagos et al., 2017).

The mixed evidence on what unions do to productivity calls for the scope of union research to extend to more countries, sectors, time periods and institutional contexts (Laroche, 2020). Unions operate in very diverse environments with respect to how institutions and legislation regulate and facilitate their activities and organization. More fundamentally, the impact of unions on firms' performance is likely to vary with the extent of unionization. Union presence may be measured along at least two important dimensions—the first being the union density (UD) at the workplace, the second the union's formal impact on the firm's organization, as measured by the presence of a collective agreement. The former dimension has been utilized in many studies, recently also in Norway (Barth et al., 2020), but less attention has been devoted to the study of collective agreements.

In countries characterized by decentralized bargaining, the introduction of a firm-level collective agreement often requires that the union wins a majority vote. In other words, collective agreements are only implemented in firms with a strong local union. In many European countries, however, there is an important distinction between having unionized workers at the plant and being covered by a collective agreement, as firms may be covered by sectoral agreements without having unionized employees in the firm (OECD, 2019). The rules for implementing a collective agreement in Norway are somewhere in between. In general, collective agreements are invoked by labour unions, but only if the UD is above the threshold determined in higher level agreements, which is usually 10 per cent of the workers in the particular bargaining area. Moreover, participation in the agreements is in principle voluntary for both parties. Both the voluntary engagement and the low threshold for invoking collective agreements, make the distinction between UD and collective agreement coverage (CAC) important in the Norwegian context.

In this article, we argue that collective agreements act as a formal recognition of the unions' right to express their views on working conditions and the organization of firms. A collective agreement thus constitutes an important organizational institution through which unions may alter industrial relations at the firm level. By not taking this dimension of unionization into account, empirical analyses of unions' impact on productivity could be biased, or at best imprecise. We contribute to the discourse on what unions do to productivity by explicitly exploring how the union-productivity relationship is altered by the presence of collective agreements. More generally, our contribution adds to the growing literature on what unions do to productivity in different contexts by providing evidence from the Norwegian labour market. Norway represents an interesting case because voluntary collective agreements are relatively more important than legislation compared to many other countries. Also, the availability of high-quality register data on all individuals enables more accurate inference. Finally, our article is an important contribution to the limited number of studies providing causal evidence on what unions do to productivity. While we are not able to fully control for the possible correlation between productivity shocks and the presence of a collective agreement, endogenous unionization is handled by instrumenting UD among workers with the UD among the workers' parents.

The remainder of the article is organized as follows. Section 2 reviews previous literature on how unions alter productivity. The section also discusses the few studies that emphasize the role of collective agreements and related labour market institutions. Section 3 then gives a brief introduction to the system and organization of unions and collective agreements in Norway. We present the data in Section 4, while Section 5 describes our methodological approach. Section 6 contains a presentation and discussion of our results. Section 7 provides a conclusion.

2 | RELATED LITERATURE

Theoretically, the influence of unions on productivity is ambiguous. In the traditional neoclassical view, unions act as monopolies that distort labour market efficiency by adding a union premium to the competitive market wage. Union presence may also limit management's flexibility in personnel decisions by introducing rules such as seniority in hiring and firing (Freeman & Medoff, 1984, p. 164). Furthermore, any form of industrial unrest will affect productivity adversely by temporarily reducing the utilization of the firm's resources and causing uncertainty about output levels (Caves, 1980; Flaherty, 1987). However, the direction of causation is not obvious, as poor labour productivity could reflect poor management, which also causes more industrial action (Addison & Schnabel, 2003, p. 123). Unions may also harm productivity by lowering investment, as shareholders' expected return is reduced by the risk of *ex post* rent-seeking by unions in the absence of binding contracts (Grout, 1984). Union rent-seeking could thus be considered a tax on the return on investments, potentially hampering innovation and technological development (Connolly et al., 1986). Finally, militant unions may disrupt industrial relations. If both employers and employees are only concerned with promoting their own interests, both may be worse off in terms of productivity and earnings than if they cooperated. In this regard, Freeman and Medoff (1984) argued that unions would only raise productivity if *'industrial relations are good, with management and unions working together to produce a bigger "pie"'* (p. 165).

However, many authors have argued that unions may promote productivity through institutional channels. Freeman (1976) and Freeman and Medoff (1984) claim that by providing workers with a means of expressing discontent through a collective voice, unions can reduce turnover and improve morale, motivation, job satisfaction and cooperation, thereby enhancing productivity. The additional information provided by a collective voice can moreover enable firms to choose a better mix of working conditions, workplace rules and wage levels (Laroche, 2020). In Norway, for example, the management and the union in firms participating in collective agreements can agree on more flexible working time arrangements than are otherwise permitted by law. A potential means of offsetting efficiency losses may thus arise if unions are able to induce managers to alter methods of production and adopt policies that improve efficiency. Unions may also give workers an increased experience of fairness because their presence reduces the potentially arbitrary nature of decisions about promotions and layoffs. That is, the union may act as '*the employees' auditor of management, checking that the employer is fulfilling his part of the labour contract'* (Pencavel, 1977, p. 141). Moreover, unions may contribute to higher productivity through the wage channel. By using their monopoly power to raise wages, unionized firms may attract more productive employees (Lazear, 2000). It is also plausible that the wage differentials between unionized and non-unionized firms may reduce turnover in unionized firms, thereby saving them potential firing and hiring costs, as well as conserving firm-specific human capital. Higher wages may give employers incentives to replace some labour by capital, which, although not socially efficient, will increase labour productivity at the firm level (Freeman & Medoff, 1984, p. 164).

Many attempts have been made to estimate empirically how unions influence productivity. The pioneering study of Brown and Medoff (1978) is one of the few studies that finds a large and positive effect of unions on productivity in the U.S. manufacturing industry. However, these estimates were later attributed to serious data limitations (Hirsch & Addison, 1986). Other studies from the United States have found both positive and negative union effects on productivity, with large variations across sectors and industries (Allen, 1988; Clark, 1980). A recent meta-analysis by Doucouliagos et al. (2017) reviews a large number of studies published over the last 30 years that address the impact of unions on productivity. The overall association between unions and productivity is shown to be near zero, but the relationship varies significantly across countries and industries. The authors indicate that the wide diversity of findings makes it hard to adopt a definitive position on what unions do to productivity: it depends on the period of analysis, the industry, the nature of the social climate in both the specific country and the firm, methods of data collection, the productivity indicator used and the econometric frameworks adopted.

It is apparent that the question of what unions do to productivity is far from resolved. To better understand the empirical ambiguity, the literature has considered various mechanisms that might be at play. There is an extensive literature examining the relationship between unionization, job satisfaction and productivity. In a meta-analysis of 235 estimates from 59 studies published between 1975 and 2015, Laroche (2016) finds an overall small negative association between unionization and job satisfaction. However, the study shows that the industrial relations climate has a positive and significant impact on the union-satisfaction effect. Moreover, when taking account of the possibility that unions often organize in firms with poor working conditions, Blanchflower and Bryson (2020) find a positive relationship between unions and several measures of worker well-being, including job satisfaction. Others have investigated how organizational commitment can be a channel through which unions affect productivity. Several studies show a positive correlation between measures of job performance and workers' organizational commitment (Jaramillo et al., 2005; Mathieu & Zajac, 1990), which has been found to be positively related to unionization in the United States and Norway (Kalleberg & Mastekaasa, 1994).

Another strand of literature has looked at how the institutional context in which unions operate affect the way they function (Blanchflower & Freeman, 1992). The focus in these studies is the institutions that enable and constrain union efforts to improve working conditions. In the United Kingdom, Bryson et al. (2006) find that employee perception of managerial responsiveness to worker voice leads to superior productivity. In France, Coutrot (1996) shows that firms with at least one union delegate in the workplace are more productive than other firms. This finding is partly confirmed by Laroche (2004). In general, several studies have shown that measures of the industrial relations climate are positively associated with better economic performance (Belman, 1992; Whitman et al., 2010). As suggested by Freeman and Medoff (1984), unions can improve the quality of labour relations by cultivating voice rather than exit.

A particular feature of the institutional context that has received less attention is the role of collective agreements. Notable exceptions are García-Serrano (2009) and Bryson et al. (2010), who separate the roles of union membership and firm-level collective agreements in their assessment of how unions affect job satisfaction in Spain and the United Kingdom. In a recent study from Belgium, Garnero et al. (2020) investigate how firm-level collective agreements affect firm performance in a multi-level bargaining system. They find that firm agreements increase both wage costs and labour productivity. However, this result must be interpreted within the context of the Belgian national bargaining system, where firm-level agreements act as supplements to agreements at sectoral level, which cover practically the entire Belgian workforce (p. 945). In another recent study, Barth et al. (2020) identify a large positive impact of UD on productivity in Norway. By exploiting exogenous variation in the rules for the tax deductibility of union membership fees, the study is one of a limited number that handle the possibly endogenous behaviour of unionization. The authors interpret the large coefficient as a threshold effect, where the union forces the employer to implement a collective agreement once the UD reaches a particular threshold. However, they do not have information in their data to further investigate this hypothesis.

Our contribution expands the current knowledge of what unions do to productivity in general, and in particular how this relationship is affected by the quality of industrial relations as measured by the presence of a collective agreement. Moreover, our article adds an important contribution to the very limited number of studies providing causal evidence of what unions do to productivity. Although we do not fully control for the possible correlation between productivity shocks and a decision to enter or exit a collective agreement, we provide a new source of exogenous variation in union memberships by utilizing information on intergenerational transmission of union preferences.

3 | UNIONIZATION AND COLLECTIVE AGREEMENTS IN NORWAY

The relationship between employers and employees in Norway is organized through an interaction between legislation and collective agreements, where the importance of the latter is relatively high compared to other countries. The labour market is characterized by strong trade unions and employer's associations. During the last decade, UD has been stable at around 50 per cent, or 38 per cent if we consider the private sector only. In the same period, the organization rate among private sector employers has been steadily increasing and amounted to approximately 71 per cent in 2019 (Alsos et al., 2021).

As shown in Figure 1, UD in Norway is high compared to most other countries in the OECD, and so is the prevalence of collective agreements.[1] About 10 per cent of Norwegian private sector firms participate in collective agreements, which accounted for 46 per cent of all private sector workers in 2018. If we include the public sector, almost 73 per cent of all workers were covered by collective agreements in 2018. However, the coverage rate is lower than in many other Western European countries, where collective agreements at sectoral level may be required by law to extend to all firms and workers irrespective of union membership.[2]

Collective bargaining in Norway has a clear hierarchical structure. As in several other Western European countries, wages in the private sector may be negotiated at three levels: central, sectoral and local. At the national level, a few major confederations determine the content of the basic agreements. The basic agreements form the basis for all lower level agreements in specific industries, set the framework for bargaining and regulate issues such as rights to information and consultation and rules for taking industrial action (most importantly strike and lock-out).

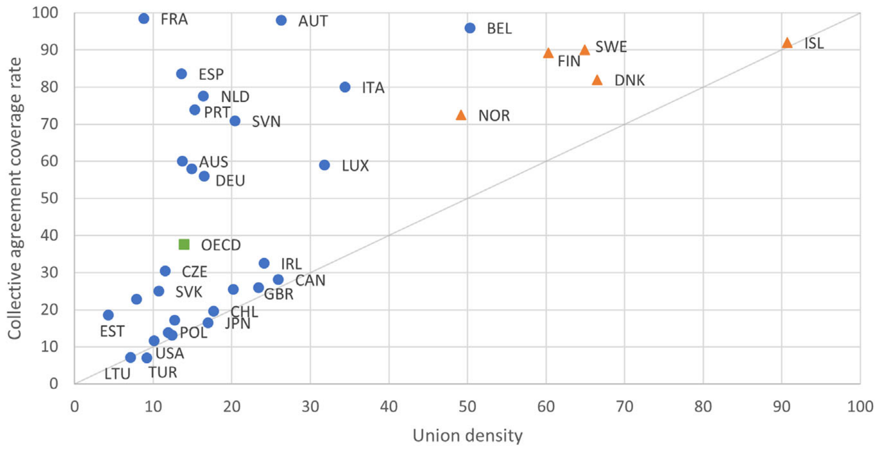

FIGURE 1 Union density and collective agreement coverage in OECD countries. 2018 or last observation. Nordic countries and OECD average are highlighted by orange triangles and a green square, respectively. *Source*: OECD databases on 'Trade union density' and 'Collective bargaining coverage' [Colour figure can be viewed at wileyonlinelibrary.com]

Moreover, the basic agreements include procedures for electing employee representatives, which are important for facilitating the firm-level relationship between employees and employers. The second level in the hierarchy consists of agreements for specific industries, often referred to as business sector agreements. Most of these agreements include the text of the corresponding basic agreement as their first section. The second part typically contains provisions regarding minimum wage and entitlements regarding working hours, overtime payment and welfare leave. Business sector agreements normally apply for 2 years at a time.

Local agreements between employers and employee representatives at company level, which are adapted to local conditions, make up the third level of the bargaining hierarchy. In contrast to sectoral agreements, local agreements automatically extend to all workers in occupations covered by the agreement, irrespective of union membership.[3] CAC in Norway thus depends on the existence of local agreements. In general, if the UD among workers within the same bargaining area in a firm is above a certain threshold, the union will demand a collective agreement. If the employer is organized in an employer's association, the agreement will be ratified more or less automatically. If the employer is not organized, the trade union will enter a direct agreement with the employer—if necessary, through the use of industrial action.

A particular feature of the Norwegian system of collective agreements is that the basic agreements include extensive provisions on co-determination. Specifically, the agreements introduce regulations designed to strengthen and further develop the collaboration between employees, their representatives and the management. Furthermore, they formalize the mutual responsibility of employer and employees for productivity growth and business development (Bergh, 2010). The presence of a collective agreement thus constitutes an important institutional feature when evaluating what unions do to productivity and other aspects of corporate performance.

In short, collective agreements in Norway are not only a means of regulating observable working conditions such as wages and hours; they also establish and codify a system of collaboration, communication and participation, with the explicit purpose of enhancing productivity. The clear focus on co-operation in the collective agreements partly reflects and partly contributes to sustaining the long Nordic tradition of close co-operation between employers' associations and trade

TABLE 1 Observations by collective agreement coverage

	Observations	Firms
Never collective agreement	969,614	158,630
Always collective agreement	88,918	9228
Firms changing status	112,138	10,520
Total	1,170,670	178,438

unions, as well as a high degree of co-determination and participation at company level. A better understanding of the interplay between unions, collective agreements and firm performance is thus paramount when investigating how unions affect productivity in Norwegian firms.

4 | DATA

The empirical analysis utilizes a matched employer–employee dataset, obtained from Statistics Norway (see Table A1 in the Appendix for descriptive statistics). The data cover the Norwegian private sector in the period 2002–2018 and consist of individual data collected by the Norwegian Tax Authorities and Social Services, matched with several other sources of register data related to both firms and employees. The most important data source for the period 2002–2014 is the State Register of Employers and Employees (the 'Aa-register'), which provides information on income, earnings, hours worked and occupation for each individual. For the remaining years, 2015–2018, our data are collected from the 'a-ordning', a coordinated service used by employers to report information about income and employees to the Norwegian Labor and Welfare Administration, Statistics Norway and the Norwegian Tax Administration. Educational statistics are attached, as well as firm-level financial data and several other characteristics. Every individual, workplace and firm has its own unique identifying number, enabling us to track the units across time.

Whether a firm participates in a collective agreement or not is obtained from membership data from the private sector collectively agreed pension scheme ('Fellesordningen for AFP'- the AFP scheme), whereby all firms that are members are also parties to a collective agreement.[4] In a model with firm fixed effects, identifying the effect of a collective agreement requires sufficient time variation in this variable. Although most firms do not change their status during the period in question, Tables 1 and 2 document substantial variation in CAC within firms. On average, 448 firms enter a collective agreement each year, while 275 firms exit. In total, this amounts to 112,138 observations of, in total, 10,520 firms changing their coverage status at least once.

Individual union membership is obtained from data on union membership fees, which are reported to the tax authorities by the unions. From the membership payments, we calculate firm-level UD as the ratio of union members to the number of employees in each firm. Figure 2 shows how the two variables UD and CAC evolve through our period of analysis. While the solid lines show unweighted firm averages, the dashed lines are weighted averages, where the number of employees in each firm are used as weights. They thus illustrate UD and CAC across firms and individuals, respectively. The differences between the weighted and unweighted means reflect the fact that UD and the prevalence of collective agreements are higher among large firms (see Figure A1 in the Appendix).

Our initial individual level dataset contains around 1.5 million yearly private sector jobs, amounting to 25.2 million observations in total for the whole period. The number of yearly jobs is not equal to the number of individuals, as one individual may have multiple jobs within the same

TABLE 2 Observations of entries and exits from collective agreements

	Entry	Exit
2002	578	552
2003	466	169
2004	409	507
2005	460	182
2006	474	388
2007	412	165
2008	453	400
2009	409	165
2010	893	363
2011	439	165
2012	413	318
2013	388	167
2014	353	307
2015	349	157
2016	335	270
2017	383	115
2018	407	286
Total	7621	4676

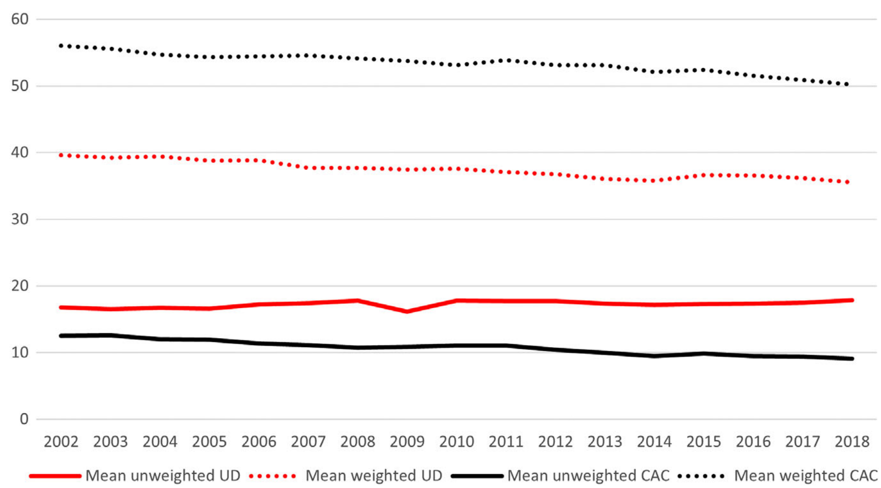

FIGURE 2 Mean union density (UD) and collective agreement coverage (CAC) unweighted and weighted by the number of employees in the firm, in our sample [Colour figure can be viewed at wileyonlinelibrary.com]

year. The total number of firms present in the initial sample is 334,511. However, we have placed some restrictions on the sample. Firms not required by law to provide financial statements, or which for other reasons do not have financial information, are excluded. This restriction leaves us with 20.9 million observations, amounting to just under 80 per cent of all private sector jobs.[5] The individual-level data are then aggregated to firm level using firm-based averages of job and worker

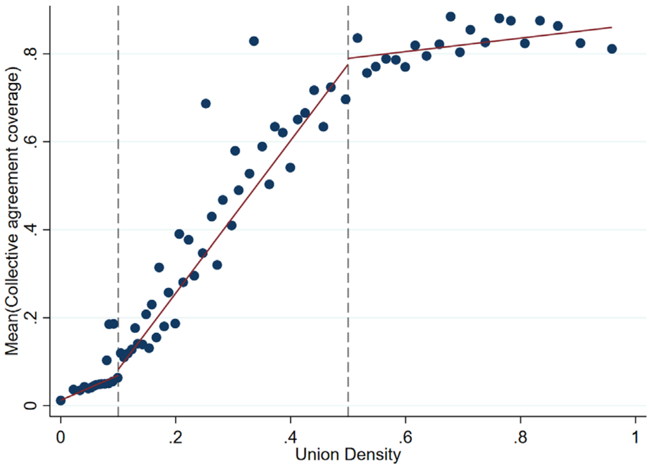

FIGURE 3 The distribution of collective agreements across union density in our sample. Firms with at least 10 employees. Binscatter, 88 bins. $N = 383{,}297$ [Colour figure can be viewed at wileyonlinelibrary.com]

information. The final estimation sample consists of 189,900 firms (corresponding to 58 per cent of all private sector firms, employing 75 per cent of all wage-earners), with a total of 1,170,670 firm-year observations. Because firms are established and dissolved throughout the period of analysis, our panel is unbalanced. Around 10 per cent of the firms are present in all 17 years, while the median number of observations per firm is 5 years. Firms with less than two observations are excluded from most estimations.

The interaction between UD and the presence of a collective agreement is of primary concern in our study. To qualify for a collective agreement, the UD among the firm's workers must exceed a certain threshold. In the largest basic agreement in Norway, this threshold is specified as 10 per cent of the workers.[6] Figure 3 illustrates the distribution of collective agreements as a function of firm-level UD in firms with at least 10 employees. The figure clearly shows a positive relation between unionization and the presence of collective agreements. The lines at 10 and 50 per cent represent two common thresholds where the union may demand a collective agreement. The relationship appears to have a steeper slope when UD passes 10 per cent, indicating an acceleration in the accumulation of collective agreements. When the firm unionization rate exceeds 50–60 per cent, most firms have implemented an agreement.

5 | METHODOLOGY

Productivity can be measured in many ways, with the various methods being confounded by a range of issues (Syverson, 2011). In the following, we use total factor productivity as our measure, in line with Barth et al. (2020). As demonstrated in the Appendix, however, our main conclusions are robust to the choice of productivity measure. As a change in total factor productivity reflects variations in output that cannot be ascribed to observable variation in factor inputs, we use a production function to estimate output conditional on the use of labour and capital. Our theoretical reference point is a skill-augmented production function specified as Cobb–Douglas, which in log-transformed notation is represented by:

$$y_{it} = \alpha + \beta_j k_{it} + \boldsymbol{l_{it}}\phi_j + \gamma_1 UD_{it} + \gamma_2 CA_{it} + \gamma_3 (UD \times CA)_{it} + \boldsymbol{X_{it}}\delta + u_i + \lambda_t + \omega_{it} + \varepsilon_{it}, \quad (1)$$

where y_{it} and k_{it} denote the value added and capital stock, respectively, of firm i in year t, both measured by their natural logarithms. Labour is divided into four skills groups determined by educational attainment, denoted by the row vector l_{it}, and is measured by the log aggregated weekly number of hours worked within each group.[7] The stock of capital and the number of hours worked both represent a measure of firm size, which is strongly correlated with the presence of a collective agreement (see Figure A1 in the Appendix).

The partial elasticities of output with respect to capital and labour are allowed to vary across industries j, as represented by the coefficient β_j and the labour coefficient vector ϕ_j. u_i denotes firm fixed effects, while λ_t represents time-specific effects reflecting both nominal and real trends common to all firms. ω_{it} represents unobservable idiosyncratic productivity shocks known to the firm, while ε_{it} represents measurement errors or random productivity shocks truly unknown to both firms and researchers, assumed to be normally distributed and i.i.d. The model equation is further augmented with our primary variables of interest, which are added successively to the estimated equation: workplace union density (UD_{it}), a dummy variable capturing the presence of a collective agreement (CA_{it}) and a term interacting UD with the presence of a collective agreement. Finally, the model is saturated with a vector of control variables (X_{it}) reflecting demographic, occupational and industry-by-year interactions.

We estimate Equation (1) to identify the impact of union presence on firm-level productivity. Our main parameters of interest are γ_1, γ_2 and γ_3. The marginal effect of an increase in UD is $\gamma_1$for firms without a collective agreement ($CA = 0$) and $\gamma_1 + \gamma_3$ in firms with an agreement ($CA = 1$). The effect of implementing a collective agreement is given by $\gamma_2 + \gamma_3 \times UD$, which may be evaluated for different values of UD.

Our strategy to identify the productivity effect of unionization is not without challenges. Any unobserved heterogeneity across firms will make the ordinary least squares (OLS) estimator inconsistent. We therefore estimate the model using the within estimator that allows for firm fixed effects. However, a key identifying assumption in the fixed effects model is the absence of any idiosyncratic productivity shocks correlated with UD or the presence of a collective agreement. This assumption is violated if, for example, the decision to implement or abolish a collective agreement is taken systematically at a specific stage of the firm's life, and if the firm is moving along a productivity path that would imply higher or lower productivity after this stage irrespective of the presence of the agreement. As the presence of a collective agreement is measured by a dummy variable, which takes on the value 0 for all years before the implementation and 1 as long as the firm participates in the agreement, any such systematic covariation will bias $\hat{\gamma}_2$.

Moreover, as first noted by Marschak and Andrews (1944), the firm's demand for factor inputs is likely to depend on idiosyncratic productivity shocks known to the firm, but unobservable to the econometrician. This is represented by the ω_{it} term in (1) and may, for example, represent the quality of machines and equipment not reflected in the book value of fixed assets. Such (to the firm) observables, and the omission of these by the econometrician, will in general make both the OLS estimator and the within estimator biased and inconsistent, as factor inputs are endogenously determined together with production. However, as proposed by Olley and Pakes (1996) and further developed by Levinsohn and Petrin (2003) and Wooldridge (2009), the issue of idiosyncratic productivity shocks may be handled by forming a control function where a polynomial in investments and/or intermediate inputs is used to proxy such unobserved productivity differences between individual firms.

A more serious problem of selection bias, however, relates to the potential endogenous determination of UD. The presence of a union is likely to not only *affect* but also *reflect* a firm's performance. The individual workers' decisions on whether or not to unionize may depend on the firm's performance in several ways (Barth et al., 2020; DiNardo & Lee, 2004). On the one hand, the scope for rent sharing is larger in highly profitable firms than in less profitable ones. On the other hand, as unions are usually considered to improve the protection of workers and workers' rights, workers may seek unionization as a matter of job security if productivity is declining.

To identify the impact of union presence on firm-level productivity, as we discuss in more detail below, we instrument UD among the workers in a workplace by the UD among their parents. As we show in Subsection 6.4, parental unionization behaviour has a strong impact on an individual's propensity to join a union. Intergenerational transmission of union membership thus provides a source of exogenous variation in analyses relating unionization to the performance of firms. It is highly unlikely for parents to unionize as a result of changes in performance at their children's workplace, and the variation in parents' union memberships could thus be considered a valid instrument for the individual's decision of whether or not to join a union. One important exception, however, is the case where parents work in the same firm as their children. In such a case, changes in the firm's performance will alter the unionization incentives of both the workers and the workers' parents in a similar manner. This situation may be of particular relevance in sparsely populated areas with one or a few major employers.

Our identification strategy rests on the assumption that any selection bias in the implementation or abolishment of a collective agreement is effectively controlled for by handling the potentially endogenous nature of unionization. In general, this assumption is not likely to hold. Although a collective agreement will often come into place following a recruitment process that results in increasing UD, this is not always the case. In some firms, the UD may be above the threshold required for the union to enter an agreement, without the workers wanting to do so. Furthermore, the decision to enter or exit a collective agreement ultimately depends on the signature of the manager, who is not obliged by law to sign the agreement. As argued in the introduction, the presence of a collective agreement must therefore be treated as a separate and independent dimension of the union's presence in the firm, as must any endogenous decision on whether or not to enter or exit an agreement. The possible selection bias arising from not fully controlling for this problem thus represents a caveat in our study.

6 | RESULTS

Table 3 summarizes the results of estimating Equation (1) by means of different estimators. In Model 1a, we employ the within transformation of Equation (1) to allow for firm fixed effects (FE), which effectively controls for any unobserved time-invariant heterogeneity across firms. In this model, we assume (for the moment) homogeneous input elasticities across industries, and union presence is measured by UD alone. As UD is measured as a rate between 0 and 1, the corresponding estimated coefficient implies a 0.11 per cent increase in productivity from a 10-percentage-point increase in UD. The effect is only significant at the 10 per cent level.[8]

In Model 1b, we include a dummy variable that captures whether the firm is engaged in a collective agreement or not, and in Model 1c we add a term for the interaction between workplace UD and the existence of an agreement. This completely alters the interpretation of how productivity is affected by the presence of a union. To facilitate interpretation, we have included the derived effects of implementing a collective agreement evaluated on average UD, as well as the marginal effects of an increase in UD with and without a collective agreement. When both variables are included in Model 1c, a 10-percentage-point increase in UD is estimated to *reduce* productivity by 0.3 per cent in the absence of a collective agreement. If the firm is covered by a collective

TABLE 3 Estimated effects of union density and collective agreements on total factor productivity

	Model 1a FE	Model 1b FE	Model 1c FE	Model 1d FE	Model 1e LPW-GMM	Model 1f LPW-GMM	Model 1g LPW-GMM
Union density (UD)	0.011	−0.013*	−0.026***	−0.025***	−0.098***	−0.130***	−0.138***
	(1.83)	(−2.29)	(−4.19)	(−4.02)	(−14.30)	(−12.87)	(−8.56)
Collective agreement (CA)		0.157***	0.117***	0.117***	0.002	−0.019*	−0.016
		(25.79)	(14.46)	(14.47)	(0.18)	(−2.22)	(−1.53)
UD × CA			0.102***	0.099***	0.159***	0.206***	0.197***
			(7.05)	(6.83)	(10.20)	(12.26)	(8.95)
Marginal effects of:							
UD for $CA = 0$			−0.026***	−0.025***	−0.098***	−0.130***	−0.138***
			(−4.20)	(−4.02)	(−14.30)	(−12.87)	(−8.56)
UD for $CA = 1$			0.076***	0.074***	0.061***	0.076***	0.059**
			(5.59)	(5.44)	(4.17)	(5.19)	(3.28)
CA for $\overline{UD}$			0.135***	0.134***	0.028***	0.024***	0.035***
			(20.14)	(20.08)	(4.07)	(3.53)	(4.52)
Test (p-value): $\hat{\gamma}_1 + \hat{\gamma}_3 = 0$[a]			0.000	0.000	0.000	0.000	0.000
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Occupation and demographics				Yes	Yes	Yes	Yes
Minimum number of employees						5	10
R^2 (within)	0.260	0.261	0.261	0.266	0.075	0.098	0.085
R^2 (between)	0.610	0.613	0.613	0.589	0.048	0.129	0.132
R^2 (overall)	0.654	0.657	0.657	0.640	0.059	0.124	0.122
N	1,109,883	1,109,883	1,109,883	1,109,842	942,084	525,791	282,417
Firms	173,257	173,257	173,257	173,247	152,683	83,536	45,168
Average observations per firm	6.4	6.4	6.4	6.4	6.2	6.3	6.3

Note: Union density measured as a rate between 0 and 1. Collective agreement measured as a dummy variable. All estimations include year dummies. Demographics include age intervals, sex and country of origin. t-Statistics are in parentheses. Models 1e and 1f use as regressand the residuals from an Levinson-Petrin-Wooldridge-GMM (LPW-GMM) estimation of value added on capital and labour inputs only. Input elasticities reported in Table A2 in the Appendix.

[a]The reported test refers to the p-value of an F-test of the sum of the coefficients on UD and UD × CA. Robust standard errors are clustered at firm level.

$^{*}p < 0.05$,

$^{**}p < 0.01$,

$^{***}p < 0.001$.

agreement, however, a similar increase in UD is estimated to *increase* productivity by 0.8 per cent. Furthermore, the implementation of a collective agreement in a firm with an average UD is estimated to increase firm productivity by 13.5 per cent. However, this estimate is likely to be biased upwards, a point we will return to below.

6.1 | Unobserved idiosyncratic productivity shocks

In Model 1d, we control for heterogeneity in workers' skills (other than educational attainment level) by including occupational shares at the 1-digit (ISCO 08) level, as well as demographic characteristics such as age, sex and immigration status. However, this enlargement of the specification has no significant effect on the estimated coefficients, which remain robust. In Table A3 in the Appendix, we demonstrate how using labour productivity as our endogenous variable produces similar results.[9] In Model 1e, however, we consider how factor inputs may be endogenously determined in the production function by allowing time-varying idiosyncratic productivity shocks, represented by ω_{it} in (1). Applying the generalized method of moments (GMM) estimator proposed in Wooldridge (2009),[10] we first estimate the production function, including capital and labour inputs only, where unobserved productivity is proxied by a third-order polynomial in intermediate inputs. We then use the residuals from this regression, which acts as a measure of total factor productivity, as regressand in the fixed-effects model. When this approach is employed, the effect of implementing a collective agreement drops sharply, suggesting that the estimated effect above is partly caused by idiosyncratic productivity shocks correlated with the decision to implement or abolish a collective agreement. However, the presence of a collective agreement still constitutes an important factor in understanding how unionization alters productivity. The implementation of a collective agreement, evaluated on average UD, is estimated to increase productivity by 2.8 per cent. Moreover, while a 10-percentage-point increase in UD is estimated to *decrease* productivity by almost 1 per cent in the absence of an agreement, a similar increase in UD in the presence of an agreement is estimated to *increase* productivity by 0.6 per cent.

The influence of collective agreements may be limited in small organizations. In Models 1f and 1g, we therefore constrain our sample to firms with at least 5 and 10 employees, respectively, to make sure our results are not driven by variation generated by small firms. The restriction is not trivial, as the models then exclude 43 and 69 per cent of the firms in our sample. Nevertheless, the results remain robust and somewhat strengthened.

6.2 | Industry heterogeneity

To control for industry heterogeneity, we start by including industry-by-year interactions to capture potential heterogeneity in technological trends across industries.[11] The results are presented in Model 2a of Table 4. In Model 2b, we expand the scope for industry heterogeneity by relaxing our previous assumption of homogeneous input elasticities. Specifically, we use the residuals from industry-specific GMM estimations as left-hand side variables in the fixed effects model, thereby recognizing heterogeneous capital and labour elasticities while assuming the impact of unions on productivity to be homogeneous. This more flexible specification changes the results slightly, but the overall pattern remains quite robust.

Finally, in Models 2c–2i, we present the results of separate GMM estimations for selected groups of industries.[12] Most noteworthy is how robust the interaction term is estimated across most industries. Although higher UD is estimated to lower productivity in the absence of a collective agreement, this effect is moderated, and in many cases becomes positive, in the presence of an agreement. Moreover, the implementation of a collective agreement is estimated to increase productivity in all industries but professional services (evaluated at average UD), with an estimated elasticity ranging from 1 to 10 per cent.

TABLE 4 LPW-GMM estimates of union density and collective agreements on total factor productivity allowing for various forms of industry heterogeneity

	Model 2a	Model 2b	Model 2c	Model 2d	Model 2e	Model 2f	Model 2g	Model 2h	Model 2i
Union density (UD)	−0.099***	−0.103***	−0.139***	−0.149***	−0.130***	−0.127***	−0.129**	−0.032	−0.135***
	(−14.43)	(−15.44)	(−5.52)	(−8.97)	(−10.17)	(−4.88)	(−3.15)	(−1.37)	(−3.78)
Collective agreement (CA)	0.002	0.018*	0.024	0.022	0.026	0.080**	−0.048	−0.037	−0.036
	(0.23)	(2.25)	(1.19)	(1.28)	(1.92)	(3.08)	(−0.84)	(−0.77)	(−0.81)
UD × CA	0.148***	0.130***	0.118**	0.199***	0.109***	0.136*	0.335**	0.0944	0.288**
	(9.60)	(8.70)	(3.05)	(5.97)	(3.92)	(2.18)	(3.17)	(1.23)	(3.29)
Marginal effects of:									
UD for $CA = 0$	−0.099***	−0.103***	−0.139***	−0.149***	−0.130***	−0.127***	−0.129**	−0.032	−0.135***
	(−14.43)	(−15.44)	(−5.52)	(−8.97)	(−10.17)	(−4.88)	(−3.15)	(−1.37)	(−3.78)
UD for $CA = 1$	0.050***	0.027	−0.021	0.050	−0.021	0.010	0.207*	0.063	0.153
	(3.44)	(1.91)	(−0.66)	(1.66)	(−0.82)	(0.17)	(2.04)	(0.83)	(1.87)
CA for $\overline{UD}$	0.030***	0.040***	0.051**	0.048**	0.040***	0.097***	0.026	−0.014	0.006
	(3.90)	(6.02)	(3.17)	(3.16)	(3.35)	(4.16)	(0.61)	(−0.41)	(0.16)
Test (p-value): $\hat{\gamma}_1 + \hat{\gamma}_3 = 0$[a]	0.001	0.056	0.507	0.097	0.414	0.864	0.041	0.404	0.062
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Occupation and demographics	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry by year dummies	Yes	Yes							

(Continues)

TABLE 4 (Continued)

	Model 2a	Model 2b	Model 2c	Model 2d	Model 2e	Model 2f	Model 2g	Model 2h	Model 2i
Heterogeneous input elasticities		Yes							
Industry	All	All	Manufactu-ring	Construc-tion	Sales & Retail	Hotel & Restau-rants	ICT	Professional service	Industrial service
N	941,995	941,969	101,571	187,111	291,121	59,801	33,950	65,923	34,472
Firms	152,658	152,651	13,363	30,272	42,934	11,327	6643	14,287	7225
Average observations per firm	6.2	6.2	7.6	6.2	6.8	5.3	5.1	4.6	4.8

Note: Union density measured as a rate between 0 and 1. Collective agreement measured as a dummy variable. Industries are divided into 19 groups. Demographics include age intervals, sex and country of origin. All models use as regressand the residuals from LPW-GMM estimation of value added on capital and labour inputs only. Standard errors of marginal effects are calculated using the delta method.

[a]The reported test refers to the *p*-value of an *F*-test of the sum of the coefficients on UD and UD × CA. *t*-Statistics are in parentheses. Robust standard errors are clustered at firm level.

*$p < 0.05$,

**$p < 0.01$,

***$p < 0.001$.

We estimate that the implementation of a collective agreement, also in manufacturing industries, has a positive and significant effect on productivity. As firms operating in manufacturing industries are generally exposed to international competition—especially in a small, open economy like that of Norway, their market power is limited. This suggests that our findings are not merely price effects caused by firms passing on the union wage premium to consumers, which is a general concern in studies using value measures of output (Freeman & Medoff, 1984, p. 167).

6.3 | Further investigation of robustness

Our estimations rely on an unbalanced panel of observations, with new firms entering and others exiting the sample along the period of analysis. On the one hand, an unbalanced panel eliminates the potential bias caused by low-productivity firms going into bankruptcy. On the other hand, the productivity effect of collective agreements and unionized workers may differ systematically between new entrants and existing firms in the market. It is therefore interesting to investigate how our results are influenced by imposing various restrictions on the sample of included firms.

Table A4 in the Appendix shows the results of estimating Model 2b with only firms present all years, only entrant firms and only entrants that remain in our sample, respectively. The effect of implementing a collective agreement evaluated at average UD is estimated to be close to zero when only firms that are present all years are considered. The estimated positive impact of collective agreements on productivity thus seems to be driven mainly by new market entrants during our sample period. However, our main result that collective agreements act as an important moderator of what unions do to productivity, remains robust across all the mentioned restrictions.

We also investigate how our results are affected by only including firms with or without changes in their CAC throughout the sample. In Table A4 in the Appendix, we first restrict the sample to firms that always or never, respectively, are covered by a collective agreement. Although the effect of collective agreements naturally cannot be identified under these restrictions, we note that an increase in UD is estimated to *reduce* productivity in firms never covered by an agreement but to increase productivity in firms that are always covered. Although the latter estimate is not significantly different from zero, the results are consistent with our prior findings. We further restrict our sample to firms that do not change coverage status and firms that do change coverage status, respectively, during our sample period. Once again, our results prove to be robust to these restrictions. Compared to the results in Model 2b, the estimated effect of implementing a collective agreement, when evaluated at average UD, is stronger when only firms that do change status are considered. This is consistent with the above finding that this effect mainly seems to be driven by new entrants to the market, as the propensity to change coverage is higher among entrant firms.

Finally, we explore the importance of the linearity assumption implicitly imposed in our estimations. In general, there is no reason why an increase in UD from 10 to 20 per cent should have the same effect on productivity as an increase from 80 to 90 per cent. In Table A5 in the Appendix, we show the results of estimating Model 2b with UD measured as a categorical variable split into five equal intervals. Each UD interval is included in the estimation, as well as the interaction between each interval and the presence of a collective agreement. This exercise reveals that UD has a nonlinear effect on productivity. The estimated effects of going from a UD below 20 per cent to a UD between 20 and 40 per cent, between 40 and 60 per cent, or between 60 and 80 per cent is in fact very similar. In other words, the negative productivity effect of unionization is estimated to be the same whether the UD goes from 0 to 20–40 per cent or from 0 to 60–80 per cent. The linearity assumption seems more reasonable when we consider the evaluated effect of implementing a collective agreement conditional on different levels of UD, which is also illustrated in Figure A2 in the Appendix. Overall, taking nonlinearity into consideration does not alter our prior findings in any significant way. If anything, our results are strengthened.

6.4 | Endogenous unionization

We may worry that our above estimates are confounded by selection bias, as unionization may be endogenously determined by the performance of the firm. To overcome this issue, we apply an instrumental variable (IV) approach, instrumenting UD at the workplace with the UD among the workers' parents. In the next section, we explore this instrument further, before we continue with the firm-level analysis in Subsection 6.3.2. Importantly, however, and as discussed in Section 5, the possible correlation between productivity shocks and the decision to enter or exit a collective agreement, remains a caveat in our study, even after controlling for fixed effects and UD.

6.4.1 | Parental influence on individual propensity to join a union

It is widely recognized that the choices of the individual are affected by intergenerational transmission of preferences regarding political orientation (Jennings et al., 2009), education (Holmund et al., 2011) and receipt of welfare insurance (Dahl et al., 2014), to mention some. This is also the case for union membership. As demonstrated in Bryson and Davies (2018), the decision of young workers in Britain of whether or not to join a union is influenced by their parents' union membership. In particular, their study reveals that young workers are 29 per cent more likely to join a union if one of their parents is a union member, and 87 per cent more likely to join a union if both are union members, compared to individuals with no unionized parents (pp. 12–13).

In our sample, the probability that a given individual was a union member in 2018 was 26 per cent higher if at least one of their parents were union members, compared to an individual with no unionized parents.[13] Note that our sample of individuals with information on parents' union memberships averages approximately 500,000 individuals per year, compared to approximately 1.6 million individuals in our full sample. This mainly reflects the fact that parents are excluded from our data when they leave the labour force. In addition, individuals working with their parents are excluded from the analysis.

To gain a better understanding of how the unionization behaviour of individuals is influenced by their parents' union memberships, we estimate a simple linear probability model, where union membership is estimated as a function of parents' union membership. We then add a list of controls, including sex, age, occupation, the industry of their current occupation, education and immigration status, as well as year dummies. We also exclude individuals co-working with any of their parents.[14] The estimated partial effect of parental union memberships on an individual's unionization behaviour is reported in Model 3 of Table 5. [15] The result shows that the probability of being unionized is 6.7 percentage points higher for individuals with at least one unionized parent, compared to an individual with no unionized parents. Evaluated at UD among individuals with no unionized parents in 2018, this amounts to a 22.3 per cent increase in the probability of being unionized, which is same order of size as found among young British workers (Bryson & Davies, 2018).

TABLE 5 Linear probability model estimates of union density as a function of parents' union memberships

	Model 3
At least one parent unionized	0.068***
	(197.26)
N	7,969,901
R^2	0.134

Note: Endogenous variable: binary variable taking the value 1 if the individual is a union member and 0 if not. Included controls: sex, age, immigration status, occupation (1-digit ISCO-08), industry of current occupation (2-digit NACE), educational attainment level (1-digit ISCED 2011) and year dummies. Individuals working together with their parents are excluded. *t*-Statistics are in parentheses. Robust standard errors are clustered at firm level.
***$p < 0.001$.

Although a rigorous analysis of intergenerational transmission of union membership should be implemented using a more sophisticated identification strategy, our aim here is limited to documenting its relevance in the Norwegian labour market. The simple analysis presented shows a strongly significant and sizeable intergenerational relationship for unionization behaviour. Admittedly, we cannot rule out the possibility that this relationship works in the reverse direction, that is, that the unionization behaviour of children affects the parents' decision on whether or not to join a union. However, our result fits into a series of studies of how the decision of parents influence the preferences and choices made by their children.

6.4.2 | UD among parents as an instrument for workplace UD

Table 6 documents the estimation results when instrumenting workplace UD with the contemporary UD among the workers' parents. Although the effect of intergenerational transmission of union memberships naturally becomes weaker when moving from individual unionization decisions to UD at the firm level, it remains highly statistically significant and passes conventional tests for weak instruments by a good margin. In Model 4a, we re-estimate Model 1e from Table 1 using two-stage least squares (2SLS). Model 4b then adds linear industry trends and allows for heterogeneous input elasticities, using the residuals from industry-specific production function GMM estimations as values for the endogenous variable (referred to as GMM-IV). Finally, Models 4c and 4d restrict the sample to firms with at least 5 and 10 employees, respectively.

Overall, the IV estimates confirm our main result: the presence of a collective agreement significantly alters what unions do to productivity. However, although the presence of an agreement moderates the negative effect of an increase in UD, the effect remains negative (though not statistically significant). Moreover, the effect of implementing a collective agreement, evaluated at average UD, is only significant (at the 10 per cent level) when we restrict the sample to firms with at least 10 employees in Model 4d. However, the estimated coefficient values in Models 4b, 4c and 4d are comparable to the above GMM estimates. It is also important to emphasize that the IV estimator identifies the local average treatment effect (LATE) of unionization among *compliers*, which in general is not equal to the average treatment effect (ATE). The results in Table 4 and 6 are thus not directly comparable, as differences may be ascribed to either selection bias or treatment heterogeneity, or a combination of the two.

TABLE 6 IV estimates of the effects of union density and collective agreements on total factor productivity

	Model 4a	**Model 4b**	**Model 4c**	**Model 4d**
	2SLS	GMM-IV	GMM-IV	GMM-IV
Union density (UD)	0.0985	−0.764***	−0.695***	−0.982**
	(0.69)	(−5.16)	(−3.45)	(−2.72)
Collective agreement (CA)	−0.131	−0.079	−0.090	−0.159
	(−1.68)	(−0.99)	(−1.09)	(−1.69)
UD × CA	0.559**	0.661**	0.658**	0.928**
	(2.67)	(3.10)	(2.84)	(3.11)
Marginal effects:				
UD for $CA = 0$	0.0985	−0.764***	−0.695***	−0.982**
	(0.69)	(−5.16)	(−3.45)	(−2.72)
UD for $CA = 1$	0.657	−0.103	−0.037	−0.054
	(3.23)	(−0.50)	(−0.17)	(−0.19)
CA for $\overline{UD}$	−0.029	0.041	0.048	0.084
	(−0.67)	(0.92)	(1.13)	(1.73)
Test (p-value): $\hat{\gamma}_1 + \hat{\gamma}_3 = 0$[a]	0.001	0.620	0.869	0.851
Year dummies	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes
Occupation and demographics	Yes	Yes	Yes	Yes
Industry trends		Yes	Yes	Yes
Heterogeneous input elasticities		Yes	Yes	Yes
Minimum number of employees			5	10
N	704,314	704,314	490,776	275,139
Firms	118,441	118,441	78,740	43,840
Average observations per firm	5.9	5.9	6.2	6.3
Test for weak instruments (F-statistics):				
Kleibergen–Paap Wald	268.2	277.9	170.9	70.6
Cragg–Donald Wald	669.4	698.0	387.6	157.8

Note: Union density measured as a rate between 0 and 1. Collective agreement measured as a dummy variable. Union density is instrumented by the contemporary union density among the workers' parents. The interaction term is instrumented with the interaction between the collective agreement dummy and the instrument. Industries are divided into 19 groups. Demographics include age intervals, sex and country of origin. Union density instrumented by union density among parents in IV estimation.
[a]The reported test refers to the p-value of an F-test of the sum of the coefficients on UD and UD × CA. t-Statistics are in parentheses. Robust standard errors are clustered at firm level.
$^{**}p < 0.01$,
$^{***}p < 0.001$.

7 | DISCUSSION AND CONCLUSIONS

Overall, our results show that the qualitative interpretation of what unions do to total factor productivity depends on whether or not the firm is covered by a collective agreement. In the absence of an agreement, increases in UD among the workers in a firm are estimated to reduce productivity. However, the implementation of a collective agreement is estimated to moderate this negative impact. Moreover, when evaluated at average UD, the implementation of a collective agreement is estimated to increase productivity in most model specifications. Our findings thus give some support to the conclusions in Barth et al. (2020), but demonstrate the importance of taking account of the industrial relations climate when evaluating the impact of unionization on firm performance.

In general, there are good reasons to believe that the institutional framework encompassed in the collective agreements contributes to improving industrial relations in a firm. In the Norwegian context in particular, the agreements formally acknowledge the importance of the workers' voice and their contributions to productivity growth by establishing a system of collaboration, communication and participation. Furthermore, they regulate issues such as the right to information and consultation, procedures for electing employee representatives and rules for taking industrial action. Collective agreements thus represent an institutionalization of a particular way of managing industrial relations. In the absence of this institution, union activity may be more poorly organized and less predictable. Similarly, it may be difficult to utilize the productivity-enhancing potential of collective agreements in the absence of union activity. Based on our findings, UD and the presence of a collective agreement represent two necessary but insufficient conditions per se for releasing the productivity-enhancing effects of unionization. However, our results indicate that a sufficiently high UD and a collective agreement combined have a positive impact on firm-level productivity.

Despite the vast body of empirical literature investigating whether unions promote or impede productivity, the evidence is mixed and inconclusive. In this article, we have demonstrated the importance of recognizing institutional contexts when answering this question. In particular, we have argued that the presence of unions can be measured along two dimensions: the density of union members among employees, and the presence of a collective agreement. Such agreements act as a formal recognition of the policy put forward by the union and constitute an important organizational institution through which unions may alter industrial relations. However, little attention has been devoted to the study of collective agreements and their influence on what unions do to productivity.

Using matched employer–employee panel data, comprising almost 21 million individual-year and almost 1.2 million firm-year observations in the period 2002–2018, we have estimated how unions alter productivity at the firm level and how this effect is influenced by the presence of a collective agreement. Our main finding, which is robust across model specifications, is that the presence of a collective agreement significantly and positively alters what unions do to productivity. In most specifications, collective agreements are estimated to increase productivity. Moreover, across all specifications, collective agreements moderate the negative impact on productivity of increases in UD found in the absence of such agreements. However, care should be taken in interpreting our results, as the possible endogenous decision to enter or exit a collective agreement may bias our findings, even when controlling for firm fixed effects and endogenous unionization.

Our findings may reflect an interdependence between UD and collective agreements with respect to how they affect productivity. Although they may have a negative or insignificant impact on productivity in isolation, our results indicate that the combination of a sufficiently high UD and a collective agreement has a positive impact on firm-level productivity. Future research should investigate this interdependence further. In particular, it would be interesting to see an explicit attempt to model this complex relationship, especially within a dynamic framework.

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DATA AVAILABILITY STATEMENT

The data that support the findings of this study are available from Statistics Norway. Restrictions apply to the availability of these data, which were used under license for this study. Researchers affiliated with an approved research institution, or a public authority can apply to data from Statistics Norway (https://www.ssb.no/en/data-til-forskning/utlan-av-data-til-forskere).

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ENDNOTES

[1] However, union density is low compared to the other Nordic countries, where trade unions have traditionally administered the unemployment benefit funds, and thus have had better recruitment opportunities.

[2] This is the case in Austria, Belgium, Finland, France, Germany, the Netherlands and Portugal (García-Serrano, 2009). A comprehensive overview of the prevalence and functioning of collective agreements in the OECD, including differences in the practice of *ergo omnes* clauses and extensions are found in the OECD report 'Negotiating Our Way Up' (2019).

[3] There is an important exception. In industries where inflows of migrant workers have led to 'social dumping', general application of collective agreements is practised. However, such extensions are 'narrow' in the sense that they only include minimum wage rates and some basic supplements. The provisions in the basic agreements about co-determination (including the election of employee representatives), do not extend to all firms in an industry unless they have a local agreement in place.

[4] Some firms in the sample are covered by collective agreements, without being members of the AFP scheme. This mainly applies to enterprises in shipping and the oil industry and privately run health and social services. The firms in question are manually coded as covered if union density exceeds 50 per cent and the number of employees is at least 25.

[5] There are only small differences between firms in the initial and the final sample in union density, collective agreement coverage, average age and distribution across sex, education levels, occupations and industries. Overall, the final sample appears to be representative of the population of private sector employees.

[6] The premise of a threshold in the union membership rate is institutionalized in the Basic Agreement between the Confederation of Norwegian Enterprise (NHO) and the Norwegian Confederation of Trade Unions (LO) (Hovedavtalen § 3-7, nr. 2). This states that employees cannot require that the enterprise become part of a collective agreement without at least 10 per cent of the employees within the particular bargaining area being members of a union.

[7] Low-skilled labour comprises workers who do not complete upper secondary school, while medium-skilled corresponds to workers who have completed upper secondary school. High-skilled labour includes workers with a degree from up to 4 years of higher education and workers with at least 120 credits without a degree. Finally, top-skilled labour includes workers who have completed more than 4 years of tertiary education.

[8] Input elasticities are omitted from Table 3 for the sake of readability and reported in Table A2 in the Appendix. The drop in the estimated coefficients of capital and labour inputs when moving from the OLS estimator to the FE estimator reflects the common issue of estimating panel data production functions using micro data (Griliches & Mairesse 1999).

[9] Table A3 in the Appendix shows the results of estimating various models using labour productivity, measured as value added per hours worked, as endogenous variable. All models are estimated with firm fixed effects, year dummy variables and controls on individual worker characteristics. Note that the hours worked by employees with different skill levels are now included as shares among the controls, in contrast to the models presented in Tables 1, 2 and 4. The model is estimated with and without controls for hours worked and capital intensity. Theoretically, the model should include the total number of hours worked, as the assumption of constant returns to scale is rejected in our models. Overall, we find that our results are robust to the choice of productivity measure.

[10] The estimator is implemented using the -prodest- command in Stata with the Wooldridge (wrdg) estimator and the gmm option specified (Rovigatti & Mollisi, 2018). The estimator proposed by Levinsohn & Petrin (2003) produces almost identical results (not reported).

[11] Specifically, we add interactions between yearly time dummies and 19 main groups of industries.

[12] Results for all 19 main groups of industries are available upon request.

[13] Figure A3 in the Appendix compares the sample's union density among workers with and without unionized parents in a given year during our sample period.

[14] This restriction barely changes the result.

[15] Full estimation results are available upon request.

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APPENDIX

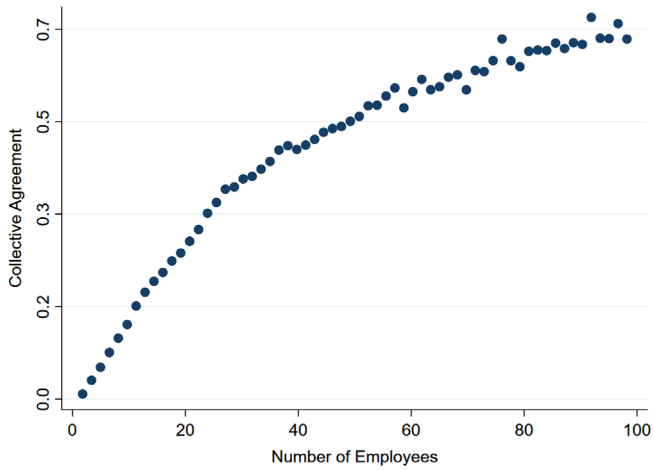

FIGURE A1 Binscatter illustrating mean collective agreement coverage as a function of the number of employees in the firm [Colour figure can be viewed at wileyonlinelibrary.com]

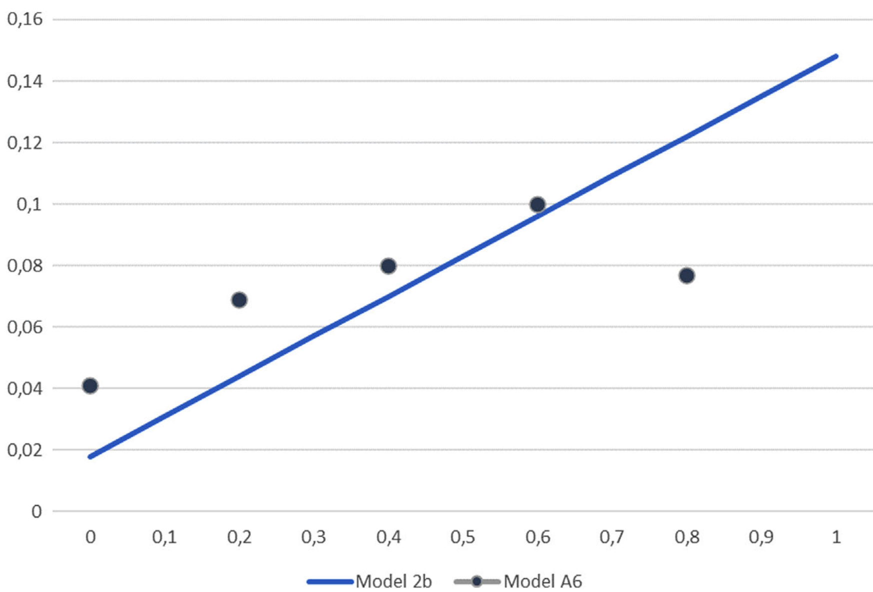

FIGURE A2 The effect of implementing a collective agreement, evaluated for different union density values [Colour figure can be viewed at wileyonlinelibrary.com]

TABLE A1 Descriptive statistics

	Observations	Firms	Average years	Min	Max	Mean	SD overall	SD within	SD between
Log value added	1,170,670	178,438	6.6	0	25.83	14.82	1.45	0.52	1.39
Log capital	1,170,670	178,438	6.6	0	25.6	12.91	2.20	0.94	2.02
Log hours, low-skilled	1,170,670	178,438	6.6	−3.98	12.33	2.58	2.30	1.02	2.05
Log hours, medium-skilled	1,170,670	178,438	6.6	−4.20	12.95	3.98	2.03	0.78	2.00
Log hours, high-skilled	1,170,670	178,438	6.6	−3.28	12.03	2.08	2.26	0.95	1.98
Log hours, top-skilled	1,170,670	178,438	6.6	−0.98	12.45	0.82	1.72	0.74	1.46
Union density (UD)	1,170,670	178,438	6.6	0	1	0.17	0.26	0.13	0.25
Collective agreement (CA)	1,170,670	178,438	6.6	0	1	0.13	0.33	0.12	0.26
Occupational share 0 and 9[a]	1,109,842	173,247	6.5	0	1	0.06	0.18	0.10	0.18
Occupational share 1–3	1,109,842	173,247	6.5	0	1	0.37	0.38	0.15	0.38
Occupational share 4–5	1,109,842	173,247	6.5	0	1	0.29	0.35	0.14	0.35
Occupational share 6–8	1,109,842	173,247	6.5	0	1	0.27	0.37	0.11	0.36
Low-skilled worker share[b]	1,161,166	176,243	6.4	0	1	0.23	0.26	0.12	0.27
Medium-skilled worker share	1,161,166	176,243	6.4	0	1	0.53	0.32	0.14	0.33
High-skilled worker share	1,161,166	176,243	6.4	0	1	0.18	0.26	0.11	0.28
Top-skilled worker share	1,161,166	176,243	6.4	0	1	0.06	0.18	0.06	0.20
Share Norwegians	1,170,670	178,438	6.6	0	1	0.80	0.28	0.12	0.30
Share immigrants from Nordic countries	1,170,670	178,438	6.6	0	1	0.04	0.12	0.06	0.13
Share immigrants from EU countries in Eastern Europe	1,170,670	178,438	6.6	0	1	0.04	0.13	0.06	0.15

(Continues)

TABLE A1 (Continued)

	Observations	Firms	Average years	Min	Max	Mean	SD overall	SD within	SD between
Share immigrants from other EU countries	1,170,670	178,438	6.6	0	1	0.03	0.10	0.05	0.11
Share immigrants from the rest of the world	1,170,670	178,438	6.6	0	1	0.08	0.19	0.08	0.22
Share low-age	1,170,670	178,438	6.6	0	1	0.10	0.18	0.11	0.18
Share medium-age	1,170,670	178,438	6.6	0	1	0.71	0.29	0.17	0.28
Share high-age	1,170,670	178,438	6.6	0	1	0.18	0.27	0.15	0.25
Share top-age	1,170,670	178,438	6.6	0	1	0.01	0.05	0.05	0.0

[a]Occupational shares are calculated at 1-digit (ISCO 08) level,
[b]2-digit NUS2000 codes, translatable to ISCED97.

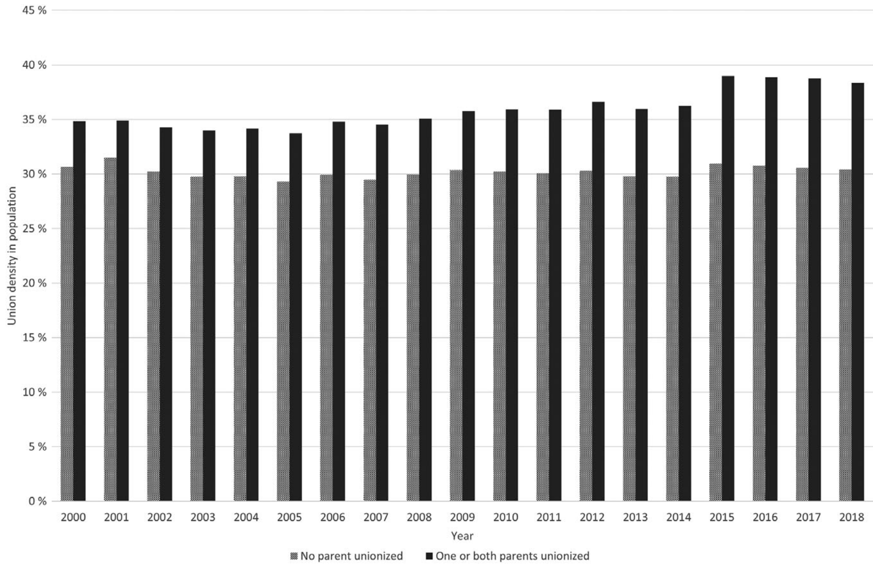

FIGURE A3 Union density in the sample and parental union membership ($N = 500{,}000$ individuals per year)

TABLE A2 Input elasticities corresponding to Table 3

	Model 1a FE	Model 1b FE	Model 1c FE	Model 1d FE	Model 1e LPW-GMM	Model 1f LPW-GMM	Model 1g LPW-GMM
Log capital	0.077***	0.077***	0.077***	0.075***	0.087***	0.087***	0.087***
	(82.64)	(82.55)	(82.55)	(81.99)	(104.54)	(104.54)	(104.54)
Log hours, low-skilled	0.080***	0.080***	0.080***	0.078***	0.108***	0.108***	0.108***
	(109.74)	(109.24)	(109.28)	(106.57)	(269.95)	(269.95)	(269.95)
Log hours, medium-skilled	0.147***	0.146***	0.146***	0.144***	0.242***	0.242***	0.242***
	(126.96)	(126.55)	(126.53)	(125.46)	(391.60)	(391.60)	(391.60)
Log hours, high-skilled	0.071***	0.070***	0.071***	0.069***	0.142***	0.142***	0.142***
	(93.64)	(93.37)	(93.47)	(91.25)	(359.91)	(359.91)	(359.91)
Log hours, top-skilled	0.059***	0.059***	0.059***	0.056***	0.160***	0.160***	0.160***
	(60.34)	(59.97)	(60.08)	(57.19)	(294.38)	(294.38)	(294.38)

Note: Union density measured as a rate between 0 and 1. Collective agreement measured as a dummy variable. All estimations include year dummies. Demographics include age intervals, sex and country of origin. *t*-Statistics are in parentheses. Models 1e–1g use as regressand the residuals from an LPW-GMM estimation of value added on capital and labour inputs only. Robust standard errors are clustered at firm level.

***$p < 0.001$.

TABLE A3 Estimation results using labour productivity (value added per hour worked) as endogenous variable

	Model A3a	Model A3b	Model A3c	Model A3d	Model A3e
$\log(C/HW)$			0.0636***	0.0633***	0.143***
			(73.86)	(74.08)	(14.72)
$\log(HW)$		−0.410***	−0.398***	−0.403***	−0.425***
		(−158.70)	(−145.18)	(−147.13)	(−23.41)
Union density (UD)	−0.0689***	−0.0386***	−0.0427***	−0.0429***	−0.0468***
	(−11.18)	(−7.02)	(−7.21)	(−7.26)	(−8.00)
Collective agreement (CA)	−0.00833	0.111***	0.0926***	0.0921***	0.0928***
	(−1.04)	(14.84)	(12.74)	(12.74)	(12.90)
UD × CA	0.0336*	0.0621***	0.0697***	0.0580***	0.0437***
	(2.19)	(4.53)	(5.17)	(4.34)	(3.32)
Marginal effects of:					
UD for $CA = 0$	−0.0689***	−0.0386***	−0.0427***	−0.0429***	−0.0468***
	(−11.18)	(−7.015)	(−7.211)	(−7.256)	(−7.997)
UD for $CA = 1$	−0.0353**	0.0235	0.0270*	0.0150	−0.00311
	(−2.410)	(1.811)	(2.137)	(1.201)	(−0.253)
CA for $\overline{UD}$	−0.00247	0.122***	0.105***	0.102***	0.100***
	(−0.381)	(19.77)	(17.58)	(17.24)	(17.01)
Test (p-value): $\hat{\gamma}_1 + \hat{\gamma}_3 = 0$[a]	0.0160	0.0702	0.0326	0.230	0.800
Industry by time dummies	No	No	No	Yes	Yes
Heterogeneous input elasticities	No	No	No	No	Yes
R^2 (within)	0.0695	0.169	0.215	0.220	0.228
R^2 (between)	0.0614	0.00941	0.0450	0.0387	0.0483
R^2 (overall)	0.0704	0.0173	0.0558	0.0504	0.0623
N	1,342,530	1,342,530	1,100,463	1,100,262	1,100,262
Firms	205,427	205,427	170,937	170,894	170,894
Average observations per firm	6.5	6.5	6.4	6.4	6.4

Note: Union density measured as a rate between 0 and 1. Collective agreement measured as a dummy variable. C and HW denote capital and hours worked, respectively. All estimations include year dummies, firm fixed effects and the following controls on individual workers' characteristics (measured as shares): education, occupation, age, sex and country of origin. t-Statistics are in parentheses.

[a]The reported test refers to the p-value of an F-test of the sum of the coefficients on UD and UD × CA.

*$p < 0.05$,

**$p < 0.01$,

***$p < 0.001$.

TABLE A4 LPW-GMM estimates of union density and collective agreements on total factor productivity with various sample restrictions

	Model 2b	Model A4a	Model A4b	Model A4c	Model A4d	Model A4e	Model A4f	Model A4g
Union density (UD)	−0.103***	−0.0730***	−0.120***	−0.100***	0.0293	−0.108***	−0.107***	−0.0737***
	(−15.44)	(−5.37)	(−14.02)	(−10.63)	(1.32)	(−15.21)	(−15.02)	(−3.61)
Collective agreement (CA)	0.0180*	−0.0152	0.0109	0.0150	–	–	–	0.0362***
	(2.25)	(−1.31)	(0.79)	(1.01)	(.)	(.)	(.)	(3.88)
UD × CA	0.130***	0.127***	0.154***	0.139***	–	–	0.156***	0.0863***
	(8.70)	(5.46)	(6.23)	(5.37)	(.)	(.)	(6.76)	(3.70)
Marginal effects of:								
UD for *CA* = 0	−0.1035***	−0.0730***	−0.120***	−0.100***				−0.0737***
	(−15.44)	(−5.37)	(−14.02)	(−10.63)				(−3.61)
UD for *CA* = 1	0.0267	0.0537	0.0344	0.0384				0.0125
	(1.91)	(0.01)	(0.15)	(0.12)				(0.66)
CA for $\overline{UD}$	0.0399***	0.00921	0.0353**	0.0375**				0.0640***
	(6.02)	(0.94)	(3.14)	(3.07)				(9.76)
Test (*p*-value): $\hat{\gamma}_1 + \hat{\gamma}_3 = 0$[a]		0.00658	0.153	0.123			0.0248	0.509
Firm presence restriction		Present all years	Enters	Enters and stays				
CA restriction					Always	Never	Always or never	Change status
N	941,969	244,407	505,937	379,206	77,173	770,598	847,771	94,198
Firms	152,651	17,614	105,200	62,052	8392	134,504	142,896	9755
Average observations per firm	6.2	13.9	4.8	6.1	9.2	5.7	5.9	9.7

Note: All models use as regressand the residuals from LPW-GMM estimation of value added on capital and labour inputs only, with heterogeneous input elasticities across 19 groups of industries. Union density measured as a rate between 0 and 1. Collective agreement measured as a dummy variable. All models include year dummies, industry by year dummies, firm fixed effects and controls on worker characteristics (occupation, age intervals, sex and country of origin). In Model 2b (our reference model), there are no restrictions on the sample. Model A4a restricts the sample to firms that were in operation throughout our entire sample period. Model A4b only includes firms that enter the market during our sample period, while Model A4c only includes those that enter the market during our sample period and stay in the market. In Models A4d and A4e, we restrict the sample of firms to those who always and those who never, respectively, have a collective agreement, while Model A4f includes all firms that do not change status during our sample period. Finally, Model A4g only includes firms that change status during our sample period (i.e. either enter or exit agreements, or both).

[a]The reported test refers to the *p*-value of an *F*-test of the sum of the coefficients on UD and UD × CA. Standard errors of marginal effects are calculated using the delta method. *t*-Statistics are in parentheses. Robust standard errors are clustered at firm level.

*$p < 0.05$,

**$p < 0.01$,

***$p < 0.001$.

TABLE A5 Nonlinear effects of unionization on total factor productivity

	Model 2b	Model A5
Collective agreement (CA)	0.018*	0.041***
	(2.25)	(5.25)
Union density (UD)	−0.103***	–
	(−15.44)	
UD × CA	0.130***	–
	(8.70)	
UD = 20–40% (UD2)		−0.052***
		(−21.55)
UD = 40–60% (UD3)		−0.056***
		(−13.70)
UD = 60–80% (UD4)		−0.061***
		(−9.12)
UD = 80–100% (UD5)		−0.001
		(−0.13)
UD2 × CA		0.028***
		(4.19)
UD3 × CA		0.039***
		(4.85)
UD4 × CA		0.060***
		(5.71)
UD5 × CA		0.036**
		(2.67)
Marginal effects of:		
CA for $\overline{UD}$	0.040***	
	(6.02)	
CA for $UD1 = 1$		0.041***
		(5.25)
CA for $UD2 = 1$		0.069***
		(9.99)
CA for $UD3 = 1$		0.080***
		(10.93)
CA for $UD4 = 1$		0.100***
		(10.54)
CA for $UD5 = 1$		0.077***
		(6.06)
N	941,969	941,969
Firms	152,651	152,651
Average observations per firm	6.2	6.2

(Continues)

TABLE A5 (Continued)

Note: Both models use as regressand the residuals from LPW-GMM estimation of value added on capital and labour inputs only, with heterogeneous input elasticities across 19 groups of industries. All models include year dummies, industry by year dummies, firm fixed effects and controls on worker characteristics (occupation, age intervals, sex and country of origin). In Model 2b, union density is measured as a rate between 0 and 1. In Model A6, union density is measured as a categorical variable taking the values {1, 2, 3, 4, 5} if the union density is within the corresponding intervals {0–0.2, 0.2–0.4, 0.4–0.6, 0.6–0.8, 0.8–1}. The first interval is used as reference category. Collective agreement measured as a dummy variable. Standard errors of marginal effects are calculated using the delta method. *t*-Statistics are in parentheses. Robust standard errors are clustered at firm level.

*$p < 0.05$,

**$p < 0.01$,

***$p < 0.001$.

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