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

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Abstract

This paper brings together the literatures on employer power and employee power by studying the effect of unions on earnings, employment, and inequality across differently concentrated markets. Exploiting national government-induced changes to union due subsidies as exogenous shocks to union density, 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.

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1 Introduction

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 evidence of 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). Employer power generates an upward-sloping labor supply curve to the firm, allowing employers to mark down worker wages 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 the level that would prevail absent such bargaining.[2] Despite centuries of theoretical discussions on the interplay of these forces in the labor market —from Smith (1776) and Robinson (1933) to Freeman and Medoff (1984) —no empirical research has been able to causally study these interactions. This is primarily due to a lack of detailed data coupled with the difficulty of obtaining exogenous variation in unionization across differently concentrated markets.

This paper brings together and bridges the literatures on employer power and employee power by empirically examining the effect of unions on earnings, employment, and inequality across differently concentrated labor markets. We are motivated to study this topic because a union's ability to influence wages may act as a countervailing force to the monopsony conditions that characterize a wide – and growing – range of labor markets. At the same time, the ability of unions to counteract 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's ability to extract these rents will be minimal (due to the lack of viable employee outside options that can be used as leverage).[3] In a more competitive market, on the other hand, there will be minimal rents, but the unions' ability to secure those rents 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 evidence of such wage-setting behavior among

[1] E.g., Schubert et al. (2020); Prager and Schmitt (2021); Caldwell and Danieli (2018).

[2] E.g., Fortin et al. (2022); Lee and Mas (2012).

[3] The idea of wages being determined, in part, by the sharing of quasi-rents has a long history, with notable examples including the works of, for example, Van Reenen (1996) and Abowd and Lemieux (1993).

employers, suggesting that the macroeconomic consequences of firm power are substantial (e.g., Schubert et al. (2020); Prager and Schmitt (2021); Stansbury and Summers (2020); Dodini et al. (2020)). In such markets, unions may be able to correct the market failure generated by imperfect competition by counter-balancing the monopsony power of employers, pushing the economy closer to the competitive equilibrium. This would generate higher worker wages and employment levels, leading to 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 such markets, union-induced changes in wage levels would cause a market failure in the absence of pure union productivity effects.[4]

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. 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 ability to extract rent is decreasing as monopsony power rises. 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 detailed information on union membership, union dues, and each worker's occupation. We then leverage national government-induced changes in the tax deduction for union dues in Norway, which led to a quadrupling of the maximum deduction 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 a tax deduction cap (Barth et al., 2020). This means that workers at firms whose union dues were high prior to the plausibly-exogenous national government changes were more intensely "treated" relative to those with lower baseline union dues. This distinction in exposure generates exogenous variation in the incentive to join a union for individual workers depending on the firm at which the worker was employed at and, therefore, different exogenous shifts in union densities across firms. By interacting

[4]Changes in wages from unionization are analogous to changes in minimum wages set by policy for low-wage workers. A lack of substantial negative employment effects (e.g., Cengiz et al. (2019; 2022)) and significant pass-through of price increases to consumers (e.g., in Harasztosi and Lindner (2019)) is suggestive of firm power either in the labor market or product market.

the exogenous shift in unionization with firm measures of local labor market concentration, which we measure via Herfindahl-Hirschman Indices, we can analyze the role of unions across markets with different degrees of labor market concentration.

The practical implementation of our empirical approach proceeds in two parts. In the first part, we model the decision-making process of individual workers and how they respond to changes in the cost of unionizing. We estimate this process as flexibly as possible, without, for example constraining individuals to remain at a particular firm since that could be part of the response. We use these results to obtain our group-level treatment intensity measure (predicted firm-level union density). In the second part, we use this intensity measure in our outcome regressions for workers within the firm. Our sequence of estimation is important because each membership decision is a function of the specific worker's own preferences and behaviors, and the firm exposure/intensity measure is thus determined by a series of individual responses to a change in incentives.

To ensure that our results are not an artifact of the particular modeling choice described above, we also present results from two alternative approaches. First, we take a "kitchen sink" approach and estimate a saturated fixed effects model that leverages only within-person, within-firm variation in individual union membership status in response to reductions in firm-specific changes in net membership costs. Similar to our main approach, firm-level union density is simply an aggregation of individual choices, and we model the effect of this density on earnings. Second, we abstract away from individual-level membership choices and instead employ a split-sample IV strategy. Here, we model firm-level union density changes in response to subsidy-induced reductions in the net union dues at the firm in one half of the sample and use the model to predict firm-level union density and the effects of union density on earnings in the other half of the sample. Each of these alternative approaches have certain advantages and disadvantages relative to our preferred specification, which we discuss in detail. While we prefer the main estimation approach because it does not restrict our identifying variation as much as the other approaches, we acknowledge that others may have a preference for one of these alternatives. Encouragingly, the results from our analysis do not depend on which of these three approaches we employ.

The core finding of our analysis across all three specifications is that high levels of unionization ameliorate the negative effects of labor market concentration on earnings. This suggests that unions play an important role in correcting market failures induced by imperfect competition. Consistent with monopsony theory, the wage effect is accompanied by positive intensive margin employment effects. Figure 1 illustrates this result, demonstrating that as predicted unionization increases in response to the government subsidies, 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 possibly having relatively lower bargaining power in these concentrated markets (Aghion et al., 1998; Yamaguchi, 2010). Unions are thus able to "level the playing field" in concentrated markets.

We present five novel 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 unionize. Specifically, increasing the annual union subsidy by NOK 1,000 leads to an increase in a worker's probability of unionizing by 9-15 percentage points and is robust across approaches. We also validate this finding externally through a survey of 5,200 Norwegian workers in which we back out the price sensitivity of union membership through hypothetical scenario analyses. The effect is considerably larger in markets that experience monopsonistic competition. 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, we find that a 1 percentage point exogenous increase in firm 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. 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. The gradient over concentration 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 identify different effects across these two sources of market power highlights that 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. In non-concentrated markets, they reduce employment on this margin. These results are 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. 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. In other words, unions have an inequality-enhancing effect within narrow sub-sectors in competitive markets, while this is not the case in concentrated markets. Thus, the role of unions in shaping aggregate income inequality is fundamentally tied to the distribution of competitiveness of the local labor markets in the country, a finding that has important implications for how we view the interplay between labor unions and overall income inequality.

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 in 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 directly measured labor market concentration and then examined how concentration affects wages and employment (e.g., Schubert et al. (2020); Azar et al. (2020b); Qiu and Sojourner (2019); Rinz (2018); Prager and Schmitt (2021); Azar et al. (2020a); Benmelech et al. (2022); Marinescu et al. (2021); Hershbein et al. (2018); Bassanini et al. (2022); Dodini et al. (2020)). On average, these studies show that labor market concentration reduces worker wages and has negative effects on workers' careers.

We advance the labor market concentration literature by demonstrating that unionization rates, as well as union wage premiums, are larger in concentrated markets. This suggests that unions may successfully act as a countervailing force to employer power. In addition, we document a positive marginal union employment effect in concentrated markets. 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, 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 recent decades.

Second, there is a small but impressive literature that causally identifies the union wage

[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.

effect through quasi-experimental research designs, using anything from regression discontinuity designs related to close union elections (e.g., DiNardo and Lee (2004); Lee and Mas (2012); Frandsen (2021); Sojourner et al. (2015)), propensity score matching techniques that directly control for endogenous selection of workers into unions (e.g., Card and De La Rica (2006); Bryson (2002)), instrumental variable methods for individual unionization based on Right-to-Work laws in the United States (Fortin et al. (2022)), and changes in national union due subsidies as a measure of unionization probability (Barth et al. (2020)).[6]

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 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 impacts not only wages but also other types of non-pecuniary benefits and social goods.

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. Closed-shop union agreements are not allowed in the country.

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 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

[6] Farber et al. (2021) has also provided very interesting survey evidence on this question by developing a new and novel historic database on union membership over the last 90 years.

(AFL-CIO) coordinates and supports union efforts across more than 50 individual unions spanning a range of professions.[7] 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.[8]

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).[9] In the local negotiations, unions and employers discuss not only 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*

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

[8]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.

[9]This is also consistent with recent work from Norway suggesting that collective bargaining agreements in themselves have no impact on wages once union density is accounted for (Blandhol et al., 2020).

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 and based on individual work performance. 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

An interesting 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 government-induced 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 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 stated above, the bargaining process in Norway can be viewed as a two-step process. In the first step, industry-wide collective bargaining agreements that 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. 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.[10] As shown in Abowd and Lemieux (1993), the solution to the bargaining problem is

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

where Φ_f is equal to ϕ_f in the strongly efficient bargaining model and to a positive fraction of ϕ_f in the weakly efficient bargaining model.[11] Written in this form, the union negotiated wage is equal to the market wage plus a fraction 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

[10] 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.

[11] In the strongly efficient bargaining model, the union sets both wages and employment. In the weakly efficient bargaining model, the union sets only wages. The derivation holds irrespective of whether one operates in the strongly efficient or the weakly efficient bargaining framework, and it is therefore not required to take a stand on which of these are true for deriving at the equations presented in this section.

about how unions 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 less 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 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)).

The above expression gives rise to two predictions crucial to our empirical analysis. First, the higher the firm's 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 (and individual's) relative bargaining strength is decreasing with the degree of monopsony power (Aghion et al., 1998; Yamaguchi, 2010; Tschopp, 2017). This is because unions can serve multiple functions that augment individual leverage, which increases with viable outside options. Unions can coordinate information on options within (e.g. across departments) and across employers, which lowers search and signal costs and induces greater individual investments within the firm, thus increasing the size of the total pie (Aghion et al., 1998). This is even more pronounced in Norway, where unions can represent workers in individual contract negotiations. Therefore, 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 ($HHI \rightarrow 1$) and one with a low level of labor market concentration ($HHI \rightarrow 0$). 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 (because individual leverage in negotiation is magnified by the union) 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. We leverage changes in tax subsidies for union members in Norway, which led to significant changes in the price of union membership for some workers, to shift U_f.

Before turning to the empirical investigation, it is important to mention 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, then $(\Phi_{if} \Pi_{if})$ 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 A6. 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 a limited set of 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.[12]

3.2 Union Dues and Tax Subsidies

To obtain exogenous variation in union density across firms, we leverage national 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 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 government-induced subsidy 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.

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. 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

[12]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.

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 of 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 by altering the occupations they decide to employ workers in or by changing the union dues directly. 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 Kroner. 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 feedback loop between the change in the law in any particular year and the imputed union dues.

With this 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 from 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. Our measure of subsidy value, therefore, captures changes that only are coming through legislative channels.

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

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

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 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 in the minimization function. 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 ($ND_{ft} = \overline{D_f^0} - Subsidy_{ft}$). We include this as an additional control in our baseline regressions. This is important because 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 unobserved firm-specific productivity difference related to their baseline occupational mix.[13] 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.

3.3 Defining Concentration

To obtain a proxy for labor market concentration, we build on Dodini et al. (2020) and take a skill requirement-based approach. To do so, we use data from the US Department of Labor's Occupational Information Network (O*NET) survey to incorporate information on occupational skill requirements into the Norwegian registers.

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

[13]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.

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.[14]

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.[15]

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 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 A9–A15). 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. (2020b;a); Benmelech et al. (2022); Marinescu et al. (2021); Qiu and Sojourner (2019); Rinz (2018); Hershbein et al. (2018))).[16] We follow this convention while at the same time considering concentration to be of independent interest.

[14]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.

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

[16]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) finds elasticities of 0.25 for the extensive margin and 0.33 for the intensive margin after accounting for frictions.

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 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 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 union that came through government-induced adjustments to the Norwegian tax code between 2001 and 2015. These changes significantly reduced the monetary cost of joining a union for workers whose union due subsidies were previously bounded by a tax deduction cap (Barth et al., 2020). Specifically, workers at firms whose union dues were high prior to the reform are more intensely "treated" relative to those with lower baseline union dues. This means that some firms and workers were more intensely treated by the reforms over time compared to others for reasons unrelated to changes in firm-specific characteristics. Intensity of treatment is identified by the combination of the base year occupation/industry mix at the firm and changes in the Norwegian tax code. This distinction generates exogenous variation in a worker's incentive to join a union depending on the firm at which the worker is employed and, therefore, different exogenous shifts in union densities across firms.

The empirical implementation of our estimation approach proceeds in two parts. In the first part, we model the decision-making process of individual workers and how they respond

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

to changes in the cost of unionizing. We want to estimate this process as flexibly as possible, without for example constraining individuals to remain at a particular firm since that could be part of the response (i.e., without including firm fixed effects). This is particularly important if a person's first (or only) year at a firm disproportionately contributes to within-person changes in union membership status. We use these results to obtain our group-level treatment intensity measure (predicted firm-level union density), which is then used in our outcome regressions. In the second part, we use these treatment intensity measures to examine how these plausibly-exogenous shifts in union density at the firm level impact wages.

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:

$$\begin{aligned} Union_{iocft} =& \beta_0 + \beta_1 Subsidy_{ft} + \beta_2 ND_{ft} + \beta_3 \overline{HHI}_f + \beta_4 \overline{HHI}_f * Subsidy_{ft} \\ &+ \beta_5 \overline{HHI}_f * ND_{ft} + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \varepsilon_{iocft}, \end{aligned} \tag{6}$$

where $Union$ is an indicator variable taking the value of one if the worker is a member of a union. $\overline{HHI}_f$ represents the firm-wide average HHI index for firm f fixed at the firm's first year in the data. We estimate this regression as individual workers may perceive differential gains to unionization as a function of the employer'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 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.[19]

The key result of this exercise is that firms with larger increases in subsidies (treatment

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

[19]When estimating our models for the average response to the subsidize, we omit HHI and its interactions from the model. We do the same when estimating earnings outcomes, as in Equation 7 below.

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 that 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:

$$\begin{aligned} Log(Earnings)_{iocft} =& \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \alpha_2 \widehat{UD}_{ft} * \overline{HHI}_f \\ &+ \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \varepsilon_{iocft} \end{aligned} \tag{7}$$

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 (i.e. conditioning on worker-firm match), as well as the difference in the marginal effects of unionization after holding constant time-invariant characteristics of the firm and individual workers. The interaction between $\widehat{UD}_{ft}$ and HHI allows the effects of union density to differentially affect earnings in concentrated markets. We take an alternative approach to modeling the union membership choice at the individual level as well as modeling firm-level union density in Section 4.1 below. These approaches point to very similar conclusions as our main approach.

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.

In terms of interpreting the interaction of predicted union density and HHI, one might be concerned that there is an underlying correlation between HHI and base dues that are driving our results rather than the actual effect of union density across the HHI distribution. We note here that we fix both the imputed union dues and HHI at baseline. The change in the subsidy is therefore coming exclusively from the interaction between the policy and these base dues. Base dues are determined by the industry-occupation mix of the workers at the firm in the base year, thus also fixed prior to the policy. If some underlying correlation between HHI and base dues were driving our results, and thus exposure intensity, it would have to come through occupation-industry compositional differences at baseline exclusively at high HHI firms. However, the correlation between baseline HHI and base imputed dues in the sample is particularly weak (approximately 0.1), meaning there is little reason to think such a dynamic is present. This is, perhaps, unsurprising since concentration is a measure concerning the spatial distribution of employment, whereas baseline dues are about within-firm occupation distributions. On a more fundamental level, it is possible that high-HHI firms have a higher union density at baseline. However, our analysis leverages the same marginal shifts in union density (1 percentage point) across the HHI distribution in Equation 7, ensuring that pre-policy-change differences are not contributing to the effects.

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 in our main models (Rubin, 1981). However, given our sample size, the standard errors do not differ in any meaningful way from our clustered standard errors. To avoid imposing unrealistic computational costs, we report our clustered standard errors throughout the paper, which, as noted in the literature on clustering, are possibly conservative already (Abadie et al., 2023).

Conditional on the composition of workers at the firm and output prices, 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 ($\overline{HHI}_f^P$):

$$Log(Earnings)_{iocft} = \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \alpha_2 \widehat{UD}_{ft} * \overline{HHI}_f + \alpha_3 \widehat{UD}_{ft} * \overline{HHI}_f^P + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \quad (8)$$

where α_1 captures the marginal effect of union density on earnings in a firm in which both industry revenue concentration (our proxy for product market power) and labor market concentration 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.[20]

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 allows 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, we must assume that firms and workers with high expected changes in subsidies (and therefore large reductions in 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 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 trends in union density at the firm (Panel A) or log earnings (Panel B). Specifically, individuals at firms with large reductions in net dues have higher earnings growth *only* during the years in which the maximum deduction changed drastically and the subsidy changes became efficacious (2003-

[20]We focus on national industry revenue shares for two main reasons. First, the Norwegian tax data from which we extract firm operating revenues do not contain information on differentiated products, so we cannot measure disaggregated product competition. Second, the data cannot capture the spatial distribution of revenues. Specifically, the location data we have are limited to the firm's tax headquarters/corporate offices, so all revenues would be allocated to the firm's corporate headquarters rather than where sales are actually occurring. This leads to a mechanical correlation between local firm employment shares and local firm revenue shares that may not reflect true underlying price dynamics.

2010), while the trends were parallel in 2001-2002 and after 2010 when the deductions were stable. This aligns extremely well with the trends in union density at the firm, which shift disproportionately for firms that had large relative reductions in net dues only during the 2003-2010 period. While additional pre-reform years would be helpful, we emphasize that our approach is not a difference-in-differences approach, and the parallel trends graphs are designed to explore the comparability of the two groups in times in which the subsidies were not changing. This is not only the pre-period (2001-2002), but also the post-period after 2010.

Given the results from the above exercise, we conclude that the differences in earnings growth we see in our final analysis across subsidy groups do not appear to be a function of differences in the direction or magnitude of demand growth across industries or occupations. Rather, they appear to be 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 experienced a large decrease in their net dues via increases in subsidies. We view this as unlikely.

4.1 Estimation Intuition and Alternative Approaches

Prior to discussing our results, we would like to provide additional intuition related to the exposure-measure estimation approach that we pursue.

In the first part, we model the decision-making process of individual workers and how they respond to changes in the cost of unionizing. We want to estimate this process as flexibly as possible, without for example constraining individuals to remain at a particular firm since that could be part of the response. In the second part, we use these results to obtain our group-level treatment intensity measure and estimate the effect of firm-level union density on outcomes. In the outcome regressions, it makes sense to include firm fixed effects because we want to isolate how the union density changes at the firm are affecting outcomes for individual workers at that firm.

Our sequence of estimation is important because each membership decision is a function of the specific worker's own preferences and behaviors, and the firm exposure/intensity measure is thus determined by a series of individual responses to a change in incentives. Because different types of workers may respond differently to the subsidy, if we were to estimate a two-stage least squares model at the individual level in which union density is the outcome of the first stage, then we could have individual workers at the same firm in the same year with different predicted union densities. At a conceptual level, individuals are not choosing union density at the firm, they are choosing union membership. Modeling group outcomes as an individual choice introduces noise.

To ensure that our results are not an artifact of the particular modeling choice made above, we will also present results from two alternative approaches. This is valuable not only because different researchers may prefer to conceptually model this decision process differently, but also because it helps ensure that our results are robust to different identifying variation.

First, we take a "kitchen sink" approach and estimate a saturated model of union membership in which we include additional fixed effects for individuals, firms, and industry-year cells in both steps of the estimation procedure. We use changes in the firm-level net union due as a shifter of individual union membership choices. This is effective because baseline dues are fixed for each firm, so changes in the net due within firms are purely functions of the change in the subsidy when we include the firm fixed effect. We aggregate predicted individual union membership to the firm-year level and then estimate union density effects on individual earnings. While this approach subsumes a large amount of the identifying variation and restricts workers' responses to particular worker-firm-industry cells, it ensures that the identifying variation is driven by within-worker changes in union membership in response to changes in the cost of joining the union *at that specific firm*. It also abstracts away from worker-firm match effects and holds constant firm worker composition.

The downside of this approach is that we remove a large number of workers from helping to predict the intensity variable and strip the model of identifying variation across industries. For example, if a worker persistently is a union member beginning in her first year at a firm, the entire worker-firm match is subsumed within the firm fixed effect even if the choice to join the union was influenced by it being relatively cheap to join the union during her first year at her new firm in comparison to her old firm. This is very important if the choice to join a union is disproportionately operating when workers change firms. The upside to this approach, however, is that this assuages concerns about unobserved characteristics of firms, individuals, and industries driving our main results.

Second, we abstract away from the individual union membership decision and instead model firm-level union density itself. We perform a split-sample IV (SSIV) in which we randomly split the private-sector sample in half. We use one-half of the sample of firms to estimate how firm-level union density responds to changes in the firm-level net union due (i.e. the net cost of joining) in a model with firm and year fixed effects. We then use the estimated parameters from this model to construct fitted values of predicted union density for firms in the other half of the sample (Angrist and Krueger, 1995). We then estimate individual-level earnings responses to this firm-level union density and include individual and firm fixed effects so as to measure within-person changes in earnings as they are affected by changes in modeled firm-level outcomes.

The first step in this approach uses only firm-level data and does not capture heterogeneous individual contributions to union density at the firm, thus generating a noisier prediction of union density and possibly inducing attenuation bias. However, the benefit of this approach is that it abstracts away from any dynamics of the worker-firm match and individual union membership choices and avoids aggregation in the first step. SSIVs are generally biased towards zero rather than towards the OLS, helping us assuage standard endogeneity concerns (Angrist and Krueger, 1995). For inference, we adjust the standard errors following the advice in Inoue and Solon (2010) to account for first-stage uncertainty.

While we prefer the main estimation approach over these two alternatives, we acknowledge that this is a subjective decision and that others may have a preference for one of these alternatives. Encouragingly, other than the results being slightly noisier when using these two approaches, the story is unaffected by the choice of estimation approach, which we present below.

5 Results

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 Equation 6. 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) thus subsume possible noise introduced by "always union" members and isolate only changes in union status within person. 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 Tables A9–A15).

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 union status when exposed to different union dues and subsidies.[21]

[21]For robustness, we also estimate this equation while excluding the net union due from the equation. The results are in Appendix Table A3. Not accounting for union dues themselves results in substantially larger estimated subsidy effects on union membership. This suggests that accounting for the remaining costs of membership is important when we are not constraining identifying variation to come exclusively from within firms.

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 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.[22]

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.

Taken together, the results shown in Table 1 point to significant price sensitivity on the part of workers to become members of a union. To provide external validation of these effects, we conducted a representative survey of 5,200 workers in Norway and asked union members and non-members how they might respond to a hypothetical change in the net cost of joining a union. For those not currently in a union, approximately 35% of workers indicated that they would reconsider joining a union if the net cost of joining were reduced by as little as 500-1,000 NOK. For current union members, more than 50% would reconsider leaving their union if the net cost increased by 500-1,000 NOK. While this hypothetical scenario asks workers about reconsidering their choices (i.e. is missing real-world stakes) and is, therefore, likely an upper bound on their price sensitivity, the results are informative. Norwegian workers, themselves, indicate high levels of price sensitivity when considering the union membership decision.

5.2 Earnings Effects

Table 3 provides estimates on the effect of union density on individual log annual earnings using the changes in tax subsidies for union members as shifter of firm-level union density. Panel A uses our full sample while Panel B restricts the sample to only the private sector. In Panels C and D, we present the results from our saturated model and our split-sample IV approach, respectively. 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

[22] 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 only from within local labor markets.

function of labor market concentration, using our preferred specification of 20 skill clusters. In column (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 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 density 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 much 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 notions that $(\frac{\Phi_f \Pi_f}{L_f})_{concentrated} > (\frac{\Phi_f \Pi_f}{L_f})_{competitive}$ and that 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 3.

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 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 also 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 3 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 the most concentrated markets is 4.8 percent. Importantly, when we take different approaches to estimation in Panels C and D, the effects closely follow those of our main approach in Panels A and B.

5.3 Source of Rents

In Table 2, 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 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.[23] Appendix Figure A3 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

[23] Products can move faster than labor across space, meaning that output competition has a larger geographic footprint than labor competition. Because competition drives price-setting power and rents in both markets are related to this price-setting power, the difference in geographic aggregation makes the two measures more conceptually comparable as suggested by their similar scales in Table A1.

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 4, 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. We do this by estimating full interactions of our HHI, subsidy, and net dues variables in Equation 6 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 7. 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.[24]

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 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 satis-

[24]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).

fying 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.[25]

The results documented above 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.

With respect to effect heterogeneity across males and females, Figure A5 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.[26]

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 4 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

[25]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, slightly less responsive to the subsidies in competitive markets and more responsive to concentration. 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.

[26]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).

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 because employers in these 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 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 A6 for further discussion).

To address this question, we conduct two analyses. First, we estimate Equation 7 at the individual level using the probability of holding a full-time position (working at least 30 hours per week) as the dependent variable. Second, we estimate Equation 7 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 with regard to employment contracts and terminations.[27] 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 predicted union density measure.

[27]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.

The results from these two sets of regressions are shown in Table 5.[28] 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 one year to materialize.

The results displayed in Table 5 are consistent with the predictions of new monopsony 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.[29] Our results align well with Azar et al. (2019), 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). For example, "crowding" effects of unions may have negative spillovers on nonunion wages in the same industries (Neumark and Wachter, 1995; Fitzenberger et al., 2013); we extend this literature

[28]The results for the private sector only are in Appendix Table A18.

[29]Theoretically, this could also reflect an efficient bargaining outcome wherein unions with more power bargain for the formalization of full-time positions at the expense of part-time positions.

to include the firm level as well. 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 across the distribution of earnings in the firm using the same approach as in our heterogeneity analysis. We do this by estimating Equation 6 while including interactions between our tax subsidy measures and indicators for a worker being below the 10th percentile, between the 10th and the 50th percentile, between the 50th and 90th percentile, and above the 90th percentile in the earnings distribution at the firm. We then calculate overall predicted firm union density using the individual predicted probabilities from this model (mean pr(union)). Then, at the firm level, we regress each of our outcomes (the percentile ratios) on the interaction of predicted union density and concentration, including firm and year fixed effects. To explore (2), we measure inequality (percentile ratios) at the local labor market level and aggregate average concentration and average predicted union density to the LLM level, weighing by total employment at each firm in the LLM. We perform the same regression at the LLM level as we did at the firm level and include LLM and year fixed effects.

Table 6 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 4 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. This suggests that union threat effects dominate crowding effects only in more concentrated labor markets, even within firms.

In Panel B of Table 6, 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.[30] Union threat effects appear to dominate crowding effects when labor markets are subject to monopsonistic competition. 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.

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 changes to the level of labor market concentration. While this exercise relies on a stricter set of assumptions and should be considered more suggestive, 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.

[30] 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)).

6 Discussion

In this paper, we examine the effects of labor unions on the dynamics of worker earnings across differently concentrated markets. Existing empirical evidence has focused either on labor market power (e.g., Schubert et al. (2020); Prager and Schmitt (2021); Caldwell and Danieli (2018); Dodini et al. (2020)) or union power (e.g., Fortin et al. (2022); Lee and Mas (2012)), without considering the causal interaction of the two. While these two strands of literature provide extremely important insights into the workings of labor markets, our lack of knowledge of how these two forces interact—monopolistic unions and monopsonistic employers—severely limits our understanding of the dynamics of labor markets and how to design optimal labor market policies.

Exploiting government-induced national tax reforms to union due deductions as an exogenous shock to unionization, we demonstrate that the price elasticity of unionization has a steep gradient over labor market concentration. We then show that there is also a gradient in the union earnings premium and that the union wage premium loads heavily on highly concentrated markets. This result holds across three distinct identification approaches. First, in our preferred approach, we flexibly model the decision-making process of individual workers within and across firms and how they respond to changes in the cost of unionizing; we then use these results to obtain our group-level treatment intensity measure (predicted firm-level union density) to estimate the effect of union density on individual outcomes within the firm. Second, we take an alternative "kitchen sink" approach in which we estimate a saturated fixed effects model that leverages only within-person, within-firm variation in individual union membership status in response to reductions in firm-specific changes in net membership costs. Finally, we perform a split-sample IV strategy at the firm level in which we model firm-level union density changes in response to reductions in the net union dues at the firm in one half of the sample and use the model to predict firm-level union density and the effects of union density on earnings in the other half of the sample.

Our main finding is consistent with the notion that the greater the market imperfection, the greater 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 an important role for unions in limiting the market failures generated by employer power. Consistent with monopsony theory, the wage effect is accompanied by positive intensive margin employment effects in concentrated markets, while it is associated with a negative intensive margin employment effect in competitive markets.

Running horse races on product market power and labor market power on the union wage premium across differently-concentrated markets suggests that unions are effective in targeting and extracting both labor and product rents. 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. That we identify different effects across these two sources of market power highlights that they are substantively different components.

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, several papers (e.g. Card et al. (2004); Neumark and Wachter (1995); Fitzenberger et al. (2013)) 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. 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 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 net benefits to union density. 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 (consistent with crowding), but reduces local inequality when markets are concentrated (consistent with threat effects). 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 sub-sectors 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 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.

We are the first to provide causal estimates of the union density wage premium in an entire country and the first to bring together the modern literatures on monopsony power and unionization in labor markets. Examining the intersection of unionization and labor market concentration allows us to substantially advance our understanding of the role of unions and their impact on the dynamics of labor markets. 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*, appears 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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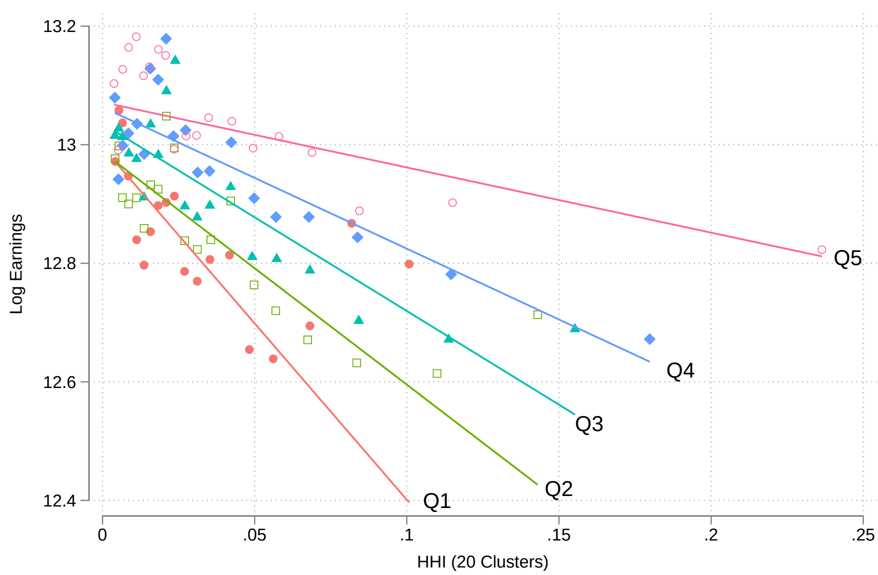
Figures

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 6 as described in the text.

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 Equation 6. 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: Union Density Effects 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

Notes: Authors' estimates as described in the text. Estimates correspond to Equation 8. 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: Effect of Union Density on Log Annual Earnings by Labor Market Concentration

	Panel A: Main Approach		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	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)
	Panel B: Main Approach, Private Sector Only		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	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)
	Panel C: Saturated Approach		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	40 Clusters
Predicted Firm Union Density	0.00702***	0.00625***	0.00619***
	(0.00209)	(0.00201)	(0.00199)
Predicted Firm Union Density * HHI		0.0203***	0.0184***
		(0.00507)	(0.00481)
	Panel D: Split-Sample IV		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	40 Clusters
Predicted Firm Union Density	0.00767*	0.00660*	0.00687
	(0.00404)	(0.00400)	(0.00439)
Predicted Firm Union Density * HHI		0.0383**	0.0305**
		(0.0188)	(0.0153)
Observations	16,181,780	16,181,780	16,181,780

Source: Authors' estimates as described in the text. Estimates for Panels A and B correspond to Equation 7.

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. All estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.

Panel C includes additional fixed effects for individuals, firms, and industry by year interactions when estimating the effect of within-firm changes in net dues on the propensity to join a union and on log earnings. Predicted individual union membership is aggregated to the firm-year level and then used in the estimates of log earnings.

The Panel D first stage is estimated on a random 50% sample of all private-sector firms. Data are structured at the firm level and estimated with firm and year fixed effects, measuring the change in union density in response to changes in net dues within firms over time. Fitted values are used to predict firm-level union density for all workers in the other half of the sample. Second-stage estimates for Panel D include individual fixed effects and measure within-person changes in earnings in response to changes in predicted firm-level union density. Standard errors in Panel D are adjusted following Inoue and Solon (2010).

Table 4: Heterogeneous Effects of Union Density on Log Annual Earnings

	Panel A: Above vs Below Firm-Occupation Median		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	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		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	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		
	(1)	(2)	(3)
VARIABLES	No HHI	20 Clusters	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 Equation 7 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 5: 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 6: 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

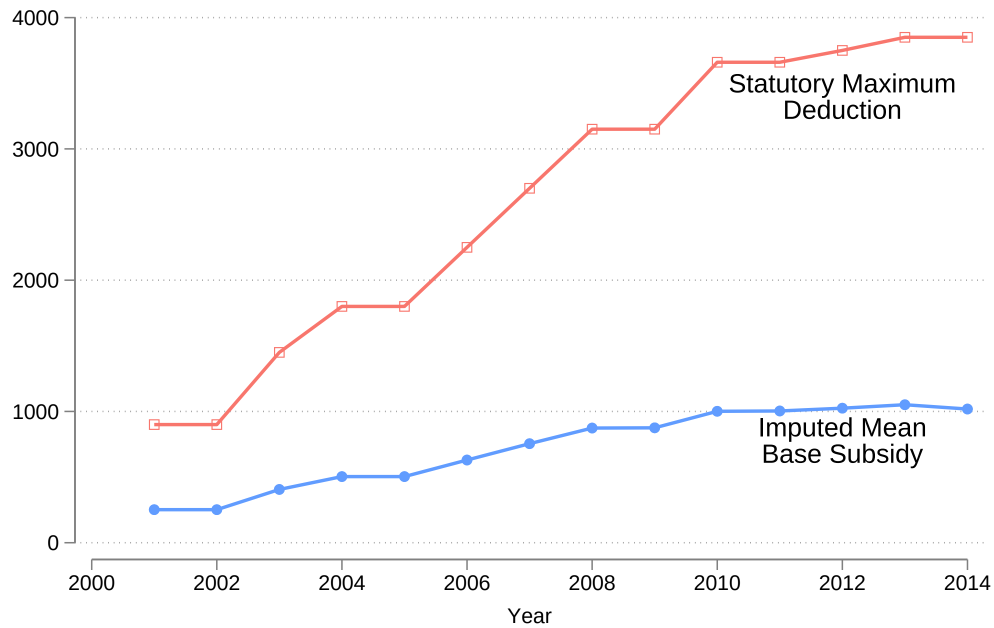
Source: Authors' estimates as described in the text.

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.

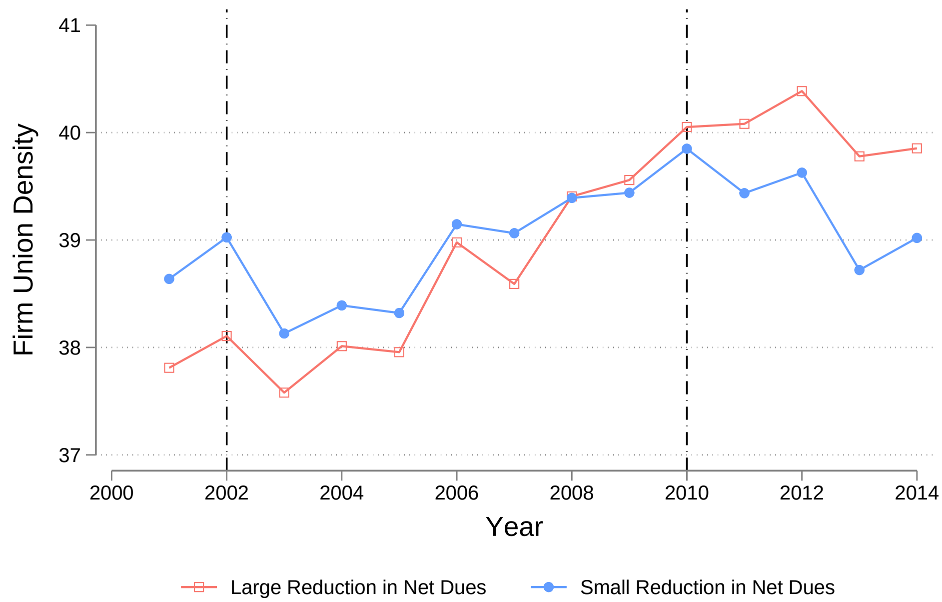
A Appendix: Additional Tables and Figures

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

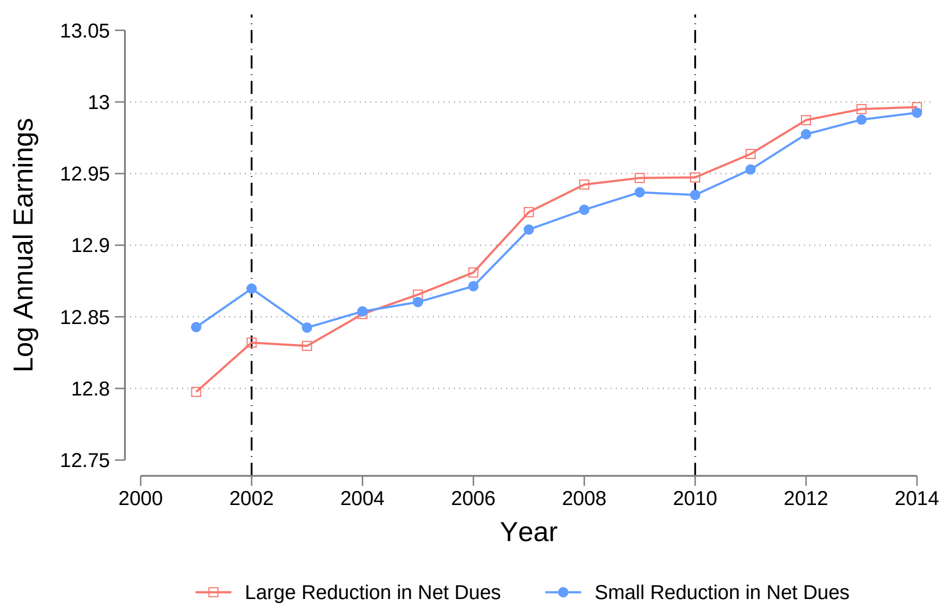


Source: Authors' calculations of Norwegian registry data.
Notes: Imputed base subsidies are calculated as the base tax rate times the lesser of imputed union dues at the occupation-by-industry cell or the statutory maximum deduction.

Figure A2: Union Density and Log Annual Earnings by Relative Subsidies
Panel A: Trends in Firm Union Density

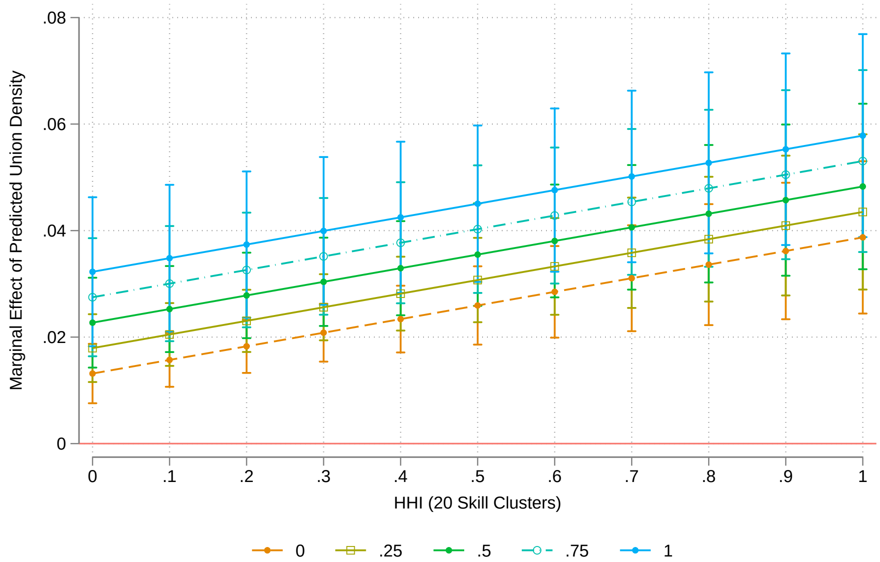


Panel B: Trends in Log Annual Earnings



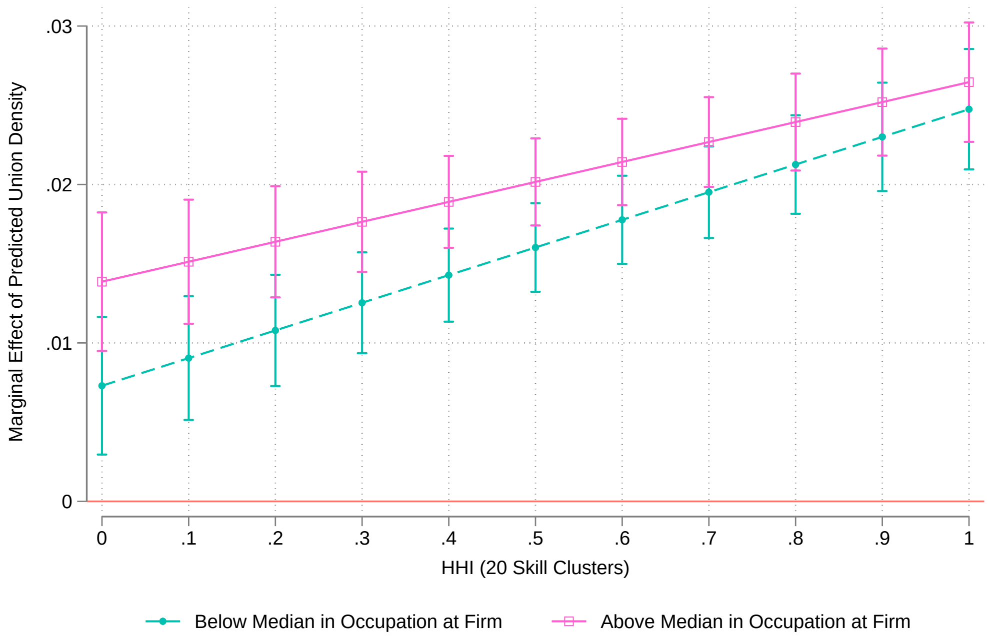
Source: Authors' calculations of Norwegian registry data.
Notes: Changes in net dues are calculated within firms from 2003 to 2010 during the period of the largest shifts in the maximum tax deduction. "Large reduction" firms are those whose 2003-2010 reduction in net dues was above the median compared to those below the median. Panel A accounts for firm fixed effects and is measured at the firm level. Panel B accounts for fixed effects for occupation by industry cells, age group, education, and firm; earnings are measured at the individual 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 2010-2011.

Figure A3: Marginal Effects of Union Density by Industry Revenue HHI



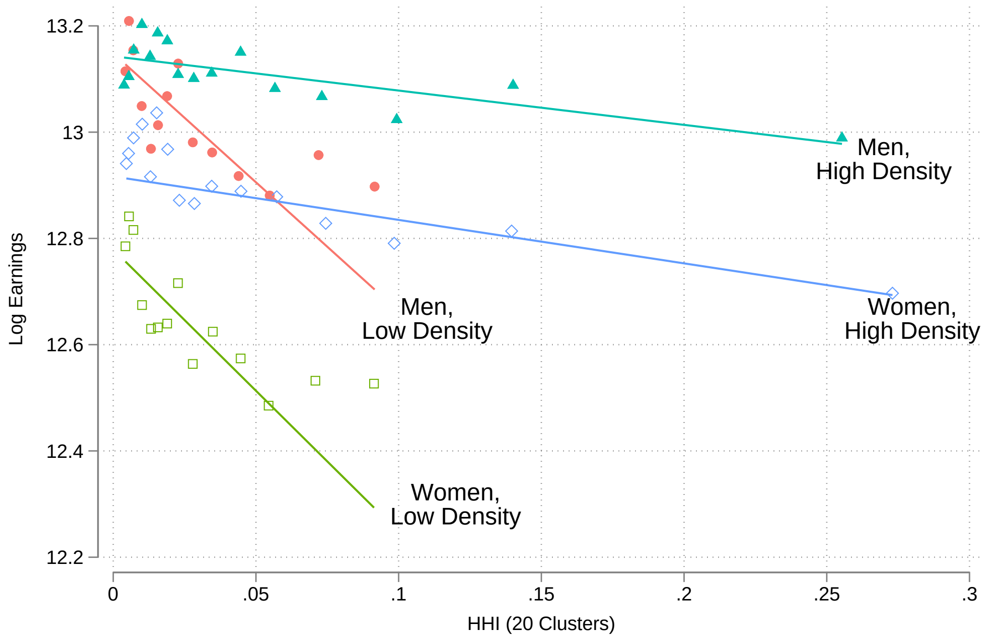
Source: Authors' calculations of Norwegian registry data as described in the text. Estimates reflect the marginal effects from Equation 8.
Notes: Standard errors are clustered at the firm level and calculated at each margin using the delta method.

Figure A4: 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 A5: Log Annual Earnings (NOK) and Labor Market HHI by Gender and Top vs Bottom Quintiles of Predicted Firm Union Density

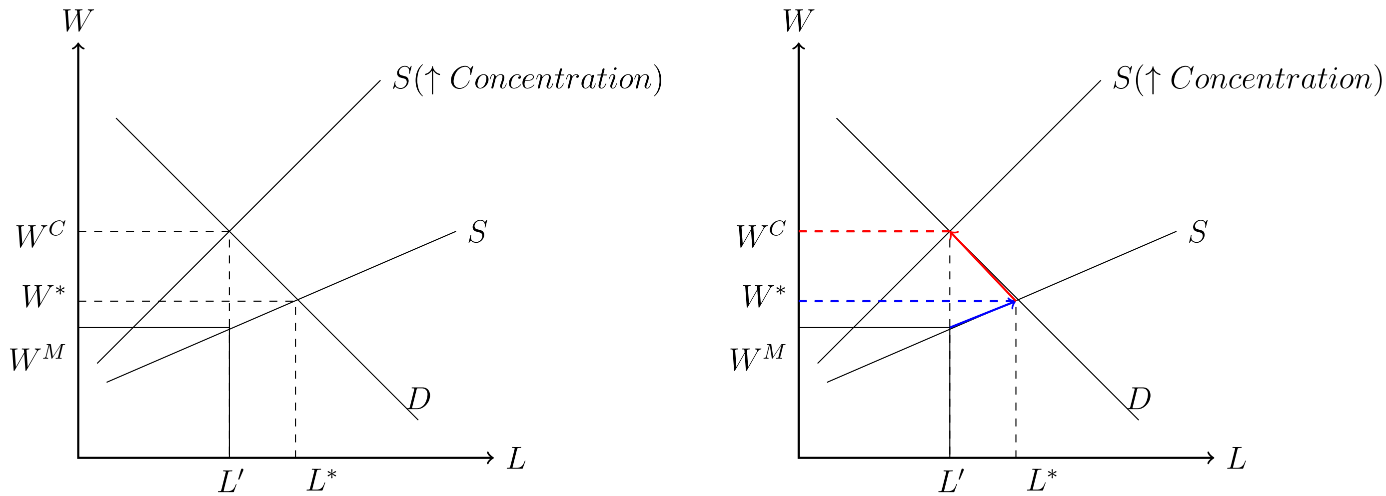


Source: Authors' calculations of Norwegian registry data.
Notes: Predicted union densities are based on average predicted unionization rates at each firm from Equation 6 as described in the text.

Figure A6: 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. (2019): 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 6 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 6 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 6 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 6 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 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, 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: 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 Equation 6.
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 A10: 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 Equation 7.
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 A11: 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 Equation 7.
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: 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 8.
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: Effect of Union Density on Log Annual Earnings by Labor Market Concentration and Firm-Occupation Median, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
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 Equation 7 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 A14: 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 Equation 7 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 A15: 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 Equation 7 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 A16: 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 A17: 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 A18: 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:

$$Exposure_{lt} = \frac{1}{L_0^l} \sum_f \frac{L_0^{fil}}{L_0^i} \Delta M_{it}, \tag{9}$$

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:

$$HHI_{ft} = \alpha_0 + \alpha_1 Exposure_{lt} + \tau_t + \phi_f + \nu_{ft}, \tag{10}$$

where all variables are defined as above.

We use predicted HHI from this equation in an equation of individual-level log earnings:

$$\begin{aligned} Log(Earnings)_{iocft} = &\alpha_0 + \alpha_1 UD_{ft} + \alpha_2 UD_{ft} * \widehat{HHI}_{ft} + \alpha_3 \widehat{HHI}_{ft} \\ &+ \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \end{aligned} \tag{11}$$

In Equation 11, 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 A16 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 A16 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 A17, 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 A16. 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 A17 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 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

[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.

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.