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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 interfirm 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 sub-samples: 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	[−0.04360, −0.02163]	−0.2859	[−0.38136, −0.19061]
	(−5.82)		(−6.68)	
2008–18	−0.0141	[−0.02658, −0.00165]	−0.0640	[−0.12039, −0.00777]
	(−2.22)		(−1.84)	

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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