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ABSTRACT

The Dynamics of Power in Labor Markets: Monopolistic Unions versus Monopsonistic Employers*

This paper brings together the modern research on employer power and employee power by empirically examining the effects of unionization on worker earnings, employment, and inequality across differently concentrated markets. Exploiting national tax reforms to union membership dues as exogenous shocks to unionization, we show that high levels of unionization mitigate the negative wage and employment effects generated by imperfect competition. We also identify considerable effect heterogeneity with respect to worker types across differentially concentrated markets, and show that this has major implications for the role of unions in shaping labor market wage inequality.

JEL Classification: J23, J24, J42, J51, J52, J63
Keywords: monopsony, skills, unions, market concentration

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1 Introduction

In a market setting, worker wages are determined through negotiations between employers and employees. These two parties have fundamentally different interests. Employees strive to get as much as possible, and employers aim to provide as little as possible. The outcome of all such negotiations therefore depends on the relative strength of the two parties.

Modern empirical research has provided extensive evidence on the role of employer power (arising from labor market concentration and/or labor market frictions) and employee power (arising from unionization or the threat of union organizing). Specifically, employer power generates an upward sloping labor supply curve to the firm, allowing wages to be marked down below the marginal revenue product of labor and negatively affect workers' welfare.[1]. Employee power, on the other hand, equips workers with monopolistic power over labor supply, enabling employees to raise wages above some level that would prevail absent such bargaining.[2] However, despite centuries of rich discussions on the interaction of these forces in the labor market —from Adam Smith (1776) and Robinson (1933) to Freeman and Medoff (1984) and Manning (2021) —very little empirical research has provided a comprehensive analysis on this topic.

The goal of this paper is to bring together and bridge the literatures on employer power and employee power by empirically examining the effect of unionization on earnings, employment, and inequality across differently concentrated labor markets. We are motivated to study this topic because a union's ability to influence the wage setting may act as a countervailing force to the monopsony conditions that characterize a wide range of particular micro labor markets. At the same time, the ability of unions to serve as a countervailing force to a firm's monopsony power is theoretically ambiguous. Specifically, in a labor market characterized by strong monopsonistic competition, there will be significant rents for unions to extract (due to the presence of abnormal profits) but the union will hold little bargaining power (due to the lack of viable employee outside options that can be used as leverage). In a more competitive market, on the other hand, there will be minimal rents, but the unions' bargaining power will be greater.

Understanding a union's ability to counteract the monopsony power of firms will improve our understanding of the dynamics of labor markets and facilitate the design of optimal labor market policies. Specifically, the wage markdown generated by imperfect competition represents a market failure in which workers are paid less than their marginal revenue product. A rapidly growing literature has provided strong empirical evidence of such wage-setting behavior among employers, suggesting that the macroeconomic consequences of firm power

[1]E.g., Dodini et al. (2020); Schubert et al. (2020); Prager and Schmitt (2021)

[2]E.g., DiNardo and Lee (2004); Card and De La Rica (2006); Barth et al. (2020)

may be substantial (e.g., Dodini et al. (2020); Schubert et al. (2020); Prager and Schmitt (2021)). In such markets, unions may be able to correct the market failure generated by imperfect competition by counter-balancing the monopsony power of employer, pushing the economy closer to the competitive equilibrium. This would result in higher worker wages and employment levels, generating a more efficient allocation of resources conducive to higher economic growth. This stands in stark contrast to a union wage premium in a perfectly competitive market in which union-induced changes in wage levels may give rise to a new imperfection in the labor market in the absence of pure union productivity effects.

We begin by presenting a conceptual framework in which the wage premium of unionized workers is the result of Nash bargaining between the employer and the employees (Abowd and Lemieux, 1993). The framework provides two important insights on the relationship between firm profit, union bargaining, and wage premiums. First, the higher the firm's profits, the higher the union wage premium will be. Second, the stronger the union's relative bargaining power at the firm, the higher the union wage premium will be. These insights deliver valuable, empirically-testable, predictions about the relationship between unions and firms across differentially concentrated markets. This is because both the relative bargaining strength of the union as well as the profits of the firm are directly related to the degree of monopsony power that the employer enjoys. However, while profits are increasing with the degree of monopsony power, the union's relative bargaining strength is decreasing with the degree of monopsony power. Our conceptual model, therefore, provides a structure for understanding which parameters determine the direction of the relationship between worker wages, employer power, and employee power, in the labor market.

To empirically investigate the effect of unionization across differentially concentrated markets, we use high-quality longitudinal Norwegian employer-employee data—including information on union membership, union dues, and each worker's occupation. We then exploit changes in tax deductions for union dues as exogenous shocks to unionization. By interacting this exogenous shift in unionization with firm measures of local labor market concentration, we can analyze the role of unions across markets with different degrees of labor market concentration.

To obtain plausibly-exogenous variation in firm-level union density, we leverage changes in tax subsidies for union members in Norway, which led to significant changes in the net price of union membership for some firms (Barth et al., 2020). Specifically, these changes significantly reduce the monetary cost of joining a union for workers whose union due subsidies were previously bounded by a tax deduction cap. In other words, workers at firms whose union dues were high prior to the reform are more intensely "treated" by the reform relative to those with lower baseline union dues. This distinction generates exogenous variation in the incentive to join a union depending on the firm at which the worker is employed and, therefore, different union densities across firms.

To obtain a proxy for labor market concentration, we build on Dodini et al. (2020) and take a skill requirement-based approach to calculate market concentration. Specifically, we combine the Norwegian register data with information on skill requirements from the Occupational Information Network (O*NET) database and implement a hierarchical clustering algorithm (an unsupervised machine learning technique) to split occupations into distinct groups that are characterized by different combinations of these skill requirements.[3] We then calculate a Herfindahl-Hirschman Index (HHI), which is the sum of squared employment shares across firms in each skill requirement cluster and labor market. We use this measure as a proxy for labor market concentration. In the appendix, we demonstrate that all our results are robust to using labor market concentration measures based on occupations as well.[4] We use commuting zones as our geography unit, of which there are 160 in Norway (Gundersen and Jukvam, 2013).

The core finding of our analysis is that high levels of unionization ameliorate the negative effects of labor market concentration on earnings which has been identified in a growing set of empirical studies over the last few years (e.g., Schubert, Stansbury and Taska 2022; Caldwell and Danielli 2022; Dodini et al., 2020). This suggests that unions may play an important role in correcting market failures induced by imperfect competition. Consistent with monopsony theory, this wage effect is accompanied by positive intensive margin employment effects. Figure 1 illustrates this result in detail, demonstrating that as predicted unionization due to the our tax reform increases, the slope of the concentration-earnings gradient becomes flatter and far less significant. This is because unions can extract more rents when labor market concentration is high, despite having relatively lower bargaining power in these concentrated markets (Aghion et al., 1998; Yamaguchi, 2010). This implies that unions are able to "level the playing field" in concentrated markets.

We present five key results in support of this finding. First, we show that changes in tax subsidies for union members in Norway have a substantial effect on workers' willingness to

[3]We consider six distinct skill requirements based on Autor et al. (2003) as well as Acemoglu and Autor (2011): non-routine cognitive analytical; non-routine cognitive interpersonal; routine manual; routine cognitive; non-routine manual, physical adaptability; and non-routine interpersonal adaptability. These skill requirements have been shown in prior work to capture the cluster structure of occupational skills well, outpacing even the principal components of the O*NET data. Occupations within each group are similar in terms of their skills requirements.

[4]The usefulness of any concentration measure is based on its ability to accurately define the relevant market. Prior research has used industry or occupation designations to define the market, but neither really captures the labor demand faced by workers due to skill transferability between sectors or occupations. Our approach allows us to capture the concentration in demand for skills, which Dodini et al. (2020) argue is a more comprehensive way to aggregate workers for the purposes of measuring labor demand.

unionize. Specifically, increasing the annual union subsidy by NOK 1,000 leads to an increase in a worker's probability of unionizing by 13-15 percentage points. The effect is considerably larger in markets that experience monopsonistic competition, increasing the effect of the subsidy on unionization by 70-100 percent in the most concentrated markets, rising from a base of approximately 9 percentage points to over 26 percentage points. These results are consistent with the notion that workers expect a higher return to union membership in concentrated markets where employers have more power.

Second, using the changes in tax subsidies as an exogenous shifter of firm-level union density, we find that a 1 percentage point increase in union density generates an increase in annual earnings of 1.8 percent. This is the first causal estimate of the union earnings premium in an entire country across all sectors and industries. Examining heterogeneous effects across labor market concentration reveals that most of the union wage premium loads upon highly concentrated markets. Specifically, a 1 percentage point increase in union density raises annual earnings by 1.1 percent in non-concentrated markets and by 2.5 percent in concentrated markets (statistically significant at the one percent level). The gradient is primarily driven by the private sector. This result supports the theory that the greater the market imperfection, the greater the amount of firm rent that unions are able to extract despite their weaker relative bargaining position in these markets.

Third, we combine our labor market data with firm-level revenue data and explore the relative impact of possible product market concentration, which we proxy with industry-level revenue shares, and labor market concentration on the union wage premium. Prior literature has provided suggestive evidence of a strong correlation between labor and product concentration at the firm level (e.g., Marinescu et al. (2021); Lipsius (2018); Qiu and Sojourner (2019)), and understanding to what extent unions extract product rent versus labor rent is of independent interest. Our results suggest that unions are successful in extracting both labor and product rents from the firm. That we are able to identify different effects across these two sources of market power highlights that even if they are correlated, they are separably accessible in union pay negotiations.

Fourth, we show that firms in concentrated markets respond to an increase in union density and worker earnings by raising employment the following year on the intensive margin, while they reduce employment on this margin in non-concentrated markets. This pattern of results is consistent with monopsony theory, which predicts that both wages and employment levels will lie below a competitive equilibrium in the presence of monopsony power.

Finally, we document important heterogeneity with respect to worker type. Specifically, we show that the modest union wage premiums that exist in competitive labor markets are reserved for high-skilled and white-collar workers. As market concentration increases, more and more of the additional rent that unions extract goes to lower-ability and blue-collar workers. This implies that unions have an inequality-enhancing effect within narrow sub-sectors in competitive markets, while this is not the case in concentrated markets.

This paper empirically brings together the modern research on labor market concentration and unionization in labor markets. This allows us to substantially advance the existing knowledge on the role of unions and their impact on the dynamics of labor markets. Our main contribution is to provide a method for identifying the causal effect of unions on worker earnings, employment, and inequality as a function of employer concentration, demonstrating that unions may offset the market failure induced by imperfect competition.[5]

We contribute to the existing literature in several ways. First, there is a rapidly-growing literature that has attempted to directly measure labor market concentration and then examine how concentration affects wages and employment (e.g., Dodini et al. (2020); Schubert et al. (2020); Azar et al. (2019a; 2020b;a); Benmelech et al. (2022); Marinescu et al. (2021); Qiu and Sojourner (2019); Rinz (2018); Hershbein et al. (2018); Bassanini et al. (2022); Prager and Schmitt (2021)). On average, these studies show that labor market concentration reduces worker wages and has negative effects on workers' careers. We advance this literature by demonstrating that unionization rates, as well as union wage premiums, are substantially larger in concentrated markets. This suggests that unions may sucesfully act as a countervailing force to employer power. In addition, we document a positive marginal union employment effect in concentrated markets. Researchers interested in accurately measuring the effects of monopsonistic competition in labor markets need therefore to carefully account for the dynamics of unions across markets that face different levels of labor concentration. More specifically, our results reveal that understanding the interplay between employer and employee power is imperative for identifying the direct impact that concentration and monopsony power may have on the dynamics of labor markets, and that one should not look at employer power or employee power in isolation. In addition, our findings help us better understand the recent macroeconomic phenomena of a decline in the share of income that is going to labor, an increase in measures of corporate valuations, a rise in average profitability even as interest rates have declined, and increases in measured markups. Specifically, based on our results, and consistent with (Stansbury and Summers, 2020), we find that a decline in relative worker power would produce predictions similar to these observed trends. This is particularly noteworthy as the average OECD country has witnessed a gradual decline in union density over the last decades.

[5]Some of the prior literature presents suggestive evidence of the relationship between wages and concentration in more vs less unionized sectors (Qiu and Sojourner, 2019; Marinescu et al., 2021; Prager and Schmitt, 2021; Benmelech et al., 2022) but does not attempt to parse the causal effect of unionization itself.

Second, there is a small but impressive literature that causally identifies the union wage effect through quasi-experimental research designs. The studies in this literature have relied either on regression discontinuity designs related to close union elections (e.g., DiNardo and Lee (2004); Lee and Mas (2012); Frandsen (2021); Sojourner et al. (2015)), or on propensity score matching techniques that directly control for endogenous selection of workers into unions (e.g., Card and De La Rica (2006); Bryson (2002)). Recently, Barth et al. (2020) has exploited changes in national union due subsidies as a measure of unionization probability, an approach we adopt in this paper as well. Finally, Fortin et al. (2022) use Right-to-Work laws in the United States as an instrument for individual unionization and find an IV wage premium that is substantially larger than typical OLS estimates suggest.

We advance the union literature by providing the first causal estimates of the average union density earnings premium in an entire country across all sectors and industries. We then further develop this literature by considering how the union wage premium differs across markets that face different degrees of concentration. The results have important implications for how we view the role of unions in labor markets, as a union wage premium in a monopsonistic market may point to unions correcting a market failure while a union wage premium in a competitive market may point to unions causing a market failure. We, therefore, see our paper as opening up a new avenue of research, exploring the dynamics of how the balance of power between employers and employees may impact not only wages but also other types of non-pecuniary benefits and social goods.

The rest of the paper proceeds as follows: In Section 2, we provide institutional background and introduce a conceptual framework which we use to guide our analysis and results. In Section 3, we provide a detailed overview of our data. In Section 4 we introduce our empirical method. In Section 5, we present all our main results. In Section 6, we conclude.

2 Background

2.1 Unions in Norway

Norway's *Working Environment Act* governs worker rights in Norway and regulates both individual employees and their contracts as well as unions and their collective bargaining agreements. Similar to other countries, the stated goal of Norwegian labor unions is to strengthen members' rights and work conditions, and they play an important role in, for example, contract negotiations. All workers in Norway have the legal right to unionize, and firms are required to enter a collective bargaining process if at least 10 percent of the workers at the firm request it (Stokke et al., 2015). On behalf of their members, unions can negotiate not only wages but also help settle legal disputes and push for better work conditions.

Unions are commonly structured by professional area or sector. Each individual local union is linked to a national federation of trade unions, and each federation is linked to one of four much larger national confederations of trade unions. The largest such employee association is the Norwegian Confederation of Trade Unions, covering approximately 50 percent of all unionized workers. This structure is similar to other countries such as the United States, where, for example, the American Federal of Labor and Congress of Industrial Organizations (AFL-CIO) coordinates and supports union efforts across more than 50 individual unions spanning a range of professions.[6] While the organizational structures of unions in Norway have changed over time, there have been no significant changes to their structure during our sample period (since 2001).

In the private sector, union density has been around 40 percent for the past several decades. In the public sector, union density is approximately 79 percent. The union density rate differs across sectors and industries, with almost 60 percent in the manufacturing sector and slightly less than 30 percent in the private services sector. More women than men are members of labor unions (57 percent versus 44 percent), partially reflecting women being more likely to sort into the public sector. The unionization rate in Norway is not particularly high relative to other OECD countries and is lower than the unionization rate in other Nordic countries such as Sweden.[7]

In terms of collective bargaining, wages can be negotiated at three different levels: the central level (between the national employee confederations and the national employer confederation), the sectoral level (employer and employee organizations in specific sectors), and the local level (company management and local trade unions). If negotiations fail, the parties are entitled to take industrial action. This usually occurs in the form of a strike (unions) or a lockout (employers). 87 percent of workers in Norway are covered by collective agreements, and approximately 79 percent of workers are employed at firms in which local bargaining takes place immediately following the national or sectoral bargaining rounds (Barth and Nergaard, 2015). Non-union employees do not have the right to bargain, and it is up to the employer to adjust the pay as they deem appropriate.

Historically, the national and sectoral wage agreements played a key role in setting worker wages. Since the late 1990s, however, these agreements primarily serve the purpose of setting industry-specific wage floors and ensuring a minimum wage increase for workers, and local negotiations now account for more than 70 percent of total negotiated wage increases (Mogstad et al., 2021). In the local negotiations, unions and employers discuss not only

[6] See https://aflcio.org/about-us/our-unions-and-allies/our-affiliated-unions (accessed January 18, 2022).

[7] One reason for this is that unemployment benefits are part of a union's purview in other Nordic countries such as Sweden, while they are governed by national law in Norway. Sweden and Denmark both have experienced declines in their unionization rates since the late 1990s.

union-wide wage increases, but also individual-specific wage increases. In other words, the bargaining process in Norway is a two-step process. In the first step, industry-wide collective bargaining agreements are established to set wage floors and some guaranteed wage increases. In the second step, local negotiations take place in which unions and employers discuss not only firm-specific wage increases for union members but also individual-specific wage increases. The local bargaining component is crucial for the purpose of our study as it enables firms and unions to adjust wages and wage demands depending on the degree of labor concentration in the market. Without this local negotiation feature, unions would not be able to adjust wage demands based on local market conditions, something that would generate an attenuation bias and work against us finding an effect.

2.2 Union Tax Deductions

A novel feature of Norwegian government policy regarding labor unions is the existence of a tax deduction for union dues that acts as a subsidy for union membership. This deduction is automatically entered on an individual's tax return, making it very salient to the worker. In the mid-2000s, the Norwegian government enacted a series of large increases in the maximum allowable tax deduction for union dues. This maximum nearly quadrupled from 2001 to 2010. The realized value of the subsidies to workers depends on the union dues required of prospective members.

Our empirical strategy exploits the national changes in the maximum allowable tax deduction for union dues. These changes significantly reduce the monetary cost of joining a union for those workers whose subsidies were previously bounded by the tax deduction cap. In other words, workers at firms whose union dues were high prior to the reform are more intensely "treated" by the reform relative to those with lower baseline union dues. This distinction generates exogenous variation in predicted unionization rates for workers and, therefore, different union densities across firms. We use this exogenous variation in union density to identify the effects of unionization on earnings for different types of workers in concentrated versus more competitive labor markets.

2.3 Conceptual Framework

In this section, we conceptualize the relationship between employer power and union power in the labor market to provide context for our empirical models and results. As we state above, the bargaining process in Norway can be viewed as a two-step process. First, industry-wide collective bargaining agreements are established to set wage floors and some guaranteed wage increases. Second, local negotiations take place in which unions and employers discuss not only firm-specific wage increases for union members but also individual-specific wage increases. We abstract away from the first step by treating the industry-wide wage floors as given and focusing on the local negotiations.

We begin by writing down a simple earnings equation for the market wage of individual i at firm f in the absence of unions at the firm:

$$w_{if}^{m} = X_i\beta + Z_f\gamma - G_f(M) + \epsilon_{if} \ , \tag{1}$$

where X_i is a vector of individual characteristics, Z_f is a vector of firm characteristics, and ϵ_{if} is an idiosyncratic error term. M is the degree of monopsony power that the firm faces, and is equal to labor market concentration (HHI) times the market-level inverse elasticity of labor supply (η). The function G_f thus explicitly allows for a wage markdown driven by imperfect competition in the labor market.

If unions are present at the firm, let the net-of-union-due wage of the worker be denoted w_{if}^{U} and be a result of Nash bargaining between the employer and the workers. In this context, the workers are interested in maximizing the wage surplus obtained through the union, w_f^U-w_f^m. The employer is interested in maximizing profits, i.e. revenues net of input costs. The threat point of the firm is thus zero profit, and the threat point of the worker is the alternative wage.

In the short run when capital inputs are fixed, the key input is labor, and the profit function can be denoted $\Pi_f = pF(L_f) - w_f^U$ (Abowd and Lemieux, 1993; Breda, 2015). Here, L_f is the firm's labor force, $F(L_f)$ is the production function, and p is a revenue shifter. We note here that profits can only arise because of market power, either through price-setting power on the product market such that p is far above the marginal production cost or else through labor market power such that w_f is far below the marginal revenue product.

The bargaining problem can be expressed as follows:

$$w_f^U = Argmax(w_f^U - w_f^m)^{\phi_f}[\Pi_f]^{1-\phi_f} \ , \tag{2}$$

where ϕ_f denotes the relative bargaining power of the union at the firm.[8] As shown in Abowd and Lemieux (1993), the solution to the bargaining problem is

$$w_f^U = w_f^m + \frac{\Phi_f \Pi_f}{L_f} \ , \tag{3}$$

Written in this form, the union negotiated wage is equal to the market wage plus a fraction

[8]There are several bargaining models that can be used at this point, based primarily on whether the firm and the union negotiate only wages or both wages and employment. Both models are compatible with this framework.

of profits, which depends on the relative bargaining power of the union.

To disaggregate the above equation to the individual level, we must invoke an assumption about how unions choose to allocate rents among their members. To this end, we follow convention and assume that unions impose an egalitarian split, with each union member receiving an equal amount. We revisit this assumption at the end of the section.

With this assumption, and denoting the union density at the firm with U_f, we can turn Equation 3 into a general wage equation at the individual level:

$$w_{if} = X_i\beta + Z_f\gamma - G_f(M) + U_f[\frac{\Phi_f \Pi_f}{L_f}] + \epsilon_{if} \tag{4}$$

The incorporation of union density (U_f) rather than a union membership dummy is based on the canonical work of Freeman and Medoff (1984). This work shows that what matters for successful union rent extraction is not whether a union is present at the firm, but how big the union density at the firm is. For example, a union covering 15% of a firm's workers has little leverage over the firm if negotiations break down, even to negotiate only for union members. The threat of adverse action such as a slowdown or walkout is minimal. If union membership at the firm covers 90% of the firm, any adverse action taken by workers has far larger consequences for the productivity and revenue of the firm. Numerous studies have confirmed the importance of union density at the firm level for successful rent extraction (see, for example, Breda (2015); Fitzenberger et al. (2013); Barth et al. (2000); Balsvik and Sæthre (2014); Barth et al. (2020)). Indeed, Norwegian studies have even found that individual membership status has no impact on wages once union density controls are included, highlighting that what matters for successful union rent extraction is not whether a union is present at the firm nor whether an individual worker is a member of the union, but on union density at the firm Barth et al. (2000).

The above expression gives rise to two predictions crucial to our empirical analysis. First, the higher the firm's pure profits, the higher should the union wage premium be. Note here that we are agnostic about whether such rents are coming from the product market (through p) or the labor market (through w^m markdowns). Second, the stronger the union's relative bargaining power at the firm, the higher should the union wage premium be.

In terms of our empirical analysis, the above discussion delivers valuable predictions about the relationship between unions and firms across markets that face different degrees of concentration. The reason is that both the relative bargaining strength of the union as well as the profits of the firm are direct functions of the degree of monopsony power that the employer faces. However, while profits are increasing with the degree of monopsony power, the union's relative bargaining strength is decreasing with the degree of monopsony power (Aghion et al., 1998; Yamaguchi, 2010; Tschopp, 2017). Thus, the direction of the relationship between monopsony power and the sum total of these two components of the wage equation is ambiguous.

To illustrate this point in greater detail, consider two markets: one with a high level of labor market concentration ($HHI \to 1$) and one with a low level of labor market concentration ($HHI \to 0$). In the first case, the relative bargaining power of the union will approach 0 while the profits of the firm rise. In the second case, the relative bargaining power of the union will approach 1 while the profits available to the firm will approach 0. The relationship between the union wage premium and labor market concentration thus depends on how much of the additional profits unions are able to extract (marginal rent extraction) as the market is becoming increasingly concentrated. If available profits increase by more than bargaining power decreases for a given change in market concentration, we would expect the union wage premium to be higher as concentration rises. If, on the other hand, the decrease in bargaining power dominates the increase in profits, we would expect to see the opposite. This is our first testable prediction, which we investigate empirically by exogenously shocking U_f across differently concentrated markets. This enables us to trace changes in wages across differently concentrated markets for the same increase in union density. If the effect of an exogenous shock to U_f is stronger in more concentrated markets, that implies that $(\frac{\Phi_f \Pi_f}{L_f})_{concentrated} > (\frac{\Phi_f \Pi_f}{L_f})_{competitive}$, and that the union premiums are higher in concentrated markets. To exogenously shock U_f, we leverage changes in tax subsidies for union members in Norway, which led to significant changes in the net price of union membership for some workers. All else equal, this leads to an increase in w_f^U, and should induce more workers to join the union by changing the threat point for workers in Equation 2.

Before turning to the empirical investigation, it should be noted that the transition from the firm-level equation to the individual-level equation above required us to invoke an assumption on how unions divide rents across their members. So far, we have followed convention and assumed that unions impose an egalitarian split, with each union member receiving an equal amount of the available rents that they secure. However, assuming that labor unions are union due maximizers (Abowd and Lemieux, 1993), unions may allocate rents strategically among their members such that they secure the most union dues possible.

This provides us with a second testable prediction: if labor unions are union due maximizers ($\Phi_{if}\Pi_{if}$) may differ across different types of workers (i), and this difference may vary across differently concentrated markets. For example, to maximize union dues, unions may decide to prioritize higher-wage earners when allocating the limited rents available in more competitive markets. First, higher-wage earners have more room to pay higher union dues, such that their union membership would generate higher dues. Second, high-skilled workers may be less likely to join unions due to outside options being better, such that unions may focus on satisfying and ensuring the continued membership renewal of high-skilled workers. Finally, higher-productivity workers also may carry more weight as representative agents in negotiations. As markets become more concentrated, however, the reduction in outside options for high-skilled workers combined with the improved rent extraction opportunities available to unions means that they may shift focus to bargaining for a more general wage increase across all worker types. If these types of non-egalitarian rent-split strategies are present in the data, they have important implications for labor market inequality and unions' role in reducing such inequalities.

As an alternative to the conceptual framework outlined above, we also provide a geometric explanation of the relationship between union-specific wage floors or negotiated increases in wages in a monopsonistic market in Appendix Figure A5. This alternative framework elicits similar questions related to the framework above but does so in a way more directly addressed by the upward-sloping labor supply curve of monopsony theory.

3 Data

3.1 Data

Our primary data come from linked employer-employee registers covering the universe of workers in Norway between the ages of 16 and 74 in the years 2001 through 2015. Using a unique individual identifier, we follow individuals over time and across registers. We obtain demographic characteristics from the central population register, we collect education information from the national education register, we use labor earnings information from the tax register, and we obtain information on contract hours, firm, and employer from the linked employer-employee register.

The linked employer-employee data allow us to identify each worker's employer and construct labor market concentration measures for each firm in the Norwegian economy. We construct these at the local labor market level, which is defined based on commuting distance. The local labor markets divide Norway into 160 regions (Gundersen and Aarhaug, 2013). By linking the unique firm identifiers to the universal firm accounting data register, we also are able to construct proxies of product market concentration for each firm in their industry, fixed at the firm's first year in the data. As we will explain in Section 4, we use this measure to run horse races between labor market concentration and product market concentration to better understand which types of rents unions extract.

Our data provide detailed earnings and employment information of each worker in the country. Labor earnings are measured as pre-tax income (income from labor and self-employment) and include taxable government transfers (parental leave, sick leave, and unemployment benefits). Employment status is defined based on the individual's status in the labor register, and full-time employment status is defined based on the number of hours the worker is registered for per week (we consider individuals with more than 30 work hours per week as full-time workers). In addition to labor market characteristics, the data give us access to a large and detailed set of demographic and socioeconomic characteristics of the individuals. These variables include gender, age, education, marital status, and place of residence and work.

Crucial to our analysis is the ability to observe individual-level union information over time. We obtain the data from a register-based union membership data set, which provides detailed information on each individual's involvement with labor unions and how much they have paid to become a union member each year.

In terms of sample construction, we impose three restrictions. First, we limit our sample to individuals who worked at least 20 hours per week on average. We impose this restriction to eliminate individuals with a weak labor market attachment and to ensure a more precise measure of the potential union wage premium. Second, we limit the sample to individuals working in firms that had at least ten workers employed each year. This excludes small family businesses and sole proprietorships. We impose this restriction to ensure that our results and concentration measures are not driven by small firms that have little impact on the larger economy. Third, we limit the sample to those with annual earnings that would qualify them for the "1G" designation in the Norwegian benefit system, which is approximately 90,000 NOK (approximately 10,000 USD) based on 2015 values. This ensures that those without meaningful attachment to the labor market do not affect our results.[9]

3.2 Union Dues and Tax Subsidies

To obtain exogenous variation in firm-level union density, we leverage changes in tax subsidies for union members in Norway which led to significant changes in the net price of union membership for some workers (Barth et al., 2020). Specifically, the maximum tax deduction for union dues nearly quadrupled between 2001 and 2010. These changes significantly reduced the monetary cost of joining a union for workers whose union due subsidies were previously bounded by the previous deduction cap. In other words, firms subject to higher union dues in 2001 could expect a substantial increase in these subsidies compared to firms with lower union dues. By construction, although workers may endogenously select into firms and occupations, the policy changes we exploit are orthogonal to changes to these firm characteristics over time and therefore represent quasi-experimental, exogenous variation in the cost of union membership to these workers.

[9] The "1G" designation (also called *Grunnbeløpet*), is used to calculate whether individuals qualify for certain government welfare payments and transfers, and how large those payments should be.

The Norwegian registers only contain information on union dues for those who are union members. We, therefore, begin by constructing a measure of union dues for those who were not part of a union had they been part of a union. To construct this measure, we take the mean union due paid by workers in each occupation-industry cell in each year and apply this to union members and non-members alike. As such, we do not use information on individual union dues or wages that may be endogenously determined by individual or firm characteristics. This imputation approach is identical to that used in (Barth et al., 2020), and has two advantages: first, it allows us to predict the average counterfactual costs of unionization faced by those who were not part of the union; second, we can abstract away from endogenous individual and firm determinants of union dues for union members. We then characterize the union dues of the firm as the average imputed union dues across all the firm's workers.

One possible concern when considering the effect of union subsidies is that firms and unions may endogenously respond to the subsidy legislation, either in terms of which occupations they decide to employ in or in terms of setting their union dues. To eliminate this issue, we fix each firm's union dues to the average imputed dues across all occupations at the firm in the first year that the firm appears in the data, which is 2001 for most firms. We then adjust for inflation forward to nominal Norwegian crowns. This approach weights the union dues for the occupational mix that existed in the firm in its first year in our data set, meaning there is no endogenous distribution of occupations in response to either union action or the legislation itself. It also ensures there is no endogenous feedback loop between the change in the law in any particular year and the imputed union dues.

Once we have obtained our baseline imputed union due measure, we calculate the value of the base subsidy for all individuals in the data set. This value is equal to the lesser of the legislated maximum deduction and the firm's imputed union due, which we multiply by the country's base tax rate (28 percent from 2001 to 2013 and 27 percent in 2014 onward). We apply the base tax rate to isolate changes in the guaranteed *statutory* subsidy from changes in the *realized* subsidy that may depend on marginal tax rates. This helps us avoid endogeneity in this policy because marginal tax rates may be determined in part by unionization and other within-firm dynamics that determine wages. Our measure of subsidy value, therefore, captures changes that only are coming through legislative channels and not changes within firms.

Our base subsidy measure for individuals at firm f at time t is expressed as follows:

$$S_{ft} = T_t * (min\{\overline{D_f^0}, MaxDeduction_t\}) \ , \tag{5}$$

where T_t is the base tax rate in year t, $\overline{D_f^0}$ is the imputed firm union due at baseline, and $MaxDeduction_t$ is the maximum statutory deduction. Identifying variation in the subsidy comes from cross-sectional differences in the occupation-industry mix of firms in their base year combined with changes in the legislated maximum deduction over time. Specifically, the base industry and occupation composition of the firm determines whether or not workers at the firm are strongly bound by the maximum deduction or not. Firms with high imputed union dues in the base year are more intensely treated when the deduction caps are relaxed over time. Changes in the net price of union membership are therefore exogenously loading on some workers and not others for reasons unrelated to labor market or firm conditions over time.

We also calculate the net-of-subsidy union due by subtracting the value of the subsidy from the gross imputed baseline union due ($ND_{ft} = \overline{D_f^0} - S_{ft}$). We include this as an additional control in all our regressions. This is important, because two workers may receive the same subsidy in a particular year yet still face different incentives and costs for membership. Specifically, two workers whose deductions have hit the ceiling of the maximum tax deduction may face different dues and therefore respond differently to the subsidy depending on what their remaining dues are. A worker whose baseline dues were far higher may respond less strongly to the subsidy and vice versa because their net costs are still higher. This is particularly important if baseline dues reflect some unobservable firm-specific productivity difference related to their baseline occupational mix.[10] Importantly, variation in the subsidies *over time* comes exclusively from the tax deduction policy because we fix baseline dues in the firm's base year. We scale our subsidy and net union dues measures to a basis of 1,000 NOK, which was approximately 120 US dollars in 2015.

Figure A1 illustrates the drastic increase in the maximum union due deduction and average imputed subsidy over our sample period. While the maximum deduction increased from just below 1,000 NOK to almost 4,000, the average imputed base subsidy went from approximately 300 NOK to approximately 1,000 NOK. Around this mean value, there is significant heterogeneity by industry and firm.

[10] For robustness, we also estimate the effects of the subsidies on the likelihood of joining a union using a subsidy ratio (subsidy divided by net union due) while controlling for the inverse net union due. This is similar to the model that is estimated in Barth et al. (2020). Our results are robust to this alternative approach. However, we prefer to include the net-of-subsidy variable as a separate control rather than relying on the subsidy ratio, because this does not constrain the effect of the net dues to have a proportional relationship with the subsidy. Estimating the equation using the ratio of the two as a single treatment variable would impose that constraint. Our approach, therefore, flexibly disentangles a potential heterogeneity in the effect of the subsidy. Our estimates also effectively replicate the pattern of findings from the manufacturing sector in Barth et al. (2020) for the entire country of Norway. See Appendix Table A4.

3.3 Defining Concentration

We follow Dodini et al. (2020) and use data from the US Department of Labor's Occupational Information Network (O*NET) survey to incorporate skills requirement information into the Norwegian registers.[11] The O*NET survey asks workers and occupational experts about the knowledge, skill requirements, and tasks associated with each occupation. We connect Norway's STYRK occupation classification system to the O*NET survey's Standard Occupation Classification (SOC) system using the crosswalk in Hoen (2016).

We focus on six skill requirement categories similar to those in Autor et al. (2003) and Acemoglu and Autor (2011). We use these skill requirements to group together occupations based on their skill content. These skills are routine, manual; non-routine, physical adaptability, manual; non-routine, interpersonal adaptability; routine, cognitive; non-routine, cognitive, interpersonal; and non-routine, cognitive, analytical skills. We focus on these skill requirements because the prior literature documents their importance in explaining labor market segmentation and wage trends over time. We create composite measures of each of these skills standardized to have a mean of zero and a standard deviation of one. We then use a Hierarchical Agglomerative Clustering (HAC) algorithm to split occupations in the Norwegian register into 20 distinct skill groups. This is an unsupervised machine learning technique in which we impose no conditions on any of the parameters other than choosing which distance measure to group clusters together after they are initially formed. As a test of robustness, we also generate estimates based on 40 clusters.[12]

The HAC clustering technique starts by treating each occupation as a separate cluster. It then non-parametrically merges the two closest occupations together into clusters based on their correlative distance, which is one minus the Pearson correlation between the two occupations based on the six skill characteristics until a full dendrogram or tree is formed. We then select a "cut point" for the tree based on the number of clusters, which we set at 20 (40). Following Dodini et al. (2020), our choice of 20 skill clusters is based on a set of validation exercises that put the data-driven "optimal" number of clusters near 20, though we show that using 40 skill clusters generates similar estimates with matching conclusions.[13]

For each occupation at the firm, we calculate a Herfindahl-Hirschman Index (HHI) of the firm's employment share in that occupation's skill requirement cluster and the worker's local labor market in each year. This measure of labor market concentration takes into account a worker's set of local counterfactual outside options that use similar sets of skills to their current occupation. This is important because a worker's skills can be transferable not

[11] The Norwegian registers do not contain information on occupation-specific skill characteristics.

[12] HAC algorithms are similar to more widely-known K-means clustering algorithms but are known to handle non-spherical cluster shapes more adeptly. The results are also more reproducible.

[13] Dodini (2022) also validates the optimal cluster number in a US context at approximately 20 clusters.

only between firms but also between occupations and industries. We argue that this makes a purely occupation-based or industry-based measure of concentration less representative of the relevant labor market. However, we emphasize that using the more conventional concentration measures calculated at the occupation level generates the same pattern of results (see Appendix Tables A10–A16).

We should note that market power in our conceptual model is defined as HHI scaled by the elasticity of labor supply to the market. A large portion of the modern empirical literature on labor monopsony abstracts away from specifying the labor supply elasticity to the market and implicitly assumes that it does not systematically vary with HHI (e.g., Azar et al. (2019a; 2020b;a); Benmelech et al. (2022); Marinescu et al. (2021); Qiu and Sojourner (2019); Rinz (2018); Hershbein et al. (2018))).[14] We follow this convention while at the same time considering concentration to be of independent interest.

The goal of any concentration measure in this literature is to identify how concentrated demand is for labor, which necessitates defining the relevant labor market by worker types. Prior research has used industry or occupation as a proxy for such groupings, but we argue these are insufficient because neither really captures the labor demand faced by workers, including possible moves across occupations and industries.[15] Our approach allows us to capture the concentration in demand for skills, which we believe is a more comprehensive way to aggregate workers for the purposes of measuring labor demand and their outside options.[16]

To characterize the overall local labor market power held by the firm and to facilitate comparisons to product market/industry revenue HHI, we generate a composite measure of concentration at the firm level by taking the average HHI across all workers at the firm in the first year in which the firm appears in the data. This enables us to leverage a single measure of labor market concentration which we then can interact with predicted union density to examine the marginal earnings effects of unionization across labor market

[14]This assumption appears warranted given the fact that labor supply elasticities in response to changes in the post-tax/transfer returns to work are low across studies with relatively little variability, suggesting overall market-level elasticities carry similar narrow bounds. Whalen and Reichling (2017) review the Frisch elasticities of labor supply in the literature and conclude that most estimates range from 0.27 to 0.53 with a central estimate of 0.4 across studies. These are used by statistical agencies such as the US Congressional Budget Office. Chetty (2012) reconciles the micro and macro literatures to find elasticities of 0.25 for the extensive margin and 0.33 for the intensive margin after accounting for frictions.

[15]Occupation classification systems across OECD countries implicitly rely on skill and task similarity to put occupational titles together as well as industry composition. Our clustering approach takes a more data-driven and less ad-hoc approach to occupational grouping than these classification systems.

[16]That our approach does this without relying on job-to-job transitions is important, since these transitions are a direct function of the underlying labor supply and demand forces in the local area. In other words, job transitions are endogenous to market conditions including monopsony power, the spatial distribution of industry activity, and product market competition.

concentration without being concerned about endogenous concentration changes (in reaction to unionization) or intra-firm occupation composition effects in any particular year. Our results are similar when using the firm's average HHI over the whole sample period.

Table A1 contains a set of basic summary statistics for our analysis sample.[17] Approximately 60% of our sample of workers are members of unions, and their earnings, on average, are approximately 463,000 NOK. The imputed base tax subsidy for our sample is on average 750 NOK over the sample period with a net-of-subsidy union due of approximately 3,200 NOK. The average labor HHI at a worker's firm in their local labor market is approximately 0.043, with a standard deviation of 0.054. Operating revenue HHI in a firm's industry (which we use as a proxy for product market power) is approximately 0.037, with a standard deviation of 0.077. To better understand these numbers, we note that the Horizontal Merger Guidelines used by the antitrust division of the US Department of Justice consider markets with HHI values of less than 0.15 as unconcentrated, markets with values between 0.15 and 0.25 as moderately concentrated, and markets with values above 0.25 as highly concentrated.

4 Empirical Strategy

Our empirical strategy relies on leveraging exogenous changes to the cost of joining a labor union that came through adjustments to the Norwegian tax code between 2001 and 2015. We use the value of the imputed base subsidy for union members and nonmembers, as well as the net-of-subsidy union due, as exogenous shifters for the incentive to join a union for each worker in the sample. These shifts come through a reduction in the cost of joining a union for those workers whose deductions were bound at the top by the deduction cap before the reform increased this cap.

The empirical method we employ relies on three key assumptions. First, it must be the case that the subsidies did, in fact, increase the rate at which workers joined labor unions in the firms affected by the maximum deduction caps before the reform. Second, it must be that the change in the subsidy is unrelated to changes in firm characteristics that may be correlated with worker productivity and/or earnings. The careful construction of our subsidy and net-of-subsidy union dues variables allow us to control for any possible endogeneity of the posted union dues in response to legislative action. This isolates variation in subsidies that come from the legislation itself.

Lastly, there is one other assumption that is implicit in our approach: that firms and workers with high expected subsidies (relative to net dues) would have had similar earnings changes to firms and workers with low expected subsidies *but for* the change in the deduction. A clear sign of this counterfactual would be if high-subsidy and low-subsidy workers had

[17] To reduce computational time, we take a 70% random subsample of workers in the data.

similar earnings trends during years in which tax deductions (and, therefore, the subsidies) were stable. We examine this in Appendix Figure A2 and conclude there is no reason to suspect diverging trends explain any of our results based either on the raw trends (Panel A) or estimated coefficients with numerous fixed effects that constitute a more formal test (Panel B). Specifically, individuals with high subsidies relative to net-of-subsidy dues have higher earnings growth *only* during the years in which the maximum deduction changed drastically and the subsidy changes became efficacious (2003-2010), while the trends were parallel in 2001-2002 and 2010-2015 when the deductions were stable. The differences in earnings growth we see in our final analysis across subsidy groups therefore do not appear to be a function of differences in the direction or magnitude of demand growth across industries or occupations but are a function of unionization itself. For an unobserved factor to bias our estimates, therefore, it must differentially affect workers and firms with high potential subsidies only during the periods in which such workers actually did receive a large increase in their subsidies.

We first show that the increase in the base subsidies had a significant effect on the probability that affected workers join a union. Our regression is expressed for individual i in occupation o, industry c, and firm f, at time t, as:

$$Union_{iocft} = \beta_0 + \beta_1 S_{ft} + \beta_2 ND_{ft} + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \varepsilon_{iocft}, \tag{6}$$

where $Union$ is an indicator variable taking the value of one if the worker is a member of a union. We include fixed effects for highest completed educational program (δ_{Ed}), which includes indicators for secondary education tracks, post-secondary majors, and tertiary concentrations; discrete age buckets (π_{Age}); occupation-by-industry fixed effects (γ_{oc}); and year fixed effects τ_t.[18] The education fixed effects allow us to non-parametrically compare workers with the same educational credentials. The age fixed effects flexibly control for differential determinants of unionization and earning over the age profile. The occupation-by-industry fixed effects control for any cross-sectional differences in baseline propensity to unionize, differences in baseline union dues for different types of workers a the firm, and other unobserved, time-invariant factors. The year fixed effects absorb any systematic changes in unionization propensity over time that concern all workers.

We next estimate a similar equation to Equation 6 but incorporate a control for the employment concentration relevant to each firm and a full set of interactions between concentration and the subsidy as well as the net-of-subsidy union due. We estimate this regression as individual workers may perceive differential gains to unionization as a function of the em-

[18]The age categories are under age 25, 25-35, 36-45, 46-55, 56-65, and 65 and over.

ployer's power over labor demand. Relating this back to the conceptual framework provided in Section 2.3, this would be the case if a worker believes that the unionized wage net of subsidized dues is greater in a concentrated market than in a competitive market, in which case a union membership subsidy would have a larger impact on a worker's willingness to unionize in a concentrated market. We estimate the following equation:

$$\begin{aligned} Union_{iocft} =& \beta_0 + \beta_1 Subsidy_{ft} + \beta_2 ND_{ft} + \beta_3 \overline{HHI}_f + \beta_4 \overline{HHI}_f * Subsidy_{ft} \\ & + \beta_5 \overline{HHI}_f * ND_{ft} + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \varepsilon_{iocft}, \end{aligned} \tag{7}$$

where $\overline{HHI}_f$ represents the firm-wide average HHI index for firm f fixed at the firm's first year in the data.

The key result of this exercise is that firms with larger increases in subsidies (treatment intensity) will have higher rates at which workers at the firm become members of a union. We, therefore, use the predictions from these regressions to calculate the predicted union density for each firm in the data in each year, which we call $\widehat{UD}_{ft}$. This is the mean of the predicted probability of union membership across all workers at the firm each year. Importantly, this predicted value jointly takes into account the individual characteristics of workers at the firm. It also means that the effect of an increase in the probability an individual joins a union on earnings only comes through changes in union density at the firm. This is important because a union's power is not contingent on a single worker's membership, but rather on the share of workers represented by the union (Freeman and Medoff, 1981). This is reflected by the U_f parameter in Section 2.3.

With predicted values of union density for each firm in the data in each year, we then estimate the effects of union density on log annual earnings for each worker:

$$Log(Earnings)_{iocft} = \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \tag{8}$$

where $\widehat{UD}_{ft}$ is the predicted union density for each firm in the data in each year based on treatment intensity calculated through Equation 6. We include a firm fixed effect such that we are comparing the effects of union density within the same firm over time, as well as the difference in the marginal effects of unionization after holding constant time-invariant characteristics of the firm and individual workers.

After estimating the average union earnings premium, we allow the effects of union density ($\widehat{UD}_{ft}$) to differentially affect earnings in concentrated markets. This is accomplished by including an interaction between the predicted firm union density from Equation 7 and our HHI measure in this equation:

$$Log(Earnings)_{iocft} = \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \alpha_2 \widehat{UD}_{ft} * \overline{HHI}_f + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \varepsilon_{iocft}, \tag{9}$$

Under the assumption that unions negotiate rent-sharing with employers, a union would have more room to bid up the wages of its workers in markets where there is substantial firm rent due to monopsonistic competition. In other words, unions have space to negotiate from the rents that the firms previously extracted from labor through monopsonistic wage setting. At the same time, the relative bargaining power of the labor union is weaker if employer power is greater because outside offers cannot be called upon in negotiations as leverage, and the threat of leaving the firm remains less credible. In terms of our conceptual model, as market concentration increases, Π_f / L_f^m goes up but Φ_f goes down. A priori, it is therefore unclear what the relationship between labor market concentration and the union wage premium is. If the change in profits/quasi-rents dominates, we would expect to see a greater union wage premium in concentrated markets than in non-concentrated markets. If the change in bargaining power dominates, we would expect to see the opposite.[19]

Conditional on the composition of workers at the firm, a union wage premium can stem from three distinct sources: recapturing rents from labor market power, capturing rents from the product market, or productivity gains. To test for the relative contributions of possible rents from labor market power as opposed to product market power, we run a horse race in which we interact our measure of predicted union density with our measure of labor concentration as well as a proxy for product market concentration—an HHI for each firm based on their share of total industry operating revenues in Norway fixed in the firm's first year in the data ($\overline{HHI}_f^P$):

$$Log(Earnings)_{iocft} = \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \alpha_2 \widehat{UD}_{ft} * \overline{HHI}_f + \alpha_3 \widehat{UD}_{ft} * \overline{HHI}_f^P + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \tag{10}$$

where α_1 captures the marginal effect of union density on earnings in a firm in which both industry revenue (our proxy for product market power) and labor market concentration

[19] Throughout the paper, we cluster our standard errors at the firm level, which is where the union density effect is allocated. One might argue that we should account for the uncertainty of our predicted union density variable. To consider this, we estimate a Bayesian bootstrap (Rubin, 1981) in our main models. However, given our sample size, the standard errors of this exercise do not differ in any meaningful way from our clustered standard errors. As such, we choose to report our clustered standard errors throughout the paper, which, as noted in the literature on clustering, are likely conservative already (Abadie et al., 2017).

are both zero. The coefficient α_2 captures the change in the marginal effect as local labor market concentration increases holding constant the differential marginal effects from industry revenue concentration. Finally, α_3 conveys the difference in the marginal effects of union density as industry revenue concentration increases after netting out differences in the marginal effects from local labor market concentration. Thus, this specification allows us to disentangle the relative importance of labor market power and product market/industry power in explaining the earnings effects of union density.

Finally, we investigate what types of workers most benefit from union density within the firm. We do this by estimating full interactions of our model variables in Equation 7 with indicators for different groups to generate predictions for firm union density based on possibly heterogeneous responses to the tax subsidy across demographic groups and labor market concentration. We then calculate predicted union density at the firm from these individual predicted values and interact these and labor market HHI measures with the same group indicators in Equation 9. In terms of our conceptual model, this exercise explores the conventional egalitarian split assumption of union wage premiums across its members. We perform this exercise for three types of workers. First, we examine heterogeneity with respect to those whose earnings are below or above the occupation-specific median earnings at the firm. Above-median earnings in the firm-occupation cell may indicate differences in productivity or skill, attachment to the firm, or labor market attachment more generally. Second, we include indicators for white-collar occupations to separate out the effects across job classes. Third, we allow the marginal effects of union density and labor market concentration to differ by gender to examine if there are differential returns to union density for men and women across markets facing different levels of labor market concentration. This exercise is important for two reasons: first, it allows for differential impacts of the subsidies across worker types; second, it allows the subsequent returns to union density to differ based on who the marginal union member is.

5 Results

In this section, we present our results on the impact of unionization as a function of labor market concentration. In Section 5.1, we show results from our estimates examining the impact of changes in tax subsidies for union members on their willingness to unionize (treatment intensity). In Section 5.2, we examine the union wage premium as a function of labor market concentration. In Section 5.3, we explore the relative impact of product market/industry power and local labor market power. In Section 5.4, we explore effect heterogeneity with respect to worker type. In Section 5.5, we examine the dynamic employment effects of union density. Finally, in Section 5.6 we explore the effect of unions on firm and overall labor market inequality as a function of market concentration.

5.1 Effects of Union Subsidies on Union Membership

Table 1 shows the effect of the Norwegian tax subsidies on workers' propensity to unionize. These results are obtained through the estimation of Equations (2) and (3). In columns (1) and (2), we look at the relationship between subsidies and unionization without taking labor market concentration into account. While the regression underlying the results in column (1) includes occupation-by-industry, education, and age group fixed effects, the regression underlying the results in column (2) further includes individual-level fixed effects. The estimates in column (2) are thus identified exclusively based on individuals who switch industry-occupation cells. In columns (3) and (4), we study the relationship between the subsidies and unionization as a function of labor market concentration, using our preferred specification of 20 skill clusters. In columns (5) and (6), we perform a similar exercise but use 40 clusters as a means to examine robustness. Results using the more conventional occupation-based concentration measure, rather than the skill-based measure, are provided in the online appendix (Appendix Tables A10–A16).

The results in column (1) demonstrate that the subsidies had a strong impact on the probability that workers unionize. Raising the subsidy by 1,000 NOK increases the probability of being in a union by 12.5 percentage points. The coefficient on the subsidy in column (2) is slightly larger. The result in column (2) thus reveals that the relationship between union subsidies and unionization probability is robust to including individual fixed effects, such that identifying variation is coming from individuals who switch industry or occupation and are therefore exposed to different union dues and subsidies.[20]

In columns (3) and (4), we allow the impact of the subsidy to vary as a function of labor market concentration. The results show that the price elasticity of unionization is considerably larger in concentrated labor markets. This implies that individuals are considerably more willing to unionize in markets where labor demand is more concentrated. This is consistent with the notion that workers may be more concerned about employers trying to set their wages below marginal productivity in imperfect markets where there are limited outside options, and that they, therefore, expect returns to unionization to be higher under those circumstances.[21]

[20]For robustness, we also estimate this equation while excluding the net union due from the equation. The results are in Appendix Table A3 and suggest that failure to account for net union dues results in substantially larger estimated effects on union membership. This suggests that accounting for the net costs of membership is essential in this context.

[21]As shown in Dodini et al. (2020), labor market concentration in Norway is higher in smaller commuting zones. To ensure that commuting zone size is not driving our results, we have replicated our main findings using models that account for local labor market fixed effects. The results are provided in Appendix Tables A5 and A6, and demonstrate that our findings are robust to restricting the identifying variation to come

In columns (5) and (6), we re-estimate the regressions underlying the results in columns (3) and (4) but use 40 skill clusters rather than 20. Consistent with the main results, we find that the price elasticity of unionization with respect to the union subsidy is considerably larger in markets that experience monopsonistic competition. This implies that our results are not driven by the particular number of skill clusters used to identify market concentration.

5.2 Earnings Effects

Table 2 provides estimates on the effect of union density on individual log annual earnings using the changes in tax subsidies for union members as a measure of treatment intensity for firm-level union density. Panel A uses our full sample while Panel B restricts the sample to only the private sector. In column (1), we study the average effect of union density on log annual earnings at the firm without taking labor market concentration into account. In column (2), we study the impact of union density on log annual earnings at the firm as a function of labor market concentration, using our preferred specification of 20 skill clusters. In columns (3), we perform a similar exercise to that in column (2) but use 40 clusters for robustness.

Focusing on our full sample in Panel A, the results in column (1) reveal that a 1 percentage point increase in firm-level union density is associated with an increase in annual earnings of approximately 1.8 percent. This coefficient on union density is nearly identical to that in Barth et al. (2020) (1.9%) for the average effect of a 1 percentage point change in union density on worker wages at select manufacturing firms. This result is of great independent value, adding to the scarce literature that has been trying to isolate the union wage premium through the use of exogenous variation in unionization. To the best of our knowledge, this is the first causal estimate of the union wage premium in an entire country across all sectors and industries.

In column (2), we allow the impact of union density to vary as a function of the labor market concentration in the market where the firm is located. The results reveal that most of the union wage premium in column (1) is restricted to highly concentrated markets. Specifically, a 1 percentage point increase in firm-level union density is associated with an increase in annual earnings of approximately 1.1 percent in non-concentrated markets, and with an increase in annual earnings of approximately 2.5 percent in concentrated markets. In column (3), we re-estimate the regressions underlying the results in columns (2) but use 40 skill clusters rather than 20. The results are robust to this adjustment. This result is consistent with the notion that the greater the market imperfection, the greater the amount of firm rent that unions can (re)extract, despite possessing theoretically weaker bargaining strength in these markets. In other words, it is consistent with that notion that

only from within local labor markets.

$(\frac{\Phi_f \Pi_f}{L_f})_{concentrated} > (\frac{\Phi_f \Pi_f}{L_f})_{competitive}$. In addition, these results are consistent with the idea that unions may be able to correct market failures caused by firm concentration by pushing wages up towards the competitive equilibrium.

To reiterate this point succinctly, Figure 1 shows that the negative correlation between earnings level and labor market concentration is strongly ameliorated by higher rates of predicted unionization from our treatment intensity measure. At a low level of labor market concentration, moving from the bottom to the top quintile of predicted union density increases earnings by approximately 10-15 log points, while the same movement at an HHI of 0.1 would increase earnings by approximately 60 log points. These visual calculations closely mirror our estimates in Table 2.

The market imperfections generated by monopsonistic power, and the rents available to unions in concentrated markets, may be significantly larger in the private sector compared to the public sector. The reason underlying this hypothesis is that bargaining in the public sector is usually done over a fixed pot of money that arrives from a government entity through the political process, which limits the terms of bargaining to be primarily about allocations. The private sector bargains over not only the allocation of money to workers but over the size of the total pot, which can include labor and product market rents. To examine this hypothesis in detail, Panel B of Table 2 replicates Panel A but restricts the sample to only the private sector. The results suggest that the relationship between union density and earnings as a function of labor concentration is considerably more pronounced in the private sector. Specifically, in the private sector, the overall union earnings premium in non-concentrated markets is 0.5 percent for a 1 percentage point increase in union density, while the return in concentrated markets is 4.8 percent.

5.3 Source of Rents

In Table 3, we combine our primary labor market data with firm-level revenue data and explore the relative impact of product market power and labor market power on the union wage premium across markets that face different labor demand concentration. Prior literature has provided suggestive evidence of a strong correlation between labor and product concentration at the firm level, and understanding to what extent unions are able to extract product rent and labor rent is of independent interest (e.g., Marinescu et al. (2021); Qiu and Sojourner (2019); Lipsius (2018)).

In column (1), we show the effect of union density on annual earnings for the sub-sample of our main analysis sample with available revenue data. In column (2), we show results from running horse races between the labor HHI and product HHI based on our preferred 20 skill cluster categorization of concentration. In column (3), we repeat the exercise from column (2) but look at 40 skill clusters rather than 20.

The results in column (2) suggest that unions are considerably more effective in extracting both labor and product market rents in concentrated markets. Specifically, the coefficient on the interaction between union density and labor HHI is similar to the coefficient on industry revenue HHI. That we are able to identify different effects across these two sources of market power highlights that they are substantively different and that unions are able to identify and separately extract rents from both sources.

Figure 2 shows the marginal effects of union density by product HHI as a function of labor HHI. The figure illustrates that the two dimensions of rents both appear to contribute to earnings gains for workers. We believe that this is a novel finding with important policy implications, alluding to the fact that unions' ability to extract rent and reallocate this rent to their members depends not only on the extent of market imperfections but also on whether these imperfections are driven by labor concentration or possible product market concentration.

5.4 Heterogeneous Earnings Effects

In Table A9, we ask if the rents that unions extract from firms are allocated differently to different types of workers, relaxing the egalitarian split assumption in our conceptual framework that is conventionally imposed in the union literature. In Panel A, we examine heterogeneity with respect to worker productivity (proxied by whether the worker earns above or below median annual earnings *within occupation* at the firm). In Panel B, we explore heterogeneity with respect to white and blue-collar workers. In Panel C, we study effect heterogeneity across men and women. In all tables, we show results without taking labor market concentration into account (column (1)), using our preferred specification of 20 skill clusters to measure labor concentration (column (2)), and using 40 skill clusters to measure labor concentration (column (3)).

Panels A and B uncover two novel sets of results. First, the panels suggest that there exist modest union wage premiums in competitive markets among high-skilled and white-collar workers, but not among lower-skilled and blue-collar workers (column (1)). Second, the panels reveal that as markets become more concentrated, more of the additional rent that unions extract goes to lower-productivity and blue-collar workers. This implies that unions have an inequality-enhancing effect on earnings within narrow sectors in competitive markets, while this is not the case in concentrated markets characterized by monopsonistic competition.[22]

While speculative, we believe that these results are consistent with unions attempting to maximize union dues (Abowd and Lemieux, 1993). This objective leads unions to prioritize

[22]The finding that workers who are more highly paid benefit more from unionization and firm-level contracts is also a core finding in work done in Spain (Card and De La Rica, 2006).

higher-wage earners when allocating limited rents among members in competitive markets. First, higher-wage earners have more room to pay higher union dues. Workers with above-median earnings in the occupation-firm cell pay approximately 750 NOK more in annual union dues than those below the median, even taking into account occupation, industry, education, and age. Second, in highly competitive markets, high-skilled and white-collar workers are less likely to join unions, and the lack of significant firm rent means that unions have to prioritize which workers to push higher salaries on. Therefore, unions focus on satisfying and ensuring the continued membership renewal of high-skilled workers who are more likely to leave the unions and avoid paying more in dues. In addition, higher-productivity workers also may carry more weight as representative agents in negotiations. For example, threatening to strike carries more weight if the strikers are the firm's most productive workers. These three characteristics make higher-productivity workers strong potential members. As markets become more concentrated, the reduction in outside options for high-skilled and white-collar workers combined with the improved rent extraction opportunities available to unions means that they can shift focus to bargaining for a more general wage increase across all worker types. As we will show in Section 5.5, there is also a positive intensive-margin effect on employment at firms with high HHI, which aligns with this proposed objective of maximizing dues over time.[23]

The results presented above align with prior work on individual bargaining and outside offers. For example, Cahuc et al. (2006) find that between-firm competition for workers is important in determining wages and that only high-skilled workers have individual bargaining power. This finding is consistent with our results in non-concentrated markets if unions provide simple scalar increases of individual bargaining power. In concentrated markets, our results suggest a disproportionate increase in bargaining power for lower-productivity workers when union density increases at the same level of competition.

The results documented above also align with the effects of the distribution of pass-through of tax credits to higher-skilled workers documented by Carbonnier et al. (2022). In that setting, a tax credit for hiring low-wage workers in France is redistributed by the firm to higher-wage and higher-skilled workers within the firm in an effort to retain these workers. In our context, unions may similarly negotiate with firms in less concentrated markets by "redistributing" union dues from lower-skilled to higher-skilled workers.

[23] These earnings effects are not reflective of differential propensities to join unions in response to the subsidies, as we show in Appendix Table A2. Above-median workers are, in fact, less responsive to the subsidies in competitive markets and more responsive to concentration, contrary to the pattern of returns. This emphasizes the need for unions to use wage returns to retain higher-skilled workers. White-collar workers are more likely to respond to the subsidies in competitive markets (and pay lower union dues and face higher labor market concentration, as we show in Appendix Table A7), but this difference goes to zero as concentration increases.

With respect to effect heterogeneity across males and females, Figure A4 suggests that high levels of unionization may disproportionately increase the earnings of women in competitive markets on average. Specifically, the vertical distance between low- and high- union density firms is larger for women when HHI is low. As HHI increases, the marginal benefit to women *increases* relative to men, suggesting that unionization reduces gender earnings gaps more in concentrated markets at the macro level. That women are unionized at far higher rates than men in Norway may reflect this understanding among male and female workers.[24]

However, men and women tend to sort into different types of occupations, industries, and firms, and prior work suggests that women in Norway, on average, encounter higher levels of labor market concentration (Dodini et al., 2020). After controlling for these differences, Panel C of Table A9 shows that there is a modest union wage premium in competitive markets among men and that this premium is slightly larger than that among women (column 1). Panel C further shows that as markets become more concentrated, more of the additional rent that unions extract goes to women. This implies that unions exacerbate the gender wage gap within narrow sub-sectors in competitive markets, while they serve to reduce the gender wage gap in markets defined by a high degree of monopsonistic competition. This finding has important policy implications, revealing a potential role for unions in reducing the persistent overall gender wage gap through targeted involvement in concentrated markets. However, unionization may exacerbate earnings gaps within firms in competitive markets.

5.5 Employment Effects

In theory, a union wage premium should generate a reduction in employment in perfectly competitive labor markets. This is because employers in perfectly competitive labor markets pay wages equal to the marginal revenue product of labor. If a union is able to leverage its bargaining power to push wages above the marginal revenue product, at the new wage level, the employer will be unable to sustain current employment levels and will reduce either the number of workers (the extensive margin) or the number of work hours (the intensive margin).

In a labor market subject to monopsony power, on the other hand, a union wage premium may have no impact—and could even increase—employment. The reason is that employers with monopsony power can hire and retain workers for wages that are below the marginal revenue product of labor. If a union is able to leverage its power to push wages above the current wage offered by the employer, such as a wage equal to the marginal revenue product, the firm would hire more workers, but their profits would be lower. In such a market, a union

[24] Appendix Table A2 shows that female workers are more likely to join a union in response to the subsidies in competitive markets, so these differential earnings effects are not a matter of male workers being more responsive to subsidies in competitive markets. Women also tend to pay less in union dues, though they do experience higher labor market concentration (see Appendix Table A7).

density wage premium could lead to an increase in employment. However, a sufficiently large union wage premium that exceeds the marginal revenue product of labor would reduce employment, even in a concentrated market (see Appendix Figure A5 for further discussion).

To address this question, we conduct two analyses. First, we estimate Equation 9 at the individual level using the probability of holding a full-time position (working at least 30 hours per week) as the dependent variable.[25] Second, we estimate Equation 9 at the firm level using the number of workers at the firm as the dependent variable. This equation includes firm and year fixed effects. While the first regression enables us to shed light on the employment effect on the intensive margin, the second regression allows us to explore the employment effect on the extensive margin. We estimate these regressions using a 1-year lag for two reasons. First, as discussed in Section 2, Norwegian labor laws are strict.[26] This makes instantaneous firm actions and adjustments to employment difficult. Second, our measure of predicted union density depends on the contemporaneous distribution of the likelihood of union membership for those employed at the firm, so estimating current employment based on this distribution could lead to endogeneity even with our instrumented union density measure.

The results from these two sets of regressions are shown in Table 4.[27] In terms of the intensive margin, the results suggest that full-time status decreases this year as a consequence of increases in union density last year in competitive markets by about 1.3 percentage points with a 1 percentage point increase in union density. The opposite is true for highly-concentrated markets, with the interaction of union density and HHI being highly statistically significant and economically meaningful: firms in concentrated markets increase the likelihood their workers have full-time status by 2.9 percentage points. This suggests that unions generate an intensive-margin increase in hours as a function of labor market concentration.

The results for the extensive margin are small and not statistically significant across market concentration. Given the fact that standard theory would predict employment *losses* after unionization in competitive markets, finding a small and not significant dynamic effect itself is notable. However, this is perhaps unsurprising, as adjustments on the extensive margin—in particular in the Norwegian labor market —likely take longer than a year to materialize.

The results displayed in Table 4 are consistent with the predictions of new monopsony

[25]We explore this outcome, rather than total hours worked, as we only have access to total hours worked measured in relatively broad intervals for a large number of years of our analysis period.

[26]For example, there is no at-will employment and there is a general requirement of a 3-month notice period in the event of job terminations. Terminated employees also can take legal action.

[27]The results for the private sector only are in Appendix Table A19.

models that firms with labor market power grant lower wages and lower levels of employment than a competitive equilibrium. A union bidding wages up more in concentrated markets changes the employment level that maximizes a monopsonist's profits. This translates into future improvements on the employment dimension for employed workers. Our results align well with Azar et al. (2019b), which documents a similar pattern of employment effects due to minimum wage regulation across differently concentrated markets, though we only find that pattern at the intensive margin. The main difference between a minimum wage and a union-negotiated wage is that unions negotiate wage floors for different types of workers at the firm, and the wage floor is applicable to different labor market segments. Our results thus enable us to generalize some of the findings in the minimum wage literature to the broader workings of labor markets when a wage floor is imposed in imperfectly competitive labor markets.

5.6 Effects on Inequality

Existing economic research on labor unions has raised the question of how unions affect inequality both within sectors (earnings inequality within the set of all unionized workers) and across sectors (the gap between non-unionized and unionized workers). Given the heterogeneous treatment effects we have documented across labor market concentration, we extend our analysis in an effort to advance the literature on unions and inequality as well. We do so by considering two types of inequality: (1) inequality within firms that are exposed to a common level of union density (within sector, within firms) and (2) inequality within local labor markets, which proxies for the net effect of within- and across-sector inequality. Within each level, we consider three measures of inequality: the 90-10 ratio, the 90-50 ratio, and the 50-10 ratio.

To study (1), we account for differential responses to the tax subsidies at the firm level by including interactions between our tax subsidy measures and individual indicators for being below the 10th percentile, 10-50th, 50-90th, and above the 90th in our treatment intensity regressions. We then calculate overall firm union density using the individual predicted probabilities from this model (mean pr(union)). We regress each of our outcomes on the interaction of union density and concentration, including firm and year fixed effects. To explore (2), we regress the same three measures of inequality at the LLM level on average measures of concentration in the LLM and average predicted union density weighted by total employment at each firm in the LLM.

Table 5 shows the results from this exercise. In Panel A, we see that for firms in the least concentrated markets, the gap between the 90th and 50th percentiles widens when union density increases (column (1)). This is consistent with our findings in Panel A of Table A9; that above-median workers in each occupation benefit the most from unionization in competitive markets. In competitive markets, there is a broad expansion of earnings inequality (5.8%), but most of the effect loads on changes above the median (3.5%). The opposite is true in concentrated markets, where union density is uniformly associated with reductions in inequality, particularly in the top half (-8.3%). Overall, within-firm inequality in firms with greater labor market shares falls when union density at the firm rises, particularly above the median.

In Panel B of Table 5, we show that a similarly clear pattern arises at the local labor market level. When local labor markets are characterized by their *overall* market concentration across all firms and workers in the area, earnings inequality increases in more competitive labor markets and falls in markets characterized by more labor market concentration when union density increases. In column (1), we see this pattern appear when considering the 90-10 earnings ratio in the LLM. A one percentage point increase in local labor market unionization in competitive markets increases the 90-10 ratio by 0.034, or by just over 1 percent relative to the mean value of 3.2. In concentrated markets, a one percentage point increase in unionization decreases this ratio by 0.051, or by approximately 1.6 percent relative to the mean. In column (2), the upper half of the distribution becomes more compressed in competitive and concentrated markets, though the effects are not statistically significant. Column (3) shows that almost all of the 90-10 ratio effect is coming from the bottom of the distribution. In concentrated local labor markets, the gains from unionization accrue to the 10th percentile as well, reducing below-median local labor market inequality. The results for local labor markets suggest that cross-firm sorting may blunt the relative effects on within-firm inequality as it translates to total local labor market inequality.

Taken together, our results suggest that the effect of union density on earnings inequality is strongly determined by the level of localized labor market concentration faced by the marginal union member. As we show above, the marginal union member is more likely to be working in concentrated labor markets, which matters when we consider comparisons of our results to other work. For example, Card et al. (2004) find that wage inequality in the United States falls both within and between sectors as unionization rises. Our work supports that result in the context of concentrated labor markets in the Norwegian context, but not in the context of highly competitive markets.[28] Thus, our findings demonstrate that the impact of unions on inequality is more nuanced than that documented in previous work, and that variation in labor market concentration is an important factor that needs to be taken into account when considering the overall impact of unions.

[28] The fact that the average labor market in the US is estimated to be relatively concentrated supports this argument (e.g., Azar et al. (2020a)).

5.7 Extension

In Section 5, we presented new evidence on the impact of unionization as a function of labor market concentration. We did this by exploiting an exogenous shift in unionization at the firm and interacting this with existing measures of labor market concentration. An alternative approach would be to utilize exogenous shifts in labor market concentration and interact this with existing levels of union density. In Appendix B, we exploit the influx of imports from China to Norway in the early 2000s as an exogenous shifter of firm labor market concentration. We then use this to measure the effects of unionization on earnings when there are plausibly-exogenous changes to the level of labor market concentration. While this exercise relies on a stricter set of assumptions, it provides a complementary approach to our main empirical strategy and helps establish the robustness of our results to shocking labor market concentration rather than unionization.

6 Discussion

In this paper, we bring together the literatures on labor market concentration and unions by empirically examining the effects of unionization on the dynamics of worker earnings across differently concentrated markets. Existing empirical evidence has focused either on market power (e.g., Dodini et al. (2020); Schubert et al. (2020); Azar et al. (2020b)) or union power (e.g., Card et al. (2004); DiNardo and Lee (2004); Lee and Mas (2012)), without considering the potential interaction effects of the two. While these two strands of literature provide extremely important insights into the workings of labor markets, the lack of understanding of how these two forces interact—monopolistic unions and monopsonistic employers—severely limits our knowledge of the dynamics of labor markets.

Exploiting tax reforms to union due deductions as an exogenous shock to unionization, we demonstrate that the price elasticity of unionization has a very steep gradient over labor market concentration. We then show that there is an equally strong gradient in the union earnings premium and that the union wage premium loads heavily on highly concentrated markets. This result is consistent with the notion that the greater the market imperfection, the greater is the amount of firm rent that unions can extract despite a potentially disadvantageous bargaining position due to the inability to leverage outside options in negotiations. It also suggests a potentially important role of unions in limiting the market failures generated by monopsonistic competition.

Running horse races on product market power and labor market power on the union wage premium across differently-concentrated markets suggest that unions are effective in targeting and extracting both labor rent and product rent. Specifically, the coefficients on the interaction between union density and labor HHI are not statistically different from the coefficient on the interaction between union density and product HHI in our main specification, though the measures are calculated at different levels, which suggests caution in overinterpreting this result. That we identify different effects across these two sources of market power highlights that they are substantively different components and that the correlations between the two are not as strong as previously thought.

We document important heterogeneity with respect to the type of worker that benefits from union membership as a function of labor market concentration. Specifically, we show that the modest union wage premiums that exist in competitive markets are loading on high-skilled and white-collar workers. As the degree of market concentration increases, more and more of the additional rents that unions extract go to lower-ability and blue-collar workers. This implies that unions may have an inequality-enhancing effect on wages in some subsectors of competitive markets, while this is not the case in concentrated markets characterized by monopsonistic competition. While speculative, we suggest that this pattern of results is consistent with unions aspiring to maximize union dues.

Related to this point, Card et al. (2004) discuss the concepts of "between-sector" versus "within-sector" wage inequality, defined as inequality measured between union and non-union workers and inequality among unionized workers, respectively.[29] Our results suggest that there is a notable, positive effect on inequality within small *sub-sectors* of the unionized sector when labor markets are competitive. In other words, when comparing unionized workers to each other within the same firm, inequality increases when outside options for the most productive and highly-paid workers are more feasible to enter. This does not appear through a redistribution of resources, but rather through unequal benefits to unionization. These findings suggest that prior work that has identified reductions in within-sector inequality as a function of unionization may be operating within concentrated markets. In the overall local economy, unionization increases total within- and across-sector inequality when labor markets are more competitive, but reduces local inequality when markets are concentrated. We believe that the results provided in this analysis introduce a new element into the debate on the relationship between unions and inequality, allowing us to deepen our understanding of how the dynamics of unions impact societal goals such as wage inequality and provide a new avenue for future research.

We find a similar result pattern when exploring effect heterogeneity between men and women, suggesting that unions exacerbate the gender earnings gap in occupation-firm subsectors within competitive labor markets, while they serve to reduce the gender earnings gap on the whole, particularly in concentrated markets. This finding has potentially important policy implications, revealing a role for unions in reducing the persistent macro gender

[29] Discussions of these effects date back to the 1950s (Friedman, 1956; Rees, 1989).

earnings gap through involvement in concentrated markets. This is particularly important as other work has found that women, on average, face more concentrated markets in their occupations (Dodini et al., 2020; Ransom and Oaxaca, 2010).[30]

In terms of employment effects, we show that firms in concentrated markets respond to an increase in union density and worker earnings by raising employment the following year on the intensive margin, with small and not statistically significant effects on the extensive margin. This employment effect is consistent with monopsony theory, which predicts that both wages and employment levels will lie below a competitive equilibrium in the presence of monopsony power.

We believe that the results from this paper have important policy implications. Specifically, our estimates suggest that the modest union earnings premium in more competitive markets comes at a modest employment cost. That there is a large sizable union wage premium in highly concentrated markets, on the other hand, may point to unions as being able to ameliorate a market failure generated by employer power. Given that our estimates are identified based on a simple policy change—modest tax subsidies for union dues —and that the workers who disproportionately benefit from unionization are those more likely to be in concentrated markets, this policy lever may serve to decrease overall earnings inequality. It is also a policy lever that, while general in *scope*, is well-targeted in *effect*. While speculative, the high unionization rates in Norway may therefore be one reason for its relatively compressed pre-tax labor earnings structure relative to countries such as the United States.

The results from this analysis also have implications for regulatory policy. According to the Horizontal Merger Guidelines used as the basis for antitrust enforcement by the US Department of Justice (DOJ), an HHI above 0.25 is considered "highly concentrated," and the US Congress has recently proposed giving the DOJ a mandate to regulate mergers and acquisitions with labor concentration in mind. However, our estimates emphasize that *unionization rates* ought to be considered as well and that certain mergers and acquisitions may not be distortionary—and could even be beneficial—in already concentrated markets as long as there is a sufficient union presence.

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[30]Women in our sample are also disproportionately likely to be in the "below-median" earner class within occupation-firm cells, with 59 percent of women falling within the "below median" earner group. Unions appear to reward workers with either high levels of productivity or attachment to the labor market or firm, which may explain part of the gender gap in returns within firms.

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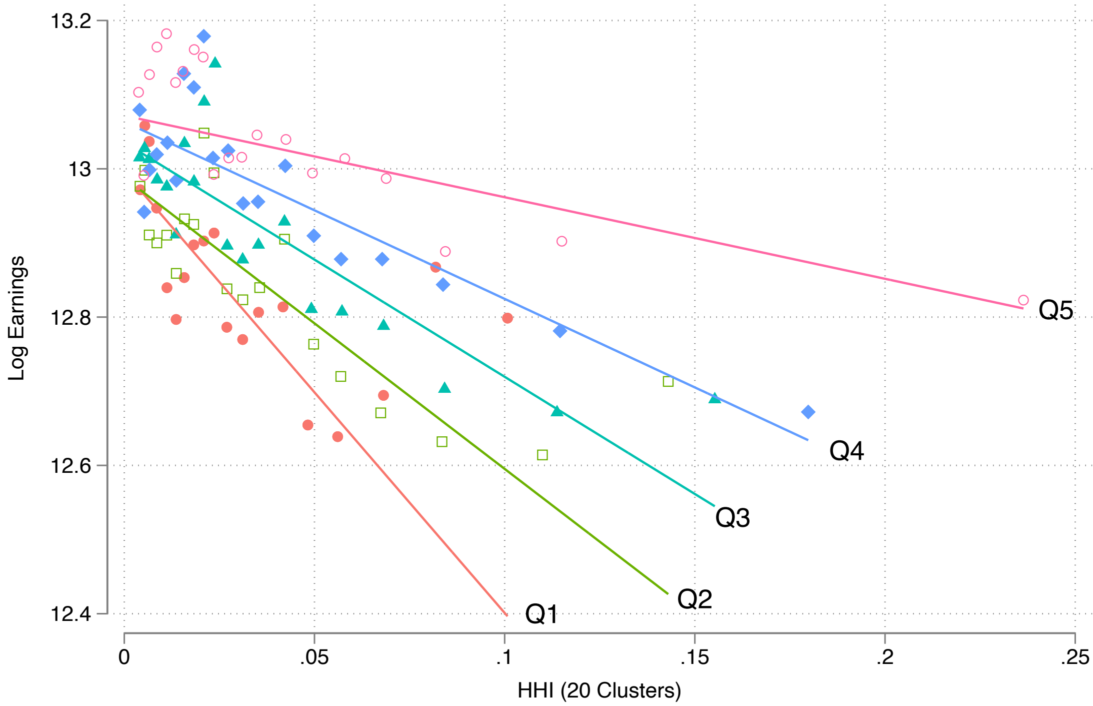
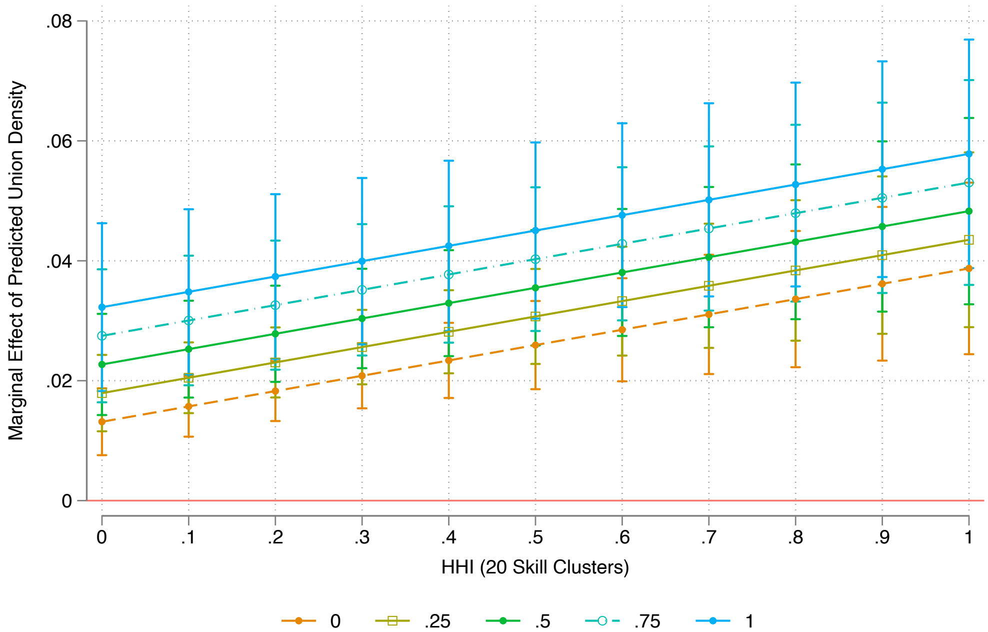
Figures

Figure 1: Log Annual Earnings (NOK) and Labor Market HHI by Quintiles of Predicted Firm Union Density



Source: Authors' calculations of Norwegian registry data.
Notes: Predicted union densities are based on average predicted unionization rates at each firm from Equation 7 as described in the text.

Figure 2: Marginal Effects of Union Density by Industry Revenue HHI



Source: Authors' calculations of Norwegian registry data as described in the text. Estimates reflect the marginal effects from Equation 10.
Notes: Standard errors are clustered at the firm level and calculated at each margin using the delta method.

Tables

Table 1: The Effect of Tax Subsidies on Propensity to Unionize

VARIABLES	(1) No HHI	(2) No HHI	(3) 20 Clusters	(4) 20 Clusters	(5) 40 Clusters	(6) 40 Clusters
Subsidy (1,000 NOK)	0.125**	0.151***	0.0926*	0.131***	0.0958*	0.135***
	(0.0517)	(0.0198)	(0.0527)	(0.0199)	(0.0528)	(0.0200)
HHI x Subsidy			0.171***	0.221***	0.109***	0.141***
			(0.0479)	(0.0294)	(0.0419)	(0.0263)
Observations	16,181,785	15,992,458	16,181,785	15,992,458	16,181,785	15,992,458
R-squared	0.232	0.739	0.234	0.739	0.234	0.739
Individual FE	No	Yes	No	Yes	No	Yes
Avg Pr(Union)	0.597	0.597	0.597	0.597	0.597	0.597
Mean Subsidy 2001 (1,000 NOK)	0.252	0.252	0.252	0.252	0.252	0.252
Mean Subsidy 2014 (1,000 NOK)	1.022	1.022	1.022	1.022	1.022	1.022

Source: Authors' estimates as described in the text. Estimates correspond to Equations 6 and 7.
Notes: Standard errors are in parentheses and are clustered at the firm level. All estimates include fixed effects for occupation by industry cells, detailed educational program, age group, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table 2: Effect of Union Density on Log Annual Earnings by Labor Market Concentration

	Panel A: Full Sample		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0181***	0.0114***	0.0107***
	(0.00219)	(0.00218)	(0.00221)
Predicted Firm Union Density * HHI		0.0141***	0.0185***
		(0.00301)	(0.00271)
Observations	16,181,780	16,181,780	16,181,780
R-squared	0.581	0.581	0.581

	Panel B: Private Sector Only		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0105***	0.00512**	0.00482**
	(0.00207)	(0.00218)	(0.00216)
Predicted Firm Union Density * HHI		0.0431***	0.0298***
		(0.00540)	(0.00560)
Observations	11,009,362	11,009,362	11,009,362
R-squared	0.593	0.593	0.593

Source: Authors' estimates as described in the text. Estimates correspond to Equations 8 and 9.

Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.

* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table 3: The Effect of Union Density on Log Annual Earnings by Labor and Product Market Concentration

VARIABLES	(1) No Labor HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0147***	0.0132***	0.0126***
	(0.00256)	(0.00285)	(0.00279)
Predicted Firm Union Density * Labor HHI		0.0256***	0.0106
		(0.00815)	(0.00752)
Predicted Firm Union Density * Industry Revenue HHI		0.0191***	0.0176***
		(0.00664)	(0.00654)
Change in ME with 10 ppt Change in Labor HHI		0.0026	0.0011
Change in ME with 10 ppt Change in Industry Revenue HHI		0.0019	0.0018
Observations	7,634,149	7,634,149	7,634,149
R-squared	0.610	0.610	0.610

Source: Authors' estimates as described in the text. Estimates correspond to Equation 10.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table 4: Employment Effects of Lagged Union Density by Concentration

VARIABLES	(1) Pr(Hours>30)	(2) Pr(Hours>30)	(3) Workers	(4) Workers
Lagged Predicted Union Density	0.00817**	-0.0128***	0.0432	0.364
	(0.00330)	(0.00399)	(0.912)	(1.093)
Lagged Predicted Union Density * HHI		0.0419***		-0.788
		(0.00522)		(2.154)
Constant	0.308	1.435***	91.48*	74.23
	(0.190)	(0.227)	(52.00)	(62.07)
Observations	14,425,353	14,425,353	221,672	221,672
R-squared	0.286	0.286	0.898	0.898

Source: Authors' estimates as described in the text.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table 5: Effect of Union Density on Inequality in Firms and Local Labor Markets

Panel A: Firm Level Inequality

VARIABLES	(1) Firm 90/10	(2) Firm 90/50	(3) Firm 50/10
Predicted Union Density	0.149***	0.0526***	0.0384***
	(0.00958)	(0.00284)	(0.00476)
Predicted Union Density x HHI	-0.213***	-0.0824***	-0.0438***
	(0.0154)	(0.00482)	(0.00780)
Dep Variable Mean	2.58	1.52	1.70
Pct Effect Union Density	5.78 %	3.46 %	2.26 %
Pct Effect Union Density x HHI	-8.26 %	-5.42 %	-2.58 %
Observations	252,363	252,363	252,363
R-squared	0.625	0.616	0.620
Firm FE	Yes	Yes	Yes

Panel B: Local Labor Market Level Inequality

VARIABLES	(1) LLM 90/10	(2) LLM 90/50	(3) LLM 50/10
Predicted Union Density	0.0339**	-0.0108	0.0319***
	(0.0156)	(0.00664)	(0.00880)
Predicted Union Density x HHI	-0.0509***	0.00160	-0.0313***
	(0.0112)	(0.00337)	(0.00709)
Dep Variable Mean	3.22	1.68	1.91
Pct Effect Union Density	1.05 %	-0.64 %	1.67 %
Pct Effect Union Density x HHI	-1.58 %	0.10 %	-1.64 %
Observations	2,396	2,396	2,396
R-squared	0.975	0.983	0.890
LLM FE	Yes	Yes	Yes

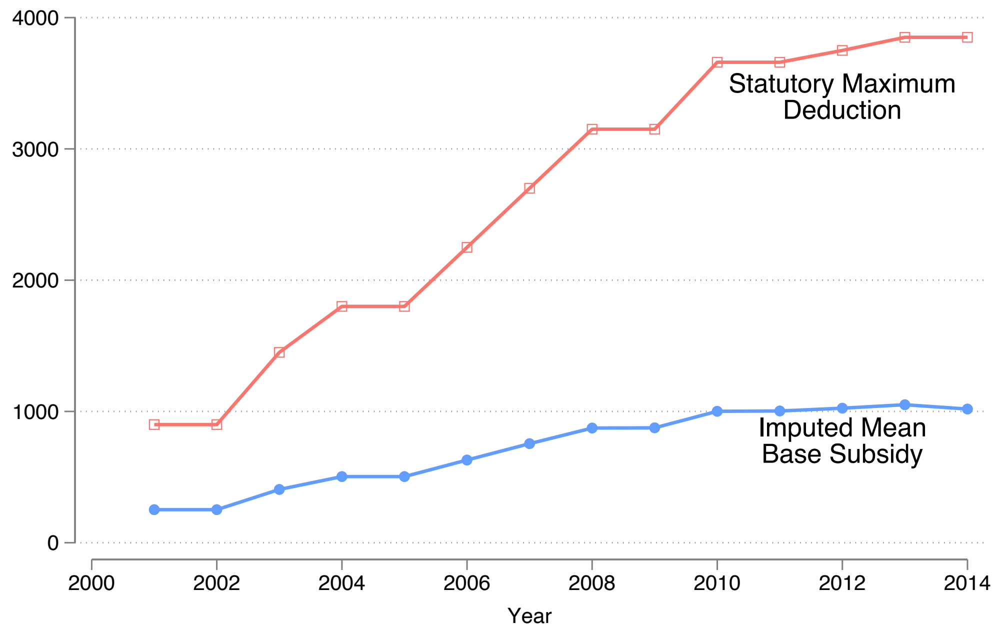
Source: Authors' estimates as described in the text.

Notes: Standard errors are in parentheses and are clustered at the firm level in Panel A and the local labor market level in Panel B. Regressions include unit and year fixed effects and are weighted by total employment at the firm or local labor market. Predicted union density in Panel A incorporates differential responses to tax subsidies by interacting the instruments with indicators for within-firm earnings percentile ranges [0,10), [10-50), [50,90), and [90,100)—the same percentiles at which we measure inequality.

* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

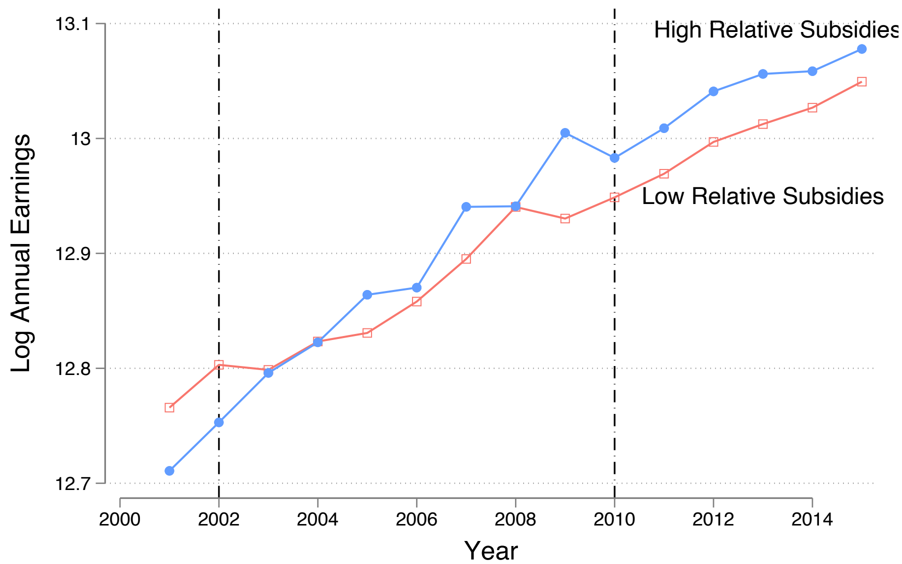
A Appendix: Additional Tables and Figures

Figure A1: Statutory Maximum Deduction and Imputed Mean Subsidy for Union Dues (NOK)

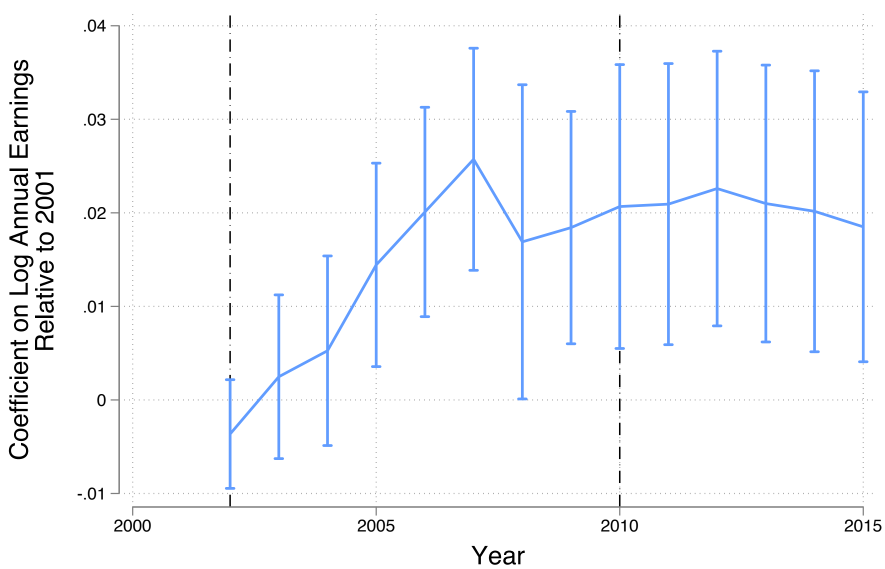


Source: Authors' calculations of Norwegian registry data.
Notes: Imputed base subsidies are calculated as the base tax rate times the lesser of imputed union dues at the occupation-by-industry cell or the statutory maximum deduction.

Figure A2: Log Annual Earnings by Relative Subsidies
Panel A: Raw Trends

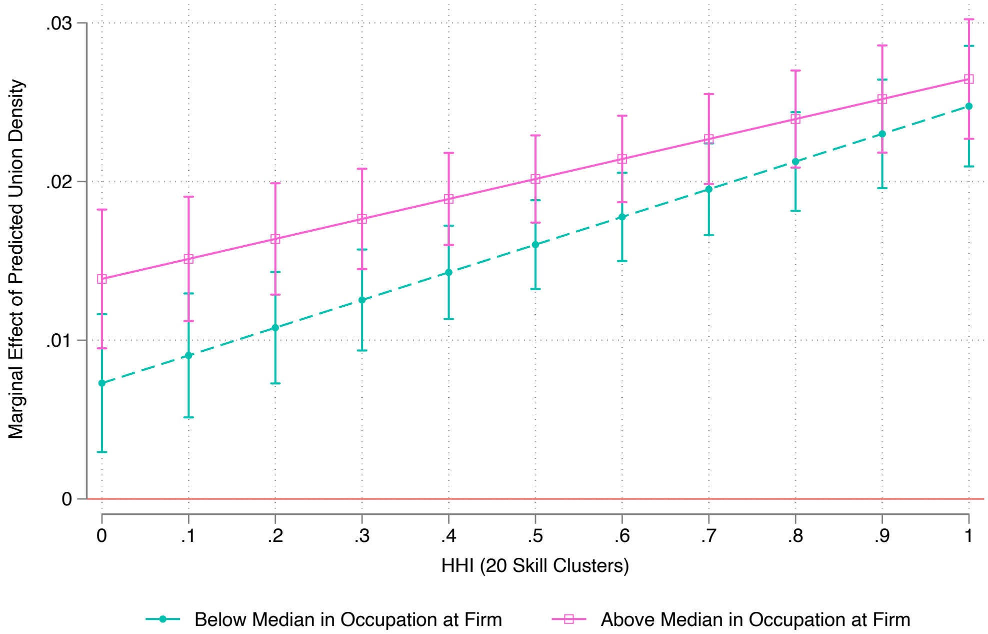


Panel B: Estimated Coefficients



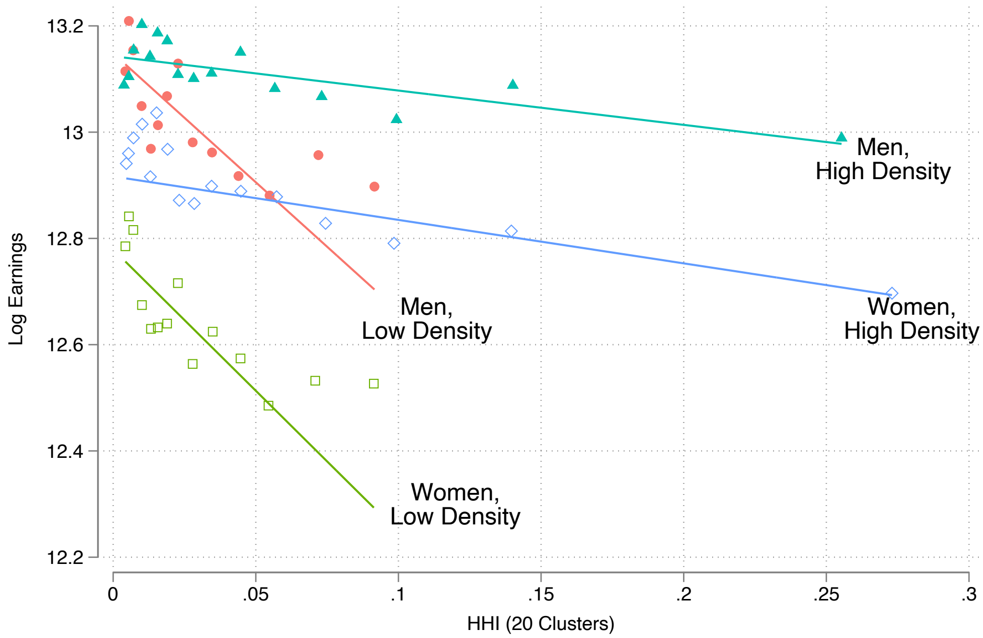
Source: Authors' calculations of Norwegian registry data.
Notes: Relative subsidies are calculated by residualizing firm-level subsidies by regressing subsidies on the net-of-subsidy union dues. "High" relative subsidies are above median subsidies in the 2003-2015 period, and "low" relative subsidies are below median.
Panel B coefficients are for an interaction between an indicator for being at a "high" relative subsidies firm and year and are relative to base year 2001. Coefficients represent differential changes relative to 2001 in "high" subsidy firms relative to "low" subsidy firms.
The regression includes fixed effects for industry by occupation cells, age group, educational attainment, and year. 95 percent confidence intervals based on standard errors clustered at the firm level.
The maximum deduction increased most substantially from 2003 to 2010 (see Figure A1). The maximum deduction was stable from 2010 onward, where the gap between the high- and low-subsidy groups stabilized. The parallel trends check, therefore, should apply to the pre-2003 period and the period after 2010.

Figure A3: Marginal Effects of Union Density by Firm-Occupation Median



Source: Authors' calculations of Norwegian registry data as described in the text.
Notes: Standard errors are clustered at the firm level and calculated at each margin using the delta method.

Figure A4: Log Annual Earnings (NOK) and Labor Market HHI by Gender and Top vs Bottom Quintiles of Predicted Firm Union Density

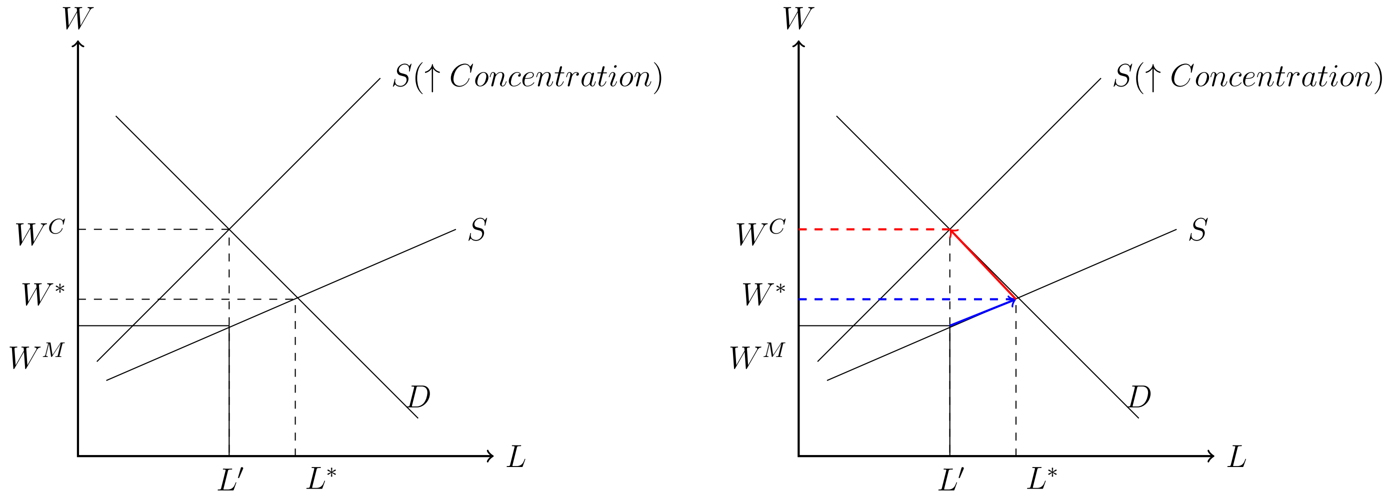


Source: Authors' calculations of Norwegian registry data.
Notes: Predicted union densities are based on average predicted unionization rates at each firm from Equation 7 as described in the text.

Figure A5: Wage Floors in a Monopsony Framework

Panel A: Monopsony Model of the Labor Market

Panel B: Wage Floor at W* (Blue) or above W* (Red)



Panel A shows a basic monopsony model, in which the firm is a price-setter in the factor market. The defining feature of a firm with monopsonistic power is the upward sloping labor supply curve that it faces, with the marginal cost of labor exceeding the opportunity cost of labor at each employment level. Under profit maximization, wages will be set below the marginal revenue product of labor, and employment will be set below the competitive equilibrium. Specifically, rather than being at a market equilibrium in a perfectly competitive setting with wage W^* and employment level L^*, workers provide labor supply to the firm at the steeper $S(\uparrow Concentration)$. This curve intersects the labor demand curve at L', resulting in monopsony wages to the workers of W^M. At this wage level, the workers are being underpaid relative to the revenue they generate to the firm.

A conventional policy tool in the presence of monopsonistic market power is the minimum wage. By restricting firms' wage-setting ability at the lower end of the wage distribution, policymakers can increase wages for low-paid workers and encourage higher wages for those just above them. In addition, modest increases in the minimum wage can lead to gains both in wages and employment. These positive employment effects are typically rationalized through a monopsony framework similar to that in Panel A and has been discussed as early as 1946 (e.g., Stigler (1946)). Panel B shows the general result of setting a wage floor where there is imperfect competition in the labor market. The important difference between a minimum wage and a union-negotiated wage is that unions negotiate wage floors for different types of workers at the firm, and the wage floor is applicable to different labor market segments. Any wage floor that sets a wage between W^M and W^* will result not only in higher wages to the worker but higher employment as well, moving along the blue arrow towards L^*. This is precisely the result found in Azar et al. (2019b): low-wage labor markets with higher concentration experience employment gains in response to minimum wage hikes. If negotiated wages are exactly at W^*, the negative wage effects (and market failure) of imperfect competition have been eliminated. However, wages above W^* may have disemployment effects relative to the competitive equilibrium as employment moves along the demand curve to the left along the red arrow. If wages are set above W^C, employment may therefore fall further.

Table A1: Key Sample Summary Statistics

VARIABLES	(1) Mean	(2) SD
Pr(Union)	0.6020	0.4895
Firm Union Density	0.5999	0.2633
Real Annual Earnings (2015 NOK)	463,060	273,601
Age	41.93	11.65
Imputed Tax Subsidy (1,000s NOK)	0.7529	0.2886
Imputed Net Union Due (1,000s NOK)	3.1825	0.5662
Labor HHI (20 Clusters)	0.0429	0.0538
Labor HHI (40 Clusters)	0.0512	0.0611
Product HHI (National Industry)	0.0369	0.0773
Public Sector Industry Worker	0.3196	0.4663
N	16,181,785	

Source: Norwegian registry data as described in the text.
Notes: The sample is limited to full-year workers at firms with at least ten workers. We take a 70% random sample of the full set of individuals to ease computational constraints.

Table A2: Heterogeneous Effects of Tax Subsidies on Propensity to Unionize by Subgroup

VARIABLES	(1) Women	(2) Above Occ-Firm Median	(3) White Collar
Subsidy (1,000 NOK) [Base]	0.0845	0.0989*	0.0638
	(0.0533)	(0.0527)	(0.0525)
Subsidy (1,000 NOK) * HHI (20 Clusters) [Base]	0.0907	0.106**	0.155
	(0.0811)	(0.0491)	(0.123)
HHI [Base]	0.740***	0.951***	0.315
	(0.246)	(0.226)	(0.374)
Subsidy * Group Interaction	0.0143***	-0.0138***	0.0293**
	(0.00542)	(0.00495)	(0.0121)
Subsidy * Group Interaction * HHI	0.0925	0.134***	-0.00231
	(0.0665)	(0.0406)	(0.120)
Group * HHI	1.071***	0.673***	1.255***
	(0.173)	(0.138)	(0.399)
Constant	0.330***	0.334***	0.331***
	(0.0452)	(0.0451)	(0.0455)
Observations	16,181,785	16,181,785	16,181,785
R-squared	0.236	0.236	0.234

Source: Authors' estimates of Equation 7 interacting right hand side variables with group indicators.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A3: The Effect of Tax Subsidies on Propensity to Unionize, Excluding Net Union Due

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 3 Digit Occupation
Subsidy (1,000 NOK)	0.260***	0.244***	0.247***	0.239***
	(0.0541)	(0.0555)	(0.0555)	(0.0552)
HHI		0.391***	0.418***	0.322***
		(0.0352)	(0.0317)	(0.0220)
Subsidy x HHI		0.0815*	0.0172	0.000260
		(0.0448)	(0.0396)	(0.0252)
Observations	8,094,215	8,094,215	8,094,070	8,094,215
R-squared	0.232	0.234	0.583	0.234

Source: Authors' estimates of Equation 7 excluding the net union due from the equation.
Notes: Standard errors are in parentheses and are clustered at the firm level. For computation ease, estimates are from a 35% random sample. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A4: The Effect of Tax Subsidies on Propensity to Unionize Using Subsidy Ratio

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 3 Digit Occupation
Subsidy Ratio	0.235	0.167	0.165	0.195
	(0.145)	(0.147)	(0.148)	(0.148)
Subsidy Ratio x HHI		0.598***	0.387***	0.340***
		(0.161)	(0.144)	(0.108)
Observations	8,094,215	8,094,215	8,094,215	8,094,215
R-squared	0.233	0.235	0.235	0.236

Source: Authors' estimates of Equation 7 replacing the raw subsidy with a subsidy ratio and the net union due with the inverse net union due per (Barth et al., 2020).
Notes: Standard errors are in parentheses and are clustered at the firm level. For computation ease, estimates are from a 35% random sample. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A5: The Effect of Tax Subsidies on Propensity to Unionize, Including Local Labor Market Controls

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 3 Digit Occupation
Subsidy (1,000 NOK)	0.113**	0.0914*	0.0935*	0.0917*
	(0.0510)	(0.0518)	(0.0522)	(0.0519)
HHI		0.807***	0.877***	0.791***
		(0.212)	(0.149)	(0.117)
Subsidy x HHI		0.143***	0.0832**	0.0587**
		(0.0462)	(0.0413)	(0.0280)
Observations	8,075,252	8,075,252	8,075,252	8,075,252
R-squared	0.241	0.242	0.242	0.242

Source: Authors' estimates of Equation 7 with the addition of fixed effects for local labor market. Notes: Standard errors are in parentheses and are clustered at the firm level. For computation ease, estimates are from a 35% random sample. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, local labor market, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A6: The Effect of Union Density on Log Annual Earnings by Labor Market Concentration, Including Local Labor Market Controls

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 3 Digit Occupation
Predicted Union Density	0.0190***	0.0128***	0.0113***	0.00712***
	(0.00239)	(0.00239)	(0.00236)	(0.00206)
Union Density x HHI		0.0161***	0.0220***	0.0104***
		(0.00319)	(0.00294)	(0.00185)
Observations	8,075,103	8,075,103	8,075,103	8,075,103
R-squared	0.585	0.585	0.585	0.585

Source: Authors' estimates of Equation 9 with the addition of fixed effects for local labor market. Notes: Standard errors are in parentheses and are clustered at the firm level. For computation ease, estimates are from a 35% random sample. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, local labor market, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A7: Summary Statistics - Means by Subgroup

VARIABLES	(1) Union Dues Paid (NOK)	(2) Subsidy (1,000 NOK)	(3) Net-of-Subsidy Dues (1,000 NOK)	(3) HHI (20 Clusters)
Men	4799	0.7467	3.2167	0.0332
Women	4347	0.7587	3.1500	0.0520
Below Occ-Firm Median	4071	0.7489	3.1800	0.0427
Above Occ-Firm Median	5016	0.7574	3.1855	0.0431
Not White Collar	5183	0.7423	3.4367	0.0324
White Collar	4383	0.7556	3.1173	0.0455

Source: Authors' estimates using Norwegian register data.

Table A8: The Effect of Predicted Labor Concentration on Union Premium

VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Average Union Density x Predicted HHI	0.0332***	0.0405***
	(0.0122)	(0.0149)
Predicted HHI	-2.014***	-2.456***
	(0.683)	(0.833)
Observations	10,351,840	10,351,840
R-squared	0.557	0.557

Source: Authors' estimates as described in the text.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A9: Heterogeneous Effects of Union Density on Log Annual Earnings

	Panel A: Above vs Below Firm-Occupation Median		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.00462**	-0.00150	-0.00239
	(0.00207)	(0.00199)	(0.00199)
Predicted Firm Union Density * HHI		0.0276***	0.0294***
		(0.00280)	(0.00249)
Union Density * Above Firm-Occ Median	0.00633***	0.00657***	0.00664***
	(4.02e-05)	(4.97e-05)	(5.03e-05)
Union Density * HHI * Above Firm-Occ Median		-0.00498***	-0.00550***
		(0.000332)	(0.000322)
R-squared	0.719	0.719	0.719
	Panel B: White Collar vs Other Occupations		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.00811***	0.00160**	0.00147*
	(0.00117)	(0.000799)	(0.000840)
Predicted Firm Union Density * HHI		0.0236***	0.0259***
		(0.00226)	(0.00222)
Union Density * White Collar	0.00206***	0.00220***	0.00226***
	(0.000326)	(0.000403)	(0.000397)
Union Density * HHI * White Collar		-0.00143***	-0.00108**
		(0.000504)	(0.000427)
R-squared	0.581	0.581	0.581
	Panel C: By Gender		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0168***	0.00857***	0.00778***
	(0.00146)	(0.00150)	(0.00147)
Predicted Firm Union Density * HHI		0.0177***	0.0211***
		(0.00261)	(0.00242)
Union Density * Female	-0.00261***	-0.00275***	-0.00275***
	(3.79e-05)	(4.68e-05)	(4.74e-05)
Union Density * HHI * Female		0.00303***	0.00248***
		(0.000362)	(0.000312)
R-squared	0.595	0.595	0.595

Source: Authors' estimates corresponding with Equations 8 and 9 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A10: The Effect of Tax Subsidies on Propensity to Unionize, Occupation-Specific HHI

VARIABLES	(1) 2-Digit Occupation	(2) 2-Digit Occupation	(3) 3-Digit Occupation	(4) 3-Digit Occupation
Subsidy (1,000 NOK)	0.0769	0.111***	0.0968*	0.127***
	(0.0523)	(0.0197)	(0.0523)	(0.0199)
HHI x Subsidy	0.180***	0.258***	0.0735***	0.115***
	(0.0398)	(0.0264)	(0.0282)	(0.0202)
Observations	16,181,785	15,992,458	16,181,785	15,992,458
R-squared	0.234	0.739	0.235	0.739
Individual FE		Yes		Yes
Avg Pr(Union)	0.575	0.575	0.575	0.575
Mean Subsidy 2001 (1,000 NOK)	0.252	0.252	0.252	0.252
Mean Subsidy 2014 (1,000 NOK)	1.018	1.018	1.018	1.018

Source: Authors' estimates as described in the text. Estimates correspond with Equations 6 and 7.
Notes: Standard errors are in parentheses and are clustered at the firm level. All estimates include fixed effects for occupation by industry cells, detailed educational program, age group, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A11: Effect of Union Density on Log Annual Earnings by Labor Market Concentration, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.0181***	0.00368**	0.00566***
	(0.00219)	(0.00171)	(0.00182)
Predicted Firm Union Density * HHI		0.0121***	0.00927***
		(0.00266)	(0.00169)
Observations	16,181,780	16,181,780	16,181,780
R-squared	0.581	0.581	0.581

Source: Authors' estimates as described in the text. Estimates correspond with Equations 8 and 9.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A12: Effect of Union Density on Log Annual Earnings by Labor Market Concentration - Private Sector Only, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.0105***	-0.000859	-5.87e-05
	(0.00207)	(0.00195)	(0.00191)
Predicted Firm Union Density * HHI		0.0217***	0.0141***
		(0.00483)	(0.00282)
Observations	11,009,362	11,009,362	11,009,362
R-squared	0.593	0.593	0.593

Source: Authors' estimates as described in the text. Estimates correspond with Equations 8 and 9.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A13: The Effect of Union Density on Log Annual Earnings by Labor and Product Market Concentration, Occupation-Specific HHI

VARIABLES	(1) No Labor HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.0147***	0.0122***	0.00922***
	(0.00256)	(0.00268)	(0.00259)
Predicted Firm Union Density * Labor HHI		-0.00421	0.00345
		(0.00696)	(0.00372)
Predicted Firm Union Density * Industry Revenue HHI		0.0199***	0.0177***
		(0.00741)	(0.00653)
Change in ME with 10 ppt Change in Labor HHI		-0.0004	0.0003
Change in ME with 10 ppt Change in Industry Revenue HHI		0.0020	0.0018
Observations	7,634,149	7,634,149	7,634,149
R-squared	0.610	0.610	0.610

Source: Authors' estimates as described in the text. Estimates correspond with Equation 10.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A14: Effect of Union Density on Log Annual Earnings by Labor Market Concentration and Firm-Occupation Median, Occupation-Specific HHI

	(1)	(2)	(3)
VARIABLES	No HHI	2-Digit Occupation	3-Digit Occupatio
Predicted Firm Union Density	0.00462**	-0.00586***	-0.00443***
	(0.00207)	(0.00153)	(0.00163)
Predicted Firm Union Density * HHI		0.0225***	0.0160***
		(0.00242)	(0.00163)
Union Density * Above Firm-Occ Median	0.00633***	0.00669***	0.00687***
	(4.02e-05)	(5.33e-05)	(5.47e-05)
Union Density * HHI * Above Firm-Occ Median		-0.00500***	-0.00510***
		(0.000304)	(0.000226)
Observations	16,181,780	16,181,780	16,181,780
R-squared	0.719	0.719	0.719

Source: Authors' estimates corresponding with Equations 8 and 9 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1 level.

Table A15: Effect of Union Density on Log Annual Earnings by Labor Market Concentration and White Collar Occupation Status, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.00811***	0.000225	8.26e-05
	(0.00117)	(0.000673)	(0.000724)
Predicted Firm Union Density * HHI		0.0147***	0.0128***
		(0.00198)	(0.00161)
Union Density * White Collar	0.00206***	0.00173***	0.00239***
	(0.000326)	(0.000433)	(0.000411)
Union Density * HHI * White Collar		-0.000145	-0.000804**
		(0.000478)	(0.000363)
Observations	16,181,780	16,181,780	16,181,780
R-squared	0.581	0.581	0.581

Source: Authors' estimates corresponding with Equations 8 and 9 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A16: Effect of Union Density on Log Annual Earnings by Labor Market Concentration and Gender, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.0168***	0.00476***	0.00593***
	(0.00146)	(0.00131)	(0.00131)
Predicted Firm Union Density * HHI		0.0126***	0.0109***
		(0.00228)	(0.00161)
Union Density * Female	-0.00261***	-0.00283***	-0.00283***
	(3.79e-05)	(4.76e-05)	(4.77e-05)
Union Density * HHI * Female		0.00292***	0.00205***
		(0.000281)	(0.000201)
Observations	16,181,780	16,181,780	16,181,780
R-squared	0.595	0.595	0.595
Occupation x Industry FE	Yes	Yes	Yes
Education FE	Yes	Yes	Yes
Age Group FE	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes

Source: Authors' estimates corresponding with Equations 8 and 9 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A17: The Effect of Exposure to Chinese Imports on Labor Concentration

VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Exposure to Chinese Imports per Worker (1,000s NOK)	-2.00e-06***	-1.64e-06***
	(1.85e-07)	(2.31e-07)
SD of HHI (full sample):	0.0538	0.0611
SD effect of 1 million NOK	-0.0372	-0.0268
Observations	145,032	145,032
R-squared	0.826	0.811
Firm Fixed Effects	Yes	Yes

Source: Authors' estimates as described in the text.
Notes: Standard errors are in parentheses and are clustered at the firm level.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A18: The Effect of Predicted Labor Concentration on Union Premium

VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Union Density	-0.000748*	-0.00158***
	(0.000403)	(0.000614)
Union Density x Predicted HHI	0.0448***	0.0546***
	(0.0115)	(0.0140)
Predicted HHI	-2.730***	-3.328***
	(0.667)	(0.813)
Observations	10,351,840	10,351,840
R-squared	0.557	0.557

Source: Authors' estimates as described in the text.
Notes: Standard errors are in parentheses and are clustered at the firm level. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A19: Employment Effects of Lagged Union Density by Concentration - Private Sector Only

VARIABLES	(1) Pr(Hours>30)	(2) Pr(Hours>30)	(3) Workers	(4) Workers
Lagged Predicted Union Density	-0.000423	-0.0281***	0.0884	-0.219
	(0.00158)	(0.00263)	(1.519)	(1.770)
Lagged Predicted Union Density * HHI		0.0733***		-2.586
		(0.00615)		(2.858)
Constant	0.831***	2.319***	81.01	104.4
	(0.0916)	(0.150)	(86.68)	(99.84)
Observations	10,031,320	10,031,320	193,250	193,250
R-squared	0.309	0.310	0.948	0.948
Occupation x Industry FE	Yes	Yes		
Education FE	Yes	Yes		
Firm FE	Yes	Yes	Yes	Yes

Source: Authors' estimates as described in the text.
Notes: Standard errors are in parentheses and are clustered at the firm level.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

B Extension: The China Shock

In Section 5, we presented new evidence on the impact of unionization as a function of labor market concentration. We did this by exploiting an exogenous shift in unionization at the firm and interacting this with existing measures of labor market concentration. An alternative approach would be to utilize exogenous shifts in labor market concentration and interact this with existing levels of union density.

In this section, we exploit the influx of imports from China to Norway in the early 2000s as an exogenous shifter of firm labor market concentration. We then use this to measure the effects of unionization on earnings when there are plausibly-exogenous changes to the level of labor market concentration. This exercise provides a complementary approach to our main empirical strategy and helps establish the robustness of our results to relying either on exogenous variation in unionization or exogenous variation in labor market concentration.

In terms of our conceptual model in Section 2.3, this complementary approach offers another advantage as well. Specifically, by exogenously shocking the labor market concentration across firms with different levels of union density, we can ask if there is a level of union density at which the wage mark-down of an increase in monopsony power ($G_f(M)$) can be completely offset by the positive wage effect of $\Phi_f \Pi_f / L_f$. That is, for what level of U_f would $G_f(M) = U_f(\Phi_f \Pi_f / L_f)$? Addressing this question not only helps us better understand the dynamics of power between employers and employees, but it also provides crucial information on the value of incentivizing and disincentivizing union membership through public policy as a means to combat market failures induced by imperfect competition.

B.1 Data and Method

We rely on a shift-share measure of import exposure where we allocate the shock to local labor markets based on baseline firm labor shares (Autor et al. (2016)). The assumption underlying this approach is that the shift induced by the Chinese imports cannot be correlated with any bias in the initial shares across our units, an assumption we believe is plausible. While we note a recent influx of studies in the econometrics literature that explore the limitations of shift-share instruments in this and other applications (e.g., Borusyak et al. (2022); Goldsmith-Pinkham et al. (2020); Adao et al. (2019); Jaeger et al. (2018)), we believe that this method provides a valuable alternative approach to our main empirical strategy. In addition, the exclusion restriction specifies that import competition should only affect the interactive effect of unions in concentrated markets through its effect on labor market concentration. Appendix Table A8 indicates that holding predicted unionization constant at a base level leads to similar estimates, meaning that import competition does not appear to change the interactive effect between union density and concentration through the union density channel. However, we emphasize that this should be considered an extension of our preferred specification, rather than a substitute to our preferred specification, that allows us to push our analysis further and explore if there is a level of U_f for which $G_f(M) = U_f(\Phi_f \Pi_f / L_f)$.

In terms of data, we follow Balsvik et al. (2015) and exploit information on the amount of imports into Norway coming from China allocated across product types to specific industry codes. This enables us to capture the size of the import shock to particular national sectors. We make use of the granularity of the Norwegian register data and use firm-specific baseline employment in affected industries to allocate the size of the shock to local labor markets.

In terms of estimation method, we measure local labor market exposure to Chinese imports, which is a per-worker measure of total firm-specific exposures in the local labor market. We define exposure at local labor market l at time t related to industry i and firm f as:

$$Exposure_{lt} = \frac{1}{L_0^l} \sum_f \frac{L_0^{fil}}{L_0^i} \Delta M_{it}, \tag{11}$$

where ΔM represents the change in total imports from China related to industry i from base year 2001 to the current year (the "shift"). The ratio $\frac{L_0^{fil}}{L_0^i}$ is the share of employment in the base year in industry i working at firm f in local labor market l (the "share"). We sum these firm-specific exposures over all firms in the local labor market and normalize the shock by the total size of the local labor market at baseline.

A firm that is not directly exposed to import competition may nonetheless be influenced at the local level by shocks to import competition through a reshuffling of labor demand across industries and occupations in the local labor market. This is, in fact, the margin at which Balsvik et al. (2015) find that Norwegian firms respond to Chinese import competition: through changes to the employment level rather than wages. We, therefore, estimate a moving value of firm-specific labor market concentration as a function of employment-weighted firm exposure to Chinese imports and include firm and year fixed effects:

$$HHI_{ft} = \alpha_0 + \alpha_1 Exposure_{lt} + \tau_t + \phi_f + \nu_{ft}, \tag{12}$$

where all variables are defined as above.

We use predicted HHI from this equation in a second-stage equation of individual-level log earnings:

$$\begin{aligned} Log(Earnings)_{iocft} =& \alpha_0 + \alpha_1 UD_{ft} + \alpha_2 UD_{ft} * \widehat{HHI}_{ft} + \alpha_3 \widehat{HHI}_{ft} \\ &+ \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \end{aligned} \tag{13}$$

In Equation 13, we use a raw value of the calculated union density UD_{ft} and interact this with predicted labor market concentration based on exogenous shifts in labor market concentration driven by the influx of imports from China to Norway in the early 2000s. Because the import data are limited in their time coverage, we measure these effects from our baseline in 2001 to 2007. In an alternative approach, we predict the probability of unionization using our various fixed effects for occupation by industry, year, age groups, and education cells and then take the firm-level mean of this predicted value. This gives us a composition-constant predicted union density for the firm that is robust to any composition changes at the firm arising from import competition. These results are provided in Appendix Table A8 and lead to the same conclusions as our main approach.[31]

B.2 Results

Table A17 shows results from estimating the impact of exposure to Chinese imports on the labor market concentration of Norwegian firms. In column (1), we show results for our preferred measure of 20 clusters. For robustness, in column (2), we show the effects on HHI calculated for 40 skill clusters.

The results in Table A17 suggest that exposure to Chinese exports has a small but highly statistically significant impact on the labor market concentration experienced by firms. Specifically, an increase in exposure to Chinese imports per worker of 100,000 NOK (approximately 12,000 USD) reduces the HHI of the firm by approximately 0.34 percent of a standard deviation. For firms at the top of the exposure distribution, with an exposure of approximately 2 million NOK per worker, the predicted effects would be nearly seven percent of a standard deviation. The F-statistics associated with the regressions underlying the results are 116 and 50, respectively. These statistics are significantly greater than the conventional rule-of-thumb values.

In Table A18, we use a raw value of calculated firm-level union density UD_{ft} and interact this with the predicted labor market concentration based on the model estimated in Table A17. Looking across the table, the results suggest a strong negative association between labor market concentration and wages. Specifically, a standard deviation change in labor market concentration is associated with a wage reduction of 15-20 percent. This is consistent with the notion that firms can leverage their labor market power to suppress wages below the competitive equilibrium.

The results in Table A18 also demonstrate that the negative impact of labor market concentration is considerably smaller in highly unionized firms. A one percentage point increase in union density increases wages by approximately 4.5 percent in the most concentrated labor markets. These estimates across definitions of HHI are remarkably consistent: according to the estimates, the negative earnings effect of labor market concentration is effectively eliminated upon reaching a union density of approximately 63 percent at the firm. This set of results highlights that unions may serve to limit the wage-setting power of monopsonistic employers and that unions may play an important role in limiting market failures generated by monopsonistic power. This result is consistent with the notion that the greater the market imperfection, the greater the amount of firm rent that unions can extract. The findings from this exercise thus provide a complementary view to our main results and help provide a better understanding of the dynamic interplay between unions and monopsonistic employers in the economy.

Given our prediction that a union density of 63 percent is sufficient to reverse the negative effects of labor market concentration at the firm, we perform a back-of-the-envelope calculation to identify what the total subsidy cost would be to incentivize workers at every firm in the Norwegian economy

[31] We also estimate the effect of import exposure on the likelihood of being in a union. When controlling for individual fixed effects, exposure to Chinese imports does not affect the probability that an individual worker is a member of a union. While there may be compositional changes that affect firm union density, the results in Appendix Table A8 indicate they do not affect our conclusions.

to reach this threshold. At the end of our sample period, approximately 37% of all workers were at firms with a predicted union density below this 63% threshold, representing 30% of all firms. On average, firms below the tipping point have predicted densities approximately 3 percentage points away from 63%. Generating a 3 percentage point change in unionization, according to our subsidy effect estimates, would require an increase in the base tax subsidy of approximately 240 NOK, or raising the deduction by approximately 889 NOK. This would induce approximately 20,560 new workers to join a union at a base cost of 4.93 million NOK. Holding constant the union membership status of those in firms already above 63%, a universal tax subsidy increase of 240 crowns per member would also result in additional payments to approximately 708,300 full-time workers totaling 170 million NOK, for a total new base subsidy value of approximately 175 million NOK (approximately $22.7 million).[32] Given the size of the workforce in our sample (approximately 1.85 million workers at the end of the sample), this amounts to a transfer of approximately 95 NOK per worker per year for the base subsidy. Furthermore, at the average labor market concentration in our sample of firms that are below 63% predicted union density, a 3 percentage point increase in unionization would also induce these firms to increase the share of workers above 30 hours by 2.5 percentage points on average. The increase in employment would also lead to an increase in the taxable income of workers, which may be used to at least partially offset the cost to the government of the tax deduction.[33]

[32]The actual size of the subsidy will be larger than the base subsidy due to Norway's progressive income tax schedule, so base subsidy costs are a lower bound.

[33]Without a full analysis of the incidence of corporate taxation on labor, the effect of lower monopsony rents on corporate profits, and changes in marginal tax rates with rising labor earnings, we cannot assess the full budgetary impacts of the tax deduction and therefore cannot infer the size of the fiscal benefits relative to the transfer costs. However, the increase in intensive-margin employment and total earnings in these firms leads to an increase in tax revenue that does appear to offset a sizable portion of the transfer cost.