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 changes to union dues subsidies as exogenous shocks to firm-level union density in Norway, 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 clear evidence on both the role of employer power (arising from labor market concentration and/or labor market frictions) as well as 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 equips workers with monopolistic power over labor supply, enabling employees to raise wages above competitive levels.[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 provided a causal framework for studying 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. 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 may 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 may 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 employers, suggesting that the macroeconomic consequences of firm power are substantial (e.g., Schubert et al. (2020);

[1] E.g., Schubert et al. (2020); Prager and Schmitt (2021); Caldwell and Danieli (2024); Dodini et al. (2024).

[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).

Prager and Schmitt (2021); Stansbury and Summers (2020); Dodini et al. (2024)). In theory, 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 introduce new inefficiencies in the absence of pure union productivity effects.[4]

To empirically investigate the effect of unionization across differentially concentrated markets, we use high-quality longitudinal Norwegian employer-employee matched data—including detailed information on union membership, union dues, and each worker's occupation. We then leverage changes in the tax deduction for union dues set by the national government of Norway, which led to a quadrupling of the maximum deduction between 2002 and 2010. These changes significantly reduced the monetary cost of joining a union for workers whose union dues subsidies were previously bounded by the deduction cap (Barth et al., 2020). This means that workers at firms whose union dues were high prior to these exogenous 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 and, therefore, different exogenous shifts in union densities across firms. By interacting the exogenous shift in unionization with firm measures of local labor market concentration, which we measure via a Herfindahl-Hirschman Index, we can analyze the role of unions across markets with different degrees of labor market concentration. As an alternative approach to our HHI-based analysis, we also estimate and use separation elasticities in response to firms' wage policies based on Bassier et al. (2022)), where separation elasticities closer to zero (inelastic) are indicative of the firm's market power over incumbent workers (Online Appendix B). Our results align well across these two approaches.

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 joining a union. 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,

[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.

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. This approach also flexibly controls for industry-specific trends and shocks. 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 then use the model parameters to predict firm-level union density and measure the effects of union density on earnings at the individual level in the other half of the sample. Each of these alternative approaches has certain advantages and disadvantages relative to our preferred specification, which we discuss in detail. 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. 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 having lower bargaining leverage in these concentrated markets (Aghion et al., 1998; Yamaguchi, 2010). Unions are thus able to "level the playing field" in concentrated markets. Consistent with monopsony theory, the wage effect is accompanied by positive intensive margin employment effects 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 ($120 in 2015) leads to an increase in a worker's probability of unionizing by approximately 10 percentage points. 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 both in the administrative data and in the survey. 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 individual annual earnings of 1.9 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 1.3 percent in non-concentrated markets and by 2.8

percent in concentrated markets. The gradient over concentration is primarily driven by the private sector. This result supports the theory that the greater the labor market imperfection, the greater the amount of firm rent unions are able to extract —despite relatively weaker union bargaining leverage in these labor markets.

Third, we combine our labor market data with firm-level revenue data and explore the relative impact of product market concentration (proxied by 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. We believe that this is a novel finding with important policy implications, alluding to the fact that unions' ability to extract rent and reallocate this rent to their members depends not only on the extent of market imperfections but also on whether these imperfections are driven by labor market concentration or by product market concentration.

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 and firm type. We show that the modest union wage premiums that exist in competitive labor markets disproportionately accrue to high-skilled and white-collar workers. As market concentration increases, more and more of the additional rents that unions extract go to lower-ability and blue-collar workers. In other words, unions have an inequality-enhancing effect within firms and narrow subsectors in competitive markets while this is not the case in concentrated markets. However, because the marginal union member is more likely to be in a concentrated market and lower in the wage distribution, we find that local income inequality falls as local union membership rates increase, particularly in concentrated markets. Thus, the role of unions in shaping aggregate income inequality is fundamentally tied to the distribution of competitiveness of local labor markets, a finding that has important implications for how we view the interplay between labor unions and overall income inequality.[5]

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

[5]We also find that if workers have low levels of social capital as measured by occupational diversity in the firm, unions extract less of the available rents in concentrated markets, highlighting the important role of decreased leverage if unions are spread thin representing multiple occupations and interests.

of the role of unions in shaping 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.[6]

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. (2024)). 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 relationship 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 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 unionization based on Right-to-Work laws in the United States (Fortin et al. (2022)), and changes in national union dues subsidies as a measure of unionization probability (Barth et al. (2020)).[7]

[6]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. Similarly, Azkarate-Askasua and Zerecero (2022) develop a structural model using French data to posit counterfactuals for output and welfare in the presence and absence of oligopsony and unions in the labor market.

[7]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.

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 and document how inequality changes at the firm, local, and national levels as union membership rates increase. The results can influence 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, as well as broader notions of labor market efficiency.

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 join a union, but they have to do so on a voluntary, individual basis. Specifically, closed-shop agreements are not allowed, and the choice to join a union is an individual one. In contrast to the private sector in the US in which firm unionization requires a majority support through a union election, and in contrast to Germany in which a firm either is covered by a union agreement or not, unions can operate and bargain on behalf of members in Norwegian workplaces as long as there is a non-zero support for the union. On behalf of their members, unions can negotiate not only wages but also help settle legal disputes and push for better work conditions.

Unions are commonly structured by professional area or sector. Each individual local union is linked to a 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 (AFL-CIO) coordinates and supports union efforts across more than 50 individual unions spanning a range of professions.[8]

In the private sector, union density has been around 40 percent for the past several decades. In the public sector, union density is approximately 80 percent. The union density rate differs across sectors and industries, with almost 60 percent in the manufacturing sector and slightly less than

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

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.[9]

In terms of collective bargaining, the most common wage determination process in Norway is a two-step bargaining procedure. In the first step, sectoral collective bargaining agreements are established at the national level, usually to set minimum wage guidelines for each occupation. Failure to reach an agreement at this stage can result in strikes or lockouts. These agreements are renegotiated every 2-4 years and are generally extended to include all individuals at firms who participate in the negotiations (including non-union workers). Firms are obligated to participate in the national collective bargaining agreements if at least 10 percent of the firm's workforce requests it. In the second step, firm-level negotiations take place in which local unions and employers discuss not only firm-specific wage increases for union members but also individual-specific wage increases.[10] These negotiations usually take place annually. Non-union employees do not have the right to participate or bargain in these local negotiations, and it is up to the employer to adjust the pay of non-union workers as they deem appropriate. This two-step process covers approximately 50 percent of the private sector workforce.

A further 30 percent of the private sector is not covered by any industry-wide collective bargaining agreement at all, and all their bargaining takes place at the firm level. Thus, for 80 percent of the private sector workforce (the 50 percent covered by the two-step bargaining process and the 30 percent not covered by any sectoral agreement at all), individual firms and local unions can have a substantial influence on wages and work conditions. The local bargaining component is crucial for 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 worker performance and bargaining power. 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.

To better understand the institutional context of our analysis and relate our results to prior literature, it should also be emphasized that Norwegian employers are legally obligated to recognize and negotiate with local unions if they are present at the workplace, even with small numbers of members. Hence, in contrast to the private sector in the US in which firm unionization requires a majority vote in an election, unions can operate in Norwegian workplaces as long as there is

[9] 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.

[10] To provide an example, every union member is invited to submit a wage and work compensation request to their local union prior to the annual negotiations, along with a justification for the request. The union will assess these requests and use them in their firm-level negotiations with the employer, prioritizing some workers over others depending on how they view the submitted requests.

non-zero support for the union. Unionization in Norway is better viewed as a continuous measure ranging from 0 (no union member at the firm) to 100 (all workers are members). The changes we exploit for identification are changes along this continuous measure, in contrast to the all-or-nothing union vote in the US context or the covered/not covered dichotomous nature of bargaining coverage in some other European countries (e.g. Germany). Union density at a firm is particularly meaningful during firm-level negotiations where firm-level union bargaining power is leveraged to extract concessions. In this regard, Norwegian unions are relatively similar to those in the UK and the rest of Scandinavia.

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. Local negotiations now account for more than 70 percent of total negotiated wage increases (Mogstad et al., 2021). In fact, recent work in Norway even goes so far as to conclude that collective bargaining agreements in themselves have no impact on wages once union density is accounted for (Blandhol et al., 2020).[11] In the local negotiations, unions and employers discuss not only union-wide wage increases but also individual-specific wage increases.

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 quadrupled from 2003 to 2010. The realized value of the subsidies to workers depends on the union dues required of prospective members.

Our empirical strategy exploits changes in the maximum allowable tax deduction for union dues enacted by the national government for identification. 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 union membership rates for workers and, therefore, different union densities across firms. We use this exogenous variation in union density to identify the effects of union density on earnings in concentrated versus more competitive labor markets. Importantly, these policy reforms were not part of larger and more general tax changes and changed only the union dues subsidies. As such, the reforms enable us to carefully isolate the effect of changes in the cost of union membership without having to worry about other policy confounders happening at the same time.

[11] We believe that the industry wage floors in themselves play a minor role in wage-setting even among our control group firms; we explore this directly in Section 5.

2.3 Conceptual Framework

Originally established as a means to correct the imbalance in bargaining power between employers and employees through the monopolization of labor supply, unions operate by restricting the supply of labor to firms in order to raise employee wages and advance the collective interests of their members. Through this bargaining process, they constitute one of the biggest departures from market wage-setting in modern economies, and they have featured prominently in policy debates across Europe and the US.

The union's ability to extract and reallocate rents from firms to workers depends on two factors. First, the firm must make supernormal profits such that there are rents available for the union to capture. For example, in a perfectly competitive market in which firms are price-takers and only earn normal profits, there would be no rents for the union to extract. Second, at any given level of union density, the union must possess some leverage over the employer such that the employer is willing to enter negotiations and yield concessions to the workers. For example, in a market with zero turnover costs, an abundance of labor with identical skills, and no laws preventing firms from replacing workers in the event of industrial action or employer-employee disputes, the leverage of unions over firms would be minimal irrespective of the available firm rents.

The two conditions required for successful union rent extraction are inversely related to each other as a function of firm labor market power and represent the core of our empirical analysis. Specifically, as the labor market power of a firm grows and the market becomes more imperfect, the firm is able to collect more abnormal profits, thereby increasing the amount of rents available for extraction.[12] At the same time, as the power of a firm grows, the leverage that unions hold over firms decreases due to the lack of viable employee outside options that can be used as credible threat points in the negotiations.[13]

The relationship between the union wage premium and labor market concentration depends on how much of the additional rents unions are able to extract (marginal rent extraction) as the market is becoming increasingly concentrated. If available rents increase by more than leverage decreases for a given change in market concentration, we would expect the union wage premium to be higher as concentration rises. This would be consistent with unions helping to combat market failures induced by imperfect competition as an increase in worker wages and employment toward the competitive equilibrium would reduce deadweight loss. If, on the other hand, the decrease in bargaining leverage dominates the increase in profits, we would expect to see the opposite:

[12] In concentrated labor markets, these rents arise from firms marking down worker wages below their marginal revenue product.

[13] Unions can serve multiple functions that enhance individual leverage, especially in more competitive labor markets. For example, unions can coordinate information on opportunities within and across employers, which lowers search and signaling costs for workers and raises the threat effect of unions. This, in turn, increases workers' leverage in negotiations within the firm (Aghion et al., 1998). This effect is even more pronounced in Norway, where unions can represent workers in individual contract negotiations and workplace disputes.

the wage gains to increased unionization would be larger in more competitive labor markets. In that case, wages would be more likely to move above a competitive equilibrium, with negative implications for employment and economic efficiency.

To illustrate this point, consider (a) a small mining town with a single employer and (b) a highly competitive metropolitan area with an abundance of employers. In the first case, the monopsony power bestowed upon the firm generates substantial supernormal profits. In the second case, the competitive nature of the market implies that there are very low rates of supernormal profits. Next, let the workforce at the firms in (a) and (b) experience an equal increase in union density. In the first case, there will be substantial rents for the union to extract, but the union's threat effect on the employer is minimal due to the lack of worker outside options. In the event of a strike, for example, striking workers would have no alternative job to transition to should the firm not yield, they would have little information on the local "market wage" for their skills, and they could face retaliation from the firm. In the second case, there will be minuscule rents for the union to extract, but the union will hold substantial power over the employer as the outside options are abundant and the threat effect substantial.[14] This suggests there may be meaningful interactions between union effects and monopsony power that should be incorporated into existing theoretical frameworks and empirical applications to ensure a complete characterization of the bargaining environment and its dynamics. Understanding these dynamics is important for knowing if union power is helping to ameliorate a market failure or if it is a neutral or even negative force for labor market efficiency.

The theoretical ambiguity related to the ability of unions to extract rent across differentially concentrated labor markets is not tied to a specific labor market model. Rather, it reflects a general phenomenon present across all labor markets and applicable to any wage-setting model trying to identify wages as a function of bargaining process between employer and employees; from the early work of Smith (1776) and Davidson (1898) to the more recent work of Abowd and Lemieux (1993) and Berger et al. (2022). The justification for incorporating this theoretical framework into existing models requires only two conditions to hold - that the rent extraction ability of unions depends both on 1) available rents and 2) leverage - but has significant and important implications for how we view labor markets and derive equilibrium conditions from bargaining models.

In existing bargaining models of wage-setting, it is assumed that the bargaining power parameters are independent of, and not endogenously related to, worker outside options or monopsony power. This applies both to papers in the standard search and matching literature, as well as to pa-

[14]Existing literature suggests that these type of dynamic trade-offs between union wage effects and employer power are present at the macro level as well. For example, recent work provides evidence that unions appear to be more powerful when labor markets are tight (Schnabel, 2013), suggesting that leverage – conditional on union density and monopsony power – plays an important role in the bargaining process. This also has been observed in trends in the period after the COVID-19 pandemic. For example, at the same time that outside options accelerated quickly in the post-pandemic labor market in the United States (Autor et al., 2023), major unions in the US were able to extract significant concessions across various industries such as automotive manufacturing, the entertainment industry, and transportation.

pers in the outside options literature and applications of the canonical Berger et al. (2022) oligopsony framework (e.g., Caldwell and Danieli (2024); Azkarate-Askasua and Zerecero (2022)). The insight from our analysis is that the wage markup potential of a union itself depends on the wage markdown potential that the firm possesses. Ignoring this endogenous relationship between the bargaining power parameters and the outside options of workers / the monopsony power of firms leads to a restricted characterization of the bargaining environment and its dynamics.

In our empirical analysis, we examine this theoretically ambiguous relationship between union rent extraction and firm market power by exogenously shocking union density 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 union density is stronger in more concentrated markets, that implies that the growth in supernormal profits exceeds the decline in union leverage over the firm. For identification, we exploit changes in tax subsidies for union members in Norway, which led to significant changes in the price of union membership for workers at some firms. Provided that union membership is a normal good, this allows us to obtain an exogenous shift in union density across firms and isolate the causal effect of union density. In an extension, we also consider the effect as we shock labor market concentration holding unionization constant, relying on a shift-share instrument based on the China shock (see Appendix C).

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 A4. 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 and Sample

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 2014. Using a unique individual identifier, we follow individuals over time and across registers. We obtain demographic characteristics from the central population register, education information from the national education register, labor earnings information from the tax register, and contract hours, occupation, 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 by occupation group level, which we explain in detail in Section 3.3. Local labor markets are geographically defined based on commuting distance. These 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. 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 for 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 (>30 hours per week).

Crucial to our analysis is the ability to observe individual-level union information over time. We obtain the data from the tax register, which provides detailed information on how much each worker has paid to become a union member each year and the value of the tax deduction they took.

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. 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.[15]

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 quadrupled between 2003 and 2010. These tax changes significantly reduced the monetary cost of joining a union for workers whose union dues 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. This policy variation in union dues is illustrated in Figures A1 and A2. In Figure A1, as the deduction caps were relaxed, those at firms whose union dues were above the old cap were "treated" by the policy, while those whose union dues were below the old cap were not exposed to the policy change. Over time, those with higher base dues on the x-axis are more intensely exposed purely through changes

[15]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.

in tax policy, and the vertical distance between net dues based on the 2001 policy and the current year's policy increases only for those with higher base 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. Because unions are typically organized around a worker's occupation and industry, we take the mean union dues paid by workers in each occupation-industry cell in each year and apply this to union members and non-members alike. As such, we do not use information on individual union dues or wages that may be endogenously determined by individual or firm characteristics. This imputation approach is identical to that used in (Barth et al., 2020), and has two advantages: first, it allows us to predict the average counterfactual costs of unionization faced by those who were not part of the union; second, we can abstract away from endogenous individual and firm determinants of union dues for union members. We then characterize the union dues of the *firm* as the average 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 (the firm) or by changing the union dues directly (the union). 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 (NOK). This approach characterizes what the typical union dues at the firm would have been holding fixed the occupational mix that existed in the firm in its first year in our data set, meaning there is no endogenous change in the 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 dues 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 dues, 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 and are a lower bound on the subsidy's value.

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

$$Subsidy_{ft} = T_t * (min\{\overline{D^0_{ft}}, MaxDeduction_t\}) \,, \tag{1}$$

where T_t is the base tax rate in year t, $\overline{D^0_{ft}}$ is the imputed firm union dues at baseline adjusted to nominal NOK in year t, 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 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 dues by subtracting the value of the subsidy from the gross imputed baseline union dues ($ND_{ft} = \overline{D^0_{ft}} - Subsidy_{ft}$). We include this as an additional control in our baseline regressions. This is important because two workers whose deductions have reached 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.[16] 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.[17]

Figure A2 illustrates the drastic increase in the maximum union dues 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.

[16]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 dues. 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 (Appendix Table A4).

[17]In our other specifications that include firm fixed effects, we use only the net union dues to predict union membership because conditional on the fixed effects, changes in the subsidy perfectly predict net dues over time.

3.3 Defining Concentration

The recent empirical literature on labor market power has established that market concentration captures important aspects of firms' labor market power, with negative effects on worker earnings and employment (e.g.Azar et al. (2020a); Marinescu et al. (2021); Dodini et al. (2024)). To obtain a comprehensive proxy for labor market concentration, we build on Dodini et al. (2024) 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 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.[18]

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. (2024), 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.[19]

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

[18]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.

[19]Dodini (2023) also validates the optimal cluster number in a US context at approximately 20 clusters.

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 A12–A16).

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. Because HHI is the result of equilibrium forces of labor supply and demand, fixing HHI at a particular point in time helps address the typical shortcomings of HHI as a conceptual measure. Fixing HHI also 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.[20]

To ensure that our results are not driven by the particular monopsony power measure that we use, we supplement the paper with the AKM approach in Bassier et al. (2022). The results from this exercise are provided in Appendix B and demonstrate that our findings are robust to using this alternative measure of labor market power.

Table A1 contains summary statistics for our analysis sample.[21] Approximately 63% of our sample of workers are members of unions, and their earnings, on average, are approximately 458,000 NOK. The imputed base tax subsidy for our sample is on average 730 NOK over the sample period with net-of-subsidy union dues 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. In 2001-2002 prior to the tax reform, approximately 59% of workers with an HHI below 0.05 were members of a union compared to approximately 79% of those with an HHI greater than 0.15.

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 2014. These changes significantly reduced the monetary cost of joining a union for workers whose

[20]Our results are similar when using the firm's average HHI over the whole sample period.
[21]To reduce computational time, we take a 70% random subsample of workers in the data.

union dues 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 exposed to the effects of the subsidy 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 to changes in the cost of joining a union. We want to estimate this process as flexibly as possible, without 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 as the treatment variable 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 individual 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:

$$\begin{aligned} Union_{iocft} = &\beta_0 + \beta_1 Subsidy_{ft} + \beta_2 ND_{ft} + \beta_3 \overline{HHI}_f + \beta_4 \overline{HHI}_f * Subsidy_{ft} \\ &+ \beta_5 \overline{HHI}_f * ND_{ft} + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \varepsilon_{iocft}, \end{aligned} \tag{2}$$

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 union membership 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 the 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.[22] 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 union membership and earnings 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, and other unobserved, time-invariant factors. The year fixed effects absorb any systematic changes in unionization propensity over time that concern all workers.[23]

The key result of this exercise is that firms with larger increases in subsidies (treatment intensity) will have higher rates at which workers at the firm become members of a union. We, therefore, use the predictions from these regressions to calculate the predicted union density for each firm in the data in each year, which we call $\widehat{UD}_{ft}$. This is the mean of the predicted probability of union membership across all workers at the firm each year. Importantly, this predicted value jointly takes into account the individual characteristics of workers at the firm. Aggregating these measures from individual propensities also means that we are capturing the average effect of an increase in union power as proxied by the share of workers represented by the union (Freeman and Medoff, 1981).

There is potentially an important trade-off between the flexibility of our main specification and the potential risk that, conditional on education and occupation-by-industry fixed effect, individuals with specific characteristics may sort into firms that experience greater exposure to the policy, and that these characteristics may also be correlated with the outcomes we examine. We conduct several supplemental analyses and alternative estimation approaches to examine this concern. First, we replace the subsidy and net dues with those based on each worker's first firm in the data rather than their current firm's first year in the data (Appendix Table A5). Second, we implement a fully saturated approach that estimates our main specification with additional fixed effects for individuals, firms, and industry-by-year interactions (Panel C of Table 2). Finally, we conduct a split-sample IV approach that estimates firm-level union density responses to the subsidy rather than aggregating individual union membership choices (Panel D of Table 2). All these approaches produce results that align closely with our baseline specification. This implies that endogenous worker sorting based on unobservables that also are correlated with our outcomes is unlikely to drive the results we find. We discuss these alternative approaches in detail in Section 4.1.

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:

$$\begin{aligned} Log(Earnings)_{iocft} = & \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \alpha_2 \widehat{UD}_{ft} * \overline{HHI}_f \\ & + \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \varepsilon_{iocft} \end{aligned} \tag{3}$$

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

[23] When estimating our models for the average response to the subsidies, we omit HHI and its interactions from the model. We do the same when estimating earnings outcomes, as in Equation 3.

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 2. 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 union density after holding constant time-invariant characteristics of the firm. The interaction between $\widehat{UD}_{ft}$ and HHI allows the effects of union density to differentially affect earnings in concentrated markets. Our alternative approaches to modeling the union membership choice are in Section 4.1 below. These approaches point to very similar conclusions as our main approach.[24]

As explained in Section 2.3, 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. At the same time, the relative bargaining leverage 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. 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 leverage dominates, we would expect to see the opposite.

We cluster our standard errors at the firm level where the union density effect is allocated. We follow the standard two-stage least squares convention and adjust the standard errors in our earnings regressions to account for the uncertainty from the union membership regression.

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 ($\overline{HHI}_f^P$):

$$\begin{aligned} Log(Earnings)_{iocft} =& \alpha_0 + \alpha_1 \widehat{UD}_{ft} + \alpha_2 \widehat{UD}_{ft} * \overline{HHI}_f + \alpha_3 \widehat{UD}_{ft} * \overline{HHI}_f^P \\ &+ \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \end{aligned} \tag{4}$$

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

[24]In Appendix Table A19, we have also estimated a version of our main model in which we leverage the leave-one-out dues for identification in which we omit firm f when calculating the occupation-industry average dues. This eliminates concerns about a mechanical relationship between the firm's initial dues and the outcome we examine. The results are not statistically significantly different from our main findings.

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.[25]

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 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 through an event study framework in Appendix Figure A3 and conclude there is no reason to suspect that diverging trends between the control and the treatment groups can explain any of our results based either on the trends in union membership (Panel A) or log earnings (Panel B).[26] Specifically, individuals at firms with large reductions in net dues have higher earnings growth *only* during the years in which the maximum deduction changed the most and the subsidy changes became efficacious (2003-2010), while the trends were parallel in 1999-2002 and after 2010 when the deductions were stable. This aligns extremely well with the trends in union membership, which began to increase disproportionately in firms that had large relative reductions in net dues only during the 2003-2010 period. These parallel trends tests are designed to explore the comparability of the two groups in times in which the subsidies were not changing.

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

[25]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.

[26]We estimate this model interacting year dummies with the total reduction in net dues within each firm before 2010, a measure of total policy exposure. We include the same fixed effects and controls as in Equation 3: occupation by industry cell, education program, age group, year, and firm.

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 constraining individuals to remain at a particular firm since that could be part of the response. 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. 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 intensity measure is therefore 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, we use this approach to better account for the fact that individuals are choosing union membership rather than firm-level union density.

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. 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 dues as a shifter of individual union membership choices. This is effective because baseline dues are fixed for each firm, so changes in the net dues 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 with the same set of controls and fixed effects. 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 is that it 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. Leveraging firm-year level observations, 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 dues (i.e. the net cost of joining) in a model with firm and year fixed effects. We 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 predicted firm-level union density in the second half of the sample. This second-stage regression mirrors Equation 3 with the addition of individual 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 others may have a preference for one of these alternatives. Encouragingly, other than the results being slightly noisier when using these two approaches, our conclusions are unaffected by the choice of estimation approach.

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 join a union. These results are obtained through the estimation of Equation 2. In columns (1) and (2), we examine the relationship between subsidies and union membership without taking labor market concentration into account. Column (1) includes occupation-by-industry, education, and age group fixed effects, while column (2) further includes individual-level fixed effects. The estimates in column (2) thus subsume possible noise introduced by "always union" members and isolate changes in union status within person. In columns (3) and (4), we study the relationship between the subsidies and union membership 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 occupation-based concentration measures, rather than the skill-based measures, are provided in the online appendix Tables A12–A16.

The results in column (1) demonstrate that the subsidies had a strong impact on the probability that workers join a union. Raising the subsidy by 1,000 NOK increases the probability of being in a union by approximately 10 percentage points. The coefficient on the subsidy in column (2) is slightly larger and demonstrates 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).[27]

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 union membership 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.[28]

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

[27]For robustness, we also estimate this equation while excluding the net union dues 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.

[28]To ensure that commuting zone size or other local factors are not driving our results, we have replicated our main findings using models that account for different local labor market fixed effects, demand shocks, and local unionization rates. The results are provided in Appendix Table A6, and demonstrate that our findings are robust to restricting the identifying variation to come only from within local labor markets. Specifically, we show using four specifications that differential shocks to demand for certain skill types or certain labor markets are not driving our results. We estimate the models with local labor market fixed effects (Panel A) and skill cluster by year dummies (Panel B). These do not appreciably affect our estimates. In Panels C and D, we use Equation 2 to estimate predicted local labor market union density (Panel C) and local labor market-cluster union density (Panel D) and include these as additional controls in order to limit the possible role of correlated union shocks or local labor market spillovers of union membership. These estimates are not statistically or economically significantly different from our baseline models and thus rule out these mechanisms.

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.[29] 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. In Panel A of Figure 2, we show that approximately 40-50% of workers indicated that they would reconsider joining (or leaving) a union if the net cost of joining were reduced (or increased) by as little as 500 NOK per month (6,000 NOK per year). While this hypothetical scenario asks workers about reconsidering their choices (i.e. is missing real-world stakes), the results are informative. Norwegian workers, themselves, indicate high levels of price sensitivity when considering the union membership decision. In addition, those working in industries in locations that are more concentrated are significantly more likely to reconsider their union membership choices in response to price changes, as we show in Panel B of Figure 2. The gradient rises from approximately 41% at an HHI of zero to approximately 60% for those with an HHI of 0.2.

5.2 Earnings Effects

Table 2 provides estimates on the effect of union density on individual log annual earnings using the changes in tax subsidies for union members as 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 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 skill clusters instead.

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 1.9 percent. The coefficient on union density is nearly identical to that in Barth et al. (2020) 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

[29] Under the assumption of declining margin effects, the impact on unionization rates post-treatment may be expected to be smaller in areas that already have a high degree of unionization. However, when we augment our core model with controls for (1) the predicted LLM - year - unionization rate and (2) the predicted LLM - cluster – year unionization rate, we obtain quantitatively similar results (Appendix Table A6). Thus, we find little evidence that the relationship between these variables is exhibiting diminishing marginal effects in our context.

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 column (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 leverage in these markets. In other words, it is consistent with the idea that unions may be able to correct market failures caused by labor market concentration by pushing wages up toward the competitive equilibrium.

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

The market imperfections generated by monopsonistic power and the rents available to unions in concentrated markets may be significantly larger in the private sector compared to the public sector. 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 2 replicates Panel A but restricts the sample to only the private sector. The relationship between union density and earnings as a function of labor concentration is 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. In Panel C, when we take a saturated approach that includes individual, firm, and industry-by-year fixed effects, the base effect is slightly smaller than those in Panel A (0.78% vs 1.3%), but there is still a significant increase in the union wage premium in more concentrated markets. In Panel D, which is comparable to Panel B using the private sector, we see slightly smaller base effects on union density but an even steeper increase in the union wage premium over HHI. Across all these approaches, the results are clear that the wage returns to union density increase when firm labor market power is higher.

One potentially important point related to our estimation framework revolves around possible spillover effects from more to less unionized firms. Specifically, as some firms become increasingly unionized and their workers experience wage gains, these more unionized firms may alter the outside options of workers in less unionized firms, thereby generating spillovers. In terms of our empirical analysis, any such spillovers should attenuate our results and lead us to estimate lower bounds of the union density effect. However, we find little evidence of such spillovers in our data. First, such spillovers could take place at the sectoral level, where the nationwide sectoral bargaining agreements are established and operate. However, concerns over such sectoral-level spillovers are assuaged in Panel C of Table 2. In this panel, we provide evidence on the union wage effect using a fully saturated version of our core model, in which we include industry-by-year fixed effects. These fixed effects absorb any wage floor effects as wage floors are set at the sectoral level, such that we are comparing more or less unionized firms holding wage floors constant in the same year.[30] Second, our results are unaffected by the incorporation of a series of additional fixed effects and controls that directly try to limit any potential spillovers. The results from this exercise are shown in Appendix Table A6. All results are obtained through the estimation of Equation 3 with additional controls and fixed effects. In Panel A, we incorporate local labor market (LLM) indicators. In Panel B, we add cluster-by-year fixed effects. In Panel C, we control for the predicted union density at the local labor market level using the predictions from Equation 2. In Panel D, we control for the predicted union density at the local labor market-by-skill cluster level. We emphasize the results in Panels C and D in particular, in which we instrument both for firm-level union density as well as for the union density in the local market. These two specifications directly incorporate the idea that the "market wage" an individual can get may depend on the unionization rates in the market. None of these estimates materially differ from our main model.

5.3 Source of Rents

In Table 3, we combine our primary labor market data with firm-level revenue data and explore the relative impact of product- 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 both product and labor rents 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 firms in our analysis sample that have 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.

[30] We also examine how the returns to union density change depending on the bindingness of the industry-level wage floor set in sectoral negotiations as measured by a Kaitz index. This exercise in Table A9 shows that the bindingness of the industry wage floors has little effect on the wage returns to union density over employment concentration.

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.[31] We believe that this is a novel finding with important policy implications, alluding to the fact that unions' ability to extract rent and reallocate this rent to their members depends not only on the extent of market imperfections but also on whether these imperfections are driven by labor concentration or possible product market concentration.

5.4 Heterogeneous Earnings Effects

We examine effect heterogeneity across three different dimensions: worker productivity (proxied by whether the worker earns above or below median annual earnings *within occupation* at the firm), worker type (with respect to white and blue-collar workers), and worker social capital. We rely on two measures to identify social capital. First, we examine heterogeneity based on within-firm occupational diversity: the number of distinct 3-digit occupations within the firm-year cell per worker in the firm. The idea behind this analysis is that greater occupational diversity will reduce the bargaining leverage of the workforce and lead to smaller union benefits due to increased coordination issues across workers and more heterogeneity in worker demands. Second, we examine heterogeneity based on within-Norway birthplace diversity using the counties that workers were born in. The idea behind this approach is that a workplace with a greater share of workers from the same geographic background may be more capable of exerting pressure in negotiations.

The purpose of the first two of these analyses is to build on prior union work and set the stage for the inequality analyses that we perform in the next section, examining differential treatment effects across the productivity and base wage dimensions. Specifically, prior work suggests that union wage effects may load on lower productivity individuals and individuals with traditionally more physically demanding jobs that are placed lower down in the income distribution (e.g., blue collar jobs). The last of these analyses builds on prior work on bargaining power (e.g., Naidu (2022)) to examine if membership in labor unions provides different bargaining options for individuals depending on their connectedness and networks.

With respect to differential treatment effects across the productivity and base wage dimensions, Table 4 uncovers two novel sets of results. First, the results 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

[31] 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.

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 firms and narrow sub-sectors in competitive markets, while this is not the case in concentrated markets characterized by monopsonistic competition.[32]

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 satisfying and ensuring the continued membership renewal of high-skilled workers who are more likely to leave the unions and avoid paying more in dues. In addition, higher-productivity workers also may carry more weight as representative agents in negotiations. For example, threatening to strike carries more weight if the strikers are the firm's most productive workers. These three characteristics make higher-productivity workers strong potential members. As markets become more concentrated, the reduction in outside options for high-skilled and white-collar workers combined with the improved rent extraction opportunities available to unions means that they can shift focus to bargaining for a more general wage increase across all worker types. As we will show in Section 5.5, there is also a positive intensive-margin effect on employment at firms with high HHI, which aligns with this proposed objective of maximizing dues over time.[33]

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 differential treatment effects across workers as a function of their social cap-

[32] 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).

[33] These earnings effects do not reflect differential propensities to join unions in response to subsidies in competitive markets. We demonstrate this directly in Appendix Table A2. Specifically, above-median workers are less responsive to the subsidies in competitive markets and more responsive in concentrated markets, which contrasts with the overall pattern of returns. This highlights the necessity for unions to leverage 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 A8), but this difference goes to zero as concentration increases.

ital, Table A10 provides a set of key findings. While social capital as proxied by within-Norway birthplace diversity has no impact on our estimates, social capital as proxied by within-firm occupational diversity has a large effect across the HHI distribution. Specifically, the greater the occupational diversity at the firm, the smaller the subsidy gradient over HHI in terms of union membership probability, and the smaller the union density gradient over HHI in terms of log annual earnings. These results are consistent with the idea that occupational diversity makes coordination issues more difficult among employees and reduces the leverage that they have over the employer in negotiations. This appears particularly true in concentrated markets, where such coordination and leverage may be more necessary in order to extract available rents. That we do not find similar effects when using birth county heterogeneity could be because it is a worse predictor of social capital, particularly in a relatively small and homogeneous country such as Norway. It is worth emphasizing that these results align well with results from large second survey on union members in Norway that we conducted in 2023. One of the questions asked in this survey is whether workers believe that unions help some workers more than others. Approximately 20 percent responded affirmatively. When then asked to specify which group of workers benefit the most, the most common answers can be summarized by "those with greater social capital," including those with better networks, native-born Norwegians, and those with a higher position in the hierarchy. While only suggestive, this survey-based evidence strongly aligns with the results in Tables 4 and A10.

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, 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 on—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 push wages above the current wage offered by the employer, such as a wage equal to the marginal revenue product of labor, the firm would hire more workers, but their marginal profits per worker would be lower (all else equal). In such markets, 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 A4 for further discussion).

To address this question, we conduct two analyses. First, we estimate Equation 3 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 3 at the firm level using the number of workers at the firm as the dependent variable. This equation includes firm and year fixed effects and omits other individual-level controls. The first regression addresses employment on the intensive margin, and the second regression addresses employment 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.[34] 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.

The results from these two sets of regressions are shown in Table 5. 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 2.3 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 nearly 4 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. However, when limiting the sample to the private sector, we do find suggestive evidence of extensive margin employment losses in competitive markets, with a positive interaction with HHI. However, the effect in competitive markets are only marginally significant, and the HHI interaction is positive but not statistically significant.

Another margin at which employment may change within the firm in response to the documented wage effects is via changes in worker composition. We explore this in Table A7 where we estimate Equation 3 with linear probability models for separation and the joint probability of separation by education cells for each individual worker. These represent changes in separation probability conditional on the observed characteristics of the worker (i.e., conditional on occupation, age, and educational program).

The results from this exercise show that retention improves slightly with union density, particularly for workers with at least a high school/vocational degree in concentrated markets. However, these effects are quantitatively small and in many instances are not statistically significant. In terms of overall worker composition, in competitive labor markets, firms tend to adjust the workforce composition to consist of more middle-skilled workers (those with a high school or vocational degree) and away from those with only secondary education and those with a college degree. In concentrated labor markets, these adjustments are minimal, consistent with there being more need

[34]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.

for firms to adjust their labor inputs in more competitive environments where they have fewer rents available. This is also consistent with our employment results in Table 5 in which firms in competitive labor markets trim their workforce in response to an increase in wages caused by union density relative to less unionized workplaces.

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.[35] 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 or broad wage increase is imposed in imperfectly competitive labor markets.[36]

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 to include the firm level 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 2 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 annual earnings distribution at the firm. We then calculate overall predicted firm union density using the individual predicted prob-

[35] 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.

[36] Because of the scaling effect, the unionization events appear to impact the entire production function of firms, with potential implications not only for input use and output produced but also for profits and product market prices. We leave this to future work but note that understanding these dynamics is important in order to obtain a complete understanding of unions' impacts on labor markets.

abilities from this model. Following this step, at the firm level, we regress each of our outcomes (the percentile ratios) on the interaction of predicted union density and concentration and include firm and year fixed effects in the model. To explore (2), we measure inequality (percentile ratios) at the local labor market level (LLM) and aggregate average concentration and average predicted union density to the LLM level from Equation 2, 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 find 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 (7.2%), but most of the effect loads on changes above the median (4.2%). The opposite is true in concentrated markets, where union density is uniformly associated with reductions in inequality, particularly in the top half (-7.2%). Part of this effect is attributable to the increase in part-time work in competitive markets that we find in Table 5: a wider variation in hours expands the dispersion of total earnings within firms. This suggests that crowding effects in competitive labor markets are operating within firms on the hours margin. Overall, within-firm inequality in firms in concentrated markets falls when union density at the firm rises, particularly above the median.

These results of increases in inequality are also consistent with our 2023 survey in which we find that approximately 20 percent of survey respondents indicated that unions prioritize some workers over others, particularly those with better network connections within the firm and those higher in the firm's hierarchy. In a competitive labor market in which higher-status or higher-productivity workers have better outside options, union bargaining strength may be disproportionately exerted to raise the earnings of these workers as a reward for loyalty to the firm, as a premium for firm-specific human capital, or as an incentive to stay in the union and pay more in union dues.

In Panel B of Table 6, we show that an opposing 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 decreases in competitive labor markets and falls by marginally more 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 decreases the 90-10 ratio by 0.05, or by just over 1.6 percent relative to the mean value of 3.2. In concentrated markets, a one percentage point increase in unionization decreases this ratio by approximately 0.06, or by approximately 1.9 percent relative to the mean, though this difference is not statistically significant. In columns (2) and (3), we find that the 90/10 ratio effect is an approximately even split between compression effects in the 90/50 and 50/10 ratios. In concentrated local labor markets, the gains from unionization accrue to the 10th percentile slightly more, reducing below-median local labor market inequality by more than above-median inequality. However, the difference in the effect in concentrated markets is not statistically significant at conventional levels. The results for local labor markets suggest that cross-firm sorting and reallocation 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 and the unit of analysis. As we show above, the marginal union member is more likely to be working in lower-paid, concentrated segments of the labor market, even within geographic areas. This 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, most specifically in the context of concentrated markets.[37]

To assess the aggregate effect of this reform and the subsequent change in union density in the economy on overall inequality, we follow DiNardo et al. (1996) by taking a semi-parametric approach to estimate the counterfactual density of annual earnings in 2014 were there to be the same distribution of predicted firm-level union density that was present in 2002 based on that year's policy environment. To do this, we estimate predicted firm-level union density for all workers and firms in the economy as in Equation 2. We then limit the sample to the years 2002 and 2014 and regress an indicator variable for 2014 on all of our control variables in Equation 3, including predicted union density and the interaction between predicted union density and HHI. We use inverse weights from the predictions from this model to reweight the counterfactual 2014 to match the environments of the 2002 policy environment. In Figure 3, we present the densities of the counterfactual 2014 and actual 2014. In Table 7, we present distributional statistics.

Our results follow much of the prior literature by suggesting that aggregate inequality decreases as union density increases. In particular, the 10th and 25th percentiles of the income distribution rose by approximately 10 percent, while the 90th percentile fell by approximately 7.5 percent. Overall, the 90/10 annual income ratio fell by 7.3 percent relative to the counterfactual. This finding suggests that despite increases in inequality within firms, because the marginal union member is more likely to be in a concentrated labor market and be lower-paid in the overall income distribution, increases in union density have a compression effect on the total income distribution in Norway.

5.7 Extensions

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 union density at the firm

[37]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)).

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 C, 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 union density 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.

6 Discussion

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 (2024); Dodini et al. (2024)) or union power (e.g., Fortin et al. (2022); Lee and Mas (2012)), without explicitly considering the dynamic 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 national reforms to tax deductions for union dues as an exogenous shock to unionization, we demonstrate that the price elasticity of union membership has a steep gradient over labor market concentration. We then show that there is 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. This gradient over market power also holds when considering two different measures of labor market power.

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 theoretically more disadvantageous bargaining position due to the inability to leverage strong 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 negative employment effects in competitive markets.

Running horse races between 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 workers who benefit from union density 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 within firms 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 firms 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. In the overall local economy, unionization decreases inequality when labor markets are more competitive and reduces local inequality by marginally more when markets are concentrated. Through our semi-parametric reweighting exercise, we show that aggregate inequality would have been significantly greater by 2014 absent the tax policy reforms for union dues deductions in the early to mid-2000s. 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 market efficiency and wage equality and provide a new avenue for future research.

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 relatively small union earnings premium in more competitive markets comes at a modest employment cost. That there are large and sizable positive wage and employment effects in highly concentrated markets, on the other hand, point to unions as being able to ameliorate market failures 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 essential 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 policies. Recently, policymakers have actively started to consider concentration in the labor market as a basis for antitrust enforcement in the US (e.g., Marinescu (2019)), the U.K. (e.g., Competition and Markets Authority (2024)), and the EU (e.g., Aresu et al. (2024)). What these discussions have in common is an exclusive focus on the firm's power, without considering the potential countermeasures provided by organized labor. Given our findings coupled with an average union density rate of 32 percent in the OECD, adding worker power into the discussion has the potential to considerably alter some of the policy recommendations made by these organizations. Specifically, 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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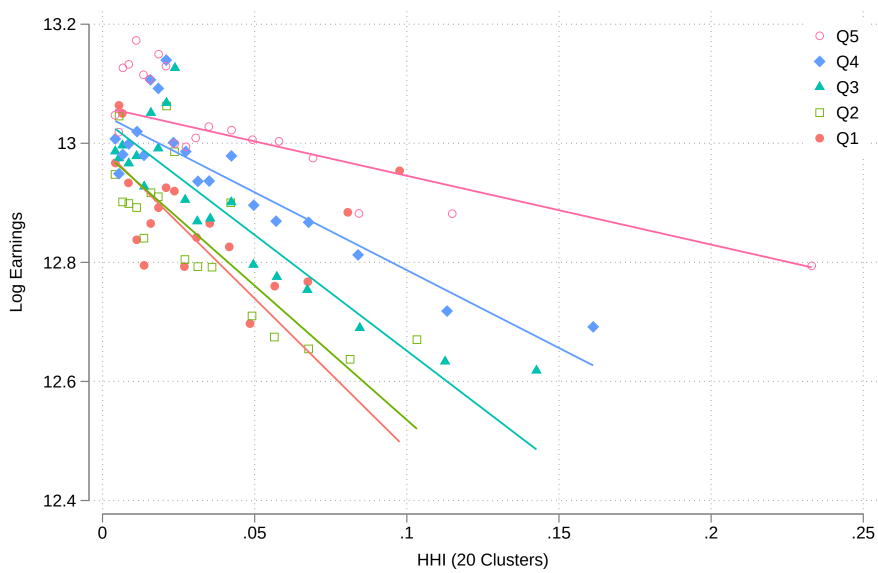
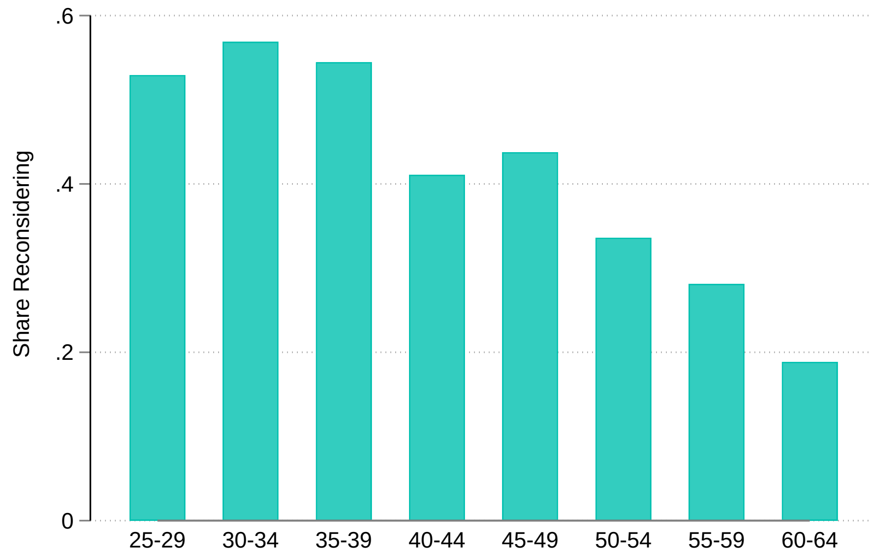
Figures

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

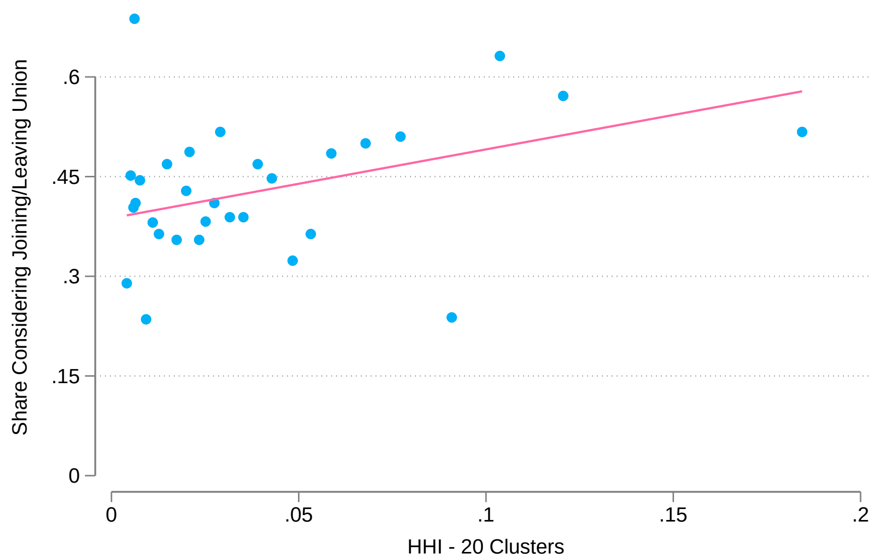


Source: Authors' calculations of Norwegian registry data.
Notes: Predicted union densities are based on average predicted unionization rates at each firm from Equation 2 as described in the text. Q1 through Q5 represent quintiles of the distribution of predicted union density.

Figure 2: Survey Responses to 500 NOK Change in Net Union Dues
Panel A: Share Reconsidering Membership by Age

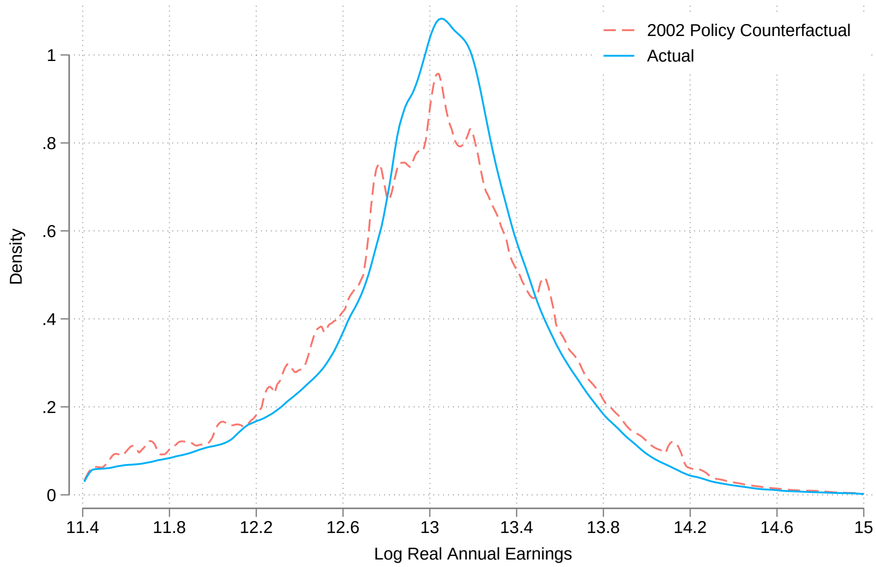


Panel B: Share Reconsidering Membership by HHI



Source: Authors' calculations of Norwegian registry data and 2022 survey.
Notes: Survey asks if union members (non-members) would consider leaving (joining) if the monthly net dues increased (decreased) by 500 NOK. Panel B shows a binned scatterplot of the relationship between the average HHI as calculated in the 2001-2014 labor registers at the industry-county level (x-axis) and the share reconsidering their membership. HHI is calculated at 20 skill clusters at the local labor market level then averaged over the county-industry groups in the registers.

Figure 3: Log Annual Earnings in 2014 vs Reweighted Counterfactual



Source: Authors' calculations of Norwegian registry data.
Notes: Using estimates of Equation 2 for all workers, we calculate predicted firm-level union density. Following DiNardo et al. (1996), we limit the sample to the years 2002 and 2014, we then estimate a dummy variable for 2014 on all of our control variables in Equation 3, including predicted union density and the interaction between predicted union density and HHI. We use inverse weights for the predictions from this model to reweight the counterfactual 2014 to match the 2002 policy environment.

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)	0.103**	0.158***	0.0723	0.134***	0.0750	0.137***
	(0.0518)	(0.0199)	(0.0530)	(0.0201)	(0.0531)	(0.0202)
HHI x Subsidy			0.133***	0.257***	0.0889**	0.182***
			(0.0443)	(0.0238)	(0.0392)	(0.0216)
HHI			1.400***	0.00364	1.242***	0.232***
			(0.246)	(0.0816)	(0.163)	(0.0520)
Observations	15,069,038	14,883,635	15,069,038	14,883,635	15,069,038	14,883,635
R-squared	0.249	0.749	0.251	0.750	0.251	0.750
Occupation x Industry FE	Yes	Yes	Yes	Yes	Yes	Yes
Education FE	Yes	Yes	Yes	Yes	Yes	Yes
Age Group FE	Yes	Yes	Yes	Yes	Yes	Yes
Individual FE	No	Yes	No	Yes	No	Yes
Avg Pr(Union)	0.631	0.631	0.631	0.631	0.631	0.631
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 2. Notes: Standard errors are in parentheses and are clustered at the firm level. All estimates include fixed effects for occupation by industry cells, detailed educational program, age group, and year. * indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

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

	Panel A: Main Approach		
	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0193***	0.0128***	0.0126***
	(0.00257)	(0.00253)	(0.00259)
Predicted Firm Union Density * HHI		0.0150***	0.0187***
		(0.00377)	(0.00347)
	Panel B: Main Approach, Private Sector Only		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0141***	0.0139***	0.0133***
	(0.00278)	(0.00329)	(0.00323)
Predicted Firm Union Density * HHI		0.0331***	0.0202***
		(0.00802)	(0.00734)
	Panel C: Saturated Approach		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.00828***	0.00777***	0.00769***
	(0.00258)	(0.00258)	(0.00256)
Predicted Firm Union Density * HHI		0.0125***	0.0120***
		(0.00404)	(0.00395)
	Panel D: Split-Sample IV		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.00704***	0.00575***	0.00586***
	(0.000569)	(0.000587)	(0.000593)
Predicted Firm Union Density * HHI		0.0452***	0.0407***
		(0.00586)	(0.00562)

Source: Authors' estimates as described in the text. Panels A and B correspond to Equation 3.
Notes: Standard errors in parentheses clustered at the firm level and adjusted for model uncertainty in Equation 2. * indicates significance at the 10% level, ** at the 5% level, and *** 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. Panel D standard errors adjusted following Inoue and Solon (2010).

Table 3: Union Density Effects by Labor and Product Market Concentration

VARIABLES	(1) No Labor HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.0178***	0.0160***	0.0160***
	(0.00287)	(0.00326)	(0.00321)
Predicted Firm Union Density * Labor HHI		0.0252**	0.00652
		(0.0103)	(0.00966)
Predicted Firm Union Density * Industry Revenue HHI		0.0247***	0.0228**
		(0.00927)	(0.00902)
Change in ME with 10 ppt Change in Labor HHI		0.0025	0.0007
Change in ME with 10 ppt Change in Industry Revenue HHI		0.0025	0.0023
Observations	7,134,338	7,134,338	7,134,338

Notes: Authors' estimates as described in the text. Estimates correspond to Equation 4. Standard errors are in parentheses and are clustered at the firm level and adjusted for model uncertainty in Equation 2. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year. * indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

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

	Panel A: Above vs Below Firm-Occupation Median		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.00351	-0.00348	-0.00415*
	(0.00278)	(0.00244)	(0.00247)
Predicted Firm Union Density * HHI		0.0346***	0.0352***
		(0.00355)	(0.00316)
Union Density * Above Firm-Occ Median	0.00597***	0.00618***	0.00625***
	(4.36e-05)	(5.02e-05)	(5.07e-05)
Union Density * HHI * Above Firm-Occ Median		-0.00448***	-0.00503***
		(0.000330)	(0.000321)
	Panel B: White Collar vs Other Occupations		
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters
Predicted Firm Union Density	0.00583***	0.00160**	0.00143*
	(0.00108)	(0.000714)	(0.000759)
Predicted Firm Union Density * HHI		0.0271***	0.0293***
		(0.00280)	(0.00280)
Union Density * White Collar	0.00238***	0.00161***	0.00175***
	(0.000418)	(0.000464)	(0.000463)
Union Density * HHI * White Collar		-0.000848	-0.000642
		(0.000518)	(0.000445)
Observations	15,069,035	15,069,035	15,069,035

Source: Authors' estimates corresponding with Equation 3 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level and adjusted for model uncertainty in Equation 2 with subgroup interactions. 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

	Individual Regressions		Firm Regressions	
	(1)	(2)	(3)	(4)
VARIABLES	Pr(Hours>30)	Pr(Hours>30)	Workers	Workers
All Sectors				
Lagged Predicted Union Density	0.00675	-0.0231***	-0.175	0.593
	(0.00417)	(0.00597)	(1.182)	(1.448)
Lagged Predicted Union Density * HHI		0.0629***		-1.630
		(0.00862)		(2.726)
Observations	14,424,842	14,424,842	206,875	206,875
Private Sector				
Lagged Predicted Union Density	-0.00566	-0.0142***	-1.884*	-2.677*
	(0.00415)	(0.00551)	(1.078)	(1.410)
Lagged Predicted Union Density * HHI		0.0395***		1.305
		(0.00905)		(2.841)
Observations	8,220,957	8,220,957	172,638	172,638
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 and are adjusted for model uncertainty in Equation 2. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year at the individual level and firm and year fixed effects at the firm level.

* 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		
	(1)	(2)	(3)
VARIABLES	Firm 90/10	Firm 90/50	Firm 50/10
Predicted Union Density	0.185***	0.0634***	0.0469***
	(0.0374)	(0.0106)	(0.00488)
Predicted Union Density x HHI	-0.284***	-0.109***	-0.0566***
	(0.0694)	(0.0202)	(0.00904)
Dep Variable Mean	2.59	1.52	1.70
Pct Effect Union Density	7.15 %	4.18 %	2.75 %
Pct Effect Union Density x HHI	-10.97 %	-7.19 %	-3.32 %
Observations	237,217	237,217	237,217
R-squared	0.627	0.639	0.624
Firm FE	Yes	Yes	Yes

	Panel B: Local Labor Market Level Inequality		
	(1)	(2)	(3)
VARIABLES	LLM 90/10	LLM 90/50	LLM 50/10
Predicted Union Density	-0.0495***	-0.0135***	-0.0153**
	(0.0164)	(0.00447)	(0.00773)
Predicted Union Density x HHI	-0.0104	-0.00107	-0.00430
	(0.00699)	(0.00208)	(0.00434)
Dep Variable Mean	3.19	1.67	1.90
Pct Effect Union Density	-1.55 %	-0.81 %	-0.80 %
Pct Effect Union Density x HHI	-0.33 %	-0.06 %	-0.23 %
Observations	2,236	2,236	2,236
R-squared	0.975	0.982	0.883
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. Both are adjusted for model uncertainty in Equation 2. 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.

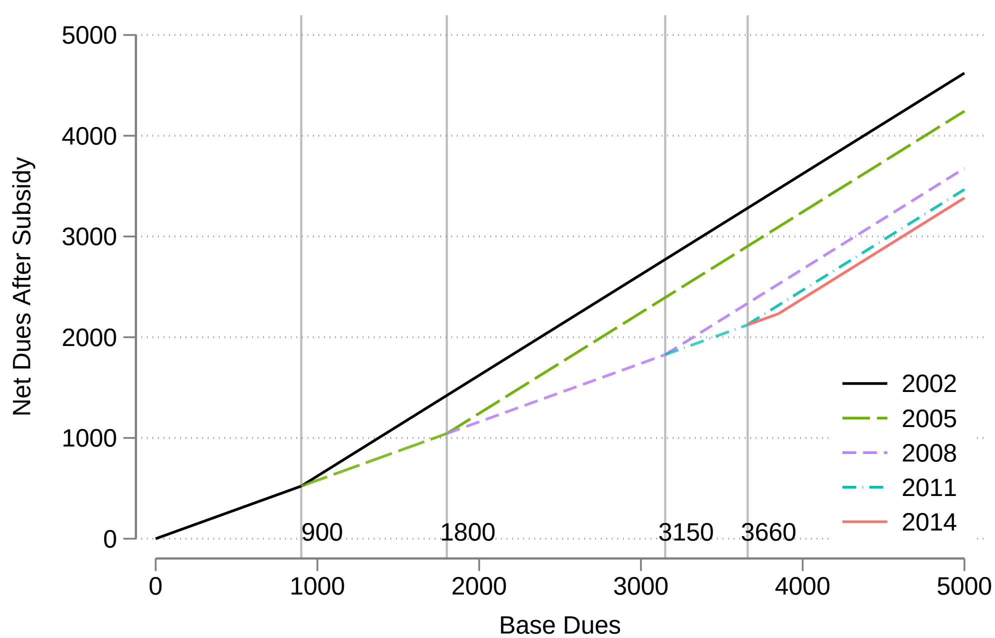
Table 7: Counterfactual Aggregate Inequality Following DiNardo et al. (1996)

	(1) 2002 Policy Counterfactual	(2) Actual 2014	(3) Gap
P10	12.243	12.345	0.102
P25	12.679	12.771	0.092
P50	13.024	13.052	0.028
P75	13.345	13.310	-0.035
P90	13.686	13.610	-0.076
90/10 Ratio	2.442	2.265	-7.3 %
90/50 Ratio	1.662	1.557	-6.3 %
50/10 Ratio	1.781	1.707	-4.1 %

Source: Authors' estimates as described in the text.
Notes: Notes: Using estimates of Equation 2 for all workers, we calculate predicted firm-level union density. Following DiNardo et al. (1996), we limit the sample to the years 2002 and 2014, we then estimate a dummy variable for 2014 on all of our control variables in Equation 3, including predicted union density and the interaction between predicted union density and HHI. We use inverse weights for the predictions from this model to reweight the counterfactual 2014 to match the 2002 policy environment.

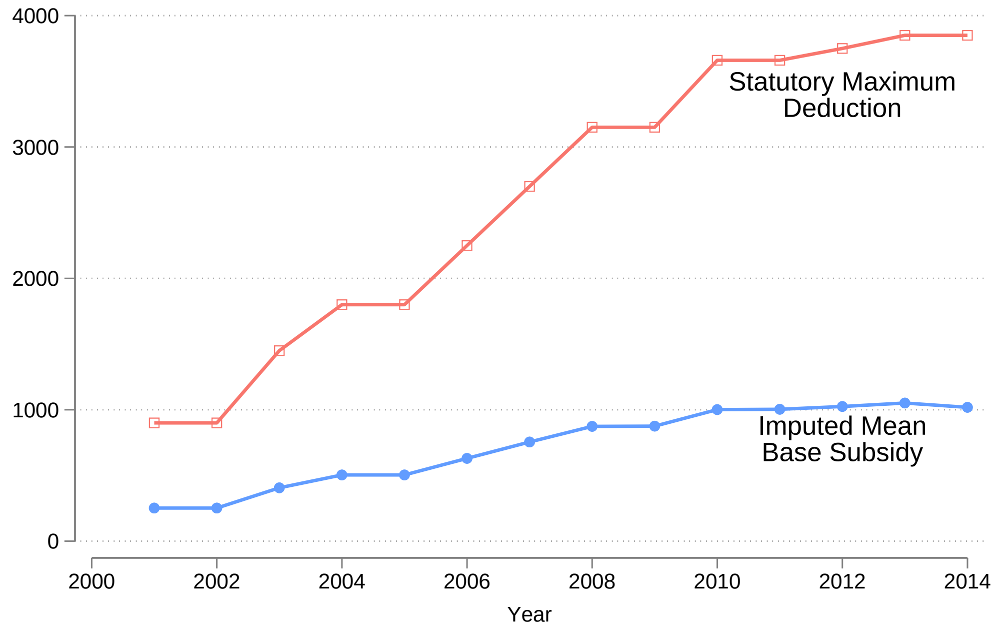
A Appendix: Additional Tables and Figures

Figure A1: Base Dues and Dues Net of Subsidies by Year



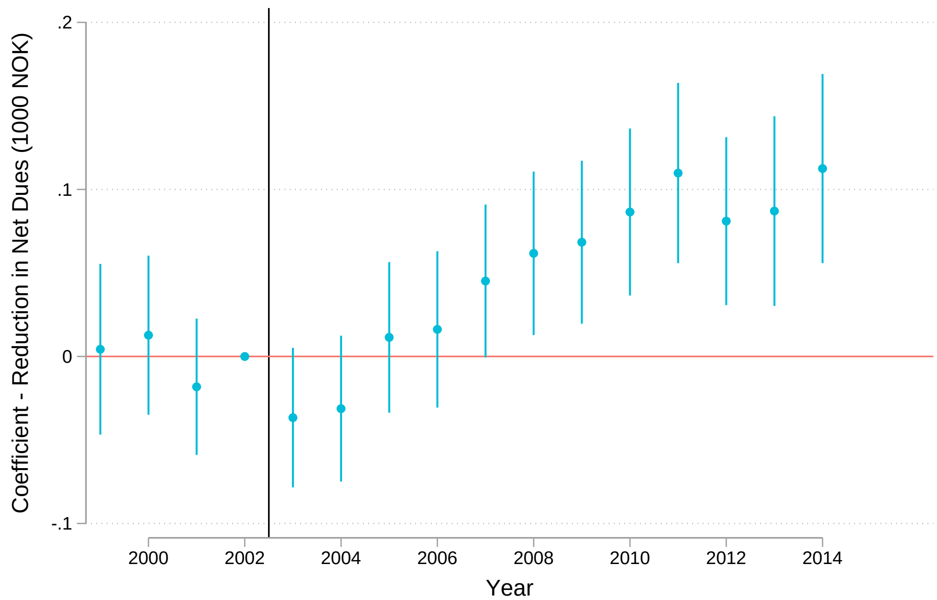
Source: Authors' illustration of base vs net union dues after tax deductions based on a 42% marginal tax rate. The vertical lines mark the value of the maximum tax deduction in particular years.

Figure A2: Statutory Maximum Deduction and Imputed Mean Subsidy for Union Dues (NOK)

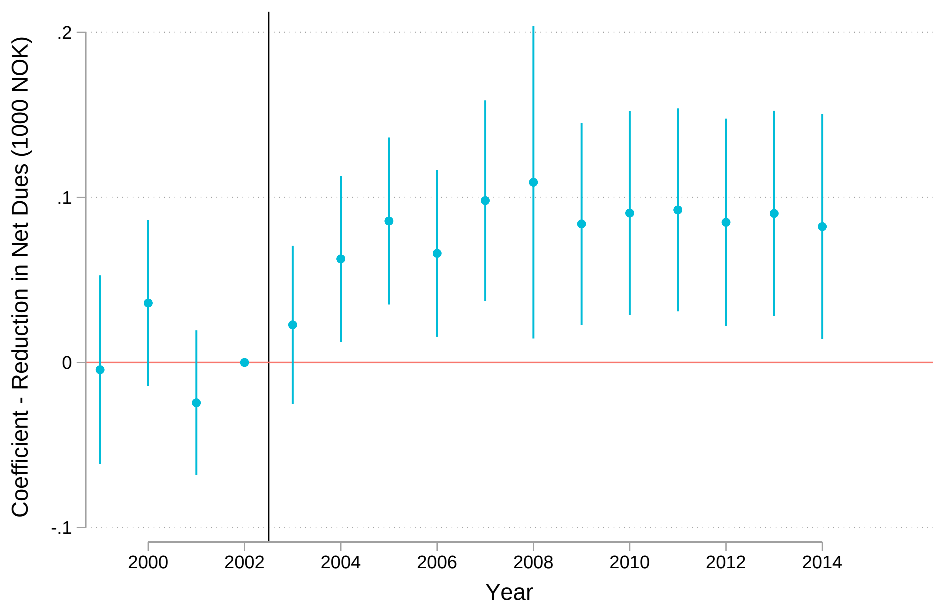


Source: Authors' calculations of Norwegian registry data.
Notes: Imputed base subsidies are calculated as the base tax rate times the lesser of imputed union dues at the occupation-by-industry cell or the statutory maximum deduction.

Figure A3: Union Density and Log Annual Earnings by Relative Subsidies
Panel A: Trends in Union Membership



Panel B: Trends in Log Annual Earnings

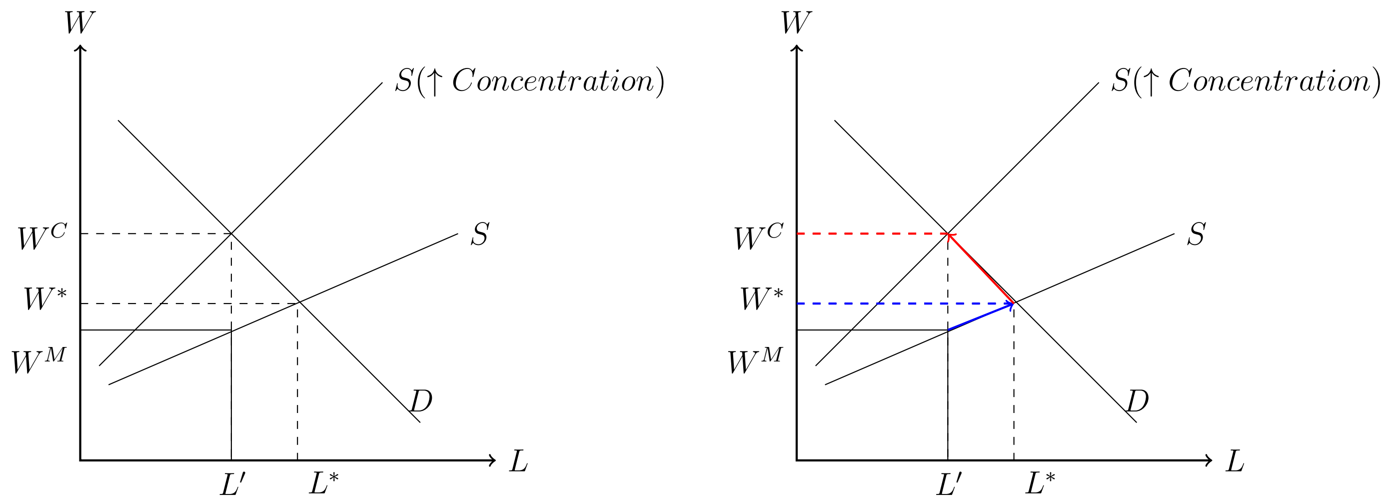


Source: Authors' calculations of Norwegian registry data.
Notes: Coefficients reflect the interaction with year dummies with the total reduction in net dues in the 2003-2010 period within each firm. Estimates include fixed effects for occupation by industry cell, educational program, age group, firm, and year. Changes in net dues are calculated within firms from 2003 to 2010 during the period of the largest shifts in the maximum tax deduction (see Figure A2). The maximum deduction was relatively stable from 2010 onward, where the gap between the high- and low-subsidy groups stabilized.

Figure A4: 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)	0.6345	0.4816
Firm Union Density	63.2869	26.6852
Real Annual Earnings (2015 NOK)	458,058	269,861
Age	41.89	11.63
Imputed Tax Subsidy (1,000s NOK)	0.7327	0.2887
Imputed Net Union Dues (1,000s NOK)	3.1591	0.5524
Labor HHI (20 Clusters)	0.0428	0.0538
Labor HHI (40 Clusters)	0.0512	0.0611
Labor HHI (2-Digit Occupation)	0.0646	0.0760
Labor HHI (3-Digit Occupation)	0.0948	0.1017
Product HHI (National Industry)	0.0370	0.0774
Public Sector Industry Worker	0.4316	0.4953
N	15,069,690	

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]	0.0660	0.0809	0.0539
	(0.0536)	(0.0530)	(0.0529)
Subsidy (1,000 NOK) * HHI (20 Clusters) [Base]	0.0397	0.0760*	0.0853
	(0.0734)	(0.0459)	(0.120)
HHI [Base]	0.690***	1.038***	0.438
	(0.264)	(0.244)	(0.404)
Subsidy * Group Interaction	0.00792	-0.0182***	0.00638
	(0.00516)	(0.00489)	(0.0115)
Subsidy * Group Interaction * HHI	0.110*	0.123***	0.0508
	(0.0587)	(0.0400)	(0.118)
Group * HHI	1.528***	0.863***	1.328***
	(0.232)	(0.144)	(0.430)
Female Indicator	-0.0114		
	(0.0164)		
Constant	0.370***	0.368***	0.368***
	(0.0464)	(0.0457)	(0.0460)
Observations	15,069,038	15,069,038	15,069,038
R-squared	0.253	0.253	0.251
F-Stat	118.6	174.8	45.33

Source: Authors' estimates of Equation 2 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 Dues

VARIABLES	(1) 20 Clusters	(2) 40 Clusters	(3) 2-Digit Occupation
Subsidy	0.225***	0.227***	0.218***
	(0.0550)	(0.0551)	(0.0546)
HHI	0.409***	0.424***	0.347***
	(0.0329)	(0.0295)	(0.0259)
Subsidy x HHI	0.0611	0.0117	0.0475
	(0.0425)	(0.0376)	(0.0322)
Observations	15,069,038	15,069,038	15,069,038
R-squared	0.250	0.250	0.250

Source: Authors' estimates of Equation 2 excluding the net union dues from the equation.
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 A4: The Effect of Tax Subsidies on Propensity to Unionize Using Subsidy Ratio

VARIABLES	(1) 20 Clusters	(2) 40 Clusters	(3) 2-Digit Occupation	(4) 3-Digit Occupation
Subsidy Ratio	0.104	0.103	0.0989	0.128
	(0.149)	(0.150)	(0.153)	(0.149)
HHI x Subsidy Ratio	0.501***	0.338**	0.502***	0.257**
	(0.150)	(0.135)	(0.127)	(0.100)
Observations	7,532,585	7,532,585	7,532,585	7,532,585
R-squared	0.252	0.252	0.252	0.253

Source: Authors' estimates of Equation 2 replacing the raw subsidy with a subsidy ratio and the net union dues with the inverse net union dues per Barth et al. (2020).
Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2. 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 Union Density on Earnings Using Workers' Base Firms to Calculate Subsidies

Panel A: Main Approach

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 2-Digit Occupation	(5) 3-Digit Occupation
Predicted Firm Union Density	0.00863***	0.00401***	0.00406***	0.00191	0.00343**
	(0.00163)	(0.00150)	(0.00149)	(0.00135)	(0.00137)
Predicted Firm Union Density * HHI		0.0129***	0.0126***	0.00823***	0.00544***
		(0.00111)	(0.00116)	(0.000996)	(0.000702)
Observations	15,069,035	15,069,036	15,069,037	15,069,038	15,069,039

Panel B: Main Approach, Private Sector Only

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 2-Digit Occupation	(5) 3-Digit Occupation
Predicted Firm Union Density	0.00200	-0.000205	-0.000305	-0.000564	-0.000751
	(0.00160)	(0.00162)	(0.00160)	(0.00157)	(0.00151)
Predicted Firm Union Density * HHI		0.0226***	0.0156***	0.00995***	0.00632***
		(0.00256)	(0.00265)	(0.00196)	(0.00123)
Observations	8,564,794	8,564,794	8,564,794	8,564,794	8,564,794

Source: Authors' estimates of Equation 2 replacing the subsidy and net dues with those based on each worker's first firm in the data rather than their current firm's first year in the data.

Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2. 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 A6: The Effect of Union Density on Log Annual Earnings by Labor Market Concentration, Including Local Labor Market Controls

			Panel A: Adding LLM Indicators			
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) No HHI - Private	(5) 20 Clusters - Private	(6) 40 Clusters - Private
Predicted Union Density	0.0206*** (0.00426)	0.0136*** (0.00340)	0.0123*** (0.00331)	0.0153*** (0.00393)	0.0140*** (0.00434)	0.0248** (0.0100)
Predicted Union Density * HHI		0.0193*** (0.00502)	0.0251*** (0.00460)		0.0370*** (0.0110)	0.0248** (0.0100)
Observations	15,031,288	15,031,288	15,031,288	8,543,965	8,543,965	8,543,965
			Panel B: Cluster by Year FE (20 Clusters)			
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) No HHI - Private	(5) 20 Clusters - Private	(6) 40 Clusters - Private
Predicted Union Density	0.0160*** (0.00241)	0.0134*** (0.00242)	0.0133*** (0.00246)	0.0130*** (0.00238)	0.0134*** (0.00286)	0.0131*** (0.00280)
Predicted Union Density * HHI		0.00616** (0.00288)	0.00917*** (0.00266)		0.0245*** (0.00631)	0.0133** (0.00601)
Observations	14,591,487	14,591,487	14,591,487	8,348,491	8,348,491	8,348,491
			Panel C: Controlling for LLM Predicted Union Density			
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) No HHI - Private	(5) 20 Clusters - Private	(6) 40 Clusters - Private
Predicted Firm Union Density	0.0200*** (0.00369)	0.0144*** (0.00315)	0.0140*** (0.00321)	0.0150*** (0.00344)	0.0151*** (0.00416)	0.0141*** (0.00400)
Predicted Firm Union Density * HHI		0.0195*** (0.00474)	0.0210*** (0.00432)		0.0386*** (0.0104)	0.0220** (0.00886)
Observations	15,069,035	15,069,035	15,069,035	8,564,794	8,564,794	8,564,794
			Panel D: Controlling for LLM-Cluster Predicted Union Density (20 Clusters)			
VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) No HHI - Private	(5) 20 Clusters - Private	(6) 40 Clusters - Private
Predicted Firm Union Density	0.0186*** (0.00369)	0.0134*** (0.00313)	0.0133*** (0.00322)	0.0148*** (0.00356)	0.0164*** (0.00433)	0.0154*** (0.00417)
Predicted Firm Union Density * HHI		0.0157*** (0.00468)	0.0192*** (0.00432)		0.0329*** (0.0107)	0.0199** (0.00914)
Observations	15,069,035	15,069,035	15,069,035	8,564,794	8,564,794	8,564,794

Source: Authors' estimates of Equation 3 with the addition of fixed effects and controls.
Notes: Standard errors are in parentheses and are clustered at the firm level and adjusted for model uncertainty in Equation 2. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year, with additional dummy variables as mentioned in each panel. Panels C and D use Equation 2 to predict local labor market-level (and LLM-cluster) union density, which are then included as an additional control in the model.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A7: The Effect of Union Density on Separation Rates and Education Composition

	Panel A: Separations, Quartiles of Annual Earnings in Firm		
	(1)	(2)	(3)
VARIABLES	Pr(Separation)	Pr(Separation*P75)	Pr(Separation*P25)
Union Density	-0.00188	0.000206	0.000112
	(0.00276)	(0.000894)	(0.000856)
Union Density x HHI	-0.00574	-0.00384***	0.00245*
	(0.00433)	(0.00144)	(0.00147)
Dep Var Mean	0.232	0.049	0.083
Observations	15,069,035	15,069,035	15,069,035

	Panel B: Separations, Education		
	(1)	(2)	(3)
VARIABLES	Pr(Separation*LTHS)	Pr(Separation*HS)	Pr(Separation*Bach+)
Union Density	0.000975**	0.00253	-0.00538***
	(0.000484)	(0.00158)	(0.00164)
Union Density x HHI	-0.000292	-0.00720***	0.00175
	(0.000911)	(0.00253)	(0.00230)
Dep Var Mean	0.042	0.099	0.088
Observations	15,069,035	15,069,035	15,069,035

	Panel C: Education		
	(1)	(2)	(3)
VARIABLES	Pr(LTHS)	Pr(HS)	Pr(Bach+)
Union Density	-0.00343***	0.0201***	-0.0155***
	(0.00109)	(0.00274)	(0.00249)
Union Density x HHI	0.00405**	-0.0173***	0.0151***
	(0.00193)	(0.00402)	(0.00374)
Dep Var Mean	0.159	0.438	0.394
Observations	15,219,857	15,219,857	15,219,857

Source: Authors' estimates of Equation 2 and 3 in the form of a linear probability model for each outcome.
Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2. Estimates include fixed effects for occupation by industry cells, age group, and year. Panels A and B include controls for detailed educational program. HHI is defined in these models at 20 skill clusters.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A8: 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	4720	0.7269	3.1935	0.0333
Women	4269	0.7383	3.1262	0.0520
Below Occ-Firm Median	4008	0.7284	3.1559	0.0426
Above Occ-Firm Median	4939	0.7378	3.1629	0.0431
Not White Collar	5104	0.7234	3.4135	0.0325
White Collar	4306	0.7351	3.0931	0.0455

Source: Authors' estimates using Norwegian register data.

Table A9: Heterogeneous Effects by Bindingness of Industry Floors

	Log Annual Earnings	
VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Predicted Firm Union Density	0.00138	0.00161
	(0.00167)	(0.00183)
Union Density x Kaitz	0.000397***	0.000405***
	(6.46e-05)	(6.95e-05)
Union Density x HHI	0.0188***	0.0198***
	(0.00347)	(0.00345)
Union Density x Kaitz x HHI	0.000339	0.000164
	(0.000459)	(0.000447)
Observations	14,431,060	14,431,060

Source: Authors' estimates of Equation 3 with the addition interactions with the Kaitz index.
Notes: Standard errors are in parentheses and are clustered at the firm level and adjusted for model uncertainty in Equation 2. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year. The Kaitz index for the industry wage floor is calculated in two steps. First, we calculate the kernel density function for annual earnings for each 2-digit industry by year cell using Epanechnikov weights. Second, we examine the density and assume the industry earnings floor is the first time the density increases by at least 0.00003 percent, which is approximately a 10-15% increase in the density at the mean. The Kaitz index is the ratio of this earnings floor to the median annual earnings in the industry-year cell. In Panel C of Table 2, we saturate our models with 2-digit by year interacted dummies, which subsumes these industry wage floors and yields similar results.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A10: Heterogeneous Effects by Measures of Social Capital

Panel A: Pr(Union)	Share of Workers with Same Birth County		Occupational Diversity	
	(1)	(2)	(3)	(4)
VARIABLES	20 Clusters	40 Clusters	20 Clusters	40 Clusters
Subsidy (1,000 NOK)	0.0584	0.0612	0.117**	0.117**
	(0.0511)	(0.0512)	(0.0467)	(0.0467)
Subsidy x Soc Capital	0.0120	0.0123	-0.0384	-0.0339
	(0.0170)	(0.0175)	(0.0393)	(0.0398)
Subsidy x HHI	0.177**	0.0991	0.369***	0.316***
	(0.0865)	(0.0751)	(0.0780)	(0.0661)
Subsidy x HHI x Soc Capital	-0.123	-0.0650	-1.168**	-1.084***
	(0.154)	(0.138)	(0.493)	(0.401)
HHI	1.457***	1.386***	1.401***	1.269***
	(0.249)	(0.156)	(0.240)	(0.154)
Observations	13,748,269	13,748,269	15,069,038	15,069,038
F-Statistic	53.11	60.83	310.1	320.2

Panel B: Log Annual Earnings	Share of Workers with Same Birth County		Occupational Diversity	
	(1)	(2)	(3)	(4)
VARIABLES	20 Clusters	40 Clusters	20 Clusters	40 Clusters
Predicted Firm Union Density	0.00703***	0.00607***	0.00124***	0.00119***
	(0.00173)	(0.00171)	(0.000277)	(0.000274)
Union Density x Soc Capital	0.000204***	0.000220***	0.00406***	0.00394***
	(3.92e-05)	(3.96e-05)	(0.000646)	(0.000644)
Union Density x HHI	0.0225***	0.0270***	0.0173***	0.0161***
	(0.00336)	(0.00344)	(0.00131)	(0.00149)
Union Density x Soc Capital x HHI	0.000217	-7.96e-05	-0.0201***	-0.0140***
	(0.000282)	(0.000261)	(0.00267)	(0.00263)
Observations	13,748,264	13,748,264	15,069,035	15,069,035

Source: Authors' estimates of Equation 3 with the addition interactions with measures of social capital.
Notes: Standard errors are in parentheses and are clustered at the firm level and adjusted for model uncertainty in Equation 2. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year. Our measure of same birth county is the share of workers in each firm in each year that share a county of birth in Norway. Those born outside of Norway are characterized as being born in the "same county" if they are both foreign-born. Occupational diversity is defined as the number of distinct 3-digit occupations within the firm-year cell per worker in the firm.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and *** indicates significance at the 1% level.

Table A11: The Effect of Predicted Labor Concentration on Union Premium, Fixed Union Density

VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Base Union Density x Predicted HHI	0.0443***	0.0525***
	(0.0121)	(0.0143)
Predicted HHI	-2.836***	-3.362***
	(0.700)	(0.830)
Observations	9,936,364	9,936,364
R-squared	0.556	0.556

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 A12: 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)	0.0595	0.115***	0.0762	0.127***
	(0.0527)	(0.0200)	(0.0526)	(0.0201)
HHI x Subsidy	0.132***	0.262***	0.0525**	0.136***
	(0.0363)	(0.0201)	(0.0261)	(0.0161)
HHI	1.518***	0.210***	1.117***	0.178***
	(0.179)	(0.0598)	(0.121)	(0.0455)
Observations	15,069,038	14,883,635	15,069,038	14,883,635
R-squared	0.251	0.750	0.252	0.750
Occupation x Industry FE	Yes	Yes	Yes	Yes
Education FE	Yes	Yes	Yes	Yes
Age Group FE	Yes	Yes	Yes	Yes
Individual FE		Yes		Yes

Source: Authors' estimates as described in the text. Estimates correspond with Equation 2.
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 A13: 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	0.0193***	0.00601***	0.00825***
	(0.00257)	(0.00210)	(0.00239)
Predicted Firm Union Density * HHI		0.0111***	0.00861***
		(0.00352)	(0.00239)
Observations	15,069,035	15,069,035	15,069,035
R-squared	0.576	0.576	0.576

Source: Authors' estimates as described in the text. Estimates correspond with Equation 3.
Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2. 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 - Private Sector Only, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.0141***	0.0119***	0.00990***
	(0.00349)	(0.00370)	(0.00344)
Predicted Firm Union Density * HHI		0.00183	0.00594
		(0.00753)	(0.00424)
Observations	8,564,794	8,564,794	8,564,794
R-squared	0.599	0.599	0.599

Source: Authors' estimates as described in the text. Estimates correspond with Equation 3.
Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2. 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 Firm-Occupation Median, Occupation-Specific HHI

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.00351	-0.00641***	-0.00510**
	(0.00278)	(0.00189)	(0.00204)
Predicted Firm Union Density * HHI		0.0262***	0.0183***
		(0.00303)	(0.00207)
Union Density * Above Firm-Occ Median	0.00597***	0.00629***	0.00645***
	(4.36e-05)	(5.39e-05)	(5.49e-05)
Union Density * HHI * Above Firm-Occ Median		-0.00447***	-0.00465***
		(0.000307)	(0.000226)
Observations	15,069,035	15,069,035	15,069,035
R-squared	0.712	0.712	0.713

Source: Authors' estimates corresponding with Equation 3 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2 with subgroup interactions. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year.
* indicates significance at the 10% level, ** indicates significance at the 5% level, and
*** indicates significance at the 1% level.

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

VARIABLES	(1) No HHI	(2) 2-Digit Occupation	(3) 3-Digit Occupation
Predicted Firm Union Density	0.00583***	0.00122**	0.000688
	(0.00108)	(0.000609)	(0.000650)
Predicted Firm Union Density * HHI		0.0165***	0.0143***
		(0.00244)	(0.00207)
Union Density * White Collar	0.00238***	0.000738	0.00164***
	(0.000418)	(0.000460)	(0.000448)
Union Density * HHI * White Collar		0.000545	-0.000334
		(0.000461)	(0.000358)
Observations	15,069,035	15,069,035	15,069,035
R-squared	0.576	0.576	0.576

Source: Authors' estimates corresponding with Equation 3 with interactions by subgroup.
Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2 with subgroup interactions. 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 A17: The Effect of Exposure to Chinese Imports on Labor Concentration

VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Exposure to Chinese Imports per Worker (1,000s NOK)	-2.03e-06***	-1.71e-06***
	(1.90e-07)	(2.35e-07)
SD of HHI (full sample):	0.0538	0.0611
SD effect of 1 million NOK	-0.0377	-0.0280
Observations	124,068	124,068
R-squared	0.832	0.813
Firm Fixed Effects	Yes	Yes

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

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

VARIABLES	(1) 20 Clusters	(2) 40 Clusters
Union Density	-0.000724*	-0.00151**
	(0.000418)	(0.000620)
Union Density x Predicted HHI	0.0444***	0.0526***
	(0.0116)	(0.0137)
Predicted HHI	-2.934***	-3.477***
	(0.701)	(0.831)
Observations	9,936,364	9,936,364
R-squared	0.556	0.556

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

Table A19: The Effect of Union Density on Earnings Using Leave-one-out Imputed Dues to Calculate Subsidies

VARIABLES	(1) No HHI	(2) 20 Clusters	(3) 40 Clusters	(4) 2-Digit Occupation	(5) 3-Digit Occupation
Predicted Firm Union Density	0.0318***	0.0173***	0.0192***	0.00724***	0.00819***
	(0.00326)	(0.00216)	(0.00237)	(0.00161)	(0.00175)
Predicted Firm Union Density * HHI		0.0345***	0.0416***	0.0140***	0.0185***
		(0.00874)	(0.0112)	(0.00500)	(0.00571)
Observations	15,069,035	15,069,035	15,069,035	15,069,035	15,069,035

Source: Authors' estimates of Equation 2 and 3 using occupation and industry mean dues excluding a worker's own firm to calculate base imputed dues and subsidies.

Notes: Standard errors are in parentheses and are clustered at the firm level and are adjusted for estimation uncertainty in Equation 2. 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.

B Alternative Measure of Market Power: Separation Elasticities

The recent literature on labor market power has shown that labor market concentration—as measured by HHI—has negative effects on the wages and employment rates of workers (e.g. Dodini et al. (2024); Azar et al. (2020a); Marinescu et al. (2021)). It is clear from this literature that concentration is capturing some important aspects of labor market power. While our measure of HHI is fixed for each firm at a particular point in time and is thus unaffected by evolving equilibrium changes in labor demand over time and space, HHI is not the only proxy for labor market power we can use. In an extension of our approach using HHI, we use separation elasticities for firms' wage policies for particular occupation groups to proxy as a measure of labor market power where separation elasticities closer to zero (inelastic) are an indication of the firm's market power over that particular occupation group.

To measure these elasticities, we follow a modified version of the Bassier et al. (2022) approach to measuring separation elasticities, which uses an AKM (Abowd et al., 1999) model to establish firms' wage policies and subsequently measures separation rates in response to these firm-level effects. To facilitate comparisons to our HHI measures, which rely on occupation definitions and skill cluster groups of occupations, we make two modifications. First, because we have data on each worker's occupation, we estimate the AKM model with fixed effects for each firm interacted with the worker's skill cluster, which allows firms to have different wage policies for each occupation group.[38] Second, we also control for individual by occupation group match effects. Thus, our AKM model more tightly controls for individual-by-occupation group matches than the standard AKM as well as capturing the fixed effects representing firm-skill-cluster wage policies.

Figure B1 presents the binned scatterplot relationship between a firm-cluster fixed effect and the separation rate from the firm for a worker relative to the overall mean rate.[39] The pattern is similar to the prior literature using this approach, though the implied elasticities are smaller than the shorter-run job transitions measured in Bassier et al. (2022) when looking at firm-level fixed effects in their split sample approach. Overall, the implied average separation elasticity is close to -1.

For our purposes, what matters most is not the magnitude of the overall separation rate per se—though it is interesting to note the relatively low elasticity when examining this level of firm-occupation wage policy and annualized job transitions—but rather how changes in union density interact with relative differences in the separation elasticities across firms. To best mirror our concentration-based measure of market power, we separately estimate separation elasticities for each local labor market-occupation group (cluster) cell for the sample period and then divide these by the sample overall separation probability. This yields a normalized separation elasticity for each LLM-occupation group. As we do in the case of concentration, we fix this measure of market power for each firm by taking the average elasticity for all LLM-occupation group cells within the firm and holding this fixed at the firm's first year in the data. This yields a single time-invariant measure of a firm's market power that we can then interact with changes in their union density. This approach also most closely matches our construction of the HHI interactions in our main approach and allows us to abstract away from any endogenous changes in the occupational composition of the firm over time.

Panel A of Table B1 shows the result of estimating Equation 2 but replacing the HHI values with separation elasticities for each occupation group definition. Panel B shows the analog of Equation 3 using the same measure. Similar to the case of measured market concentration, separation elasticities closer to zero (a positive change in the separation elasticity) are associated with a higher probability of joining a union overall and a greater sensitivity to the subsidy, though this interaction is not statistically significant. Carrying forward the predicted firm-level union density from the estimates into Panel B, we show that as separation elasticities

[38] We also estimate these with 2-digit occupation effects as an additional robustness check.

[39] Following Bassier et al. (2022), we trim the top and bottom 2 percentiles to avoid significant outliers.

move closer to zero (i.e. firms possess greater market power), measured earnings returns to union density increase. Thus, our findings with regard to labor market concentration are supported by another measure of labor market power from the literature.

Figure B1: Separation Rate vs Firm-Skill Cluster Fixed Effects

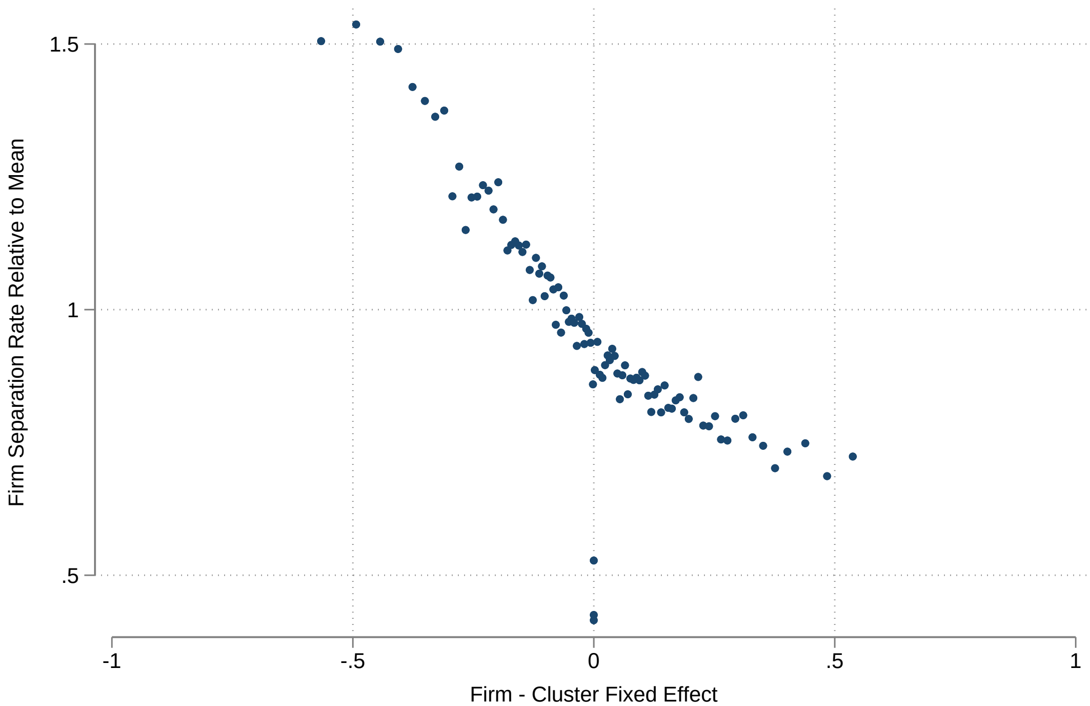


Table B1: Using Separation Elasticities as a Measure of Market Power

	Panel A: Pr(Union)	
VARIABLES	(1) 20 Clusters	(2) 2-Digit Occupation
Subsidy (1,000 NOK)	0.104*	0.0878*
	(0.0533)	(0.0528)
Separation Elasticity (<0)	0.248***	0.184***
	(0.0473)	(0.0413)
Elasticity x Subsidy	0.0110	-0.0121
	(0.0119)	(0.0112)
Observations	15,013,782	15,013,980
F-Statistic	16.19	19.91

	Panel B: Log(Earnings)	
VARIABLES	(1) 20 Clusters	(2) 2-Digit Occupation
Predicted Firm Union Density	0.0234***	0.0218***
	(0.00355)	(0.00470)
Union Density * Separation Elasticity (<0)	0.00463**	0.00428**
	(0.00192)	(0.00201)
Observations	15,013,779	15,013,977

Source: Authors' estimates as described in the text. Estimates correspond with Equation 3 but replacing measured HHI with measured separation elasticities.

Notes: Standard errors are in parentheses and are clustered at the firm level and adjusted for estimation uncertainty in Equation 2, but replacing HHI with the separation elasticity. Estimates include fixed effects for occupation by industry cells, detailed educational program, age group, firm, and year. Separation elasticities are less than zero, so higher numbers relate to lower separation elasticities, i.e. greater firm market power.

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

C 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, 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 can be completely offset by the positive wage effect of union monopoly power. 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.

C.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 A11 indicates that holding unionization constant at the firm's 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:

$$Exposure_{lt} = \frac{1}{L_0^l} \sum_f \frac{L_0^{fil}}{L_0^i} \Delta M_{it}, \tag{5}$$

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:

$$HHI_{ft} = \alpha_0 + \alpha_1 Exposure_{lt} + \tau_t + \phi_f + \nu_{ft}, \tag{6}$$

where all variables are defined as above.

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

$$\begin{aligned} Log(Earnings)_{iocft} =& \alpha_0 + \alpha_1 UD_{ft} + \alpha_2 UD_{ft} * \widehat{HHI}_{ft} + \alpha_3 \widehat{HHI}_{ft} \\ &+ \delta_{Ed} + \pi_{Age} + \gamma_{oc} + \tau_t + \phi_f + \eta_{iocft}, \end{aligned} \tag{7}$$

In Equation 7, 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 A11 and lead to the same conclusions as our main approach.[40]

C.2 Results

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

The results in Table A17 suggest that exposure to Chinese exports has a small but highly statistically significant impact on the labor market concentration experienced by firms. Specifically, an increase in exposure to Chinese imports per worker of 100,000 NOK (approximately 12,000 USD) reduces the HHI of the firm by approximately 0.37 percent of a standard deviation. For firms at the top of the exposure distribution, with an exposure of approximately 2 million NOK per worker, the predicted effects would be nearly seven percent of a standard deviation. The F-statistics associated with the regressions underlying the results are 116 and 50, respectively. These statistics are significantly greater than the conventional rule-of-thumb values.

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

The results in Table A18 also demonstrate that the negative impact of labor market concentration is

[40] 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 A11 indicate they do not affect our conclusions.

considerably smaller in highly unionized firms. A one percentage point increase in union density increases wages by approximately 4.4 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 66 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 66 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 with at least 10 workers in the Norwegian economy to reach this threshold. At the end of our sample period, approximately 43% of all workers were at firms with a predicted union density below this 66% threshold, representing 49% of all firms with at least 10 workers. On average, firms below the tipping point have predicted densities approximately 2.9 percentage points away from 66%. Generating a 2.9 percentage point change in unionization, according to our subsidy effect estimates, would require an increase in the base tax subsidy of approximately 285 NOK, or raising the deduction by approximately 889 NOK. This would induce approximately 14,500 new workers to join a union at a base cost of 4.15 million NOK. Holding constant the union membership status of those in firms already above 66%, a universal tax subsidy increase of 285 crowns per member would also result in additional payments to approximately 646,000 full-time workers totaling 170 million NOK, for a total new base subsidy value of approximately 190 million NOK (approximately $24.5 million).[41] Given the size of the workforce in our sample (approximately 1.14 million workers at these firms at the end of the sample), this amounts to a transfer of approximately 167 NOK per worker per year for the base subsidy. Furthermore, at the average labor market concentration in our sample of firms that are below 66% predicted union density, a 2.9 percentage point increase in unionization would also induce these firms to increase the share of workers above 30 hours by 1.96 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.[42]

[41] 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.

[42] Without a full analysis of the incidence of corporate taxation on labor, the effect of lower monopsony rents on corporate profits, changes in prices, 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.

Survey of Norwegian Workers

[INTRO1] This is a survey that Norstat conducts on behalf of the Norwegian School of Economics and Business Administration. The results will be used in a research project.

All information collected through the survey is anonymized and will not be disclosed to any third party. As part of scientific publishing, anonymised data may be shared in open scientific repositories.

If you want more information about the project, you can choose the option below. If you want to start the survey, you choose it.

[R1] I want more information
[R2] I want to start the survey

[R1] Information and declaration of consent

Purpose of the project

We want to understand how individuals in Norway value their work environment and how they view unions. The results of the study will increase our understanding of workplace preferences and their relative importance.

Who is responsible for the project?

The Norwegian School of Economics (NHH) is the responsible institution for the project. Alexander Willen, professor at NHH, is the project manager. The other project members are Kjell G. Salvanes, professor at NHH, Samuel Dodini, postdoctoral fellow vid NHH, and Julia Zhu, postdoctoral fellow at NHH. If you have any questions about the project, you can contact NHH via Alexander Willen (alexander.willen@nhh.no).

What does participation mean for you?

If you choose to participate in the project, you will be asked to answer a survey by completing an online questionnaire. It takes about 7 minutes. The survey includes questions about your work situation, union status, and your job preferences. In addition, we will ask some basic demographic questions about, for example, age and gender. Participation in the survey is voluntary and you can withdraw your consent at any time without giving any reason. All information collected through the survey is anonymized and will not be disclosed to any third party. As part of scientific publishing, anonymised data may be shared in open scientific repositories. There will be no negative consequences if you choose not to participate or decide to withdraw at a later date.

Declaration of consent

I have received and understood information about the survey and hereby consent:

• to participate in the online survey.
• to enable researchers to process my anonymised data and use them for publications in scientific journals and other scientific dissemination.

[R2] Survey

[Age] What is your age?
[Gender] Are you male or female?
[Zip code] What is your zip code?
[Fylke] Which county do you live in?

What is your highest completed education?
[R1] Primary school/primary school
[R2] Upper secondary school (incl. former vocational school)
[R3] Vocational school, trade certificate/journeyman's certificate and other 1-2 year education after upper secondary school
[R4] University/college up to 3 years (Bachelor's degree)
[R5] University/college 4 years or more (Master's degree and higher)
[R98] Other

Where were you born?
[R1] Norway
[R2] Outside Norway
[R3] Don't want to answer

Can you state which country you were born in?

At what age did you move to Norway?

How many years of full-time work experience do you have?

Are you currently in part-time or full-time work?
[R1] Part-time (less than 30 hours per week)
[R2] Full-time (at least 30 hours per week)
[R3] Not working

What industry is your main job in?

Do you work in the public or private sector?
[R1] Public sector
[R2] Private sector

How many people work at your workplace?
Row:
[R1] 1-5
[R2] 6-10
[R3] 11-50
[R4] 51-100
[R5] More than 100
[R6] Don't want to answer

Rank the following job characteristics based on importance to your future career and well-being: Salary, Job Safety, Promotion Potential and Work Environment Quality.
Here we ask you to award 100 points across the four categories. You can assign anything between 0 and 100 to any of the categories, as long as the total amount of points for all four categories is 100.
Row:
[R1] Salary: Everything associated with the financial payment of your work (base salary, bonuses, overtime pay, generosity with retirement plans, etc.)
[R2] Job security: Protection and support (legal and otherwise) against being laid off and fired, both in the event of mass closures and individual layoffs (wrongful or not)
[R3] Promotion potential: Potential to move up the career ladder in the company
[R4] Work environment quality: The day-to-day quality of your work environment, including physical environment (e.g. equipment and facilities), company culture (e.g. support, feedback, collaboration, potential to influence) and working conditions (e.g. workplace safety, conditions employment, work-life balance)

Are you a member of a trade union?
[R1] Yes
[R2] No
[R3] Don't want to answer

For how many years have you been a member?

Have you been a member continuously during that time, or have you changed in and out of membership over the years?
[R1] Continuous
[R2] Not continuously

How important do you think the union is to improving your pay, job security, promotion potential and work environment quality?
0 means "not at all" and 100 means "entirely". The total for all four need NOT be 100.
[R1] Monetary compensation
[R2] Job security
[R3] Promotion potential
[R4] Working environment quality

Compared to members, the extent to which do you think nonmembers in your workplace can benefit from the presence of unions along these four dimensions
0 means "not at all" and 100 means "complete". The total for all four need NOT be 100.
[R1] Monetary compensation
[R2] Job security
[R3] Promotion potential
[R4] Working environment quality

Have you found a union membership useful for receiving non-work benefits such as lower mortgage rates, access to cheaper/better insurance, etc.?

How important has this been for your decision to join a union?

If your after-tax dues for union membership increased by [XYZ] dollars, would you reconsider the decision to join a union?
Row:
[R1] Yes
[R2] No

The purpose of this question is to understand the reason why you do not join a union. Check all the boxes that apply.
Row:
[R1] I don't want to spend so much money being a union member
[R2] I don't think unions can affect my work situation
[R3] I find that unions focus on dimensions of the workplace that are not important to me.
[R4] I don't think I need to be a member of a union to take advantage of the influence unions have on my work situation and well-being
[R5] Other reason, note:

If your after-tax dues for union membership were reduced by [XYZ] NOK, would you reconsider your decision to join a union?
Row:
[R1] Yes
[R2] No

Second Survey

Our second survey of union members in Norway recruited 960 respondents. Below are the questions asked in the survey that inform the analysis in the text:

How do you think union members are treated at your workplace by the local union?

[R1] I think all union members are treated equally by the local union
[R2] I think some members are prioritized over other members by the local union

[If answer R2 in last question]
Who do you think is being prioritized by your local union?

[Open text response]