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ABSTRACT

We study the labor market impacts of unions by accounting for their effects on employers' insurance provision and examine how social insurance policies, in turn, affect unionization and labor market outcomes. We document that unions increase employer-sponsored insurance provision and that expansions in social insurance reduce unionization in the United States. We then develop and estimate an equilibrium labor search model where unionization, wages, and non-wage benefits are endogenously determined. We demonstrate that unionization, as well as the threat of unionization, increases employer-sponsored insurance provision in both unionized and nonunionized firms. We find that social insurance policies can affect labor market inequality through (de)unionization, and inequality may increase or decrease depending on how social insurance is targeted. Social insurance expansions, along with technological changes, contribute to the long-term decline of unions in the U.S. in the last fifty years. Although historical deunionization driven by these forces increased overall welfare, we find that subsidizing unions in the current economy can improve welfare.

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1 Introduction

Labor unions in the United States have steadily declined over the past few decades. In 1955, about 36% of the private sector workers were unionized; today, it is less than 10%. This trend has renewed interest in the effects of unions on labor markets and led to various policy interventions. Although most of the research focuses on the effects of unions on wages, inequality and labor demand, unions also influence workers by increasing employers' provision of insurance benefits and job security (Freeman and Medoff, 1984). Indeed, unions' insurance effect is highlighted as one of the primary reasons to promote unions and collective bargaining in the recent Executive Order on April 26, 2021 (E.O. 14025) by the Biden administration.

However, the insurance role of unions is deeply intertwined with the design of social insurance systems. If union-negotiated employer-provided insurance benefits are a primary reason for workers to join unions, the availability of affordable insurance options outside of the unions, through the government or other sources, can dilute the value of union membership. Consequently, social insurance expansion can reduce the unionization rate, and, in turn, affect labor market inequality and various other labor market outcomes.[1] This channel is particularly relevant in the U.S., where, unlike in many European countries, employers determine access to many essential insurance benefits, and union formation is decided at the employer level. Thus, understanding the equilibrium and welfare effects of unions requires a detailed examination of the mechanisms through which unions interact with social insurance systems.

In this paper, we study how labor unions shape employers' insurance provision and how social insurance policies, in turn, affect unionization and labor market outcomes. We begin by documenting empirical evidence on the relationship between unionization, employer-provided insurance, and social insurance expansions. We then develop and estimate an equilibrium model of labor unions that incorporates the roles of unions in wage compression and insurance provision. Using the model, we study several central questions regarding unions and social insurance, including the mechanisms through which unions affect wages and non-wage benefits, the effects of social insurance policies on unionization and wage inequality, the forces underlying the long-run decline in unionization, and the welfare consequences of union decline and policies that affect unionization.

We first document how unionization relates to insurance provision and how social insurance affects unionization. We show that unionized firms tend to be larger and more likely to provide various non-wage benefits, including employer-sponsored health insurance (ESHI), and to offer

[1]The connection between social insurance and the labor movement is well known. Otto von Bismarck, the German Chancellor in the late 19th century, tried to undermine socialist organizations and trade unions when in 1889 he introduced the world's first old age social insurance program to "promote the well-being of workers, and to stave off calls for more radical socialist alternatives." See US Social Security Administration.

better job security, consistent with the findings in Freeman and Medoff (1984). We also provide evidence suggesting that an important part of the workers' valuation of unions derives from the provision of union-negotiated non-wage benefits. Next, by exploiting the introduction of Medicare and Medicaid in the 1960s, we show that both programs reduced unionization rates. Exploiting variations across states and over time in Medicaid expansion under the Affordable Care Act (ACA) and in the generosity of state unemployment insurance (UI) benefits, we further show that more recent expansions of social insurance programs also lowered unionization rates.

To interpret the above evidence and study the aggregate and welfare implications of union decline, social insurance expansions, and union policies, we develop a model of labor unions that builds on the standard search-and-matching framework of Pissarides (2000), where search frictions give firms monopsony power in the labor market and, thus, create a role for unions as a countervailing force. Our model jointly incorporates two key ingredients. First, following Taschereau-Dumouchel (2020), we endogenize firm size and union formation, where unionization in each firm reflects the endogenous preferences of the workers for collective representation. In unionized firms, wages are determined through collective bargaining between the union and the firm, whereas in nonunionized firms, each worker engages in individual bargaining with the firm. Second, our model endogenizes the provision of non-wage benefits and job security. Firms choose whether to provide non-wage benefits in order to attract workers, following the framework of Hwang et al. (1998) and Aizawa and Fang (2020). We show that unions increase the provision of non-wage benefits by facilitating more efficient sharing between firms and workers of the cost of benefit provision.

We estimate the model using microdata on union status, labor market outcomes, demographics, and non-wage benefit provision. Motivated by our empirical evidence and the fact that health care accounts for a sizable part of the U.S. economy (Hall and Jones, 2007), we consider health insurance the main non-wage benefit in our empirical specification and model various health insurance programs. The estimated model successfully accounts for the relationship among unionization status, insurance provision, wages, worker skills, and firm size.

Importantly, our estimates indicate that the threat of union formation significantly increases employers' insurance provision, even among nonunionized firms. This insight is revealed through a counterfactual experiment in which union threat costs were removed (i.e., firms are not constrained by workers' preferences regarding unionization); we find that because low-skill workers tend to favor unionization, nonunionized firms would increase their hiring of low-skill workers to economize on the union threat cost in this counterfactual, which in turn results in a substantial decline in the insurance provision of nonunionized firms. The reason is that nonunionized firms and workers engage in individual bargaining where the firm splits the surplus with each individual worker separately; individual bargaining makes it difficult to share insurance costs efficiently among heterogeneous workers, thus lowering the insurance provision in nonunionized firms as they hire more low-skill workers in the counterfactual.

Using the estimated model, we analyze how social insurance policies affect unions and labor market outcomes. Policies that increase insurance provision outside of union jobs, such as subsidies or mandates for employer-provided insurance and public programs such as Medicaid and Medicare, reduce workers' incentives to seek union jobs for insurance access, and thereby lower the unionization rate. Because unions also compress wages, these policies can affect wage inequality in equilibrium. For example, we find that subsidies for non-wage benefits, such as the tax exemption of ESHI premiums,[2] reduce unionization as nonunionized firms expand insurance provision, and thus increase wage inequality. At the same time, the effects of social insurance expansions on labor market outcomes depend on how they are targeted. We find that expanding public health insurance only to low-skill unemployed workers, akin to a significant expansion of Medicaid, lowers the unionization rate but raises average wages and reduces wage inequality, although it also decreases insurance coverage among high-skill workers.

Next, we examine the extent to which social insurance expansions help explain the long-run decline, beginning in the mid-1950s, of private-sector labor unions in the U.S. Existing explanations for this decline typically emphasize skill-biased technological change (Acemoglu et al., 2001), which shifted labor demand away from low-skill workers who tend to favor unionization, and the implementation of right-to-work (RTW) laws (Farber, 2005), which made unions less sustainable by allowing workers to opt out of paying union dues. At the same time, social insurance programs expanded substantially in the U.S. We extend our model to incorporate the existing factors *and* the expansion of various social insurance programs. Through simulations, we find that technological change and the implementation of RTW laws account for about 32% and 1% of the observed union decline between 1955 and 2007, respectively; interestingly, we find that social insurance expansions through the introduction and expansion of multiple health insurance programs contributed to about 16% of the overall decline in that time period.

Our framework also allows us to evaluate the welfare implications of union decline. We find that declines in unionization driven by social insurance expansions and technological change are welfare-enhancing in the following sense: welfare is higher when the unionization rate adjusts endogenously to these forces than when it is fixed at its mid-1950s level. Social insurance expansions reduce the need for unions to provide insurance benefits, while skill-biased technological change shifts employment toward high-skill workers whom nonunion firms are more likely to hire.

[2]Coincidentally, the year 1954 when the U.S. Congress enacted legislation exempting employer-sponsored health insurance from federal income taxation was also the year with the highest union density, at almost 36%, among American workers.

However, this finding does not imply that unions should be eliminated. We also find that in the current economy subsidizing unions still improves overall welfare, suggesting that the current level of unionization is below the social optimum. The current social insurance system remains incomplete in that a sizable fraction of employed individuals lack health insurance. Although unions can distort wages, their role in insurance provision remains valuable because many workers still lack adequate health insurance. Thus, historical deunionization was an efficient response to the technological changes and social insurance expansion, yet those same forces have pushed union density to a level at which a marginal increase in unionization through subsidization can be welfare-improving.

Union formation and social insurance systems, as well as the bargaining protocols between unions and firms, vary substantially across countries; therefore, our quantitative findings are specific to the U.S. Nevertheless, our framework may also be useful outside the U.S., where unions similarly play dual roles in wage compression and the provision of various non-wage amenities. We conclude by discussing how sectoral bargaining interacts with insurance provision and what our framework implies for the long-run evolution of unionization.

Related Literature. This paper contributes to several strands of literature. First, it relates to the literature on unions and labor markets. A large empirical literature studies the effects of unions on wages and wage inequality (e.g., DiNardo et al., 1996; DiNardo and Lee, 2004; and Farber et al., 2021; see also Jäger et al., 2025 for a recent comprehensive survey). A smaller literature examines the effects of unions on non-wage benefits (e.g., Freeman and Medoff, 1984; Buchmueller et al., 2002; Knepper, 2020; and Lagos, 2021). Our paper is most closely related to a growing macro-labor literature that studies the equilibrium effects of unions on labor markets. For example, Acemoglu et al. (2001) develop a model in which unions affect redistribution, wage insurance, and investment, and argue that skill-biased technological change shapes the decline of unions. More recent work examines the broader macroeconomic implications of unions (e.g., Açıkgöz and Kaymak, 2014; Dinlersoz and Greenwood, 2016; Krusell and Rudanko, 2016; Taschereau-Dumouchel, 2020; Alder et al., 2023; and Pickens, 2025). However, this literature has largely abstracted from the effects of unions on non-wage benefits. Our paper builds on Taschereau-Dumouchel (2020), which develops a model with endogenous union formation, by incorporating employer-provided non-wage benefits and social insurance. This extension allows us to study the interactions among unions, employer-provided insurance, and social insurance in a general equilibrium labor market framework.

Second, by focusing on the role of non-wage benefits, our paper is related to the literature studying equilibrium labor market impacts of non-wage benefits. Recent studies have emphasized the allocative function of non-wage benefits and their heterogeneity across firms (e.g., Sorkin, 2018;

Taber and Vejlin, 2020; Lamadon et al., 2024; Morchio and Moser, 2026; Mas, 2025; and Bagga et al., 2025). We add to this literature by showing how unions interact with firms in determining the provision of non-wage benefits.

Third, our paper contributes to the large literature on the labor market and welfare effects of social insurance. Empirically, many studies examine how social insurance affects private insurance coverage (e.g., Cutler and Gruber, 1996; Shore-Sheppard et al., 2000; and Garthwaite et al., 2014). In structural settings, a large body of work evaluates the welfare effects of social insurance using life-cycle models (e.g., French and Jones, 2011; De Nardi et al., 2010; and Low and Pistaferri, 2015). A smaller but growing literature studies social insurance within equilibrium labor market models. For example, Dey and Flinn (2005), Aizawa (2019), and Aizawa and Fang (2020) examine health insurance; Mitman and Rabinovich (2015) and Chodorow-Reich et al. (2019) study unemployment insurance; and Cole et al. (2019), Aizawa et al. (2025), and Lise et al. (2024) analyze disability insurance and related policies.[3] We contribute to this literature by providing evidence that social insurance can crowd out labor unions when unions partly function as an important instrument for insurance provision, and by studying the interactions among labor market institutions, equilibrium outcomes, and social insurance.

2 Background

This paper focuses on the private sector labor unions in the United States. In this section, we summarize the key features of unions and insurance in the U.S.

In the U.S., workers can form a union to collectively bargain with their employers over compensation and benefits under the National Labor Relations Act (NLRA). To organize a union, workers first need to collect union authorization cards or petitions from at least 30% of their coworkers. Workers can then file a petition for a union election with the National Labor Relations Board (NLRB), and a union is formed if more than 50% of workers who vote are in favor of unionization.[4]

Once a union is formed, collective bargaining covers all workers in the bargaining unit. The NLRA stipulates that an appropriate bargaining unit is a group of two or more employees who share a community of interest, and the determination of a bargaining unit is left to the discretion of the NLRB. In practice, most of the bargaining takes place at the *enterprise* level.[5] Once a union is organized, all workers at the same workplace are covered by collective bargaining even if they

[3] See Fang and Krueger (2022) for a survey of recent macroeconomic research on health policy.

[4] For more details, see the NLRB website https://www.nlrb.gov/about-nlrb/rights-we-protect/the-law/employees/your-right-to-form-a-union.

[5] According to the OECD/AIAS ICTWSS database, collective bargaining in the U.S. occurs at the company or enterprise level for more than two-thirds of union coverage.

are not union members. Operating a union incurs costs, and typically union dues are automatically withheld from the wages of all covered workers. However, some states have passed RTW laws, allowing non-members to avoid paying union dues while still being covered by collective bargaining agreements.

In theory, forming a union is up to the employees in the firm, but in practice, firms play a crucial role. Firms often use anti-union tactics to dissuade workers from unionizing (Dickens 1983, Bronfenbrenner 2009).[6] Consequently, unionization is determined not only by workers' preferences for unions but also by how costly it is for firms to prevent unionization through various tactics.

In addition, employers play an important role in insurance provision in the U.S. For example, ESHI is a dominant source of insurance coverage for working adults, covering more than 60% of them (Aizawa and Fang, 2020). The government provides public insurance through Medicaid and Medicare, but only low-income adults and the elderly, respectively, are eligible. Moreover, since U.S. employment protection is weaker than that of European countries, employers have a direct role in determining workers' layoff risks (Cahuc and Postel-Vinay, 2002).

These features are substantially different from those in many European countries, where unions are organized and collective bargaining takes place at the *sectoral* level (Jäger et al., 2025).[7] For example, while union density and collective bargaining coverage are roughly equivalent in the U.S., France exhibits a stark contrast, with union density below 20% despite nearly universal collective bargaining coverage. In addition, in Europe, the government often provides various insurance benefits through sectoral labor unions. Thus, employers in Europe play a limited role in both union formation and insurance provision at the firm level. In the U.S., employers play an essential role in the interactions between labor unions and social insurance.

3 Empirical Evidence

To motivate our focus on the relationship among employer-provided insurance benefits, unions, and social insurance, this section provides a variety of empirical evidence. First, we document that workers in unionized firms are more likely to receive a variety of employer-provided insurance benefits, as well as better job security. Then, we provide new evidence on the effects of social insurance expansion on the unionization rate.

[6] These tactics include both lawful actions (e.g., hiring anti-union consultants) and unlawful actions (e.g., threats, interrogations, and harassment); see Bronfenbrenner (2009) for more details.

[7] See the OECD/AIAS ICTWSS Database (https://www.oecd.org/employment/ictwss-database.htm) for the level at which collective bargaining takes place in various OECD countries.

3.1 Unionization and Insurance Provision by Employers

Firm Heterogeneity in Unionization. First, we present several descriptive patterns using establishment-level data from the 1997 Robert Wood Johnson Foundation Employer Health Insurance Survey, which provides detailed information on establishment characteristics, ESHI offerings, and unionization. We restrict the sample to private establishments and exclude the agriculture/forestry and fisheries industries. Table 1 provides summary statistics of the establishment-level outcomes. First, unionized establishments tend to be larger, provide higher wages, and be 25% more likely to provide health insurance to their workers. Second, unionized establishments are larger and more likely to provide health insurance within *any* industry, although wages are not necessarily higher in unionized establishments within industries. For example, unionized establishments are 11%–38% more likely to provide health insurance to their workers, conditional on industry. In Appendix Table A.2, we also show that unionization is positively associated with ESHI provision, even after controlling for the size and industry of the establishment. Although union wages are much higher in the construction sector, they are similar to nonunion wages in other industries and even lower than nonunion wages in the manufacturing industry. Third, some industries are more likely to be unionized; specifically, the construction and manufacturing industries have more unionized establishments than the service industries.

Table 1: Summary Statistics of Establishment-Level Variables

			Estab. Size		Wage ($1,000)		ESHI Offer (%)	
Industry	Unionized (%)	Empl. Share (%)	Union	Nonunion	Union	Nonunion	Union	Nonunion
All	3.2	100.0	55.4	13.9	28.7	25.2	76.0	49.6
Construction	9.4	5.3	25.0	8.6	37.6	25.4	82.0	42.3
Mining and manufacturing	4.4	22.0	150.0	30.6	29.1	29.0	86.6	62.6
Wholesale/retaile trade	2.2	21.4	30.1	12.8	19.5	20.0	69.8	44.6
Finance/insurance/real estate	2.2	15.1	28.7	10.7	27.5	29.3	60.2	54.6
Other services	2.8	36.2	65.2	13.3	27.7	25.4	77.8	48.3

Note: This table reports the summary statistics of establishment-level variables from the 1997 Robert Wood Johnson Foundation Employer Health Insurance Survey. Other services include Transportation/communications/other public utilities; Professional services; Other services in the original data. The agriculture/forestry & fisheries industry is dropped.

Employer-Provided Insurance Benefits. Next, we provide further evidence on the effects of union status on various employer-provided insurance benefits from workers' perspectives. We use the household survey data from the Health and Retirement Study (HRS, 1992-2019), which contains rich measures of various non-wage benefits for individuals aged 50 or older and their spouses. We restrict our sample to private-sector workers aged 65 or less and regress indicators for various insurance coverages on union membership and several demographic variables. Specifically, we examine (i) ESHI coverage, (ii) pension from the current job, (iii) life insurance coverage,

Table 2: Union Membership and Insurance Coverage

	ESHI	Pension	Life Ins.	LTC Ins.
	(1)	(2)	(3)	(4)
Union	0.056***	0.186***	0.039***	0.008
	(0.018)	(0.018)	(0.013)	(0.015)
Mean outcome	0.72	0.68	0.84	0.10
Observations	30,192	30,361	30,314	29,823

Note: This table reports the estimation result of equation (1). The sample consists of workers aged 65 or less in the HRS 1992-2019. The time-varying covariates include quadratic polynomials of age, the log of the number of people in the same workplace, the log of earnings, dummies for occupations, industries, and four census regions. Year fixed effects and individual fixed effects are also controlled. Person-level analysis weights are used. Standard errors are clustered at the individual level. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

and (iv) long-term care (LTC) insurance coverage, and estimate the following regression equation:

$$y_{it} = \beta \cdot \textit{Union}_{it} + x'_{it}\gamma + \alpha_i + \lambda_t + \epsilon_{it}, \tag{1}$$

where i is the individual, t is the year, y_{it} is a binary (0 or 1) indicator of the insurance coverage of i in year t, $\textit{Union}_{it}$ is a binary indicator for whether i is a union member in year t, x_{it} is a vector of time-varying covariates, α_i and λ_t are, respectively, individual and time fixed effects, and ϵ_{it} is an error term. The coefficient β represents how much insurance coverage is related to i's union status. Since we control for individual fixed effects, we exploit changes in the union status of the same individuals over time to identify the coefficient β.

Table 2 shows that union membership is associated with better access to health insurance, pension, and life insurance. Access to LTC insurance is weakly correlated with union membership, although the coefficient is not statistically significant.[8]

Although we control for worker fixed effects and a rich set of covariates, including firm, industry, and occupation characteristics, interpreting the estimate as the causal effect of unions may still be subject to a possible concern arising from selection into unionization status based on unobserved changes in demand for insurance benefits. Moreover, the HRS sample consists primarily of older individuals. In Appendix A.5, we follow Fortin et al. (2022) and exploit the differential exposure to RTW laws across industries as a source of variation in unionization. We use the Current Population Survey (CPS) and restrict our sample to private-sector workers aged 22-65. Using this strategy, we find that unionization, instrumented by RTW exposure, has a positive effect on ESHI coverage.

Job Security. Unions can also provide workers with insurance in the form of better protection against layoffs. We investigate how union membership is related to subsequent job loss using the Survey of Income and Program Participation (SIPP, 1996-2008 panels), which contains more

[8] In Table A.1, we gradually add covariates to clarify which confounders explain variations in ESHI coverage.

accurate measures of worker's job turnovers. Here, we summarize the main findings and relegate the details to Appendix A.2. We find that, first, the monthly job loss probability is smaller for union workers than for nonunion workers; second, the decline in job loss probability from unionization is much larger for low-skill workers than for high-skill workers.

Insurance Demand and Union Preferences. Our analysis above suggests that unionized workers have better access to insurance benefits. However, this alone does not imply that such benefits are an important motive for union membership. We therefore examine both workers' stated motivations for supporting unions and whether workers with greater ex-ante insurance demand are more likely to be unionized. Using data from the 1977 Quality of Employment Survey, Kochan (1979) shows that nonunion workers dissatisfied with fringe benefits are significantly more likely to support union representation, while union workers rank fringe benefits slightly above wages among unions' priorities. More recently, Gallup's 2022 Work and Education survey reports that better pay and benefits, together with job security, are among the most frequently cited reasons for supporting unions.[9]

We also examine whether workers with characteristics associated with greater expected health care needs are more likely to be unionized. Using CPS data for 2001–2019, we regress union membership on demographic characteristics associated with health care needs. We find that older workers, parents, and, to a lesser extent, married workers are significantly more likely to be union members. Full results are reported in Appendix A.3.

3.2 Effects of Social Insurance Expansions on Unionization

We now provide evidence on the effect of the expansion of social insurance programs on unions. The key premise is that if unions increase employers' insurance provision, and if employer-provided insurance and publicly provided insurance are substitutes from the workers' perspective, then expansions in social insurance may reduce the demand for unionization.[10] We begin with some aggregate data patterns. First, Figure A.1a in Online Appendix displays the national union membership density for private sector workers from 1948 onward, based on Farber et al. (2021). Union density was around 35% during the 1950s, began to decrease around the mid-1950s, and stood at less than 10% after 2010.[11] Interestingly, the union decline occurred in all four census regions from 1964 onward (Figure A.1b). Second, Figure A.1c shows government spending on

[9] A summary of the survey findings is available at https://news.gallup.com/poll/398303/approval-labor-unions-highest-point-1965.aspx.

[10] See Cutler and Gruber (1996), Rust and Phelan (1997), French and Jones (2011), and Garthwaite et al. (2014) for early contributions studying substitution between various forms of employer-provided insurance and public insurance.

[11] Union density is highly heterogeneous across sectors, and large sectoral mobility occurred over the second half of the twentieth century (Lee and Wolpin, 2006), but we confirm in Online Appendix A.1 that such sectoral mobility is not a major factor behind the decline in unions.

the three major social insurance programs, namely Medicaid, Medicare, and Social Security, as a percentage of U.S. GDP. In contrast to the trend in union density, government spending on social insurance programs has steadily increased over the same time period. Before 1965, neither Medicare nor Medicaid existed; however, spending on each program has risen to around three percent of GDP in recent years.[12]

Although aggregate trends suggest that social insurance expansions may crowd out labor unions, they alone do not establish causality.[13] As emphasized in the literature (Acemoglu et al., 2001; Farber et al., 2021), union decline has also been driven by factors such as technological changes (particularly since the 1980s) and the adoption of state-level RTW laws. Political factors could also play a role; for example, union density is lower in the South, where social insurance programs are generally less generous (Figure A.1b). Furthermore, several alternative mechanisms could potentially lead to a positive correlation between unions and social insurance expansion.[14] Below, we exploit plausibly exogenous variations across time and space in social insurance programs to estimate their impact on unionization.

3.2.1 Introduction of Medicare

First, we look into the introduction of Medicare, which was enacted into law on July 30, 1965, and implemented from July 1, 1966. Medicare is a large public social insurance program that provides almost universal health insurance coverage mainly for Americans aged 65 or older. It can affect union density for the following reasons. Before the implementation of Medicare, individuals had to rely on private insurance to cover the health risks associated with old age. Due to the lack of well-functioning individual markets, workers needed to rely on ESHI, which often included post-retirement coverage. As such, unions played a crucial role in providing retiree health insurance coverage, which could have incentivized workers to seek union jobs to secure access to insurance. The implementation of Medicare delinked the retiree insurance coverage from unions. Indeed, employer-provided retiree health insurance has largely disappeared: while more than 90% of large firms offer ESHI, fewer than 20% of those firms provide retiree health insurance benefits today (KFF, 2023). Thus, by lowering the demand for employer-provided retiree health insurance negotiated by unions, the introduction of Medicare may crowd out unions.

[12]Social Security was introduced in the late 1930s following the passage of the Social Security Act in 1935, and union density *increased* during this period; however, the rise in union density in the late 1930s was likely driven by the National Labor Relations Act of 1935, which guaranteed private-sector employees the right to organize unions and engage in collective bargaining (Farber et al., 2021).

[13]Early studies, such as Neumann and Rissman (1984) and Moore et al. (1989), documented a time-series association between government welfare expenditures and union density, concluding that higher social program spending correlated with lower union density in the late 20th century.

[14]For example, unions may encourage workers to utilize unemployment insurance, as shown by recent evidence from Lachowska et al. (2025), suggesting that social insurance spending could be endogenous to union density.

To identify the effect of Medicare on unions, we follow the empirical strategy of Finkelstein (2007) and exploit geographic variation in pre-1965 health insurance coverage for retirees. Before the introduction of Medicare, the private health insurance coverage rates for the elderly differed across regions, and the introduction of Medicare increased coverage, almost uniformly, to 100%. A region with a higher pre-reform coverage rate would be more affected by (or exposed to) the introduction of Medicare because access to Medicare substantially lowers the need for workers to rely on unions to gain private retiree health insurance coverage.[15]

For this analysis, we use the data on state-level union density since 1964 from Hirsch et al. (2001),[16]. In some analysis, we also use the Current Population Survey (CPS), and information on state-level political environments from Klarner Politics and the National Conference of State Legislatures.[17]

We first look at raw trends in union density among the group of high pre-reform insurance coverage (i.e., high policy exposure) states and the group of low coverage (i.e., low policy exposure) states over the years. As shown in Figure A.5 in the Online Appendix, both groups moved in parallel before 1966, but then the union density decreased from 1966 onward only for the high policy exposure group.

Given this differential pattern across the exposure groups, we estimate the following exposure-based difference-in-differences (DID) regression equation:

$$\log\left(\textit{union}_{st}\right) = \beta \times \textit{Exposure}_s \times \textit{Post}_t + x_{st}'\gamma + \alpha_s + \lambda_t + \epsilon_{st} \tag{2}$$

where the outcome variable $\log\left(\textit{union}_{st}\right)$ is the log of union membership density in state s at year t.[18] The treatment variable $\textit{Exposure}_s$ is the exposure to the Medicare introduction measured by the fraction of the elderly in state s covered by private retiree insurance in 1963 (prior to the introduction of Medicare);[19] $\textit{Post}_t$ is an indicator that takes value 1 if $t > 1965$; x_{st} is a vector of

[15]Figure A.4 in the Online Appendix also shows that the pre-reform coverage rate is positively correlated with state-level union density prior to the introduction of Medicare. In Appendix A.4, we provide further details about the role of unions in providing employer-sponsored retiree (post-65) health insurance coverage.

[16]Although we cannot distinguish between the public and private sectors in the state-level data by Hirsch et al. (2001), we supplement our analysis by using the election data from the NLRB, which covers private-sector union elections.

[17]We obtained the data on partisan balance in early years from https://www.klarnerpolitics.org/datasets-1 (Last accessed March 11, 2024) which is based on Klarner (2003), and we obtained the data for recent years from the National Conference of State Legislatures.

[18]Union membership density and union coverage are very close in the U.S., in contrast to many European countries where collective bargaining agreements often extend beyond union members (Krusell and Rudanko, 2016; Jäger et al., 2025). In our setting, changes in union membership would therefore closely track changes in union coverage.

[19]Note that this variable captures total insurance coverage, rather than coverage provided specifically through unions. Although this choice is partly due to data limitations, we believe total coverage remains an appropriate measure for the policy exposure because insurance provision by nonunion firms may also reflect union threat effects. In particular, nonunion firms may expand insurance provision to discourage unionization, as discussed in Section 6.1.

time-varying state-level covariates described later in detail; and α_s and λ_t are the state and year fixed effects, respectively. Standard errors are clustered at the state level.

As discussed in Callaway et al. (2024), the parameter β identifies a weighted average of the average causal responses, under a strong parallel-trends assumption. Specifically, it assumes that the average evolution of outcomes for the entire population if all states experienced $\mathit{Exposure} = c$ is equal to the actual evolution of outcomes for states with $\mathit{Exposure} = c$.[20] Essentially, this assumption requires that the differential effect of Medicare implementation across states arise only from state-level differences in pre-Medicare retiree health insurance coverage levels *Exposure* and not from other time-varying unobserved factors in each state that may be correlated with the policy effect of interest ($\mathit{Exposure}_s \times \mathit{Post}_t$). To assess the plausibility of this identifying assumption, we conduct several diagnostic exercises. First, as discussed above, outcomes evolved similarly across states before 1966. Second, we control for a rich set of time-varying state characteristics x_{st}, including demographic composition (the shares of workers with some college education, female workers, and workers aged 40 or older), political environment (an indicator for a Democratic governor and third-order polynomials in Democratic representation in the state Senate and House), local labor market dynamics through state-level unemployment rates, and differential exposure to trade and automation through manufacturing employment shares interacted with year fixed effects.

The pooled specification captures only the average post-Medicare effect and offers no direct test of the parallel-trends assumption. To address both, we also estimate a dynamic event-study version of the DID that interacts pre-Medicare coverage ($\mathit{Exposure}_s$) separately with each year relative to 1965:

$$\log\left(\mathit{union}_{st}\right) = \sum_{\tau=-1,\tau\neq 0}^{5} \beta_\tau \cdot \mathit{Exposure}_s \cdot \mathbb{1}\{t = \tau + 1965\} + x'_{st}\gamma + \alpha_s + \lambda_t + \epsilon_{st}, \quad (3)$$

where year 1965 is the omitted reference.

Results. Table 3 reports the DID estimates of the effect of Medicare introduction on union density based on equation (2). The coefficients are negative and statistically significant, suggesting that the states more exposed to the Medicare introduction experienced a greater decline in union density. Quantitatively, column (1) shows that moving from the state with the minimum Blue Cross coverage in 1963 (12%) to that with the maximum (51%), or an exposure increase of 0.39, implies

[20]Formally, in the language of the potential outcomes framework, if all states experienced the exposure level of c, their average change in outcomes Y_s from pre- to post-policy periods must satisfy:

$$\mathbb{E}[Y_{s,post}(c) - Y_{s,pre}(c)] = \mathbb{E}[Y_{s,post}(c) - Y_{s,pre}(c) | \mathit{Exposure} = c]$$

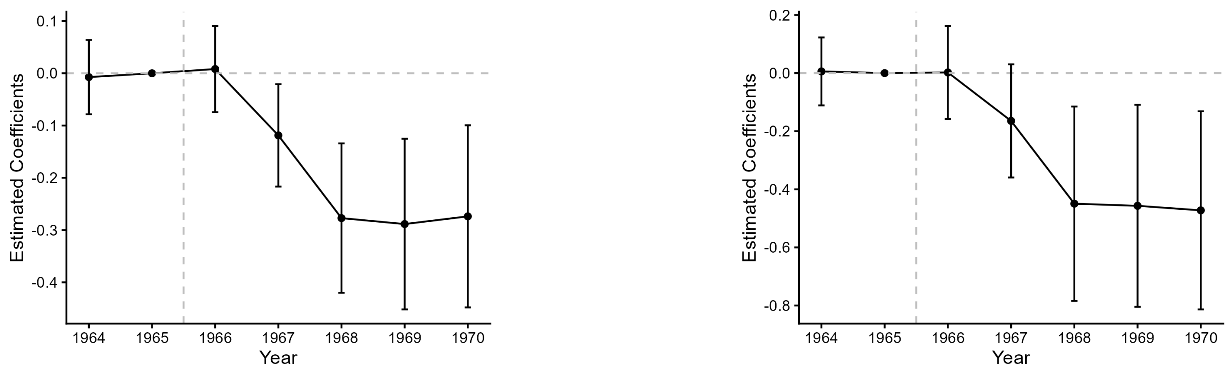
where $Y_{s,t}(c)$ is an outcome for state s in period t if it experienced the exposure level c. The left-hand side is the average evolution of the outcomes if all states experienced the exposure level c, and the right-hand side is the average for states that actually experienced the exposure c.

Table 3: Medicare Impact on Union Density: DID Estimate

	log(union density)	
	(1)	(2)
Exposure × Post	-0.174***	-0.300**
	(0.057)	(0.122)
Coverage	Blue Cross	Any
Observations	343	343

Note: This table reports the estimation results of equation (2). Data on state-level union density are from Hirsch et al. (2001); the Medicare exposure measure is from Finkelstein (2007). The treatment variable *Exposure*$_s$ is based on Blue Cross insurance coverage in 1963 in column (1) and on any insurance coverage in column (2). State population in 1960 is used as weights. Standard errors are clustered at the state level. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

Figure 1: Estimated Impact of Medicare Introduction on Unions



(a) Blue Cross Coverage

(b) Any Insurance Coverage

Note: This figure displays the estimated coefficients of equation (3). Data on state-level union density are from Hirsch et al. (2001); the Medicare exposure measure is from Finkelstein (2007). Panel (a): *Exposure*$_s$ is Blue Cross insurance coverage in 1963. Panel (b): *Exposure*$_s$ is any insurance coverage in 1963. State population in 1960 is used as weights. The error bars represent 95% confidence intervals based on standard errors clustered at the state level.

an additional decline in log union density of roughly 6.8%. Using "Any coverage" in column (2), the analogous range (43% to 68%, an exposure increase of 0.25) implies about a 7.5% decline in log union density. The two specifications yield a consistent picture: the most exposed states experienced 7–8% larger post-Medicare declines in union density than the least exposed.[21]

Figure 1 reports the estimates of β_τ from equation (3), normalized to zero in 1965. Consistent

[21]Following Callaway et al. (2024), we also estimate a non-parametric specification that allows flexible interactions between exposure and time dummies, and we construct average causal responses weighted by the sampling distribution. We replace the linear Exposure$_s$ × Post$_t$ term with interactions between Post$_t$ and a third-order polynomial in Exposure$_s$, leaving the other controls and fixed effects unchanged, and aggregate the implied treatment-effect function across states using the empirical (population-weighted) distribution of Exposure$_s$ in 1963. The resulting estimate is -0.236 in the Blue Cross specification, which is larger in magnitude than the linear DID coefficient of -0.174 in column (1) of Table 3, indicating that the linear specification, if anything, understates the average post-Medicare decline in union density attributable to exposure.

with the pooled DID estimates, the estimated coefficients show that states with higher pre-Medicare retiree insurance coverage experienced larger declines in union density after Medicare's introduction. While data limitations prevent us from examining long-run pre-trends, we find no evidence of differential pre-trends in the observed pre-treatment period.[22]

Finally, we interact the treatment with the pre-Medicare share of workers aged 40 or older and find a negative interaction, implying that more exposed states with relatively older workforces experienced larger declines in unionization (Table A.3). This finding is consistent with our mechanism, as older workers had more to gain from Medicare and were also more likely to value retiree health benefits traditionally associated with unionized employment. We report the specification and full results in Appendix A.7.

In Online Appendix, we confirm that this result is robust to controlling for state-level Medicaid implementation that occurred mostly between 1966 and 1972 (Figure A.6). We also provide additional supporting evidence using data on NLRB elections in Appendix A.6.

3.2.2 Effects of Other Social Insurance Programs

Here we summarize the main findings from our analysis of three other social insurance policy reforms and relegate all the details to Appendix A. First, we consider the impact of Medicaid implementation on unions. Specifically, although Medicaid was signed into law in July 1965, the timing of its implementation was left to each state. As a result, some states implemented the program earlier than others. We exploit this staggered treatment timing and estimate a difference-in-differences specification following the approach of Sun and Abraham (2021). The estimate suggests that union density was reduced by approximately 5% two years after implementation.

Second, we examine the effect of insurance expansions under the 2010 Affordable Care Act (ACA). One of the key provisions of the ACA is a state-based expansion of Medicaid, which provides Medicaid coverage to anyone whose income is below 138% of the Federal Poverty Line (FPL). We exploit the variation in the timing of the ACA Medicaid expansion across states and find that it slightly lowers union membership on average, but it lowers the unionization rate much more significantly for less-educated workers, as one would expect from the fact that Medicaid is targeted toward low-income individuals.

Furthermore, we consider the effect of more generous UI benefits. The UI provides temporary benefits to individuals who have lost their jobs, which may substitute for the union's role in job protection. Importantly, each state can adjust UI generosity, including the amount of benefits. We use variation in UI replacement rates across states and over time to estimate the impact of UI gen-

[22] We also check the robustness of our result to the potential violation of the parallel-trends assumption in the spirit of Rambachan and Roth (2023). Appendix Figure A.7 shows that the estimated effect remains statistically significant even when the post-treatment violation of parallel trends is allowed to be up to 1.2 times larger than the pre-treatment violation.

erosity. We find that a more generous UI replacement are lowers the individual unionization rate.

3.2.3 Discussions

Our empirical analyses thus far suggest that expansions in social insurance programs may have contributed to the decline in unions; however, they are not sufficient to fully quantify the extent of the contribution. For example, our Medicare analysis compares the evolution of union density across states with different levels of exposure to policy change. Because Medicare was introduced nationally, the analysis does not compare treated and untreated states and therefore does not directly identify the aggregate effect of Medicare on union density. Moreover, it is difficult to infer the long-run effects from our quasi-experimental estimates alone.

These limitations make it difficult to assess the quantitative importance of the social insurance expansion channel compared to other proposed explanations for union decline, such as technological change. In addition, understanding the welfare consequences of declining unionization requires a clearer understanding of the *mechanisms* through which unions affect insurance provision. These considerations motivate us to develop and estimate an equilibrium labor market model with endogenous union formation and employer-provided insurance, which we describe below.

4 The Model

4.1 Environment

We consider a discrete-time, infinite-horizon model. There is a unit mass of risk-averse workers with skill types indexed by $x \in \mathcal{X} = \{1, \ldots, X\}$, where N_x denotes the fraction of type x. Workers consume wages w and amenities $a \in \mathcal{A}$ (or non-wage benefits, including insurance products), where $\mathcal{A}$ is a finite set, and $a = 0$ denotes no benefits.[23] Wages can vary across employees within a firm, but amenities cannot.[24]

Firms are risk-neutral and heterogeneous in their production technologies indexed by $y \in \mathcal{Y} = \{1, \ldots, Y\}$. A firm's union status is $k \in \{u, n\}$: unionized ($k = u$) or nonunionized ($k = n$). Each firm uses only labor to produce output according to the production function $F_y(\boldsymbol{g})$, with $\boldsymbol{g} = (g_1, \ldots, g_X)$ where g_x denotes the measure of type-x workers it hires (see equation (8) for details). The measure of type-y firms is given by M_y, and the total measure of firms is $M = \sum_{y \in \mathcal{Y}} M_y$.[25]

[23] Although amenities here are only for employed workers, we later extend the model by introducing retirement and post-retirement amenities to study Medicare. See Section 4.8 for details.

[24] Some amenities, such as workplace safety, are inherently determined at firm-level, while for other amenities, such as health insurance and workplace accommodation, anti-discrimination laws prohibit firms from providing different levels to different workers.

[25] Although the model features a fixed measure of (multi-worker) firms rather than free entry, changes in firm profitability still affect the labor market through endogenous vacancy creation. Acemoglu and Hawkins (2014), for example, show that the response of unemployment to productivity shocks is quantitatively similar between models with a fixed number of firms and those with free entry.

Both workers and firms discount the future at a common factor $\gamma \in (0, 1)$. We assume that workers cannot save or borrow, and that they have no access to individual insurance.[26] In what follows, we focus on a steady state.

Labor Market. There is a frictional labor market for each skill type x, where firms can post multiple vacancies. In the sub-market for type x, matches are created according to a matching function $m(s_x, \nu_x)$, where s_x and ν_x denote the measure of unemployed workers and vacancies, respectively. We assume m is strictly concave and strictly increasing in each argument, and homogeneous of degree one. We let $\theta_x = \nu_x/s_x$ denote labor market tightness for type-x workers. The vacancy-filling probability is then $q(\theta_x) = m(1/\theta_x, 1)$, and the job-finding probability is $p(\theta_x) = m(1, \theta_x)$. Matches are destroyed at the end of each period with probability $\delta_{x,k}$, which depends on worker skill type $x \in \mathcal{X}$ and firm union status $k \in \{u, n\}$. There is no on-the-job search.

Timing. The timing of events in each period is as follows: (i) Firms' union status is endogenously determined; (ii) Firms choose vacancy postings in each sub-market and amenity provision.; (iii) Vacancies and unemployed workers are randomly matched in each sub-market; (iv) Wage bargaining and production take place, and then wages and amenities are provided; (v) A fraction $\delta_{x,k}$ of jobs is destroyed for each x and k.

An important timing assumption is that amenities are determined at the recruiting stage and are not renegotiable ex post. This assumption is plausible for many job attributes, such as health insurance, job flexibility, and workplace safety, and is standard in models that incorporate job amenities (e.g., Aizawa and Fang, 2020, Morchio and Moser, 2026).[27]

4.2 Worker's Problem

Preference. If a type-x worker receives wage w and amenity a, then the worker obtains utility $u_x(w, a)$, increasing and concave in the first argument. An unemployed worker obtains $u_x(b_x, 0)$ where b_x denotes unemployment benefits (and/or home production).

Value Function. The value for a type-x worker employed by a type-y firm with union status k that offers a compensation package (w, a) this period is given by

$$V^E_{x,y,k}(w, a) = u_x(w, a) + \gamma\Big[\delta_{x,k}V^U_x + (1 - \delta_{x,k})V^E_{x,y,k}(w_{x,y,k}, a_{x,y,k})\Big]. \tag{4}$$

[26] Adding these features is conceptually straightforward but would substantially complicate the model and its computation. For example, if we allow workers to save, the wealth distribution of a firm's employees—an infinite-dimensional object—becomes a state variable. This is because firms in our model employ multiple workers, and each worker's wealth affects their outside option in bargaining. Section 6.3.2 examines the robustness of our findings to this assumption. Moreover, the individual health insurance market was relatively small during our sample period, limiting the importance of this channel.

[27] See also static monopsony models of the labor market featuring wage competition across firms (e.g., Volpe, 2024). An important feature of these models is that firms may differentiate themselves through amenities, but these amenities are taken as fixed and predetermined when firms choose wage offers.

The value of employment consists of flow utility from the package (w, a) plus the discounted continuation value.[28] With probability $\delta_{x,k}$, the job is destroyed and the worker receives the unemployment value V_x^U described below. When current wages are bargained, future equilibrium wages and amenities $(w_{x,y,k}, a_{x,y,k})$ are taken as given.

The unemployment value for a type-x worker, V_x^U, is given by

$$V_x^U = p(\theta_x) \overbrace{\mathbb{E}\left[\max\{V_{x,y,k}^E(w_{x,y,k}, a), u_x(b_x, 0) + \gamma V_x^U\}\right]}^{V_x^M} + (1 - p(\theta_x))[u_x(b_x, 0) + \gamma V_x^U] \quad (5)$$

where V_x^M is the worker's expected value upon receiving a job offer. With probability $p(\theta_x)$, the worker meets a firm and with the remaining probability, the worker remains unmatched. The expectation is taken over the *equilibrium* distribution of vacancies posted in the sub-market for type-x workers, distinguished by the firm type $y \in \mathcal{Y}$ that posts the vacancy, as well as the wage, amenity, and union status associated with the vacancy.[29]

4.3 Cost of Unionization and Union Prevention

Although in theory a firm is expected to unionize if the majority of workers favor it, the reality is more nuanced. As discussed in Section 2, firms often resort to various strategies to prevent unionization. To fully capture both the costs of preventing and promoting unionization, we assume that firms determine their union status, but the costs of doing so are influenced by the endogenous collective "preferences" of their workers for or against unionization.[30] Consequently, although a firm can always choose to remain nonunionized, doing so may not be profitable if its workers have a strong collective preference for unionization.

To derive the endogenous preference for unionization, we denote by $\mathcal{W}_{x,y,n}(\boldsymbol{g}, a) \in \mathbb{R}$ the *willingness to pay for unionization* of a type-x worker in a type-y nonunionized firm with workforce composition $\boldsymbol{g}$ and amenity a. It measures how strongly a worker favors unions, expressed in consumption units, and tends to be positive for low-skill workers and negative for high-skill workers. Its formal derivation is in Appendix B.2 (see equation (A13)); intuitively, $\mathcal{W}_{x,y,n}(\boldsymbol{g}, a)$ represents the consumption amount a type-x worker must be paid to remain in a nonunionized firm. The firm-level "*collective preference*" for unionization, is the aggregation of $\mathcal{W}_{x,y,n}(\boldsymbol{g}, a)$ over $x \in \mathcal{X}$ using $\boldsymbol{g}$ as the weights, and it is denoted by $\mathcal{W}_{y,n}(\boldsymbol{g}, a)$ (see equation (A14) for its derivation). The

[28] The right-hand side of the equation above implies that the firm chooses the same union status period after period; this holds in a steady state; we elaborate in Section 4.4, which describes the firm's problem.

[29] The precise expression is given by equation (A12) in Online Appendix B.1.

[30] In Appendix B.3, we model workers' voting decisions explicitly and show that, under additional assumptions, the specification above is equivalent to one in which the cost depends on the outcome of majority voting.

firm's cost of preventing unionization, which we term the *union threat cost*, is then given by:[31]

$$C_{y,n}(\boldsymbol{g}, a) = c_0 \max\{0, \mathcal{W}_{y,n}(\boldsymbol{g}, a)\} \tag{6}$$

where $c_0 > 0$ reflects the cost of various strategies a firm may use to counteract unionization.[32] If the employees' aggregate willingness to pay for unionization is positive, a firm must incur this cost to prevent unionization, and the more eager workers are to form a union, the larger the cost. The importance of the union threat cost is governed by the parameter c_0, which parametrizes the firm's role in the eventual unionization outcome. Note from (6) that the union threat cost depends on both $\boldsymbol{g}$ and a, which has two implications. First, if the willingness to pay for unionization varies across worker types, firms may distort their workforce composition $\boldsymbol{g}$ to reduce the union threat cost (Taschereau-Dumouchel, 2020). Second, a nonunionized firm may choose $a = 1$ to save on the union threat cost, which underlines our empirical finding below that the union threat can increase health insurance offerings even in nonunionized firms.

We define an analogous *union maintenance cost* function for the case in which a firm prefers unionization but its workers might oppose it. Given the employees' aggregate *willingness to accept nonunionization*, $\mathcal{W}_{y,u}(\boldsymbol{g}, a)$, formally derived in Appendix B.2, the total cost to a type-y unionized firm of maintaining unionization is given by:

$$C_{y,u}(\boldsymbol{g}, a) = FC^{union} + c_0 \max\{0, \mathcal{W}_{y,u}(\boldsymbol{g}, a)\} \tag{7}$$

where $FC^{union} > 0$ is the fixed cost of unionization that a firm needs to pay regardless of whether workers agree on unionization;[33] and $c_0 > 0$ represents the marginal cost of counteracting de-unionization.

[31] An advantage of using the flexible cost function as a penalty function rather than imposing a hard constraint with $c_0 \to \infty$ (e.g., Taschereau-Dumouchel 2020) is numerical tractability. With a hard constraint, there is a cutoff for $\hat{\alpha}$ such that there cannot be a solution to the hiring problem of nonunionized firms with $\alpha < \hat{\alpha}$, while some firms find it optimal to prevent unionization if $\alpha \geq \hat{\alpha}$. As a result, we encounter a discontinuity in the union probability at $\hat{\alpha}$, which hampers the convergence of an iterative algorithm. This also generates a counterfactual pattern where smaller firms (with smaller α) all become unionized. See Section 6.3.2 for robustness checks under a different functional form.

[32] One could ask why this cost takes the form of resource destruction rather than a state-contingent transfer to workers conditional on remaining nonunionized. We make this modeling choice for two reasons: (i) our wage-setting environment we describe later assumes that firms cannot commit to unionization-contingent payments before union status is realized; once it is realized, wages are set by the bargaining protocol corresponding to that status, and (ii) even if such conditional transfers were allowed, they would be undone in the subsequent bargaining, since within-firm transfers do not change the total surplus to be split. Studying conditional transfers as a meaningful margin would therefore require a fundamentally different wage-setting mechanism.

[33] See Section 4.7 for an interpretation of FC^{union}.

4.4 Firm's Problem

Firms produce consumption goods using only labor inputs. A type-y firm's production depends on worker composition $\boldsymbol{g} = (g_1, \ldots, g_X)$ and is given by

$$F_y(\boldsymbol{g}) = A_y \left(\sum_{x \in \mathcal{X}} z_{x,y} g_x^{\frac{\sigma-1}{\sigma}} \right)^{\frac{\sigma}{\sigma-1}\alpha_y}, \tag{8}$$

where A_y is the type-y firm's total factor productivity (TFP), α_y is its returns to scale, σ is the elasticity of substitution between skills, and $z_{x,y}$ is the relative skill intensity satisfying $\sum_{x\in\mathcal{X}} z_{x,y} = 1$ for each firm type y. We assume decreasing returns to scale, $\alpha_y < 1$, for all firms. The parameter α_y may differ across firms and is drawn from a distribution with CDF G_α.[34]

The current-period profit function of a type-y firm with union status k is given by

$$\pi_{y,k}(\boldsymbol{g}, a) = F_y(\boldsymbol{g}) - \sum_{x\in\mathcal{X}} \left[w_{x,y,k}(\boldsymbol{g}, a) + c_x(a) \right] g_x - FC^a_{y,k}(a) - C_{y,k}(\boldsymbol{g}, a). \tag{9}$$

The first term is revenue from output. The second term is the compensation cost of hiring workers: $w_{x,y,k}(\boldsymbol{g}, a)$ is the wage schedule and $c_x(a)$ is the type-specific expected per-worker cost of amenities. The third term, $FC^a_{y,k}(a)$, is the firm-level fixed cost of providing amenity a, with $FC^a_{y,k}(a) > 0$ if $a > 0$ and $FC^a_{y,k}(0) = 0$.[35] We assume that $FC^a_{y,u}(a) \leq FC^a_{y,n}(a)$ and specify the cost for unionized firms as $FC^a_{y,u}(a) = \widetilde{FC}^a(K_y)$, where K_y denotes the measure of unionized type-y firms providing amenities *in equilibrium* and $\widetilde{FC}^a(\cdot)$ is a decreasing and convex function of K_y.[36] This specification allows for potential cost pooling among unionized firms. We discuss the mechanism and its institutional motivation in Section 4.7 and provide the precise functional form in Section 4.8. The fourth term, $C_{y,k}(\boldsymbol{g}, a)$, is the union cost function, either (6) or (7), as defined in Section 4.3.

Given $(\boldsymbol{g}, a)$ and k, a type-y firm posts, in each sub-market for skill type x, vacancies ν_x at a

[34] This is motivated by our empirical evidence in Section 3.1, which suggests that variation in unionization across firms may reflect heterogeneity beyond productivity differences. We formally discuss the theoretical implications in Section 4.7.

[35] The fixed cost encapsulates various costs associated with amenity provision that remain invariant with respect to firm size — for example, the costs of operating a benefits office or the transaction costs of contracting with insurance providers. In the case of health insurance, insurance companies often impose administrative loading over anticipated claims costs. Karaca-Mandic et al. (2011) report that firms with up to 100 employees face loading fees of approximately 34%, while firms with more than 10,000 employees face only 4%; this type of size-dependence maps into the fixed cost in our model.

[36] Since the fixed cost of amenity provision enters profit with a minus sign, convexity in K_y makes the cost-pooling benefit concave in K_y, and the marginal cost reduction from an additional unionized amenity provider shrinks as K_y grows. This diminishing-returns property dampens the feedback from K_y to firms' amenity-provision incentives and helps rule out multiple equilibria.

cost of $\kappa > 0$ per vacancy, to maximize the discounted sum of profits:

$$J_{y,k}(\boldsymbol{g}, a) = \max_{\nu_1,\ldots,\nu_X \geq 0} \pi_{y,k}(\boldsymbol{g}', a) - \kappa \sum_{x\in\mathcal{X}} \nu_x + \gamma J_{y,k}(\boldsymbol{g}', a), \tag{10}$$

$$\text{subject to} \quad g'_x = (1-\delta_{x,k})g_x + \nu_x q(\theta_x) e_{x,y,k,a}, \quad x = 1,\ldots,X, \tag{11}$$

where $q(\theta_x)$ is the vacancy-filling probability defined in Section 4.1 and $e_{x,y,k,a}$ is worker's decision of whether to accept a job offer from this firm.[37, 38] The first term in the law of motion for the firm's worker composition (11) is the number of workers who are not hit by the exogenous separation shock from the firm, while the second term is the number of new hires.[39] Note that $\delta_{x,k}$ varies with x and k, capturing two forces: first, workers of different skill types may be subject to different separation rates; second, unions can affect job security and the impact may differ by skill type. We include in the fixed cost of unionization FC^{union} the cost of offering different degrees of job security on the firm, due, e.g., to not being able to flexibly adjust to macroeconomic disturbances.

In a steady state, a firm of type (y, k) enters each period with each worker type's employment reduced to a fraction $1 - \delta_{x,k}$ of the previous period's level because of exogenous job destruction. Under linear vacancy posting costs, the firm immediately restores employment to its optimal level, and the constraint $\nu_x \geq 0$ is never binding. As we discuss in the next paragraph, each firm in steady state makes the same unionization and amenity choices in every period, so k and a are constant over time. Following Taschereau-Dumouchel (2020), we obtain the following simplified expression for the firm's hiring objective function:

$$\hat{\pi}_{y,k}(\boldsymbol{g}, a) = \pi_{y,k}(\boldsymbol{g}, a) \; - \; \psi_{y,k}(\boldsymbol{g}), \tag{12}$$

$$\text{where} \quad \psi_{y,k}(\boldsymbol{g}) = \kappa \sum_{x\in\mathcal{X}} \frac{g_x}{q(\theta_x)} \; - \; \kappa\gamma \sum_{x\in\mathcal{X}} (1-\delta_{x,k}) \frac{g_x}{q(\theta_x)}. \tag{13}$$

Equation (12) states that the firm's per-period payoff is current-period profit (9) net of the vacancy cost $\psi_{y,k}(\boldsymbol{g})$. The first term in equation (13) is the gross cost of posting enough vacancies to attain employment g_x for each type x, given the vacancy-filling rate $q(\theta_x)$ and the unit posting cost κ. The second term captures the continuation value of today's hires: a fraction $1 - \delta_{x,k}$ of newly hired workers carry over to the next period and reduce next period's hiring needs, and this anticipated saving is subtracted from the gross cost.

Hiring, Amenity Provision, and Unionization. Firms draw choice-specific shocks for amenities $\boldsymbol{\epsilon} = \{\epsilon_a\}_{a\in\boldsymbol{\mathcal{A}}}$ and union formation $\boldsymbol{\varepsilon} = \{\varepsilon_k\}_{k\in\{u,n\}}$ that are independent across firms but fixed

[37] Recall that $\theta_x = \nu_x/s_x = \sum_{y'\in\mathcal{Y}} \nu_{x,y'}/s_x$. We assume that each type-y firm is infinitesimally small so its choice of $\nu_{x,y}$ does not impact θ_x.

[38] The right-hand side of equation (10) implies that the firm chooses the same union status, which is true in a steady state. We clarify this point near the end of this subsection.

[39] Although vacancies are filled randomly, the law of large numbers makes the number of new hires is deterministic.

over time, implying that, in a steady state, each firm maintains the same union status and amenity provision over time. We assume these shocks are unobservable to workers and cannot be bargained over; consequently, they do not affect wage functions.

Given firm type y, amenity provision a and union status k, a firm chooses its hiring profile $\boldsymbol{g}_{y,k}(a)$ to maximize its steady-state profit flow (12):[40]

$$\boldsymbol{g}_{y,k}(a) = \arg\max_{\boldsymbol{g}} \hat{\pi}_{y,k}(\boldsymbol{g}, a). \tag{14}$$

Given the optimal hiring choices above, a firm's value of choosing amenity level a is given by the discounted sum of steady-state profits $\hat{J}_{y,k}(a) = \hat{\pi}_{y,k}(\boldsymbol{g}_{y,k}(a), a)/(1-\gamma)$. For each y and k, a firm's amenity choice problem is given by $J_{y,k}(\boldsymbol{\epsilon}) = \max_{a\in\boldsymbol{\mathcal{A}}} \left\{\hat{J}_{y,k}(a) + \epsilon_a\right\}$. Then the probability that a type-y firm provides amenity a conditional on the union status $k \in \{u, n\}$, which we denote by $P_{y,k}(a)$, and the probability that a type-y firm unionizes, which we denote by $\mathcal{Q}_y$, are given by:[41]

$$P_{y,k}(a) = \Pr(\hat{J}_{y,k}(a) + \epsilon_a = \max_{a'\in\mathcal{A}}\{\hat{J}_{y,k}(a') + \epsilon_{a'}\}) \tag{15}$$

$$\mathcal{Q}_y = \Pr(J_{y,u}(\boldsymbol{\epsilon}) + \varepsilon_u \geq J_{y,n}(\boldsymbol{\epsilon}) + \varepsilon_n). \tag{16}$$

Equation (16) shows that forming a union is up to firms, but they cannot ignore workers' preferences. The reduced-form union threat cost (6) implies that if workers have strong preferences for unionization, firms cannot profitably prevent unionization and likely end up with unionized workers; likewise, the union maintenance cost (7) implies that if workers have strong preferences for nonunionization, firms cannot profitably unionize the workers and therefore likely end up with nonunionized workers.

4.5 Wage Bargaining

Wages are determined by Nash bargaining between an employer and its workers. The bargaining protocol differs between unionized and nonunionized firms as described below; both take the hiring profile $\boldsymbol{g}$ and amenity provision a as given.

Individual Bargaining in Nonunionized Firms. In individual bargaining, the firm bargains with each worker separately. Due to decreasing returns,the firm's surplus from an agreement depends on whether the worker is treated as marginal or infra-marginal. We follow Stole and Zwiebel (1996) and Brügemann et al. (2019), treating every worker as a marginal worker.[42]

[40] The choice of $g_x(a)$ in equation (14) is not independent of the firm's vacancy choices in the original problem (10). Given the workforce $g_{-1,x}(a)$ entering the period, the law of motion of employment induces a one-to-one mapping between $\nu_x(a)$ and $g_x(a)$. Combined with linear vacancy posting costs, this allows us to recast the firm's problem as a choice over $g_x(a)$, with $\psi_{y,k}(\boldsymbol{g})$ capturing the present value of the required vacancy costs.

[41] Both (15) and (16) appear in equation (A12) in Online Appendix B.1.

[42] The same approach is adopted by, for example, Elsby and Michaels (2013), Acemoglu and Hawkins (2014), and Taschereau-Dumouchel (2020).

Since bargaining takes place after the hiring decision, it does not take into account the vacancy posting cost needed to hire the worker. In addition, because bargaining takes place after union status has been determined and the firm has already paid the union threat cost to remain nonunionized, that cost is sunk at the time of nonunion bargaining.[43] Accordingly, the marginal gain to the firm from an extra worker of type x considered in the bargaining is obtained by differentiating equation (12) ignoring the first term of (13) and the union threat cost (6):

$$\Delta_{x,y,n}(\boldsymbol{w}, \boldsymbol{g}, a) = \frac{\partial F_y(\boldsymbol{g})}{\partial g_x} - w_{x,y,n}(\boldsymbol{g}, a) - c_x(a) - \sum_{x' \in \mathcal{X}} \frac{\partial w_{x',y,n}(\boldsymbol{g}, a)}{\partial g_x} g_{x'} + \frac{\gamma \kappa (1 - \delta_{x,n})}{q(\theta_x)}. \quad (17)$$

The individual bargaining problem is then given by, for each $x \in \mathcal{X}$:

$$\max_{w(\boldsymbol{g},a)} \quad \left[V^E_{x,y,n}(w_{x,y,n}(\boldsymbol{g}, a), a) - u_x(b_x, 0) - \gamma V^U_x\right]^\beta \left[\Delta_{x,y,n}(\boldsymbol{w}, \boldsymbol{g}, a)\right]^{(1-\beta)}, \quad (18)$$

where $\beta \in (0, 1)$ is the worker's bargaining power. The first bracket captures the individual worker's net surplus, while the second bracket captures the net surplus from hiring a marginal worker of type x. The bargaining problems in (18) need to be solved simultaneously for all $x \in \mathcal{X}$.

Collective Bargaining in Unionized Firms. Following Taschereau-Dumouchel (2020), we model collective bargaining as a multi-player Nash bargaining problem between the firm and all its workers represented by their union. We assume that a breakdown of collective bargaining entails a cost ϕ for firms, which generates a union wage premium.[44] Specifically, the vector of wage functions $\boldsymbol{w}(\boldsymbol{g}, a) = \{w_x(\boldsymbol{g}, a)\}_{x \in \mathcal{X}}$ is simultaneously determined by

$$\begin{aligned} \max_{\boldsymbol{w}(\boldsymbol{g},a)} & \left[\prod_{x \in \mathcal{X}} \left(V^E_{x,y,u}(w_x(\boldsymbol{g}, a), a) - u_x(b_x, 0) - \gamma V^U_x\right)^{\frac{g_x}{\sum_{x' \in \mathcal{X}} g_{x'}}}\right]^\beta \\ \times & \left[F_y(\boldsymbol{g}) - \sum_{x \in \mathcal{X}} (w_x(\boldsymbol{g}, a) + c_x(a)) g_x - FC^a_{y,u}(a) + \kappa\gamma \sum_{x \in \mathcal{X}} \frac{(1 - \delta_{x,u}) g_x}{q(\theta_x)} - (-\phi)\right]^{(1-\beta)}. \end{aligned} \quad (19)$$

The first bracket captures the workers' collective net surplus, which can be interpreted as the union's surplus and is constructed as an employment-weighted geometric average of individual worker surpluses, while the second bracket captures the firm's net surplus. The latter does not include the union maintenance cost because it has already been paid to achieve unionization and is sunk at the time of bargaining; it does, however, include the break-up cost ϕ, which is the source of the union wage premium in our model.[45] Unlike under individual bargaining, the fixed cost of

[43] As directly implied by bargaining, firms cannot commit to future transfers at the time when union status is determined. This rules out the possibility of side payments from firms to workers in the promise of nonunionization.

[44] Although not modeled explicitly, ϕ can be interpreted as capturing the various costs firms bear when bargaining fails, including reputational damage from labor disputes, legal expenses associated with labor proceedings, and losses arising from strikes or other forms of work disruption.

[45] An alternative approach is to assume that unionized workers have greater bargaining power than nonunion workers, i.e., $\beta_u > \beta_n$. However, we find that, as long as workers' outside options (the unemployment value) are calibrated

amenities $FC^a_{y,k}(a)$ enters the collective bargaining problem since it is part of the firm's *overall* profit but not part of each worker's *marginal* contribution. As we show in Section 4.7, this difference could give unionized firms stronger incentives than nonunionized firms to offer amenities.

4.6 Equilibrium

We pin down the vector of market tightness $\theta = (\theta_1, \ldots, \theta_X)$ by equalizing two steady-state relationships between unemployment and market tightness. On the firm side, firms' hiring decisions determine unemployment for each skill type $\mathcal{U}^{JC}_x(\theta)$ given θ, where superscript JC stands for "Job Creation." On the worker side, equating the flows into and out of unemployment also determines unemployment for each skill type $\mathcal{U}^{BC}_x(\theta)$ given θ, where superscript BC stands for "Beveridge Curve." We pin down θ so that

$$\mathcal{U}^{BC}_x(\theta) = \mathcal{U}^{JC}_x(\theta) \quad \text{for all } x \in \mathcal{X}. \tag{20}$$

See Appendix B.5 for the derivations of $\mathcal{U}^{JC}_x(\theta)$ and $\mathcal{U}^{BC}_x(\theta)$.

A *steady-state equilibrium* consists of a set of value functions $\{V^E_{x,y,k}, V^U_x\}$, hiring functions $\{\boldsymbol{g}_{y,k}\}$, wage schedules $\{w_{x,y,k}\}$, amenity provision functions $\{P_{y,k}\}$, unionization probabilities $\{\mathcal{Q}_y\}$, market tightness $\{\theta_x\}$ such that: (i) the value functions solve the Bellman equations (4) and (5); (ii) the hiring functions solve the optimal hiring problem (14); (iii) the wage schedules solve the bargaining problems (18) and (19); (iv) the amenity provision functions are determined by (15); and (v) the unionization is determined by (16); and (vi) the market tightness satisfies (20). Although the general model requires a numerical solution to find an equilibrium, several key properties can be characterized analytically in a simplified version, which we summarize below.[46]

4.7 Mechanisms: Wages, Unionization, and Amenity Provision

We highlight several mechanisms affecting wages, unionization, and amenity provision.

Illustration of Wage Functions. To illustrate how wages relate to skills, amenity costs, and union status, we consider a static setting with risk-neutral workers, i.e., $u_x(w,a) = w + v_x(a)$, take a to be binary with $v_x(0) = c_x(0) = 0$, and impose $\phi = 0$. Solving the first-order conditions of the bargaining problems yields closed-form expressions for the union and nonunion wage functions, which we derive in Appendix B.7. Two observations follow from these expressions.

to standard values in the literature (typically 50–90% of wages), the surplus available for bargaining remains small, and even a substantial gap between β_u and β_n fails to generate a wage premium as large as that observed in the data.

[46] The numerical algorithm for solving the general model is provided in Appendix B.9. In all cases, the algorithm converges to a unique equilibrium quickly.

First, holding $\boldsymbol{g}$ fixed, the effect of amenity provision on wages is:

$$w_{x,u}(\boldsymbol{g},1) - w_{x,u}(\boldsymbol{g},0) = \underbrace{-v_x(1) + \beta\bar{v}(1)}_{\text{Compensating diff.}} - \underbrace{\beta\left[\bar{c}(1) + \frac{FC^a}{\sum_{x'} g_{x'}}\right]}_{\text{Cost pass-through}}, \tag{21}$$

$$w_{x,n}(\boldsymbol{g},1) - w_{x,n}(\boldsymbol{g},0) = \underbrace{-(1-\beta)v_x(1)}_{\text{Compensating diff.}} - \underbrace{\beta c_x(1)}_{\text{Cost pass-through}}, \tag{22}$$

where $\bar{v}(a) = \frac{1}{\sum_{x'} g_{x'}} \sum_{x'} v_{x'}(a) g_{x'}$ and $\bar{c}(a) = \frac{1}{\sum_{x'} g_{x'}} \sum_{x'} c_{x'}(a) g_{x'}$ are workforce averages. Three features of collective bargaining stand out: compensating differentials are partly pooled (via $\beta\bar{v}(1)$), the variable cost of amenities is pooled (via $\bar{c}(1)$), and a share of the fixed cost is passed through to workers in proportion to β.

Second, fixing $(\boldsymbol{g}, a)$, the union wage premium $\Delta w_x(\boldsymbol{g}, a) = w_{x,u}(\boldsymbol{g}, a) - w_{x,n}(\boldsymbol{g}, a)$ decomposes as:

$$\Delta w_x(\boldsymbol{g},a) = \beta \underbrace{\left[\frac{F(\boldsymbol{g})}{\sum_{x'} g_{x'}} - \varkappa \frac{\partial F(\boldsymbol{g})}{\partial g_x}\right]}_{\text{Avg. prod. vs. scaled marginal prod.}} + \beta\left[c_x(a) - \bar{c}(a) - \frac{FC^a}{\sum_{x'} g_{x'}}\right] + \beta(\bar{v}(a) - v_x(a)), \tag{23}$$

where $\varkappa = \frac{1}{1-(1-\alpha)\beta} \geq 1$. The first bracketed term captures the wedge between the average product of labor and a scaled version of its marginal product. Individual bargaining ties each worker's compensation to her marginal product while collective bargaining ties it to the average, creating a preference for unionization that strengthens as returns to scale fall. With skill heterogeneity, the same averaging implies *cross-skill transfers* under collective bargaining, so workers with higher marginal product receive a smaller union wage premium than workers with lower marginal product. Collective bargaining therefore redistributes rents from high- to low-skill workers, and reduces wage inequality. Appendix B.7 formalizes both points.

Union Threat Cost. The union threat cost (equation (6)) forces nonunionized firms to account for their workers' endogenous preferences for unionization. As shown above, these preferences are heterogeneous across skill types — low-skill workers prefer unions because collective bargaining redistributes rents in their favor — while, under decreasing returns to scale, firms themselves prefer individual bargaining since the sum of marginal contributions falls short of total output.[47] Nonunionized firms may respond to the threat by distorting their skill mix away from the optimum, but doing so entails production inefficiency. When the union threat cost or these production losses are sufficiently large, some firms find it optimal to unionize. Firms with larger returns to scale are more likely to do so, since a less concave production function reduces the surplus advantage of individual over collective bargaining. Because firms with higher α_y tend to be larger, the model

[47] In a simple case with risk-neutral workers, no amenity provision, no union effect on job security, no break-up cost ϕ, and no union threat cost, firms always choose nonunionization (Taschereau-Dumouchel, 2020).

generates a positive correlation between firm size and unionization, consistent with the data. Heterogeneity in TFP A_y alone would not produce this pattern: different TFP levels affect production scale but not the relative surplus extraction under individual versus collective bargaining.

Hold-up Problem. Individual bargaining creates a hold-up problem that makes amenity provision costly for nonunionized firms. The fixed cost $FC^a_{y,k}$ is sunk at the time of individual bargaining and is borne entirely by the firm, whereas under collective bargaining it enters the firm's surplus from reaching an agreement and a fraction of the fixed cost is passed through to the unionized workers. This gives amenity-providing firms an additional incentive to unionize. This channel is especially relevant in our quantitative analysis because we take health insurance as the non-wage benefit, which involves sizable fixed costs (Karaca-Mandic et al., 2011 and Aizawa and Fang, 2020).[48]

Efficient Cost Sharing. Collective bargaining allows amenity costs to be shared among heterogeneous workers through wage adjustments, which can also give unionized firms an advantage in amenity provision. To see this formally, we consider the same static setup as above, but assume a utility function $u(w) + v_x(a)$ that is concave in the wage. As shown in Appendix B.8, the first-order condition of the collective bargaining problem equalizes worker surplus in marginal-utility units across types. Let $c = \frac{1}{\sum_{x'} g_{x'}}[\sum_{x'} c_{x'}(1)g_{x'} + FC^a]$ denote the per-worker cost of providing amenities, under CRRA utility with curvature θ the wage response is

$$\frac{\partial w_x}{\partial c} = -\frac{\beta}{1-\beta}\frac{w_x}{w_x + \theta S}\left(1 - \frac{\chi}{1+\chi}\right), \tag{24}$$

where $S > 0$ is the firm's surplus per worker scaled by relative bargaining power and $\chi > 0$ is a type-invariant term (both defined in Appendix B.8). In the risk-neutral case ($\theta = 0$), the wage response reduces to $-\beta$ uniformly across types; with concave utility ($\theta > 0$), wages fall more sharply for high-wage workers. Collective bargaining thus implements efficient cost sharing, with costs disproportionately absorbed by workers with lower marginal utility of consumption.

This has an important implication for amenity provision. Under individual bargaining, amenity costs are not shared across skill types, and the presence of a consumption floor may prevent firms from offering amenities because low-skill workers are unwilling to pay for them.[49] Under collective bargaining, by contrast, the cost-sharing rule described above can make amenity provision viable even when low-skill workers' willingness to pay falls short of the cost, since high-wage workers absorb a disproportionate share through wage adjustments. As a result, unionized firms may provide amenities that nonunionized firms do not. This channel also helps explain why union-

[48] In Appendix A.11, we provide supportive evidence for this fixed-cost channel. Specifically, we use union election data and show that the Medicare introduction significantly reduces union elections among small establishments, where the fixed cost matters, while it does not reduce those among large establishments.

[49] In the context of health insurance, a consumption floor leads low-income workers to have a smaller willingness to pay for employer-sponsored health insurance.

ization tends to compress wages and reduce earnings inequality across skill types.

Cost Pooling Across Unionized Firms. Beyond the within-firm cost sharing discussed above, our model allows for the possibility that the fixed cost of amenity provision is lower for unionized firms than for nonunionized firms, and that this fixed cost for union firms decreases as more type-y firms unionize and offer the amenity *in equilibrium*. This gives unionized firms an additional incentive to provide amenities, i.e., as more firms unionize and offer the amenity, the per-firm fixed cost falls, making amenity provision more attractive for unionized firms; interestingly, this positive externality can also be a source of a socially sub-optimal level of unionization.

This formulation is empirically grounded. For example, unionized employers negotiate health insurance (ESHI) on behalf of a larger and more concentrated group of workers, which gives them greater bargaining power against insurers and providers. This manifests in lower loading fees and administrative costs (Karaca-Mandic et al., 2011) and in the ability to negotiate volume discounts and direct provider contracts that nonunionized employers cannot access individually (Lechner and Draper, 2014). Moreover, the Taft–Hartley Act of 1947 permits joint employer-union benefit trusts through which multiple unionized employers contribute to collectively administered benefit plans covering health insurance and other insurance and welfare benefits. These arrangements institutionalize cost pooling across unionized firms.

Job Security. Finally, with a lower rate of job destruction in unionized firms ($\delta_{x,u} < \delta_{x,n}$), as empirically documented in Section 3.1, firms can provide better job security to workers through unionization. Better job security increases the duration of a match, generating a larger surplus for the firm by reducing the future hiring cost needed to maintain the same number of workers. This provides another incentive to unionize. However, better job security should come at a cost. In particular, it could result in costly labor hoarding when a firm is hit by a negative productivity shock and wants to scale down. Although we abstract from productivity shocks in the model, the fixed cost FC^{union} captures such a cost in a reduced-form way.[50, 51] We provide microfoundation for this interpretation using a simplified setting in Appendix B.4.

4.8 Quantitative Specification and Extension

Although our model considers general non-wage benefits, given our empirical evidence in Section 3 and the fact that healthcare consists of a sizable part of the U.S. economy (Hall and Jones, 2007), we now focus on health insurance as the amenity; as such, a is now binary: $a = 1$ if a worker is insured and $a = 0$ otherwise. Each worker faces a medical expenditure shock m drawn

[50]Since firms are risk-neutral, whether FC^{union} is a one-shot cost or a lump-sum cost does not matter for this interpretation.

[51]It is worth noting that the discussions above hold conditional on the hiring profile $\boldsymbol{g}$; therefore, they are not the result of the over-hiring inefficiency implied by Stole and Zwiebel (1996) bargaining.

from a skill-specific distribution H_x and has utility:

$$u_x(w,a) = \int \log C(w,a,m) dH_x(m), \tag{25}$$

where $C(w,a,m)$ is the consumption level as a function of wage w and insurance a provided by the firm and medical cost m.[52] The consumption level is given by $C(w,a,m) = \max\{w - OOP(m;a), \underline{c}\}$ where $\underline{c}$ is the consumption floor, and $OOP(m;a)$ is the out-of-pocket medical expenditure that depends on a worker's health insurance status.

We set $X = 2$ skill types: low-skill workers ($x = 1$) have at most high school education, and high-skill workers ($x = 2$) have at least some college education.

Given our interests in social insurance, we model the public health insurance system more realistically. Specifically, we model Medicaid as follows: The fraction p_x^{Med} of type-x workers become eligible for Medicaid upon unemployment, and remains eligible until reemployment. The ex-ante unemployment value is now given by $V_x^U = p_x^{Med} V_x^U(1) + (1 - p_x^{Med}) V_x^U(0)$ where $V_x^U(i) = p(\theta_x) V_x^M + (1 - p(\theta_x))[u_x(b_x, i) + \gamma V_x^U(i)]$ denotes the unemployment value with ($i = 1$) and without ($i = 0$) Medicaid coverage.

We also impose additional assumptions on the firm side. Motivated by our theoretical discussion of firm size and unionization in Section 4.7, the benchmark specification assumes that firms differ in their returns-to-scale parameters α_y but not in TFP A_y: in Appendix D.2.2, we also show that our main findings are robust to the extended model with TFP heterogeneity. We assume that α_y is drawn from a (transformed) Beta distribution, *Beta*(a,b), on the support $[0.5, 0.9]$. To capture the empirical sorting between skill and firm size, we let skill intensity depend on α_y: $z_{1,y} = \bar{z} - a_z(\alpha_y - \bar{\alpha}), z_{2,y} = 1 - z_{1,y}$ where $\bar{z}$ is the average low-skill intensity across firms and a_z captures the extent to which firms with higher returns to scale rely on low-skill labor. The cost shocks for amenities $\boldsymbol{\epsilon} = \{\epsilon_a\}_{a \in \{0,1\}}$ and for unionization $\boldsymbol{\varepsilon} = \{\varepsilon_k\}_{k \in \{u,n\}}$ are drawn from the type-I extreme value distributions with scale parameters σ_a and σ_{union}. The amenity fixed cost, which incorporates the cost-pooling benefit of unionization, takes the following form:

$$FC_{y,k}^a = \frac{FC^a}{1 + \mathbb{1}\{k = u\} \lambda K_y}, \tag{26}$$

where K_y is the measure of type-y unionized firms offering insurance. For nonunionized firms ($k = n$), the fixed cost is FC^a; for unionized firms ($k = u$), it decreases in K_y, capturing cost pooling among union firms. The parameter for the cost pooling intensity, λ, is estimated.

Later in the decomposition analysis in Section 6.3, we also introduce retiree ESHI coverage

[52] We adopt log utility to impose lower risk aversion than the estimates or calibrated values commonly used in the literature, e.g., French and Jones (2011). The reason is that our model abstracts away from a self-saving technology, thus we would like to avoid overstating the value of the insurance channel permitted in the model by restricting ourselves to the log utility function.

and Medicare. Workers retire with exogenous probability p^R and exit the model with probability p^D. The retirement value is the expected discounted sum of post-retirement utilities, $V_x^R(a) = \frac{u_x(c_x^R, a)}{1-\gamma(1-p^D)}$, where a denotes pre-retirement ESHI coverage. Retirement shocks arrive after job destruction shocks. The values of employment $V_{x,y,k}^E$ and unemployment V_x^U are modified to incorporate transitions into retirement, and unemployment value further splits according to Medicaid eligibility; the full expressions are given in Appendix C.1.

These modifications enter the firm's problem in three ways. First, the effective separation rate becomes $\delta_{x,k}^{\text{eff}} = \delta_{x,k} + (1-\delta_{x,k})p^R$, as retirement is an additional source of separation. Second, the wage bargaining problem is unchanged except that worker values now also reflect the future gains from employer-provided retirement insurance. Third, the Beveridge curve is slightly modified because each retiring worker is replaced by an identical unemployed worker (see Appendix C.1).

In the extended model, firms also bear the cost of providing ESHI to retired workers. To preserve the structure of the firm's problem, we replace $c_x(a)$ with an *effective per-period cost* $\tilde{c}_x(a)$ paid only during employment but chosen so that, in expectation, it covers both the worker's current and post-retirement ESHI costs. As derived in Appendix C.1, this yields

$$\tilde{c}_x(a) = \left[1 + \frac{\gamma p^R}{1-\gamma(1-p^D)}\right] c_x(a). \tag{27}$$

The scaling factor $\frac{\gamma p^R}{1-\gamma(1-p^D)}$ reflects the discounted probability of retirement and the expected duration of retirement: a higher p^R, or a smaller p^D, raises the share of post-retirement ESHI cost amortized across each period of employment. Without retirement ($p^R = 0$), the scaling factor vanishes, $\tilde{c}_x(a) = c_x(a)$, and the firm's problem reduces to the baseline.

5 Estimation

5.1 Externally Set or Estimated Parameters

We target the 2007 U.S. economy. Each model period is a quarter. Several model parameters, summarized in Table A.11 in the Online Appendix, are taken directly from the literature or estimated outside the model. The discount factor is set to $\gamma = \frac{1}{1+r}$ with $r = 1.05^{1/4} - 1$, corresponding to an annual interest rate of 5%. We set the measure of firms to $M = 0.042$, matching the average firm size of 22.56 derived from the Census Business Dynamics Statistics (BDS).[53] The elasticity of substitution between skill types in the production function (8) is set to $\sigma = 1.5$ (Johnson, 1997). The consumption floor per quarter $\underline{c}$ is set to \$1,000 (French and Jones, 2011). We specify the matching function as $m(s,\nu) = s\nu/(s+\nu)$, following Den Haan et al. (2000). The matching

[53]The average firm size in the model $\frac{1-\mathcal{U}}{M}$ also depends on the endogenous unemployment rate $\mathcal{U}$. We plug in the targeted unemployment rate from the estimation to calculate this number.

efficiency is normalized to 1 as this parameter and the vacancy creation cost κ cannot be separately identified using the unemployment rate. b_x includes both unemployment insurance benefits and other sources of non-labor income. For each skill type, we set b_x to 70% of the average wage, which corresponds to the lower end of the relative consumption during unemployment reported in Chodorow-Reich and Karabarbounis (2016). We set worker's bargaining power to $\beta = 0.5$.

Job destruction rates are allowed to depend on both skill type and union status. Using SIPP data, we estimate the effect of union status on subsequent job loss and apply the estimate to adjust the separation probabilities in the CPS sample. For unionized workers, we set $\delta_{1,u} = 0.0549$ and $\delta_{2,u} = 0.0276$; for nonunionized workers, $\delta_{1,n} = 0.0639$, and $\delta_{2,n} = 0.0313$.

The medical-expenditure distribution $H_x(m)$ is parameterized as a mixture of a log-normal component $LN(\mu_{H,x}, \sigma^2_{H,x})$, and a point mass $p_{0,x}$ at zero, and is estimated the 2007 Medical Expenditure Panel Survey (MEPS). Note that $OOP(m; a)$ depends on the characteristics of the insurance contract. Following Aizawa (2019), we use the characteristics of representative employer-sponsored plans reported by Sommers and Crimmel (2008), setting the annual deductible at \$714 and the coinsurance rate at 18%. We compute the average per-worker insurance cost $c_x(\cdot)$ using the estimated medical-expenditure distribution $H_x(m)$ and the contract parameters above.

We calibrate Medicaid eligibility p_x^{Med} using the type-specific Medicaid coverage rates in the CPS. Among low-skill workers, 16% are covered by Medicaid, compared with 9% among high-skill workers. In practice, not all eligible workers take up coverage. Using the average take-up rate of 0.52 reported by Davidoff et al. (2005), we obtain $p_1^{Med} = 0.31$ and $p_2^{Med} = 0.17$.

5.2 Internally Estimated Parameters

We identify and estimate the remaining parameters jointly within the model. These parameters are: the TFP A; the Beta distribution parameters *Beta*(a, b) for returns to scale α_y; the mean low-skill intensity $\bar{z}_1$ and its slope a_z with respect to α_y; the baseline fixed cost of insurance FC^a; the union cost pooling intensity λ; the scale parameter σ_a of the insurance cost shocks; the fixed cost of unionization FC^{union}; the scale parameter σ_{union} of the union cost shocks ε_u and ε_n; the marginal cost c_0 in the union threat and maintenance cost functions; the vacancy posting cost κ; and the firm's break-up cost ϕ of collective bargaining. We first present heuristic identification arguments based on variations in unionization, firm size, compensation, and employment, and then examine moment informativeness.

Insurance-Related Parameters. The first set of parameters governs insurance provision: $(FC^a, \sigma_a, \lambda)$. To illustrate identification, consider a simplified static version of the firm's insurance decision that abstracts from dynamics, bargaining, and skill heterogeneity. A type-y firm of size n_y earns profits $\pi_{y,k}(a) = F_y(n_y) - n_y \cdot w_{y,k}(a) - \left(n_y \cdot c + FC^a_{y,k}\right) \cdot \mathbf{1}\{a = 1\}$ where $a \in \{0, 1\}$ denotes insurance provision, and the firm chooses a after drawing Type-I extreme value shocks (ϵ_0, ϵ_1) with scale σ_a. Since workers value insurance, firms can lower wages when offering coverage. Let $\Delta w_{y,k} = w_{y,k}(0) - w_{y,k}(1)$ denote the wage reduction enabled by providing coverage: workers' risk aversion implies $\Delta w_{y,k} > c$. Hence, $\pi_{y,k}(a = 1) - \pi_{y,k}(a = 0) = n_y \cdot (\Delta w_{y,k} - c) - FC^a_{y,k}$, and the implied insurance provision probability is

$$P_{y,k}(a = 1) = \Lambda\left(\frac{n_y \cdot (\Delta w_{y,k} - c) - FC^a_{y,k}}{\sigma_a}\right),$$

where $FC^a_{y,n} = FC^a$ and $FC^a_{y,u} = FC^a/(1 + \lambda K_y)$, and $\Lambda(\cdot)$ is the logistic function. We denote by $n^*_{y,k} = FC^a_{y,k}/(\Delta w_{y,k} - c)$ the indifference firm size at which expected profits from providing and not providing insurance are equal. This indifference size $n^*_{y,k}$ is increasing in $FC^a_{y,k}$.

The model predicts that insurance provision rises with firm size. The overall level of ESHI offer rates identifies FC^a: a higher fixed cost shifts the insurance threshold toward larger firms and lowers offer rates at every size. The slope of the size gradient identifies σ_a: smaller values of σ_a generate a sharper transition in coverage around the threshold, while larger values flatten it. We therefore identify (FC^a, σ_a) using ESHI offer rates at small and large firms.

The parameter λ is identified from the size-adjusted union effect on ESHI coverage. The unconditional union-nonunion gap reflects two channels: selection (since unionized firms are larger) and the pooling advantage captured by λ. We isolate the latter using the size-controlled union effect in Section 3. Holding (FC^a, σ_a) fixed, a larger λ lowers the fixed cost for unionized firms to offer insurance, thus shifting the insurance threshold $n^*_{y,u}$ leftward for unionized firms, and increasing ESHI provision at any given firm size.

Union-Related Parameters. The second set of parameters governs unionization: $(c_0, FC^{union}, \sigma_{union}, \phi)$. As discussed in Section 4.7, firms have no incentive to unionize in the absence of the union threat and the cost pooling benefits of unionization. When $c_0 > 0$, firms may unionize to avoid the cost $C_{y,n}(\boldsymbol{g}, a)$. Because this cost rises linearly with firm size while profits grow sublinearly under decreasing returns to scale, c_0 primarily disciplines unionization among large firms. In contrast, FC^{union} governs unionization among small firms, while σ_{union} smooths the relationship between firm size and unionization. We separately identify these parameters using overall union density and the joint distribution of unionization and firm size.

The breakup cost ϕ governs the union wage premium. Conditional on the bargaining weight β, moments capturing union-nonunion wage differentials identify ϕ.

Other Parameters. Since α_y shapes firm size, the firm size distribution identifies the Beta distribution parameters (a, b). TFP A is identified by average wages. The skill intensity parameters $\bar{z}_1$ and a_z are identified from the relative wages across skills and the relationship between firm size and

Table 4: List of Internally Estimated Parameters

Parameters	Description	Estimate	SE
A	TFP	40.3	1.04
$\mathit{Beta}(a,b):a$	Returns to scale distribution	0.72	0.19
$\mathit{Beta}(a,b):b$	Returns to scale distribution	0.71	0.05
$\bar{z}_1$	Low-skill worker relative skill: average	0.48	0.04
a_z: slope	Low-skill worker relative skill: slope	1.51	0.34
FC^a	Fixed cost of insurance provision	9.0	1.58
λ (1e-3)	Union ESHI cost advantage	1.20	0.52
σ_a	Std. dev. of insurance cost shock	9.04	1.46
FC^{union}	Fixed cost of unionization	21.7	7.63
σ_{union}	Std. dev. of union cost shock	4.02	1.61
c_0 (1e-2)	Marginal cost of union threat	1.99	0.37
κ	Vacancy posting cost	2.99	0.78
ϕ	Break-up cost	15.9	5.1

Note: This table reports the estimated model parameters and standard errors. Monetary values are 2007 USD.

skill composition. Finally, the vacancy posting cost κ is identified using the unemployment rate.

Estimation Strategy. We estimate these parameters via the minimum distance estimator. Following the identification arguments above, we select targeted moments that include (i) cross-sectional moments on ESHI coverage, wages, union density, firm size, and skill composition; (ii) employers' ESHI offer rates; and (iii) regression coefficients capturing the effects of unions on ESHI and wages. See Table 5 for the full list.[54] We minimize the objective function $Q(\vartheta) = [\hat{\boldsymbol{m}} - \boldsymbol{m}(\vartheta)]'\boldsymbol{W}[\hat{\boldsymbol{m}} - \boldsymbol{m}(\vartheta)]$, where ϑ is a vector of parameters to be estimated (listed in Table 4), $\boldsymbol{m}(\vartheta)$ is a vector of model moments based on ϑ, and $\hat{\boldsymbol{m}}$ is a vector of empirical moments. $\boldsymbol{W}$ is a diagonal weighting matrix whose entries are the diagonal elements of the inverse of the empirical-moment covariance matrix. We compute standard errors based on the asymptotic variance.

5.3 Estimation Results

Parameter Estimates. Table 4 reports the estimated parameters of the model. We estimate the TFP A at 40.3, which implies that, based on our production function (8), a firm employing one low-skill and one high-skill worker produces \$40,300 per quarter. The Beta distribution parameters of α_y are estimated at 0.72 and 0.71, translating into average returns to scale of approximately 0.75. This estimate is in line with the values reported in the literature (e.g., Elsby and Michaels 2013; Cooper et al. 2015), although these studies rely on different models and identifying moments. The average skill intensity of low-skill workers is 0.48, with an estimated slope of 1.51. Together, these

[54] The moments in (i) are drawn from the CPS, while those in (ii) are from the Robert Wood Johnson Foundation Employer Health Insurance Survey. For the regression coefficients, the estimated union coefficient for ESHI is reported in Table 2, while the union wage differential is estimated from the regression $w_{it} = \beta \cdot \mathit{Union}_{it} + x_{it}'\gamma + \lambda_t + \epsilon_{it}$ using CPS data.

parameters imply that the skill intensity of low-skill workers ranges from 0.26 in the largest firms (those with the highest α_y) to 0.71 in the smallest firms.[55] The baseline fixed cost of insurance FC^a is about \$9,000 per quarter, while the standard deviation of the insurance cost shock σ_a is estimated to be of similar magnitude. The estimated intensity of the pooling benefit among unionized firms, λ, implies that the average union firm incurs only 28% of the insurance fixed cost borne by its nonunionized counterparts. The fixed cost of unionization FC^{union} is approximately \$21,700 per quarter, with a cost shock standard deviation σ_{union} of \$4,020. The marginal cost of the union threat is estimated at $c_0 = 0.02$, meaning that for every \$100 of workers' aggregate willingness to pay for unionization, firms must incur \$2 to suppress it. The vacancy posting cost is estimated at \$3,000. Finally, the firm's (one-shot) break-up cost of union collective bargaining, ϕ, is estimated at approximately \$16,000, which is equivalent to roughly 4-5 months of average wages.

Sensitivity of Estimates. In Appendix C.2, we complement the identification discussion with an informativeness analysis in the spirit of Honoré et al. (2020). The results are consistent with the identification arguments developed above. For example, we find that the estimated effect of unions on ESHI provision is especially informative about the parameter capturing the union cost-pooling intensity λ.

Model Fit. Table 5 shows the fit of the estimated model. As shown, the model can account for various important patterns in the data, including the positive relationships between union status and insurance coverage, as well as between firm size and union status.[56]

It should be noted that our model generates reasonable union wage premiums for each skill type using a single parameter ϕ. In our CPS sample, the union wage premium is 0.12 for low-skill workers and 0.04 for high-skill workers, controlling for firm size and industry. The overall magnitude is consistent with Farber et al. (2021), who report an average union wage premium of around 0.1 in recent years. Note also that although collective bargaining involves transfers from high-skill workers to low-skill workers, high-skill workers still receive a positive union wage premium, consistent with Beauregard et al. (2025), who use Canadian matched employer-employee data to estimate union wage premiums.

6 Counterfactual Experiments

In this section, we conduct various counterfactual experiments to understand the mechanisms of how insurance provision interact with labor unions, and how social insurance policies affect

[55] A skill intensity of low-skill workers above 0.5, exceeding that of high-skill workers, does not imply that low-skill workers are more productive (i.e., have a higher marginal product) due to production complementarity.

[56] The model predicts a firm-level unionization rate of 3.5%. The corresponding one in the 1997 Robert Wood Johnson Foundation Employer Health Insurance Survey is 3.2% (Table 1), despite not being targeted in the estimation.

Table 5: Model Fit

Moment	Model	Data	Moment	Model	Data
(a) Cross-sectional moments					
ESHI coverage			*Firm size*		
ESHI coverage: union	0.87	0.90	Emp. share 10+: union	0.98	0.95
ESHI coverage: nonunion	0.71	0.72	Emp. share 10+: nonunion	0.82	0.85
ESHI coverage: low skill	0.66	0.62	Emp. share 100+: union	0.77	0.80
ESHI coverage: high skill	0.77	0.82	Emp. share 100+: nonunion	0.48	0.59
Wages and labor market			*Skill composition*		
Average wage: low skill	8.27	8.41	Emp. share 10+: low skill	0.70	0.82
Average wage: high skill	14.8	14.8	Emp. share 10+: high skill	0.92	0.89
Union density	0.10	0.10	Emp. share 100+: low skill	0.40	0.55
Unemployment rate	0.05	0.05	Emp. share 100+: high skill	0.58	0.66
(b) ESHI offer rates					
ESHI offer rate: small (<50)	0.35	0.48	ESHI offer rate: large (50+)	0.91	0.95
(c) Regression coefficients					
Union ESHI effect	0.05	0.06	Union wage premium: low skill	0.12	0.12
			Union wage premium: high skill	0.04	0.04

Note: This table reports the targeted data moments and their simulated counterparts. "Emp. share x+: union" ("nonunion") is the share of union (nonunion) workers employed at firms whose total workforce is at least x where the numerator sums employment across unionized (nonunionized) firms with $\geq x$ total employees and the denominator sums employment across all unionized (nonunionized) firms. "Emp. share x+: low (high) skill" is defined analogously, restricting the numerator and denominator to low- (high-)skill employment, while the size cutoff x continues to refer to the firm's total employment across both skill groups.

unionization rates and other labor market outcomes. We also use our model to show that social insurance expansions could be an important factor in understanding union declines in the U.S.

6.1 Mechanisms: Equilibrium Effects of the Union Threat

We use the estimated model to explore the role of the union threat cost for nonunionized firms $C_{y,n}(\boldsymbol{g}, a)$ as in equation (6), by setting it to zero. To isolate the impact of how the union threat cost affects the behavior of firms *conditional on* their union status, we fix the union status of each firm at the baseline level in line with Taschereau-Dumouchel (2020).

Table 6 shows the results of this counterfactual experiment. Without the union threat cost, nonunionized firms no longer hesitate to hire low-skill workers for fear of unionization. This expansion in low-skill hiring by nonunionized firms also shows up in skill-specific unemployment rates: the low-skill unemployment rate falls significantly.

A notable impact of eliminating the union threat cost appears in the ESHI coverage. Without the cost of union threat, the overall ESHI coverage decreases by about 18 percentage points; interestingly and surprisingly, this decrease is primarily driven by the lower ESHI offerings from nonunionized firms. This results from how the change in skill composition of workers in nonunionized firms towards low-skill workers interacts with the bargaining protocol. As discussed in 4.7,

Table 6: Impact of Removing the Union Threat Cost

	Baseline	No Union Threat
ESHI coverage (%)		
Overall	72.9	55.0
Unionized firms	86.6	77.7
Nonunionized firms	71.3	53.4
Unemployment rate (%)		
Overall	4.7	3.1
Low skill	9.1	5.6
High skill	1.6	1.4
Labor productivity (% change)	-	-0.7
Average wage (% change)	-	1.9
Skill wage gap (log pts change)	-	-1.7

Note: This table reports the impact of removing the union threat cost by setting $c_0 = 0$. The union status of each firm is fixed at the baseline.

individual bargaining in nonunionized firms precludes efficient sharing of the insurance cost among heterogeneous workers. It is particularly costly for nonunionized firms to offer insurance to low-skill workers because they cannot pass some of the cost on to high-skill workers, whose marginal utility of income is lower than that of low-skill workers. The expansion of nonunionized firms in hiring low-skill workers, interacted with individual bargaining, reduces the ESHI access for all workers in nonunionized firms.[57]

Note that the increase in low-skill employment translates into lower labor productivity due to the compositional change among employed workers. However, the average wage increases in the counterfactual experiment, partially because workers are compensated for the reduction in ESHI coverage, and partially because removing the union threat cost raises firms' profitability, which is then passed on to workers through higher wages. The increased demand for low-skill labor also pushes up wages for low-skill workers, contributing to a smaller skill wage gap in the counterfactual.

6.2 Social Insurance Design and Labor Unions

In this section, we study how social insurance programs affect unionization and labor market equilibrium. In the U.S., many such programs operate either by subsidizing or mandating employer-provided insurance (e.g., ESHI, workers' compensation, or employment protection policies) or through the direct provision of public insurance. Below, we consider stylized versions of each of these programs.

[57]ESHI coverage also declines among union workers, albeit to a lesser extent. Absent threat costs, nonunion firms expand by hiring more low-skill workers, thereby increasing equilibrium hiring costs for union firms and inducing them to downsize, which in turn reduces ESHI provision.

Table 7: Counterfactual Policy Simulation: Insurance Policies

	(1) Baseline	(2) Insurance subsidy	(3) Targeted SI
Union density (%)	10.30	6.62	9.55
ESHI coverage (%)			
Overall	72.9	80.8	70.8
Union	86.6	92.4	83.7
Nonunion	71.3	80.0	69.5
Low skill	66.2	73.9	64.3
High skill	77.1	85.2	74.9
Unemployment rate (%)			
Overall	4.7	4.3	5.7
Low skill	9.1	8.5	11.5
High skill	1.6	1.5	1.7
Output per capita (% change)	-	0.26	-0.65
Labor productivity (% change)	-	-0.10	0.42
Average wage (% change)	-	-0.25	0.38
Skill wage gap (log pts change)	-	1.37	-1.77

Note: This table reports the general equilibrium impacts of each policy change. Column (2) is the economy with free public health insurance only for low-skill unemployed workers. In column (3), firms receive subsidies for offering insurance to their workers.

6.2.1 ESHI Cost Subsidy

First, we consider the effect of subsidies for employer's insurance provision. In the U.S., many employer-sponsored insurance benefits are tax exempt, which incentivizes firms to provide those benefits. To examine the equilibrium effects of this form of social insurance policies on unionization, we model subsidies as reductions in both the fixed and flow costs of providing insurance. Specifically, we reduce the fixed cost by 20%, and the flow cost by 20% for high-skill workers and by 10% for low-skill workers. The subsidies are modeled as direct reductions in firms' insurance costs and are not tax-financed, which isolates their equilibrium effect on unionization from the distortions of how they are funded. This differential treatment is intended to capture, in a simple way, the higher marginal tax advantage for high-wage workers relative to low-wage workers.

Column (2) in Table 7 shows the outcomes under the subsidy for the firm's provision of insurance for their workers. Recall that unionized firms have an advantage in insurance provision in terms of both fixed cost and flow cost (see Section 4.7). By reducing the union advantage in insurance provision, the insurance subsidy reduces the union density by 3.7 p.p.[58] As a result, the insurance subsidy, which may be intended to help workers, has unintended consequences. By reducing the coverage of collective bargaining, the subsidy reduces the average wage by 0.25% and widens the wage inequality as the skill wage gap increases by 1.37 log points. The sub-

[58]The deunionization effect of the ESHI subsidy is robust to specifications in which the subsidy affects only the flow cost or only the fixed cost.

sidy increases the ESHI coverage of nonunionized workers more than that of unionized workers, shrinking the difference in ESHI coverage between unionized and nonunionized workers.

This finding has several implications. First, policies such as the tax deductibility of employer-sponsored insurance benefits can lower unionization by weakening the union's advantage in insurance provision. Second, subsidizing insurance provision can also contribute to wage inequality. This result complements the existing arguments that the tax exemption of these benefits has regressive effects when income tax is progressive, because our finding suggests that even *pre-tax* income could be affected, leading to further consumption inequality between the skilled and the less skilled workers.

In Appendix D.3, we examine additional employer-based social insurance policies, including employer mandates for ESHI provision and increased job security for all workers. Consistent with the policies discussed above, these interventions increase employers' provision of insurance, thereby reducing unions' comparative advantage in providing such benefits. As a result, union density declines and the skill wage gap widens.

6.2.2 Targeted Social Insurance Policies

We next examine the effect of government-provided insurance. Many of these programs target low-income population (e.g., Medicaid). We consider social insurance provision targeted only at low-skill unemployed workers financed by a uniform payroll tax on firms.

Column (3) in Table 7 shows the equilibrium impact of the above targeted social insurance for low-skill workers. By providing public insurance outside their jobs, this policy reduces the low-skill workers' incentives to obtain health insurance through unions, reducing the union density by 0.75 p.p. The higher marginal hiring cost increases the unemployment rate by 1.0 p.p., reducing total output by 0.65%. The decline in unions comes with a slight improvement in labor productivity of 0.42% due to the compositional change in the skills of employed workers.

The decline in unions is also damaging to high-skill workers in terms of their ESHI coverage, reducing their coverage by 2.2 p.p. Although low-skill workers see a similar decline in ESHI, 11.5% of them are unemployed and therefore receive free public insurance under the targeted social insurance provision.

Although the union decline reduces the coverage of collective bargaining, the policy change directly increases the wages of low-skill workers by improving their outside options, suppressing wage inequality. In total, the effect of improved outside options dominates the impact of the union decline, reducing the skill wage gap by 1.77 log points.

Our results highlight the difference between the insurance subsidy that targets all workers regardless of their skill type, and the targeted social insurance provision for low-skill unemployed workers only. Although both reduce the union density, their impacts on wage inequality also differ sharply: the insurance subsidy increases the skill wage gap by 1.37 log points through the decline of unions, while targeted social insurance reduces it by 1.77 log points by directly improving the outside options of low-skill workers. However, the decline in unions under targeted social insurance reduces the ESHI coverage for high-skill workers, in contrast to that under insurance subsidy

6.3 Long-run Implications

We now use our model to quantify the relative importance of historical expansions of social insurance policies with other relevant factors that led to the decline in unions. Motivated by our empirical evidence, we focus on the implementation and expansion of Medicare and Medicaid.[59] Moreover, following the literature, we evaluate skill-biased technological changes that favor high-skill workers and the adoption of the Right-to-Work laws.

6.3.1 Decomposition of the Union Decline in the U.S. During 1955-2007

We simulate the expansion of social insurance, skill-biased technological change, and the adoption of RTW laws between 1955 and 2007. We begin from the 2007 economy—where social insurance has expanded, skill-biased technological change has taken place, and RTW laws have been more widely implemented—and then examine counterfactual outcomes in which each factor is set to its 1955 level.[60] We briefly outline their implementation below and defer additional details to Appendix D.1.

In the 2007 economy, we assume that all retired workers have access to government-provided health insurance through Medicare, while unemployed workers have partial access to public coverage through Medicaid. To quantify the impact of Medicare and Medicaid, we consider a counterfactual in which unemployed workers have no access to public insurance and retired workers rely solely on retiree health insurance provided by their former employers.

We model skill-biased technological change by adjusting the productivity of high-skill workers, $z_{2,y}$, while keeping that of low-skill workers fixed.[61] To discipline the magnitude of the change, we follow the literature (e.g., Katz and Autor, 1999; Krusell et al., 2000), which finds that skill-biased technological change accounts for a substantial but incomplete share of the rise in the skill

[59]Note that we underestimate the effect of social insurance programs because we do not account for many programs that expanded over the last half century, including, but are not limited to, the establishment of OSHA in 1970 and various subsidy programs for employer-provided benefits, including the tax exemption of ESHI premiums.

[60]This approach follows Braxton et al. (2024), who use a similar strategy to evaluate the long-run effects of credit market expansion. Alternatively, we calibrate the model to the 1955 economy and examine the impact of each factor in a forward simulation. While this approach is intuitive, a key limitation is the lack of detailed micro-level data for 1955. As a result, some targeted moments must be calibrated by extrapolating from later years. As reported in Appendix D.1.2, the main results remain both qualitatively and quantitatively similar.

[61]We allow the endogenous variables to change, including unionization, in response to the change in $z_{2,y}$. Therefore, not all changes in the skill premium are attributed to the technological change because some part is attributable to changes in endogenous variables, including unionization in particular.

Table 8: Deunionization Analysis: Contributions

	Social Insurance	Tech Change	RTW Laws
Contribution to union decline (%)	15.8	31.9	0.9

Note: This table reports the fraction of the decline in union density between 1955 and 2007 explained by social insurance, skill-biased technological changes (Tech change), and RTW laws.

premium. Specifically, we use the estimate of Krusell et al. (2000), who attribute approximately 56% of the observed rise to skill-biased technological change, as an external benchmark, and choose $z_{2,y}$ so that the model reproduces this targeted change, as detailed in Appendix D.1.1. In our framework, skill-biased technological change generates two opposing forces. From the firm's perspective, the relative decline in the productivity of low-skill workers disproportionately affects unionized firms, which employ more low-skill workers, thereby contributing to union decline. From the workers' perspective, technological change widens wage inequality across skill groups, increasing the incentives of low-skill workers to seek unionization and raising the threat costs faced by firms attempting to deter it.

RTW laws undermine union sustainability by allowing workers to benefit from collective bargaining without paying union dues. We capture this in reduced form through a cost parameter, c_{RTW}: in the RTW economy, the firm's unionization decision in (16) becomes $J_{y,u}(\boldsymbol{\epsilon}) - c_{RTW} + \varepsilon_u \geq J_{y,n}(\boldsymbol{\epsilon}) + \varepsilon_n$, so that a positive c_{RTW} lowers the probability that a firm unionizes. We calibrate c_{RTW} so that it reduces union density by 2 p.p., which is in line with estimates in Fortin et al. (2022). See Appendix D.1.1 for more details.

We ask how much each factor separately accounts for the observed decrease in union density from about 35% in 1955 to 10% in 2007. Table 8 reports the contribution of each factor to the decline in union density. Social insurance accounts for a sizable 16% of the decline.[62] Technological change explains 32%, while the implementation of RTW laws accounts for 1%.[63]

6.3.2 Robustness

We also examine the robustness of our findings to several modeling assumptions; details are provided in Appendix D.2. First, our baseline abstracts from consumption and saving decisions, which may overstate the insurance value of unions and social insurance, as well as their trade-off. In the Appendix, we show that the quantitative effect of social insurance expansions on union decline remains similar when workers have greater access to self-insurance against medical expen-

[62] Relative to our quasi-experimental estimates in Section 3.2—8% greater Medicare-induced decline in the highest- versus lowest-exposure state and an approximately 5% short-run decline from Medicaid expansion—the 15.8% contribution of social insurance to the 1955–2007 union decline implied by our structural model represents a reasonable magnitude of deunionization.

[63] The small impact of RTW laws partly reflects the fact among states with RTW laws, only 6 introduced RTW laws between 1955 and 2007.

Table 9: Deunionization Analysis: Outcomes

	Social Insurance	Tech Change	RTW Laws
Endogenous union:			
ESHI rate (p.p. change)	-4.6	-2.6	0.0
Skill wage gap (log points change)	-2.0	5.5	0.1
Welfare (% change)	0.9	4.9	0.0
Fixed unionization:			
ESHI rate (p.p. change)	0.6	1.8	-
Skill wage gap (log points change)	-2.1	3.4	-
Welfare (% change)	0.7	1.9	-

Note: This table reports the changes in ESHI rates, skill wage gap, and the welfare between 1955 and 2007 explained by social insurance, skill-biased technological changes (Tech change), and RTW laws.

diture shocks, implemented through a higher consumption floor.[64] Second, our results are robust to allowing for stronger complementarities between total factor productivity and firm scale in the production function. Third, the results are also robust to alternative specifications of union threat costs.

6.4 Welfare Implications

Our findings on the decline in unionization raise two normative questions: (i) are union declines welfare improving responses to skill-biased technological changes and social insurance expansions in the last half century? and (ii) under the current economic environment, should labor unions be taxed or subsidized? This section addresses these questions.

6.4.1 Welfare Impacts of Declining Unionization

We begin by examining the welfare implications of declining unionization in response to social insurance expansions and technological change. To isolate the role of unions, we compare our baseline economy—where unionization rates are endogenously determined—to a counterfactual in which the unionization decision at each firm is held fixed. We measure worker welfare using consumption-equivalent variation (CEV), defined as the percentage change in consumption in the baseline economy that makes workers indifferent between the baseline and the counterfactual equilibrium levels prior to the introduction of the respective changes.[65]

Table 9 reports the main results. For social insurance expansions, overall welfare is higher when unionization adjusts endogenously: welfare increases by 0.7% when unionization is fixed, compared to 0.9% when it responds endogenously. The endogenous decline in unionization is accompanied by a 4.6% reduction in the ESHI rate. As access to health insurance outside the labor

[64] This exercise varies an exogenous self-insurance parameter and does not capture endogenous changes in precautionary savings. We discuss in Appendix D.2 why this concern is likely modest in our setting.

[65] Our welfare analysis abstracts from transitional dynamics, which may differ under labor market mobility frictions depending on whether deunionization occurs through declines in unionized firms or through worker reallocation away from unionized firms.

market via social insurance improves, the demand for ESHI decreases, reducing the incentive to unionize. The skill wage gap increases, but its effect is modest.[66] Although, *ceteris paribus*, deunionization would increase the skill wage gap, this effect is attenuated by shifts in labor demand arising from a weaker union threat. Because Medicare and Medicaid disproportionately expand insurance access for low-skill workers who are less likely to receive health insurance through employers either during employment or in retirement, their willingness to pay for unionization declines more sharply, reducing firms' union threat costs. This induces nonunionized firms to hire more low-skill workers, which raises their wages and partially offsets the direct effect of deunionization on the increase in the skill wage gap.

For skill-biased technological change, the difference is more pronounced. When unionization is fixed, welfare increases by 1.9%, whereas it rises by 4.9% when unionization adjusts endogenously. Skill-biased technological change raises relative demand for high-skill labor, pushing up wages for high-skill workers in nonunionized firms. As unionization declines, the skill wage gap widens more sharply because, under individual bargaining, the wider gap in marginal productivity passes through more directly to wages, widening the skill wage gap. This is again accompanied by a decrease in the ESHI rate. Despite the increase in wage inequality, overall welfare is higher under endogenous unionization due to gains in productive efficiency from less distorted labor demand.

RTW laws have only a small aggregate impact: the skill wage gap widens slightly (by about 0.1 log points) under endogenous unionization, while the effect on the ESHI rate is negligible. This limited impact reflects the relatively small shift in union density induced by RTW laws.

Taken together, these results suggest that the decline in unionization driven by social insurance expansions and technological change constitutes an overall welfare-enhancing response.

6.4.2 Subsidizing Labor Unions

Next, we examine the welfare impacts of subsidies for unionization covering one-third of the union's fixed cost to enhance unions. The welfare effect on workers is measured by the percent change in consumption in the baseline economy that makes a worker indifferent between the baseline and a counterfactual economy. The effect on social welfare is measured by changes in workers' welfare while uniformly redistributing the change in firms' profits and government revenue to workers. The union subsidy is not financed through distortionary taxes; instead, its budgetary cost is accounted for in this measure of social welfare through the change in government revenue, which falls by the cost of the subsidy and is subtracted from workers' consumption.

Table 10 reports the welfare effects of union subsidies. The policy increases union density by approximately 24.8 p.p. Notably, union subsidies generate welfare gains for workers across the

[66] The result is not driven by how taxes are collected. Table A.27 shows the impact of social insurance expansion without the tax adjustment and we obtain very similar impacts on all outcomes.

Table 10: Welfare Impact of Union Subsidies

	Changes from baseline
Union density (p.p.)	24.8
Worker welfare (%)	
All workers (ex-ante)	1.8
Low skill	3.4
High skill	0.6
Social welfare (%)	0.8

Note: This table reports the impact of the union subsidy on union density, worker welfare, and social welfare.

skill distribution. The welfare of low-skill workers rises by 3.4%, reflecting both wage compression and improved insurance provision through unions. High-skill workers also benefit, albeit more modestly (0.6%), partly due to enhanced insurance.

On the firm side, higher unionization can reduce productive efficiency, as not all firms optimally operate under unionization in this framework. In particular, unions may charge excessive wage premiums, reducing labor demand. At the same time, equilibrium unionization may be inefficiently low due to positive externalities associated with unions, particularly through insurance cost pooling. Taking these forces together, overall social welfare increases by 0.8%.

These findings suggest that the equilibrium level of unionization is below the social optimum. While unionization and its threat can entail production inefficiency, accounting for their effects on non-wage benefit provision implies that expanding unionization can be welfare-enhancing.

6.5 Implications Beyond the U.S.

Although our main finding is based on the U.S., unionization rates also declined in several other countries.[67] Our finding of the role of social insurance expansions as a cause of union decline in the U.S. complements the existing arguments. For example, the United Kingdom and Australia experienced large union declines despite adopting universal social insurance long before these declines began; however, their declines accelerated primarily after 1980, consistent with skill-biased technological change (Acemoglu and Autor, 2011). Canada provides another interesting case. Although the public-sector union in Canada has been highly stable, the private one has steadily declined since the mid-twentieth century.[68] In particular, as shown in Troy (1992), the private-sector union density declined relatively steeply from 34% in 1958, when universal health care was introduced, to 31% in 1965, and further down to 26% in 1975, which is consistent with our

[67]In contrast, many European countries maintain stable union density, especially in terms of collective bargaining coverage, even though they have experienced similar technological changes; this difference may reflect institutional features such as sectoral or national bargaining systems, which weaken the link between union membership, social insurance, and technological change.

[68]Morissette (2022) uses the Labor Force Survey and reports that, in 2022, union density is 61.6% in the public sector while it is 15.2% in the private sector.

argument about the deunionization effect of social insurance expansion. Additionally, in line with skill-biased technological change, the private-sector union density in Canada has also decreased in recent years, although at a slower pace than in the U.S. (Morissette, 2022).

Our framework also provides insight into economies with sectoral bargaining or more extensive public insurance. As discussed by Jäger et al. (2025), unions in many countries provide various insurance benefits, and our mechanism rationalizes this as cost-sharing through collective bargaining and across unionized firms. The model also clarifies the role of union threat effects. As shown in Section 6.1, without this channel, nonunionized firms provide less insurance. This force is likely weaker under sectoral bargaining, where individual firms cannot substantially affect unionization incentives. A direct implication is that sectoral systems may have lower insurance provision in nonunionized sectors, strengthening the role of public insurance.

7 Conclusion

In this paper, we study the equilibrium interactions between labor unions and social insurance. We provide evidence that expansions of social insurance programs reduce unionization in the United States. Then, we develop an equilibrium model in which union formation and employer-provided insurance are jointly determined. Unions, together with the threat of unionization, substantially increase employer-provided insurance in both unionized and nonunionized firms. Counterfactual analyses show that social insurance policies significantly affect unionization, ESHI provision, and wage inequality, with distributional effects depending on policy targeting. We find that the historical expansion of social insurance programs accounts for about 16% of the union decline in the United States. From a welfare perspective, the resulting decline in unionization is welfare-improving, although subsidizing unionization in the current economy can still raise welfare, suggesting that the current level of unionization may be lower than the social optimum.

We believe that our framework may be useful for studying other interactions among unions, non-wage benefits, and social insurance systems. In addition to employer-provided insurance, unions may affect other workplace amenities (e.g., female-friendly amenities, see Corradini et al., 2025) and access to public insurance programs, including unemployment insurance under the Ghent system (Kjellberg, 2011). Extending the framework to incorporate these channels, as well as firms' technology adoption and entry decisions (e.g., Holmes, 1998), may further improve our understanding of unions' long-run economic effects.

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Online Appendix (Not For Publication)

A Additional Empirical Evidence

A.1 Sectoral Shifts and Union Declines

This section calculates how much sectoral changes can account for the decline in unions. We calculate the counterfactual union density given by $\textit{union}_t^{CF} = \sum_{i\in\mathcal{I}} \omega_{i,1983} \times \textit{union}_{i,t}$ where $\omega_{i,1983}$ is the employment share of sector i (defined at the one-digit industry) in 1983 and $\textit{union}_{i,t}$ is the union membership density in sector i in year t. $\textit{union}_t^{CF}$ represents the counterfactual aggregate union density calculated as if the employment share of each sector had remained constant at its 1983 level. We use workers in the private sector from the CPS 1983-2019.

Figure A.3 displays the actual and counterfactual union density over time. The counterfactual union density is higher, suggesting that the sectoral employment shifts have indeed contributed to the decline in unions. However, the figure also shows that the contribution is quantitatively small. For example, in 2019, the difference is just 0.4 p.p.

A.2 Labor Unions and Job Security

We study how the union membership of a worker is related to subsequent job loss using a sample of workers from the SIPP. The sample is restricted to private-sector workers aged 22-65. We estimate the following regression equation:

$$\textit{Job loss}_{it} = \beta \cdot \textit{Union}_{it} + x_{it}'\gamma + \alpha_{s(i)} + \lambda_t + \epsilon_{it}, \tag{A1}$$

where the outcome variable $\textit{Job loss}_{it}$ is an indicator that takes the value 1 if worker i loses a job from month t to month $t+1$. We are interested in the coefficient β of $\textit{Union}_{it}$, which is an indicator for union membership in month t. Although we observe employment status in each month, union membership is recorded only once at the end of each four-month wave. A worker reports union status in a firm for which the worker worked the longest hours during a wave. We control for demographic variables such as age, sex, race, and education. We also control for state fixed effects $\alpha_{s(i)}$ and time fixed effects λ_t. ϵ_{it} is an error term.

Table A.6 reports the estimated coefficients. The first two columns suggest that the monthly job-loss probability is smaller for union workers by 0.2 p.p., which is sizable given the overall monthly job-loss probability of 0.7%. Columns (3) and (4) suggest that the coefficient is much larger (in absolute value) for low-skill workers.

A.3 Worker Demographics, Insurance Demand, and Union Membership

In this section, we ask whether workers with greater *ex-ante* insurance demand are disproportionately more likely to be union members. The idea is that, if access to employer-sponsored insurance is one reason workers value union representation, demographic characteristics associated with higher expected medical spending such as age, family structure, and health status should correlate with union membership in the cross-section.

We use data from the Current Population Survey, combining the Outgoing Rotation Groups (ORG) and the Annual Social and Economic Supplement (ASEC) files for 2001–2019, and restrict the sample to private-sector employed workers aged 22-65. We estimate

$$\mathit{Union}_{ist} = x_{ist}'\beta + z_{ist}'\gamma + \alpha_s + \lambda_t + \epsilon_{ist}, \tag{A2}$$

where Union_{ist} is an indicator equal to one if worker i in state s and year t is a union member; x_{ist} contains the variables of interest including log age and indicators for being married, having a child, and self-reported bad health, and z_{ist} represents other controls. The baseline specification controls only for demographics (sex, education, race). The richer specification additionally controls for firm size, three-digit occupation, three-digit industry, and log earnings. All specifications include year and state fixed effects, and robust standard errors are reported in parentheses.

Table A.4 reports the estimated coefficients. In column (1), which conditions only on basic demographics, union membership is positively and significantly associated with each of the characteristics most directly tied to expected medical spending: log age, being married, and having a child. These patterns are consistent with the idea that workers with higher insurance demand sort disproportionately into union jobs, which are more likely to provide health insurance. Column (2) adds the covariates (firm size, three-digit occupation, three-digit industry, and log earnings). The coefficients on log age and on having a child remain positive and significant at the 1 percent level, although their magnitudes are attenuated; the marital-status coefficient becomes economically and statistically negligible.

The one apparent exception is self-reported health: the coefficient on bad health is negative and statistically significant in column (1) and remains negative, though statistically insignificant, in column (2), suggesting that workers in poor health are not disproportionately sorting into unionized jobs. We view this variable as more difficult to interpret as a proxy for ex-ante insurance demand because health status is jointly determined with employment and insurance access. In particular, prior work (e.g., Aizawa and Fang, 2020; Aizawa, 2019) documents a positive correlation between health and ESHI, whereby healthier workers are more likely to hold jobs offering ESHI. Moreover, better insurance access itself may improve subsequent health outcomes and labor-force attachment. Both of these forces push the bad-health coefficient toward negative values, while the insurance-demand mechanism emphasized in our paper pushes it toward positive. The attenuation toward zero in column (2) therefore suggests a more nuanced interpretation in which the coefficient reflects the net effect of these opposing forces; the near-zero estimate does not necessarily indicate the absence of an insurance-demand channel, but may reflect partial cancellation of these forces once firm-level sorting is partially controlled for. For these reasons, we place greater weight on demographic measures such as family structure and age, which more directly capture variation in insurance demand ex ante.

A.4 Health Insurance After Retirement

We use the HRS to examine the relationship between union membership and retirement coverage of an ESHI plan at the individual level in recent years. Using the sample of employed workers aged 65 or younger, we estimate the same linear probability model as in Section 3.1. The sample period is now restricted to 2000 onward because the post-retirement coverage variable is available only from that year.

Table A.7 reports the estimated coefficients. In the sample, 15.2% of workers have an ESHI plan that provides retirement coverage after age 65. Columns (1) and (2) show that union workers are 12.5 p.p. more likely to be covered by such a plan without covariates and 10 p.p. more likely with covariates. Column (3) shows that once we control for individual fixed effects, the coefficient is statistically insignificant. This might partly be due to the limited mobility of workers between union and nonunion firms for relatively old workers in the HRS. Furthermore, since we rely only on variation in the union status of workers moving between union and nonunion firms, once we control for individual fixed effects, there might be some selection issues. For example, union workers might be willing to move to nonunion firms only if access to insurance is guaranteed. To explore this further, we make a distinction between the move from union firms to nonunion firms and from nonunion firms to union firms. Column (4) indicates that the move from union to nonunion firms is not necessarily associated with a loss of coverage, whereas the move from nonunion to union firms is associated with a statistically significant 7.9 p.p. increase in coverage. This result suggests that the decline of unions could reduce a worker's opportunities to gain access to retirement coverage through switching to unionized employers, which is consistent with the decline in retirement coverage in recent years.

A.5 Union Effect on ESHI Provision: RTW Laws as an Instrument

In this section, we use RTW laws as an instrument for unionization, and examine the effect of unions on ESHI provision. In doing so, we follow Fortin et al. (2022) and exploit differential exposure to RTW laws by industry. Fortin et al. (2022) estimate the impact of unionization on wages using RTW laws as an instrument for it. We follow their specification but use ESHI coverage as an outcome. We take this approach of exploiting cross-industry variation instead of only using variation in RTW laws across states over time because (i) there are only a small number of states that implemented RTW laws in recent years, and (ii) many of them implemented RTW laws around the same time as the ACA Medicaid expansions, which would be a major confounder for our analysis. Our estimation sample comes from the CPS 2001-2019; we restrict it to private-sector workers aged 22-65 who reported their union status, and exclude individuals who were out of the labor force at the time of the survey.

In the first stage, we estimate

$$\mathit{Union}_{ist} = \sum_{j=1}^{J} \mathbb{1}[\mathit{Industry}_{ist} = j] \cdot (\theta_j^u + \phi_j \cdot \mathit{RTW}_{st}) + x'_{ist}\gamma_1 + \alpha_{1,s} + \lambda_{1,t} + \epsilon_{1,ist}, \tag{A3}$$

where i is individual, s is state, and t is time. θ_j^u captures baseline union density in industry j common across states, and ϕ_j captures the impact of RTW laws on union membership in industry j. We use ten broad industry groups ($J = 10$) and absorb differences in industry composition within the broad industry groups with 3-digit industry fixed effects. The differential exposure to RTW laws across industries is captured by ϕ_j. As discussed in Fortin et al. (2022), at the heart of the identification is the variation in the within-sector difference in the *levels* of unionization rates, which we highlight in Figure A.2.[1]

In the second stage, we estimate

$$\mathit{ESHI}_{ist} = \beta \cdot \mathit{Union}_{ist} + \sum_{j=1}^{J} \mathbb{1}[\mathit{Industry}_{ist} = j] \cdot \theta_j + x'_{ist}\gamma_2 + \alpha_{2,s} + \lambda_{2,t} + \epsilon_{2,ist}. \tag{A4}$$

β is our coefficient of interest. The implementation uses the interactions between the RTW dummy and the broad industry dummies as instruments for union membership.

Table A.5 reports the estimates based on the sample of employed workers (column 1) and the sample of employed workers either covered by ESHI or not covered at all (column 2), and we find

[1]Roughly speaking, near-zero-unionization sectors serve as a control group against sectors where RTW laws matter (Fortin et al., 2022).

a significant and large effect of unions on ESHI. Note that the magnitude is larger than the one we find in the HRS: this could reflect the possibility that our IV estimate is local[2]; another possibility is a difference in sample selection because the HRS primarily covers older individuals.

A.6 Results from Election Data

We provide additional evidence on the impact of Medicare using data on NLRB elections obtained from Sojourner and Yang (2022).[3] Union elections primarily capture new union formation at the establishment level and therefore reflect flows into unionization rather than the stock of unionized firms. In this respect, the election data allow us to examine whether the decline in union density may partly operate through reduced new union formation.[4]

We first use the same exposure-based DID specification as the regression equation (2) in the main text but use the log of the number of elections as the outcome variable.[5] The election data is available starting from 1962. Table A.8 reports the estimated coefficients. The estimated coefficients on $Exposure_s \times Post_t$ are negative and statistically significant in both specifications, mirroring the pattern documented for union density: states more exposed to the Medicare introduction experienced a larger post-Medicare decline in union organizing activity. Moving from the minimum Blue Cross coverage in 1963 (12%) to the maximum (51%), an exposure increase of 0.39, implies an additional decline of roughly 24% in the number of union elections; the analogous range along the Any-coverage distribution (43% to 68%, an exposure increase of 0.25) implies a decline of about 33%.

To complement the pooled estimates and assess the timing of the response, we again estimate the dynamic event-study counterpart, given by equation (3), with log elections as the outcome variable. Figure A.8 displays the estimated coefficients β_τ. Panel (a) reports the case in which exposure is defined as Blue Cross coverage in 1963, and panel (b) reports the case in which exposure is defined as any insurance coverage in 1963. Consistent with the pooled estimates in Table A.8 and with the main-text findings for union density, the post-1965 coefficients are negative in both panels, indicating that states with higher pre-Medicare insurance coverage experienced sharper

[2]Indeed, our model suggests that large firms are likely to be unionized anyway, and therefore RTW would affect medium-sized firms that are at the margin of unionization. Given the insurance fixed cost, such medium-sized firms are also likely to be close to the margin of whether to offer ESHI or not, and our IV estimate might be picking up the response of such firms.

[3]Sojourner and Yang (2022) provide the data at https://www.dropbox.com/s/m92aw4vb1u3qy0n/nlrb_elections_19622009.dta?dl=0 (last accessed June 7, 2026).

[4]The number of unionized firms may also decline because existing unionized establishments contract or exit, which is not fully captured by election data.

[5]In the regression using union density, we control for the share of workers with college education from the CPS. Since education information is not available for the year 1963, we do not control for it here. Other covariates are the same as in the regression for union density.

post-reform declines in union elections. We note, however, that the estimated coefficients in 1962 and 1964 are also statistically distinguishable from zero, suggesting the presence of potential confounding factors that challenge the strong parallel-trends assumption underlying this specification.

To further address these potential confounding factors, we estimate the same event study separately for the low- and high-exposure groups, using Blue Cross coverage to define exposure. Restricting each subsample to states with more similar levels of observed characteristics (pre-reform insurance coverage) mitigates the risk that differential trends across exposure groups confound the estimates, though at the cost of reduced within-group variation in exposure. The low-exposure group consists of states with Blue Cross coverage in 1963 below the median, while the remaining states are the high-exposure group. Figure A.9 shows that in each group, there are no significant pre-trends, and we also detect significant policy impacts in each group.

A.7 Heterogeneity by Age Composition in the Medicare Effect

This appendix provides details behind the age-composition result discussed in Section 3.2.1. If union-provided insurance was particularly valuable to workers approaching retirement age or to those who placed greater value on retiree health insurance, then the introduction of Medicare should have reduced union density more in states with relatively older workforces. To examine this possibility, we exploit geographic variation in workforce age composition and estimate

$$\log\left(\textit{union density}_{st}\right) = \beta_1 \textit{Treatment}_{st} + \beta_2\, \textit{Treatment}_{st} \times \textit{Old}_{s,1964} + x'_{st}\gamma + \alpha_s + \lambda_t + \epsilon_{st}, \quad \text{(A5)}$$

where $\textit{Old}_{s,1964}$ is the share of workers aged 40 or older in 1964, a pre-treatment period, and $\textit{Treatment}_{st}$ is the exposure to Medicare interacted with the post-treatment indicator $\textit{Exposure}_s \times \mathbb{1}[t \geq 1966]$. The covariates x_{st}, state effects α_s, and time effects λ_t are the same as in the baseline specification. The coefficient of interest is β_2, which captures how the treatment effect varies with the share of older workers.

Table A.3 reports that $\hat{\beta}_2$ is negative in both columns and statistically significant in column (2): states more exposed to Medicare experienced a larger reduction in union density when they also had relatively older workforces. This pattern is consistent with our hypothesis that union-provided insurance was more valuable for older workers, who were closer to Medicare eligibility and more likely to value retirement-related insurance benefits. The results therefore suggest that public insurance provision through Medicare reduced the demand for unionization more strongly in states where such workers represented a larger share of the workforce.

A.8 Introduction of Medicaid

This section estimates the impact of Medicaid implementation on unions. To begin with, we note that there are both labor supply and demand mechanisms through which Medicaid may reduce union density. First, without Medicaid, individuals may strongly prefer to work (Garthwaite et al., 2014) and to join a union to gain access to ESHI. Second, to the extent that Medicaid increases the value of unemployment for low-skill workers, the introduction of Medicaid makes it more costly for firms to hire low-skill workers. By lowering labor demand for low-skill workers, Medicaid may shift the worker composition away from low-skill workers to high-skill workers, who tend to be less favorable toward unions.

One complication of staggered treatment timing is that it makes the standard difference-in-differences estimates hard to interpret. Furthermore, most states quickly implemented the program within a few years, and there is only a small group of states belonging to "not-yet-treated" group if we aim to estimate dynamic effects over a long period. As a compromise, we take a short time window.[6]

We begin with an interaction-weighted event-study specification following Sun and Abraham (2021). Specifically, we estimate cohort-by-event-time effects:

$$\log(\textit{union}_{st}) = \sum_{e\in\mathcal{E}} \sum_{\tau\neq -1} \beta_{e,\tau} \mathbb{1}\{E_s = e\}\, \mathbb{1}\{t-e=\tau\} + \; x'_{st}\gamma + \alpha_s + \lambda_t + \epsilon_{st}, \tag{A6}$$

where E_s is the year at which state s implemented Medicaid and $\mathcal{E}$ denotes the set of adoption years observed in the sample.[7] We omit $\tau = -1$ as the reference period. We report event-time effects as weighted averages of the cohort-specific estimates. We use the sample until $t = 1968$, control for the same set of time-varying covariates as in equation (3), and cluster standard errors at the state level.

One important concern is that the timing of Medicaid implementation may be correlated with other state-level labor market shocks that also contributed to declines in unionization. To address this concern, as in our Medicare analysis, we augment our specification by controlling for additional confounders, including state-year manufacturing employment shares interacted with year fixed effects. These controls are intended to capture differential exposure across states to trade shocks, automation, and other structural changes affecting manufacturing-intensive states.

Figure A.10 displays the estimated coefficients of equation (A6). The estimate suggests that union density is reduced by about 5% two years after implementation. We do not detect significant

[6] Specifically, we stop at the year 1968, which ensures that more than 10 states belong to the control group.
[7] See Gruber (2003) for the timing of implementation by each state.

pre-trends.

Given the impact of the introduction of Medicare we found in the previous subsection, one concern is that some of the effects of Medicaid could be confounded by the introduction of Medicare. We address this issue by controlling for Medicare exposure. Specifically, we include $\mathbb{1}\{t > 1965\} \times \textit{High Exposure}_s$ where $\textit{High Exposure}_s$ is an indicator that takes the value 1 if the Blue Cross coverage rate among retirees in state s is greater than the national median prior to 1965. Figure A.11 confirms that this result is robust to controlling for Medicare exposure.

Placebo test. A potential concern with our staggered-adoption design is that early-adopting states may have been on differentially declining unionization trajectories before actually implementing Medicaid, which would confound the estimated treatment effect with pre-existing trends. To address this concern, we estimate a standard two-way fixed-effects event-study specification with the actual adoption dates and with placebo adoption dates shifted backward by $k = 1, 2, 3$ years:

$$\log(\textit{union}_{st}) = \sum_{\tau \in \mathcal{T}, \tau \neq -1} \beta_\tau \, \mathbb{1}\{t - (E_s - k) = \tau\} + x_{st}'\gamma + \alpha_s + \lambda_t + \epsilon_{st}, \tag{A7}$$

where E_s is the year when state s implements Medicaid, $k \in \{0, 1, 2, 3\}$ is the placebo shift (with $k = 0$ corresponding to actual adoption dates), and $\mathcal{T}$ is the set of event-time indices, with $\tau = -4$ and $\tau = 3$ binning all earlier and later event times, respectively. The reference period $\tau = -1$ is omitted. We use the same sample window, covariates, fixed effects, and weighting as in equation (A6), and cluster standard errors at the state level. We estimate equation (A7) separately for each value of k.[8]

Under the concerning scenario, the placebo "treatments" would generate negative estimated post-treatment effects, since the placebo "post-treatment" years coincide with actual pre-treatment years in which the steeper decline would have been occurring. Figure A.12 reports the results. Panel (a) shows the actual estimates, which confirm a negative post-treatment effect. Panels (b)–(d) show that the placebo estimates show no evidence of differential pre-trends; if anything, early adopters exhibit a slight upward trend in union density relative to late adopters, which works against the concern that they were experiencing a steeper decline even before implementation.

A.9 ACA Medicaid Expansion

We use the CPS sample and the variation in the ACA Medicaid expansion across states to estimate the impact of the expansion on union membership. We focus on states that expanded

[8]The actual-date estimates ($k = 0$) also serve as a robustness check on the Sun–Abraham specification: they confirm the negative post-treatment effect under a standard TWFE estimator.

Medicaid in January 2014 or never expanded during the sample period. Our empirical specification is

$$\textit{Union}_{ist} = \beta \cdot \textit{(ACA Medicaid)}_{st} + x'_{ist}\gamma + \alpha_s + \lambda_t + \epsilon_{ist}, \tag{A8}$$

where i is the individual, s is the state, t is the year, $\textit{Union}_{ist}$ is an indicator that takes the value 1 if individual i in state s is a union member at t, $\textit{(ACA Medicaid)}_{st}$ is an indicator that takes the value 1 if state s has expanded Medicaid coverage in t. x_{ist} is a vector of time-varying covariates, including age, education, gender, race, year-specific dummies for industries and occupations, the same set of political variables used in the analysis of Medicare and Medicaid introductions, and indicators for the periods before and after the passage of RTW laws. α_s and λ_t are the state and time fixed effects. To focus on those who are likely to be affected by the expansion, we split the sample into individuals with low education (high school or less), who are more likely to be eligible due to low income, and those with high education, who are less likely to be eligible.

Table A.9 reports the estimation results. In Column (1), we report the result using all individuals in the sample. About 12% of individuals are union members, and the ACA Medicaid expansion decreased union membership by 0.3 percentage points, although the coefficient is not statistically significant. Column (2) shows that the expansion had a statistically significant impact on low-education individuals, decreasing union membership among them by 0.5 p.p., which is about a 5% decrease in union membership given that 10% of individuals in this sample were union members. In contrast, the expansion had almost no impact on high-education individuals, as indicated by the last column. Figure A.13 shows an event study plot consistent with these results, which also shows there is no pre-trend.

A.10 Unemployment Insurance

We use variation in UI generosity across states and over time to estimate the its impact on union membership. We use the CPS 2000-2019 to estimate the following specification:

$$\textit{Union}_{ist} = \beta \cdot (\textit{Replacement rate})_{ist} + x'_{ist}\gamma + \alpha_s + \mu_t + \epsilon_{ist}, \tag{A9}$$

where i is the individual, s is the state, t is the year, $\textit{Union}_{ist}$ is an indicator that takes the value 1 if individual i in state s is a union member at t, $\textit{Replacement rate}_{ist}$ is the UI replacement rate, calculated as the weekly benefit amount divided by the weekly wage, for worker i in state s at time t, x_{ist} is a vector of time-varying covariates, and α_s and μ_t are state and year fixed effects, respectively. ϵ_{ist} is an error term.

Table A.10 reports the estimation results of equation (A9). We find a statistically significant impact of the UI replacement rate on union membership. Specifically, a 10 p.p. increase in the UI replacement rate makes an individual 2.1 p.p. less likely to be a union member. Columns (2)-(4) indicate that this pattern remains even after we control for UI maximum duration, RTW laws, and the same political variables used in our Medicare analysis.

A.11 Testing the Fixed-Cost Mechanism: Heterogeneity by Firm Size

Union firms have an advantage in ESHI offering partly through the insurance fixed-cost channel discussed in Section 4.7. The mechanism therefore predicts that, when social insurance expansions reduce the value of employer-provided insurance, the unionization response should be concentrated among smaller establishments, while large establishments, where the fixed cost is already a small share of total insurance costs, should respond less. This appendix tests this prediction by exploiting the election data described above.

Following the approach in Appendix A.6, we estimate the dynamic event-study specification separately for high- and low-exposure states to address potential confounding factors. However, this time, within each subsample, we further split establishments into "small" and "large" groups, defined as below or above the median number of voters in the election data. Since this finer split yields cells with zero elections, we use the inverse hyperbolic sine (IHS) transformation of the number of elections, rather than logs, as the outcome variable.

Figure A.14 reports the results. Across panels, large establishments show little response to the introduction of Medicare, whereas small establishments exhibit a negative response. The figure therefore provides direct evidence that the union response to Medicare was concentrated among smaller establishments, consistent with the fixed-cost mechanism formalized in Section 4.7 of the main text.

A.12 Impact of Right-to-Work Laws on Unionization

This section first presents the effect of RTW laws on union membership by following Fortin et al. (2022) and then presents its impact on union elections.

A.12.1 Individual Union Membership

To examine the impact of RTW laws on union membership, we estimate the following event-study specification using the CPS data:

$$Union_{ist} = \sum_{\tau=-5,\neq-1}^{4} \beta_\tau \mathbb{1}_{\{t-E_s=\tau\}} + \beta_{-6}\mathbb{1}_{\{t-E_s\leq-6\}} + \beta_{+5}\mathbb{1}_{\{t-E_s\geq5\}} + x'_{ist}\gamma + \alpha_s + \epsilon_{ist} \tag{A10}$$

where $Union_{ist}$ is an indicator for union membership for individual i in state s at time t. E_s represents the timing of the event (i.e., the passage of RTW laws) in state s. x_{ist} is a vector of covariates, including age, education, sex, race, year-by-industry dummies, year-by-occupation dummies, and month fixed effects. We control for an indicator for ACA Medicaid expansion in state s and for the same political variables as in the Medicare analysis. α_s are state fixed effects. We cluster standard errors at the state level.

Panel (a) of Figure A.15 displays the estimated coefficients and the 95% confidence intervals. Overall, the RTW laws reduced union membership by 2 p.p., consistent with Fortin et al. (2022) although our specification is not exactly the same since we control for the ACA Medicaid expansion and variables capturing state political environment. We detect a slight indication of pre-trend in four years before the event.

A.12.2 Union Elections

Next, we investigate the impact of RTW laws on union elections using the NLRB election data. Specifically, we estimate the following event-study specification:

$$y_{st} = \sum_{\tau=-5,\neq-1}^{4} \beta_\tau \mathbb{1}_{\{t-E_s=\tau\}} + \beta_{-6}\mathbb{1}_{\{t-E_s\leq-6\}} + \beta_{+5}\mathbb{1}_{\{t-E_s\geq5\}} + x'_{st}\gamma + \alpha_s + \lambda_t + \epsilon_{st} \tag{A11}$$

where y_{st} is the outcome in state s at time t. E_s represents the timing of the event. x_{st} is a vector of state-level covariates, including an indicator for ACA Medicaid expansion and the political variables. α_s and λ_t are state fixed effects and year fixed effects. We cluster standard errors at the state level. We use the inverse hyperbolic sine (IHS) transformation of the number of elections in state s at time t as the outcome y_{st}.[9]

Panel (b) of Figure A.15 displays the estimated coefficients and their 95% confidence intervals. The estimated coefficient at $\tau = 0$ suggests that RTW laws reduce the number of union elections by 25% upon passage, but the estimated coefficients are not significant after that. This result thus suggests that the negative effect on union density may operate through the deunionization of unionized firms, instead of a reduction in union formation.

[9] In the analysis of the introduction of Medicare and Medicaid, we used the log of the number of elections since there were no zeros in the data during that time period. Since the data for recent years have zeros for some states, we use the IHS transformation.

B Model Appendix

B.1 Expected Value of a Match for Workers

In the main text, we mentioned that the expected value of a match depends on the equilibrium distribution of vacancies posted by different firms. Formally, the value of meeting a firm is given by

$$
\begin{aligned}
\mathbb{E}\left[\max\{V^E_{x,y,k}(w_{x,y,k},a),V^U_x\}\right] = \sum_{y\in\mathcal{Y}}\Omega_{x,y}\sum_{a\in\mathcal{A}}\Big[&\mathcal{Q}_y P_{y,u}(a)\max\{V^E_{x,y,u}(w_{x,y,u},a),V^U_x\} \\
&+(1-\mathcal{Q}_y)P_{y,n}(a)\max\{V^E_{x,y,n}(w_{x,y,n},a),V^U_x\}\Big],
\end{aligned}
\tag{A12}
$$

where $\Omega_{x,y}=\nu_{x,y}M_y/\sum_{y'\in\mathcal{Y}}\nu_{x,y'}M_{y'}$ is the fraction of vacancies in sub-market x posted by type-y firms. A worker of type x meets a vacancy posted by a firm of type y with probability $\Omega_{x,y}$; among them, the fraction $\mathcal{Q}_y$ given by equation (16) is unionized while the remainder $1-\mathcal{Q}_y$ is not unionized; $P_{y,k}(a)$ is the fraction of type-y firms providing amenity a with union status k, given by equation (15).

B.2 Aggregate Willingness to Pay for (Non)Unionization

This section provides formal definitions of the aggregate willingness to pay for (non)unionization that enters the cost functions (6) and (7) in the main text. We first define the willingness to pay for unionization or nonunionization for each worker type and then aggregate them to the firm level.

First, let $w_{x,y,u}(\boldsymbol{g},a)$ and $w_{x,y,n}(\boldsymbol{g},a)$ be the union and nonunion wage schedules for a type-x worker in a type-y firm with amenity a, determined in the bargaining problems (19) and (18), respectively.

We now define each worker's *willingness to pay for unionization*. Let $\mathcal{W}_{x,y,n}(\boldsymbol{g},a)$ denote the willingness of a type-x worker in a type-y nonunionized firm with amenity a to pay for unionization. It is implicitly determined by

$$
V^E_{x,y,u}(w_{x,y,u}(\boldsymbol{g},a),a)=V^E_{x,y,n}(w_{x,y,n}(\boldsymbol{g},a)+\mathcal{W}_{x,y,n}(\boldsymbol{g},a),a) \tag{A13}
$$

where $V^E_{x,y,k}(\cdot,\cdot)$ for $k=u,n$ is defined in (4). $\mathcal{W}_{x,y,n}(\boldsymbol{g},a)$ can be either positive or negative, and it gives the amount of consumption a type-x worker needs to be compensated for staying nonunionized in a type-y firm with amenity a. The firm-level aggregate willingness to pay for unionization in a nonunionized firm is then given by

$$\mathcal{W}_{y,n}(\boldsymbol{g}, a) = \sum_x \mathcal{W}_{x,y,n}(\boldsymbol{g}, a) \times g_x. \tag{A14}$$

Similarly, let $\mathcal{W}_{x,y,u}(\boldsymbol{g}, a)$ denote the amount of consumption a type-x worker in a unionized firm needs to be compensated if the union were to be disbanded, which we refer to as the *willingness to accept deunionization*. It is defined implicitly by

$$V^E_{x,y,u}(w_{x,y,u}(\boldsymbol{g}, a) + \mathcal{W}_{x,y,u}(\boldsymbol{g}, a), a) = V^E_{x,y,n}(w_{x,y,n}(\boldsymbol{g}, a), a). \tag{A15}$$

Again, it can be either positive or negative. Then, we define the firm-level aggregate willingness to accept deunionization for all the workers in a unionized firm as

$$\mathcal{W}_{y,u}(\boldsymbol{g}, a) = \sum_x \mathcal{W}_{x,y,u}(\boldsymbol{g}, a) \times g_x. \tag{A16}$$

B.3 Alternative Formulation of the Union Cost Function

Given that a voting outcome, in reality, depends on whether a simple majority is in favor of a union, a natural alternative approach would be to let the cost function depend on the excess number of voters in favor of the union. We argue below that this approach yields the same cost function as in the baseline case under some additional assumptions.

Suppose workers draw i.i.d. taste shocks υ for unionization from a CDF H. Let $\Delta_{x,y,n}(\boldsymbol{g}, a) = V^E_{x,y,u}(w_{x,y,u}(\boldsymbol{g}, a), a) - V^E_{x,y,n}(w_{x,y,n}(\boldsymbol{g}, a), a)$ be the value gain from unionization. Note that if the utility function is quasilinear in w, then $\Delta_{x,y,n}(\boldsymbol{g}, a)$ equals the willingness to pay for unionization $\mathcal{W}_{x,y,n}(\boldsymbol{g}, a)$. To see this, letting $u_x(w, a) = w + f_x(a)$ and using expression (4), we can rewrite equation (A13) as

$$\begin{aligned} V^E_{x,y,u}(w_{x,y,u}(\boldsymbol{g}, a), a) = w_{x,y,n}(\boldsymbol{g}, a) + \mathcal{W}_{x,y,n}(\boldsymbol{g}, a) + f_x(a) \\ + \gamma[\delta_{x,n} V^U_x + (1 - \delta_{x,n}) V^E_{x,y,n}(\tilde{w}_{x,y,n}, \tilde{a}_{x,y,n})]. \end{aligned} \tag{A17}$$

where the right-hand side is the value of nonunion employment and $(\tilde{w}_{x,y,n}, \tilde{a}_{x,y,n})$ are the equilibrium future wages and amenities. This implies $\mathcal{W}_{x,y,n}(\boldsymbol{g}, a) = \Delta_{x,y,n}(\boldsymbol{g}, a)$.

A type-x worker votes for unionization if $\Delta_{x,y,n}(\boldsymbol{g}, a) - \upsilon > 0$. Then, the excess number of workers in favor of unionization is given by

$$N_{y,n}(\boldsymbol{g}, a) = \sum_x g_x [H(\Delta_{x,y,n}(\boldsymbol{g}, a)) - 0.5], \tag{A18}$$

where $H(\Delta_{x,y,n}(\boldsymbol{g}, a))$ is the fraction of type-x workers in favor of unionization. Note that $N_{y,n}(\boldsymbol{g}, a) \geq 0$ means a majority of workers are in favor of unionization. Given this, the alternative formulation of the union threat cost is $\tilde{C}_{y,n}(\boldsymbol{g}, a) = \tilde{c}_0 \max\{0, N_{y,n}(\boldsymbol{g}, a)\}$.

Note that the difference between this version and the baseline specification is the difference between $H(\Delta_{x,y,n}(\boldsymbol{g}, a)) - 0.5$ and $\mathcal{W}_{x,y,n}(\boldsymbol{g}, a)$. The two cases are equivalent if the utility function is quasilinear in w and if we impose the following linear approximation to the CDF H: $H(x) = hx + 0.5$. With these assumptions, we have $H(\Delta_{x,y,n}(\boldsymbol{g}, a)) = h\mathcal{W}_{x,y,n}(\boldsymbol{g}, a) + 0.5$, and then $\tilde{C}_{y,n}(\boldsymbol{g}, a) = \tilde{c}_0 h \max\{0, \sum_x g_x \mathcal{W}_{x,y,n}(\boldsymbol{g}, a)\}$. Letting $c_0 = \tilde{c}_0 h$, this is equivalent to the baseline formulation of the union threat cost. In other words, the baseline specification can be interpreted as a special case of the voting model above with a linear-approximation assumption on the distribution of workers' taste shocks.

B.4 Labor-Hoarding Costs of Lower Job Destruction Rates

In our baseline model, the exogenous job-destruction rate δ differs across union and nonunion firms, with union firms facing a lower rate. If having a lower destruction rate is not costly, then it is hard to justify nonunion firms having higher δ because firms would like to retain workers, and workers value job security. In this appendix, we argue that a lower δ is not necessarily desirable from the firm's perspective because stronger job security could generate labor-hoarding costs when the firm is hit by an adverse productivity shock and would like to scale down employment. In the full model, we abstract from explicit productivity shocks and capture these costs in reduced form through the fixed cost of unionization FC^{union}. The simple one-period example below illustrates the underlying mechanism.

Consider a firm endowed with an initial workforce g_0 of homogeneous workers. Output is produced according to a concave production function

$$y = zg^{\alpha}, \qquad \alpha \in (0, 1), \tag{A19}$$

where g is the level of employment and z is a stochastic productivity draw. The firm can expand employment by posting vacancies $\nu \geq 0$ at unit cost c, but it cannot freely contract employment: separations occur only at the exogenous, predetermined rate δ. Letting q denote the vacancy-filling probability and assuming a constant wage w, the firm's profit-maximization problem is

$$\pi(g_0, z) \;=\; \max_{\nu \geq 0} \; zg^{\alpha} - wg - c\nu, \tag{A20}$$

subject to the law of motion for employment

$$g \;=\; q\nu + (1-\delta)g_0. \tag{A21}$$

The non-negativity of vacancies, $\nu \geq 0$, translates via (A21) into a lower bound on employment: the firm cannot operate below the carry-over workforce $(1-\delta)g_0$. Substituting (A21) into (A20) to eliminate ν yields

$$\pi(g_0, z) \;=\; \max_{g \geq (1-\delta)g_0} \; zg^\alpha - wg - \frac{c}{q}(g - (1-\delta)g_0)\,. \tag{A22}$$

The first-order condition for the unconstrained problem delivers the firm's frictionless desired employment

$$g^*(z) \;=\; \left(\frac{\alpha z}{w + c/q}\right)^{1/(1-\alpha)}, \tag{A23}$$

so that optimal employment is

$$g(g_0, z) \;=\; \max\{(1-\delta)g_0,\; g^*(z)\}\,. \tag{A24}$$

The firm engages in inefficient labor hoarding whenever the realized productivity draw is low enough that $g^*(z) < (1-\delta)g_0$. A lower destruction rate δ raises the cutoff $(1-\delta)g_0$ at which hoarding kicks in, both increasing the probability that the firm is stuck above its frictionless optimum and enlarging the difference $(1-\delta)g_0 - g^*(z)$ in those states. Expected labor-hoarding losses are therefore decreasing in δ: greater job security imposes greater adjustment costs on the firm.

In a model with time-varying productivity, lower δ is not always desirable for the firm: it raises the expected cost of being forced to retain inefficiently many workers when productivity is low. In the full model, we capture these labor-hoarding costs in reduced form through the fixed cost of unionization FC^{union}, which allows us to keep the wage-setting and bargaining environment tractable while preserving the economic logic that stronger job security can be costly to firms.

B.5 Equilibrium Condition for Labor Market Tightness

This section describes how labor market tightness θ is determined in a steady-state equilibrium. First, note that, given tightness θ, firms decide on the optimal hiring, which leads to the following total mass of workers hired by firms:

$$\bar{g}_x(\theta) = \sum_{y=1}^{Y} M_y \sum_{a \in \boldsymbol{\mathcal{A}}} [\mathcal{Q}_y P_{y,u}(a) g_{x,y,u}(a;\theta) + (1-\mathcal{Q}_y) P_{y,n}(a) g_{x,y,n}(a;\theta)],$$

where we now let $g_{x,y,k}$ explicitly depend on θ. The optimal hiring decisions of firms give us a relationship between a mass of unemployed workers and market tightness: $\mathcal{U}_x^{JC}(\theta) = N_x - \bar{g}_x(\theta)$ for each x. We use the superscript JC (shorthand for "job creation") to emphasize that $\mathcal{U}_x^{JC}(\theta)$ reflects the optimal job creation decisions.

To express the flow conditions consistently with the law of motion in equation (11), it is useful to define the vacancy-weighted average acceptance probability faced by a type-x job seeker. Let $\mu_{x,y,k}(a;\theta)$ denote the share of type-x vacancies posted by firms of type (y,k,a).The vacancy-weighted average acceptance probability is then $\bar{e}_x(\theta) = \sum_{y,k,a} \mu_{x,y,k}(a;\theta)\, e_{x,y,k,a}$.[10]

On the labor supply side, let $s_x(\theta)$ be the steady-state mass of type-x job seekers at the beginning of a period. For each x, we have $\sum_{k\in\{u,n\}} \delta_{x,k}\bar{g}_{x,k}(\theta) = s_x(\theta)p(\theta_x)\bar{e}_x(\theta)$ where the left-hand side is the flow into unemployment and the right-hand side is the flow-out of unemployment.[11] $\bar{g}_{x,k}$ is the mass of workers hired by firms with union status k. They are given by $\bar{g}_{x,u}(\theta) = \sum_{y=1}^{Y} M_y \mathcal{Q}_y \sum_{a\in\mathcal{A}} P_{y,u}(a) g_{x,y,u}(a;\theta)$ and $\bar{g}_{x,n}(\theta) = \sum_{y=1}^{Y} M_y(1-\mathcal{Q}_y) \sum_{a\in\mathcal{A}} P_{y,n}(a) g_{x,y,n}(a;\theta)$. Given $s_x(\theta)$, we obtain the mass of unemployed workers (after firms make their hiring): $\mathcal{U}_x^{BC}(\theta) = (1-p(\theta_x)\bar{e}_x(\theta))s_x(\theta) = \frac{1-p(\theta_x)\bar{e}_x(\theta)}{p(\theta_x)\bar{e}_x(\theta)} \sum_{k\in\{u,n\}} \delta_{x,k}\bar{g}_{x,k}(\theta)$. The function $\mathcal{U}_x^{BC}(\theta)$ represents the mass of unemployed workers of skill x that equalizes flows into and out of unemployment given tightness θ, and BC is shorthand for "Beverage curve." Note that both $\mathcal{U}_x^{JC}(\theta)$ and $\mathcal{U}_x^{BC}(\theta)$ are the mass of unemployed workers after matches are formed in the frictional labor markets and before jobs are destructed at the end of a period. Equilibrium market tightness is pinned down by $\mathcal{U}_x^{BC}(\theta) = \mathcal{U}_x^{JC}(\theta) \quad$ for all $x \in \mathcal{X}$.

B.6 Illustration of Wage Functions in a Simple Setting

In this section, we describe some properties of the wage functions in a simple setting that allows us to derive analytical expressions for the wage functions. To obtain analytical tractability, we assume that the utility is linear in wages and given by $w + v_x(a)$ where $v_x(a)$ is the indirect utility from amenity. To simplify the exposition, we consider a static case by setting $\gamma = 0$ (or $\delta = 1$), and set the break-up cost ϕ to 0. To further simplify notation, let $\bar{C}(\boldsymbol{g}, a) = \sum_x c_x(a) g_x - FC_u^a(a)$ be the total amenity cost.

[10] In a steady-state equilibrium, firms only post vacancies whose expected match surplus is positive, and Nash bargaining then implies a positive worker surplus at any posted vacancy. Therefore $e_{x,y,k,a} = 1$ for every (x,y,k,a) with $\nu_{x,y,k}(a;\theta) > 0$, so $\bar{e}_x(\theta) = 1$ and the acceptance term drops out of the equilibrium conditions. We retain $\bar{e}_x(\theta)$ in the derivations for internal consistency with the law of motion in equation (11).

[11] One can get this by imposing the steady state condition on $s'_x = (1-p(\theta_x)\bar{e}_x(\theta))s_x + \sum_k \delta_{x,k}\bar{g}_{x,k}(\theta)$ where s'_x is the mass of job seekers in the next period.

B.7 Analytical Wage Functions

In this subsection, we derive analytical expressions for the wage functions used in the main-text discussion (Section 4.7).

Collective bargaining. The collective bargaining problem is given by

$$\max_{w_x(\boldsymbol{g},a)} \left[\prod_x \Big(w_x(\boldsymbol{g},a)+v_x(a)-b_x\Big)^{\frac{g_x}{\sum_{x'} g_{x'}}}\right]^{\beta} \times \left[F(\boldsymbol{g})-\sum_{x'} w_{x'}(\boldsymbol{g},a)g_{x'}-\bar{C}(\boldsymbol{g},a)\right]^{1-\beta}, \tag{A25}$$

given $\boldsymbol{g}$ and a. The first-order condition is:

$$\frac{\beta}{\sum_{x'} g_{x'}}\left[F(\boldsymbol{g})-\sum_{x'} w_{x'}(\boldsymbol{g},a)g_{x'}-\bar{C}(\boldsymbol{g},a)\right]=(1-\beta)\Big(w_x(\boldsymbol{g},a)+v_x(a)-b_x\Big). \tag{A26}$$

Multiplying both sides by g_x and summing them over x gives:

$$\beta\Big[F(\boldsymbol{g})-\bar{C}(\boldsymbol{g},a)\Big]-\beta\sum_{x'} w_{x'}(\boldsymbol{g},a)g_{x'}=(1-\beta)\sum_x g_x\left(w_x(\boldsymbol{g},a)+v_x(a)-b_x\right), \tag{A27}$$

which can be rearranged to:

$$\sum_x w_x(\boldsymbol{g},a)g_x=\beta\Big[F(\boldsymbol{g})-\bar{C}(\boldsymbol{g},a)\Big]-(1-\beta)\sum_x g_x\left(v_x(a)-b_x\right) \tag{A28}$$

Plugging this back into the first-order condition yields

$$\frac{\beta(1-\beta)}{\sum_{x'} g_{x'}}\left[F(\boldsymbol{g})-\bar{C}(\boldsymbol{g},a)+\sum_{x'} g_{x'}\left(v_{x'}(a)-b_{x'}\right)\right]=(1-\beta)\left(w_x(\boldsymbol{g},a)+v_x(a)-b_x\right), \tag{A29}$$

which can be rearranged to the wage function:

$$w_{x,u}(\boldsymbol{g},a)=b_x-v_x(a)-\beta\Big(\bar{b}-\bar{v}(a)\Big)+\frac{\beta}{\sum_{x'} g_{x'}}\Big(F(\boldsymbol{g})-\bar{C}(\boldsymbol{g},a)\Big). \tag{A30}$$

Individual bargaining. Individual bargaining is given by

$$\max_{w_x(\cdot,a)} \Big(w_x(\boldsymbol{g},a)+v_x(a)-b_x\Big)^{\beta} \times \left[\frac{\partial F(\boldsymbol{g})}{\partial g_x}-\left(w_x(\boldsymbol{g},a)+c_x(a)\right)-\sum_{x'}\frac{\partial w_{x'}(\boldsymbol{g},a)}{\partial g_x}g_{x'}\right]^{1-\beta}, \tag{A31}$$

given a. Taking the first-order condition yields

$$\beta\left[\frac{\partial F(\boldsymbol{g})}{\partial g_x}-\left(w_x(\boldsymbol{g},a)+c_x(a)\right)-\sum_{x'}\frac{\partial w_{x'}(\boldsymbol{g},a)}{\partial g_x}g_{x'}\right]=(1-\beta)\Big(w_x(\boldsymbol{g},a)+v_x(a)-b_x\Big). \tag{A32}$$

This static case is a special case of Lemma 3 of Taschereau-Dumouchel (2020), which, with the boundary condition $\lim_{g_x \to 0} w_{x,n}(\boldsymbol{g}, a) g_x = 0$, gives

$$w_{x,n}(\boldsymbol{g}, a) = (1-\beta)(b_x - v_x(a)) + \frac{\beta}{1-(1-\alpha)\beta} \frac{\partial F(\boldsymbol{g})}{\partial g_x} - \beta c_x(a). \tag{A33}$$

B.8 Cost Sharing Under Collective Bargaining

To discuss the cost sharing, we keep the previous assumptions except that utility is now given by $u(w) + v_x(a)$, which is not necessarily linear in wages. The first-order condition with respect to the wage of worker type x is

$$\frac{u(w_x(\boldsymbol{g}, a)) + v_x(a) - u(b_x)}{u'(w_x(\boldsymbol{g}, a))} = \frac{\beta}{1-\beta} \frac{1}{\sum_{x'} g_{x'}} \left[F(\boldsymbol{g}) - \sum_{x'} w_{x'}(\boldsymbol{g}, a) g_{x'} - \bar{C}(\boldsymbol{g}, a) \right], \tag{A34}$$

for $x = 1, \ldots, X$. The left-hand side is the worker's match surplus normalized by the marginal utility of consumption, while the right-hand side is proportional to firm surplus per worker. Importantly, the right-hand side does not depend on worker type x, implying that collective bargaining equalizes worker surplus, normalized by marginal utility, across all types.

Amenity cost pass-through. We explore how an increase in the amenity cost is passed on to workers, given the worker composition $\boldsymbol{g}$. Fix the worker composition $\boldsymbol{g}$ and consider a firm providing amenity $a = 1$. Define $\bar{c} = \frac{1}{\sum_{x'} g_{x'}} \bar{C}(\boldsymbol{g}, a)$ as the total amenity cost per worker. The wage-setting condition can then be written as

$$u(w_x) + v_x(1) - u(b_x) = \frac{\beta}{1-\beta} u'(w_x) \left[\frac{F(\boldsymbol{g})}{\sum_{x'} g_{x'}} - \frac{1}{\sum_{x'} g_{x'}} \sum_{x'} w_{x'} g_{x'} - \bar{c} \right], \tag{A35}$$

where we suppress the dependence of w_x on $(\boldsymbol{g}, a)$ for notational simplicity.

To characterize how wages respond to a marginal increase in amenity costs, we totally differentiate this condition with respect to $\bar{c}$:

$$\left[u'(w_x) - S u''(w_x) \right] \frac{\partial w_x}{\partial \bar{c}} = -\frac{\beta}{1-\beta} u'(w_x) \left(1 + \frac{1}{\sum_{x'} g_{x'}} \sum_{x'} \frac{\partial w_{x'}}{\partial \bar{c}} g_{x'} \right), \tag{A36}$$

where

$$S = \frac{\beta}{1-\beta} \left(\frac{F(\boldsymbol{g})}{\sum_{x'} g_{x'}} - \frac{1}{\sum_{x'} g_{x'}} \sum_{x'} w_{x'} g_{x'} - \bar{c} \right) \tag{A37}$$

is the firm surplus per worker scaled by the relative bargaining power, which is nonnegative for the match to be formed. Rearranging yields

$$\frac{\partial w_x}{\partial \bar{c}} = -\frac{\beta}{1-\beta}\frac{u'(w_x)}{u'(w_x) - Su''(w_x)}\left(1 + \frac{1}{\sum_{x'} g_{x'}}\sum_{x'}\frac{\partial w_{x'}}{\partial \bar{c}}g_{x'}\right). \tag{A38}$$

To solve for $\frac{\partial w_x}{\partial \bar{c}}$, let $\chi = \frac{\beta}{1-\beta}\frac{1}{\sum_{x'} g_{x'}}\sum_{x'}\frac{u'(w_{x'})}{u'(w_{x'})-Su''(w_{x'})}g_{x'} \geq 0$, which does not depend on x. Substituting and rearranging, we obtain

$$\frac{\partial w_x}{\partial \bar{c}} = -\frac{\beta}{1-\beta}\frac{u'(w_x)}{u'(w_x) - Su''(w_x)} \cdot \frac{1}{1+\chi}. \tag{A39}$$

Since $u'' < 0$, we have $u'(w_x) - Su''(w_x) > 0$, so this derivative is strictly negative for any $\beta > 0$: an increase in the amenity cost lowers wages for all worker types.

If utility is CRRA, $u(w) = \frac{w^{1-\theta}}{1-\theta}$ with $\theta \geq 0$, the expression simplifies considerably. We have $u'(w_x) = w_x^{-\theta}$ and $u''(w_x) = -\theta w_x^{-\theta-1}$, so

$$\frac{u'(w_x)}{u'(w_x) - Su''(w_x)} = \frac{w_x^{-\theta}}{w_x^{-\theta} + \theta S w_x^{-\theta-1}} = \frac{w_x}{w_x + \theta S}. \tag{A40}$$

The wage response therefore becomes

$$\frac{\partial w_x}{\partial \bar{c}} = -\frac{\beta}{1-\beta}\frac{w_x}{w_x + \theta S} \cdot \frac{1}{1+\chi}, \tag{A41}$$

where χ is now given by $\chi = \frac{\beta}{1-\beta}\frac{1}{\sum_{x'} g_{x'}}\sum_{x'}\frac{w_{x'}}{w_{x'}+\theta S}g_{x'}$.

Two cases are instructive. First, in the risk-neutral case ($\theta = 0$), the expression reduces to

$$\frac{\partial w_x}{\partial \bar{c}} = -\frac{\beta}{1-\beta} \cdot \frac{1}{1+\chi} = -\beta, \tag{A42}$$

which is identical across all worker types: each worker's wage falls by the same amount β in response to a unit increase in the amenity cost per worker.

Second, with concave utility ($\theta > 0$), the magnitude of the wage response $\left|\frac{\partial w_x}{\partial \bar{c}}\right|$ is increasing in w_x, since $\frac{w_x}{w_x+\theta S}$ is increasing in w_x. This means that wages fall by more for high-wage workers than for low-wage workers in response to an increase in the amenity cost. Collective bargaining therefore implements an efficient sharing rule in which amenity costs are disproportionately absorbed by workers with lower marginal utility of consumption, i.e., those with higher wages.

Implications for amenity provision. The differential pass-through derived above has direct implications for firms' incentives to provide amenities. A firm's incentive to provide amenities depends on how much additional surplus it can extract from workers through wage adjustments. Under individual bargaining, amenity costs are not shared across skill types: a firm can only recoup the cost of an amenity from the workers who value it, and the presence of a consumption floor may prevent low-skill workers from accepting lower wages in exchange for amenities, making provision unprofitable. Under collective bargaining, by contrast, the cost-sharing rule described above allows the firm to adjust the wages of all workers, with high-wage workers absorbing a disproportionately large share of the cost. This makes it easier for the firm to recoup the total amenity cost even when some workers have low willingness to pay, and can therefore make amenity provision viable under collective bargaining when it would not be under individual bargaining.[12]

B.9 Numerical Algorithm

In this section, we lay out our numerical algorithm to solve for the equilibrium.

1. Provide an initial guess of a vector of market tightness, θ, wages $w_{x,y,k}(a)$ for each x, y, k, a, union probability $\mathcal{Q}_y$, insurance provision probability given union status $P_{y,k}$.
2. Solve for worker value functions by the value function iteration.[13]
3. Solve firm problems for each firm type and get $(w^*_{x,y,k}(a), \mathcal{Q}^*_y, P^*_{y,k}, g^*_{x,y,k}(a))$:

   a. Solve the bargaining problems. Discretize the space of $\boldsymbol{g}$ and approximate the partial derivatives by finite differences. Iterate the first-order conditions until wages converge. Obtain $w^*_{x,y,k}(\boldsymbol{g}, a)$ and $w^*_{x,y,k}(\boldsymbol{g}, a)$.
   b. Given the numerically solved wage functions, solve the firm hiring problem for each union status and insurance status. Obtain $g^*_{x,y,k}(a)$ and $w^*_{y,x,k}(a)$, which is the wage functions evaluated at the optimal hiring.
   c. Compute insurance provision probability and union probability $(\mathcal{Q}^*_y, P^*_{y,k})$.

4. Update wages, union probability, and insurance provision based on the solution in 3 as follows: $w^{new}_{x,y,k}(a) = \omega_w w^*_{x,y,k}(a) + (1-\omega_w) w_{x,y,k}(a)$, $g^{new}_{y,x,k}(a) = \omega_g g^*_{x,y,k}(a) + (1-\omega_g) g_{x,y,k}(a)$, $\mathcal{Q}^{new}_y = \omega_Q \mathcal{Q}^*_y + (1-\omega_Q)\mathcal{Q}_y$, and $P^{new}_{y,k} = \omega_P P^*_{y,k}(a) + (1-\omega_P) P_{y,k}(a)$, where $\omega_w, \omega_g, \omega_Q, \omega_P \in (0,1]$ are weights for facilitating convergence.
5. Compute $\mathcal{U}^{BC}_x(\theta)$ and $\mathcal{U}^{JC}_x(\theta)$ based on $(w^{new}_{x,y,k}(a), \mathcal{Q}^{new}_y, P^{new}_{y,k}, g^{new}_{x,y,k}(a))$.
6. Update market tightness. Increase θ_x if $\mathcal{U}^{BC}_x(\theta) > \mathcal{U}^{JC}_x(\theta)$ and otherwise decrease θ_x. Specifically, $\log \theta^{new}_x = \log \theta_x + \omega_\theta(\mathcal{U}^{BC}_x(\theta) - \mathcal{U}^{JC}_x(\theta))$ where $\omega_\theta > 0$ is a pre-specified constant chosen for facilitating convergence.

[12] The same argument applies when there are insurance loading costs: collective bargaining can make employer-sponsored insurance viable by spreading the loading cost across workers in proportion to their ability to bear it.

[13] In our estimated structural model, workers face medical expenditure shocks. We numerically integrate to calculate the worker's expected utility.

Importantly, the model incorporates sufficient shocks and heterogeneity. It helps us account for the observed heterogeneity in data and makes our algorithm very stable across different parameter configurations.

C Additional Details about Structural Estimation

C.1 Details about Quantitative Extension

In the decomposition analysis in Section 6.3, we introduce retirement ESHI coverage and Medicare. To capture retirement coverage, we introduce exogenous retirement that occurs stochastically with probability p^R, and assume that retired workers exit the model with probability p^D. The retirement value for a worker is given by $V_x^R(a) = \frac{u_x(c_x^R, a)}{1-\gamma(1-p^D)}$, which represents the expected discounted sum of post-retirement utilities $u_x(y_x^R, a)$, where y_x^R denotes income after retirement, which we set to b_x, and a denotes ESHI coverage prior to retirement. We assume that retirement shocks arrive after job destruction is realized. Under this timing assumption, the value of employment becomes

$$\begin{aligned} V_{x,y,k}^E(w,a) = u_x(w,a) + \gamma \Big[& p^R V_x^R(a_{x,y,k}) \\ & + (1-p^R)\delta_{x,k} V_x^U + (1-p^R)(1-\delta_{x,k}) V_{x,y,k}^E(w_{x,y,k}, a_{x,y,k}) \Big] \end{aligned} \tag{A43}$$

and the value of unemployment becomes

$$V_x^U = p_x^{Med} V_x^{U,I} + (1 - p_x^{Med}) V_x^{U,i} \tag{A44}$$

where the value of unemployment $V_x^{U,i}$ for each Medicaid eligibility status $i = I$ (eligible) and $i = N$ (noneligible) is given by

$$V_x^{U,i} = p(\theta_x) V_x^M + (1 - p(\theta_x)) \Big[u_x(b_x, a^i) + \gamma \Big\{ p^R V_x^R(0) + (1 - p^R) V_x^{U,i} \Big\} \Big] \tag{A45}$$

with $a^I = 1$ and $a^N = 0$. The firms' optimization and wage bargaining problems require slight modifications as well. Specifically, the job destruction rate $\delta_{x,k}$ is replaced by $\delta_{x,k} + p^R - \delta_{x,k} p^R$, since retirement constitutes an additional source of job destruction. The wage bargaining expressions are otherwise unchanged from the baseline model, except that worker values in bargaining now account for the future insurance gains from firm coverage. Finally, whenever a worker retires, an identical worker type enters the labor market as unemployed, which slightly modifies the Beveridge curve relationship between market tightness and unemployment. Specifically, we have $\mathcal{U}_x^{BC}(\theta) = \frac{1-p(\theta_x)}{p(\theta_x)} \sum_{k=u,n} (p^R + \delta_{x,k} - \delta_{x,k} p^R) \bar{g}_{x,k}(\theta)$.

In the extended model, firms also bear the cost of providing ESHI to retired workers. To preserve the structure of the firm's problem, we replace $c_x(a)$ with an *effective per-period cost* $\tilde{c}_x(a)$ paid only during employment, but chosen so that, in expectation, it covers both the worker's current and post-retirement ESHI costs. Note that, for a newly retired worker, the expected present-discounted sum of the costs, which we denote by $EC_x^R(a)$, is given by

$$EC_x^R(a) = \frac{c_x(a)}{1-\gamma(1-p^D)}. \tag{A46}$$

For a type-x worker still in the labor market, the expected present-discounted cost $EC_{x,k}(a)$ incorporates retirement shocks, job destruction shocks, and survival:

$$EC_{x,k}(a) = \frac{c_x(a) + \gamma p^R \, EC_x^R(a)}{1-(1-p^R)(1-\delta_{x,k})}. \tag{A47}$$

In the baseline model, firms paid $c_x(a)$ only while a worker was employed. In the extended model, this must be replaced by an *effective per-period cost* $\tilde{c}_x(a)$ that takes into account the expected future cost of the worker's post-retirement ESHI. Since $\tilde{c}_x(a)$ is paid only while the worker is employed, it must satisfy

$$EC_{x,k}(a) = \frac{\tilde{c}_x(a)}{1-(1-p^R)(1-\delta_{x,k})}. \tag{A48}$$

Equating (A47) and (A48) yields

$$\tilde{c}_x(a) = \left(1 + \frac{\gamma p^R}{1-\gamma(1-p^D)}\right) c_x(a). \tag{A49}$$

Note that if there is no retirement as in the baseline model ($p_R = 0$), the scaling factor reduces to 1 and the firm problem reduces to the one in the original model.

C.2 Informativeness of Moments

In this section, we investigate the informativeness of estimation moments following the approach by Honoré et al. (2020). It measures the informativeness of a moment by computing the proportional change in the asymptotic variance of estimated parameters from excluding the moment. Specifically, we use E_4 in Honoré et al. (2020), which is computed as follows: for each moment k, compute

$$I_k = \frac{\text{diag}(\tilde{\Sigma}_k - \Sigma)}{\text{diag}(\Sigma)} \tag{A50}$$

where Σ is the estimated asymptotic variance-covariance matrix and $\tilde{\Sigma}_k$ is the one obtained by removing k-th moment.

Table A.12 reports the results. We highlight several key patterns that are consistent with the identification arguments discussed in the main text.

For the insurance-related parameters, we find that the size gradient in insurance provision is particularly informative about the dispersion of the ESHI fixed cost, σ_a. Dropping either the ESHI offer rate among small firms or the ESHI offer rate among large firms substantially increases its asymptotic variance. The baseline fixed cost of insurance provision, FC^a, is primarily informed by overall insurance coverage rates, particularly the insured rate among low-skill workers. For the parameter governing unions' pooling benefit, λ, consistent with our identification argument, the estimated effect of unions on ESHI provision is especially informative.

For the union-related parameters, we find that moments describing the joint distribution of unionization and firm size are particularly informative. The marginal threat cost, c_0, is mainly informed by the employment shares of large union and large nonunion firms. The dispersion parameter, σ_{union}, is informed by a similar set of moments, although the relative importance of the employment shares of large and small firms differs. In contrast, the fixed cost of unionization, FC^{union}, is more heavily informed by overall union density. Finally, the break-up cost, ϕ, is primarily informed by the low-skill union wage premium and the overall wage level of low-skill workers.

D Additional Details about Counterfactual Experiments

D.1 Analysis of the Union Decline

We adopt two approaches to quantify the contributions of technological change, social insurance, and RTW laws to the decline in unionization between 1955 and 2007. The first approach, which we refer to as the backward simulation, fits the model to the 2007 economy—when technological change has already occurred, social insurance has expanded, and RTW laws have been partially implemented—and then removes each of these factors. The second approach, which we refer to as the forward simulation, fits the model to the 1955 economy—prior to major technological change, the expansion of social insurance, and the widespread adoption of RTW laws—and then introduces each of these factors. We provide detailed descriptions of both approaches below.

D.1.1 Backward Simulation

The 2007 baseline consists of two separately calibrated economies that differ by right-to-work status. The RTW economy represents workers in RTW states, accounting for 37.4% of workers, and the non-RTW economy represents the remaining 62.6%. Each economy is estimated separately, targeting the same set of moments as in Section 5, computed separately for RTW and non-RTW states except for the following: (i) the union ESHI effect based on the HRS, due to the absence of state identifiers in the public-use file; (ii) the union wage premium, due to the loss of precision from splitting the estimator; and (iii) the ESHI offer rates from the RWJ Employer Health Insurance Survey, also due to the absence of state identifiers in the public-use file. These three moments are computed on the pooled sample and matched in both economies. Table (A.13)-(A.18) report the externally set parameters, the estimated parameters, and the model fit separately for the RTW economy and the non-RTW economy. Each counterfactual simulation is performed in both economies, and the aggregate counterfactual moment is formed with the same weights.

Technological change simulation. We model skill-biased technological change by reducing the productivity weight of high-skill workers, $z_{2,y}$, uniformly across all firm types in both sub-economies, while keeping that of low-skill workers fixed. For each economy, we choose the reduction so that the model-implied aggregate skill wage gap falls by the target amount described below. We measure the skill premium in each period as the coefficient on a college indicator in a log-wage regression from the sample of full-year workers with positive earnings. The 2007 value is computed from the CPS Annual Social and Economic Supplement, and the mid-1950s value from the pooled 1950 and 1960 censuses; we rely on these two censuses for the early period because individual-level wage and education data are not available for 1955 itself. The change in the college coefficient between the two periods gives the observed increase in the skill premium of roughly 10 log points. Not all of this increase is due to technology: the existing literature finds that skill-biased technological change accounts for a substantial but incomplete share of the rise in the skill premium (e.g., Katz and Autor, 1999; Krusell et al., 2000). To set a reasonable target, we follow Krusell et al. (2000), who find that, had capital equipment continued to grow at its pre-1975 average rate, the skill premium would have risen by about 8% relative to 1963, compared with the 18% increase implied by the observed path of equipment. We therefore attribute a fraction $1 - 8/18 \approx 0.56$ of the observed rise to skill-biased technological change, giving a target reduction in the aggregate skill wage gap of 5.5 log points. While this estimate is obtained in a different quantitative framework, we use it as a disciplined conservative external benchmark to avoid attributing the entire observed increase in the skill premium to technological change. In each economy, we search over a grid of uniform reductions in the high-skill productivity weight $z_{2,y}$ and select the value at which the model-implied aggregate skill wage gap falls by 5.5 log points, while simultaneously scaling the high-skill unemployment benefit b_2 down by 5.5% to keep unemployment consumption fixed relative to wages.

Social insurance simulation. To simulate the absence of Medicare and Medicaid — both enacted in 1965 — we remove both programs from the model. Specifically, we set the Medicaid take-up probability to zero ($p^{Med} = 0$) and also remove Medicare coverage. We then recompute the equilibrium holding all other parameters at their 2007 calibrated values.

RTW law simulation. Between 1955 and 2007, the share of workers in RTW states rose from 18.3% to 37.4%, implying that 19.1 percentage points of workers enacted RTW laws over this period. As discussed in Section 6.3, we capture the effect of RTW law in reduced form through a cost parameter, c_{RTW}: in the right-to-work economy, the value a firm obtains from operating as a unionized firm, $J_{y,u}(\epsilon)$, is reduced by c_{RTW} relative to an otherwise identical economy without RTW laws. A positive c_{RTW} lowers the value of operating as a unionized firm in the firm's unionization decision (16) and thereby reduces union density. To quantify the union-density effect of this expansion, we first back out the implied RTW cost in the non-RTW economy: we find the value of the RTW parameter c_{RTW} such that imposing RTW laws in the non-RTW economy reduces union density by 2 percentage points, consistent with empirical estimates of the union density effect of RTW laws in the 2010s (Fortin et al., 2022). We then form the counterfactual that removes the RTW expansion. Workers are partitioned into three groups: always-RTW states, always non-RTW states, and the switching states that adopted RTW laws between 1955 and 2007. For the switching states we use the RTW sub-economy with the backed-out RTW cost removed, x_{switch}, representing their union density had they not adopted RTW laws. Formally, the counterfactual union density x is:

$$x = p_{\text{RTW}}\, x_{\text{RTW}} + p_{\text{noRTW}}\, x_{\text{noRTW}} + p_{\text{switch}}\, x_{\text{switch}},$$

where $p_{\text{RTW}} = 0.183$ is the always-RTW share, $p_{\text{noRTW}} = 0.626$ is the always non-RTW share, and $p_{\text{switch}} = 0.191$ is the employment share of switching states.

D.1.2 Forward Simulation

As a robustness check, we also conduct a forward simulation starting from the 1955 calibrated economy and introducing each factor in the direction of 2007. The 1955 baseline again consists of two separately calibrated economies that differ by right-to-work status. The RTW economy represents workers in RTW states, accounting for 18.3% of workers in 1955, and the non-RTW economy represents the remaining 81.7%. Each counterfactual simulation is performed in both economies, and the aggregate counterfactual moment is formed with the same weights.

One challenge is that there are not readily available data in 1955 for constructing moments. We construct some moments using extrapolation whenever it is possible while we use the same moments as in the 2007 economy when it is difficult.

First, the following information is available in the 1950s: wages, employment status, and education at the individual level. More specifically, these variables are not available in 1955, but they are available in the 1950 and 1960 censuses. Using this information in the 1950 and 1960 censuses, we calculate average wages for each education, the overall unemployment rate, and the fraction of workers of each skill type in 1950 and 1960 and interpolate them. Aggregate union density at the national level is available in 1955 in Farber et al. (2021), but we need to obtain union density in states with and without RTW laws separately. To do that, we first calculate the relative union density between RTW states and no RTW states in 1964 using state-level union density in Hirsch et al. (2001), and combine it with national-level union density to calculate union density in RTW states and no RTW states separately, assuming that the relative density is similar between 1955 and 1964.

Second, the following information is not available in the 1950s: insurance status, union status, firm size, and medical expenditure at the individual level. We need those variables to construct the moments, such as ESHI rate by education/union status, firm size by union status, and the distribution of medical expenditure by education. We use the data in the CPS in 1980 onward to calculate those moments, and then extrapolate them to obtain the moments in 1955 except for ESHI rates.[14]

As for ESHI rates, there are a few sources on the overall ESHI rates in the early years. First, the figure in page 11 of Health Insurance Association of America (1965) shows that the insured rate in 1954 is slightly above 60% while that in 1954 is about 70%. Another source is Cohen et al. (2009), which reports that 69.1% of people under age 65 were covered by hospital insurance from 1958 to 1960. From those numbers, we assume that 65% of employed workers were covered by ESHI in 1955. However, this number alone is not enough to obtain targeted moments for the ESHI rate, conditional on union status or skill types. To proceed, we assume that the relative ESHI rate between union workers and nonunion workers is similar over time. We also assume the relative ESHI rate between low-skill workers and high-skill workers is similar over time. We combine

[14]Extrapolations are linear in year except for employment shares, and the fraction of zero medical costs. In these cases, we make sure the values are between 0 and 1 by regressing $\log\left(\frac{y}{1-y}\right)$ on years where y is the variable of interest.

the relative ESHI rates in later years in the CPS and the aggregate ESHI rate of 65% in 1955 to calculate the ESHI rates conditional on either union status or skill types.

Tables A.19 and A.20 report the externally set parameters in states with RTW laws and without, respectively, while Tables A.21 and A.22 report the parameters internally estimated to match the extrapolated moments in states with and without RTW laws. The data moments and the simulated moments are reported in Tables A.23 and A.24.

Technological change simulation. This step mirrors that of the backward simulation, with the sign reversed. We raise the productivity weight of high-skill workers, $z_{2,y}$, uniformly across all firm types and scale up the high-skill unemployment benefit b_2 by 5.5%, choosing the increase in $z_{2,y}$ so that the aggregate skill wage gap rises to 5.5 log points above its post-social-insurance level. We then recompute the equilibrium under the updated parameters. See the corresponding paragraph in Section D.1.1 for the construction of the target.

Social insurance simulation. To simulate the introduction of Medicare and Medicaid — both enacted in 1965 — we add both programs to the 1955 economy. Specifically, we set the Medicaid take-up probability to $p^{Med} = (0.16, 0.09)/0.52$ and enable Medicare coverage. We then recompute the equilibrium holding all other parameters at their 1955 calibrated values.

RTW law simulation. Between 1955 and 2007, the share of workers in RTW states rose from 18.3% to 36.6%, implying that 18.3 percentage points of workers enacted RTW laws over this period. We reuse the RTW cost c_{RTW} backed out in the backward simulation, rescaled to the 1955 level by dividing it by the ratio of average firm value between the 2007 and 1955 economies, and impose it on the switching states. Formally, the counterfactual union density x is:

$$x = p_{\text{RTW}}\, x_{\text{RTW}} + p_{\text{noRTW}}\, x_{\text{noRTW}} + p_{\text{switch}}\, x_{\text{switch}},$$

where $p_{\text{switch}} = 0.183$ is the employment share of switching states, $p_{\text{RTW}} = 0.183$ is the always-RTW share, $p_{\text{noRTW}} = 0.634$ is the always non-RTW share, and x_{switch} is the union density in the non-RTW economy with the RTW cost imposed.

Main Findings. Table A.25 reports the forward simulation that simulates each channel forward from the 1955 economy. The estimated contribution of each factor to the decline in union density is quantitatively very close to our main backward-simulation estimates: social insurance accounts for 15% of the decline (compared to 16% in Table 8), skill-biased technological change for 33%

(compared to 32%), and RTW laws for 6% (compared to 1%). [15] The central conclusion that social insurance explains a sizable part of the union decline is therefore robust to the direction in which the counterfactuals are constructed.

D.2 Robustness Analysis

In this section, we provide robustness checks of our model analysis of the union decline in Section 6.3. Specifically, we explore how sensitive our results are to (i) the degree of self-insurance, (ii) additional firm heterogeneity in terms of TFP, and (iii) the specification of the union threat cost.

D.2.1 Robustness with Respect to Self-Insurance Channel

Due to computational constraints, we abstract away from self-insurance through savings in our model. Without buffer-stock savings, the model might implicitly assign a high marginal value of employer-provided insurance, which may inflate workers' willingness to pay for union membership and affect the tradeoff between unions and social insurance. To partially address this concern, our baseline model features a consumption floor $\underline{c}$ that guarantees workers a subsistence level of consumption in the face of large medical expenditure shocks, limiting the extent to which the absence of savings overstates the value of insurance. The presence of implicit insurance through uncompensated care is also considered to be an important determinant of the demand for health insurance (Finkelstein et al., 2019). To further assess the sensitivity of our results to this assumption, we explore how the findings change when workers have access to a stronger consumption buffer. Specifically, we increase $\underline{c}$ by 50% relative to the baseline, recalibrate the union-related parameters to match the observed union density, and quantify the contribution of each factor.

We emphasize that this exercise varies an exogenous self-insurance parameter and therefore does not capture endogenous changes in precautionary savings in response to social insurance expansions. We expect such behavioral responses to be modest in our setting: Aizawa and Fang (2020) discusses that the literature generally finds little evidence of precautionary saving responses to health risks for working adults, and uninsured individuals tend to hold very little financial wealth,

[15] The smaller RTW response in the backward decomposition relative to the forward simulation reflects the curvature of the logit unionization choice. Letting p be the logit unionization probability of a firm, the marginal effect of a change in the value/cost of unionization on p scales with $p(1-p)$, which is largest where the firm is closest to indifference between union and nonunion ($p = 0.5$) and small in the tails. In 1955, when baseline union density was high, a substantial mass of firm types sat near this indifference point, so the same nominal RTW shock moved unionization sharply. By 2007 most firm types lie deep in the nonunion tail of the choice, where the marginal response is mechanically small. This concern is specific to the RTW channel: the shifts in social insurance and technological change are large enough that the implied union-density response in the backward simulation is sufficiently dominated by the high-slope region close to the indifference point, and the forward and backward results for those two channels are comparable in magnitude.

limiting the scope for self-insurance through assets.[16] Together with our log-utility specification (see footnote 52 in Section 4.8), which is on the low end of values used in the literature and so already implies a small precautionary motive, this leads us to view the absence of savings as unlikely to overturn our main quantitative conclusions.

D.2.2 Model with TFP Heterogeneity

The firm-side heterogeneity in our baseline model is in terms of the returns to scale α_y and the relative skill intensity $z_{x,y}$. Another source of firm-side heterogeneity often considered in the literature is productivity. We assume that firms are heterogeneous in TFP A_y. In addition, we consider the case where TFP A_y is positively related to the returns to scale α_y. This case is of particular interest because, when more productive firms also tend to be more scalable, attributing the prevalence of unionization among large firms solely to returns to scale may overstate the role of production function concavity. Specifically, in our baseline model, firms with larger returns to scale are more likely to unionize, since a less concave production function reduces the gain from individual bargaining relative to collective bargaining. However, if productivity and returns to scale are positively correlated in reality, this sorting pattern reflects a joint force—firms unionize not only because they are more scalable, but also because they are more productive. Allowing for such a correlation therefore disciplines the model's predictions about which firms sort into union status, and may also alter how social insurance expansions affect union membership and wage inequality. Specifically, we let firm-level TFP depend linearly on the returns-to-scale parameter so that the firms with the largest returns to scale have TFP roughly 10 percent above the smallest returns-to-scale ones.

D.2.3 Robustness with Respect to Union Cost Functions

In our baseline specification, the union threat cost faced by a nonunion firm is assumed to be linear in the aggregate willingness to pay of its workers for unionization. While we provide theoretical justification for this functional form in Appendix B.3 through a voting model with a linear approximation, the linearity assumption directly governs how responsive firm-level unionization is to worker preferences, and therefore could influence the predicted magnitude of union decline following social insurance expansion. To assess the sensitivity of our results to this assumption, we consider a convex function of workers' aggregate willingness to pay and examine how adding the curvature affects our quantitative findings. Specifically, we raise the exponent on the union-threat cost from 1.0 in the baseline to 1.5.

[16] For example, the median asset value among working-age individuals without health insurance is only $619 in the 2013 Survey of Consumer Finances (Aizawa and Fang, 2020).

D.2.4 Results

Table A.28 reports a series of robustness checks in which we modify the model along three dimensions: raising the consumption floor, introducing a positive correlation between firm-level TFP and the curvature of the production function, and imposing convexity in the threat-cost function. In each case we re-fit the union-threat cost so that the model continues to match the 2007 union density target and then repeat the decomposition exercise. Across all three cases, the qualitative ranking and rough magnitudes of the three forces are essentially unchanged: technological change remains the most dominant driver of deunionization explaining one third of the union decline, the decline of social insurance accounts for a sizable share, and the spread of RTW laws contributes only marginally. Our main conclusions are therefore not driven by the particular structural assumptions adopted in the baseline.

D.3 Assessing Other Social Insurance Policies

In this section, we present simulation results for additional policy experiments.

Employer Mandate. First, we consider a policy that mandates firms to provide health insurance. Column (2) of Table A.26 reports the equilibrium effects of this policy change. When all firms are required to offer insurance, unionized firms are not as attractive to workers as in the baseline, raising hiring costs for union firms and reducing union density. The average wage falls and the skill wage gap widens, partly due to the decline in unionization, but also reflecting compensating differentials: since all workers now receive employer-provided insurance, they are willing to accept lower wages, and because insurance access improves disproportionately for low-skill workers, this effect further contributes to the widening of the skill wage gap.

Job Security. Second, we examine regulations that strengthen job security at nonunionized firms. Specifically, we set the job destruction rate of nonunionized firms to the midpoint between the baseline values $\delta_{x,u}$ and $\delta_{x,n}$, where $\delta_{x,u} < \delta_{x,n}$.[17,18] This experiment is intended to capture the effects of stricter workplace safety enforcement, such as through Occupational Safety and Health Administration (OSHA) regulations or more stringent employment protection legislation.

Column (3) of Table A.26 reports the equilibrium effects of this policy change. As with the counterfactual exercises in Section 6, the policy reduces union density by diminishing the relative attractiveness of unionized firms to workers. It also lowers the average wage and raises wage

[17] This corresponds to approximately a 7% reduction in the job destruction rate for nonunionized firms.

[18] To account for the potential costs to firms of providing enhanced job security, we assume that nonunion firms bear 5% of the union fixed cost FC^{union} through reduced output. We confirm that the qualitative results are not sensitive to this assumption.

inequality through deunionization. While better job security reduces unemployment and increases output, labor productivity declines, driven by a proportionally larger reduction in unemployment among low-skill workers.

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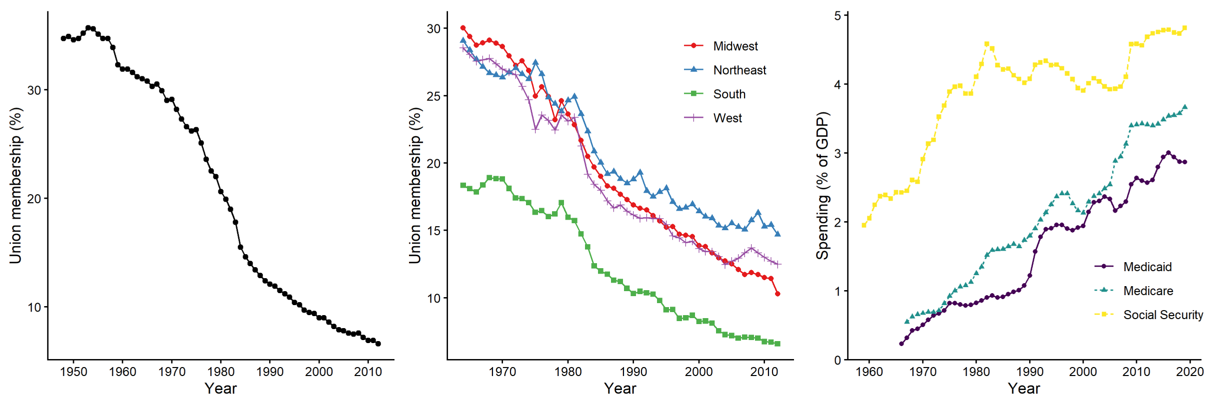
E Additional Figures and Tables

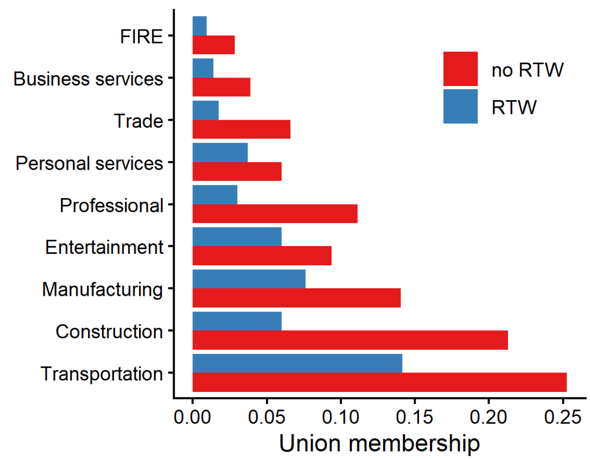
Figure A.1: National / Regional Trend in Union Membership and Spending on Social Insurance

(a) Union Membership (b) Regional Union Membership (c) Spending on Social Insurance



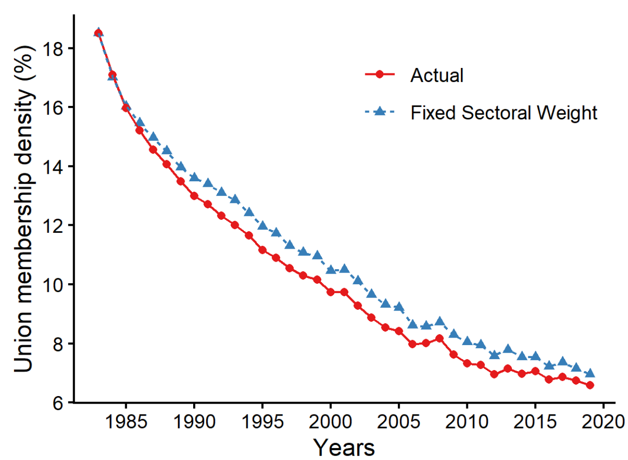
Note: Panel (a): Data are from Farber et al. (2021). The union density before 1983 is based on the survey conducted by the BLS while the data from 1983 onward are from the CPS. See Farber et al. (2021) for more detail. Panel (b): Data are from Hirsch et al. (2001). Panel (c): Data on the government spending on each social insurance program are from Federal Reserve Economic Data (FRED).

Figure A.2: Differential Exposure to RTW Laws by Industry



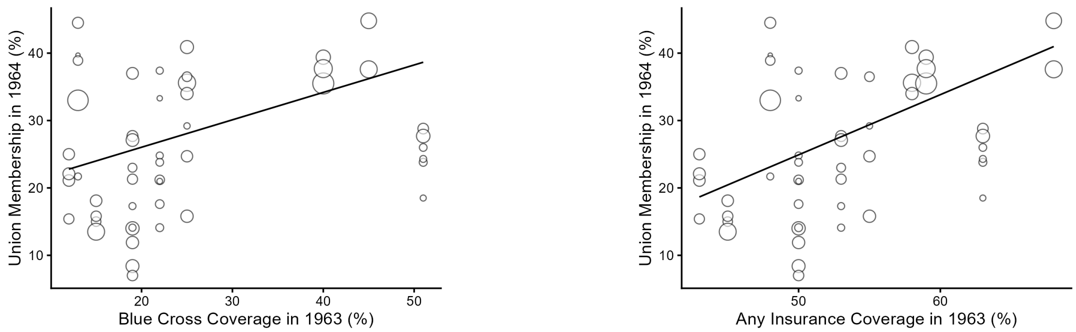
Note: This figure displays the exposure of each industry to RTW laws. Each bar represents the union membership in each industry, separately for states with RTW laws and those without. Data come from the CPS 2001-2019.

Figure A.3: Union Density with Fixed Sectoral Share



Note: This figure reports the actual union density (red solid line) and the counterfactual union density with fixed sectoral employment share (blue dashed line) based on equation $union_t^{CF} = \sum_{i \in \mathcal{I}} \omega_{i,1983} \times union_{i,t}$, where $\omega_{i,1983}$ is the sectoral employment share in 1983. Data come from the CPS 1983-2019.

Figure A.4: Private Insurance Coverage and Union Density Prior to Medicare

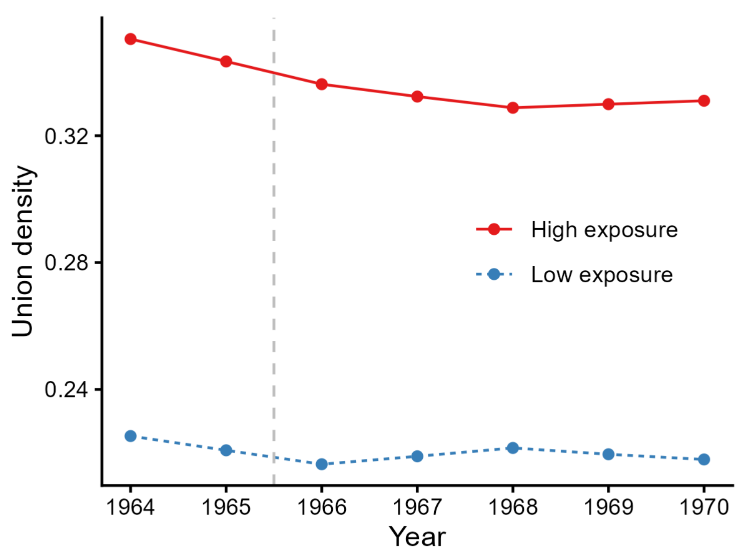


(a) Blue Cross Coverage

(b) Any Coverage

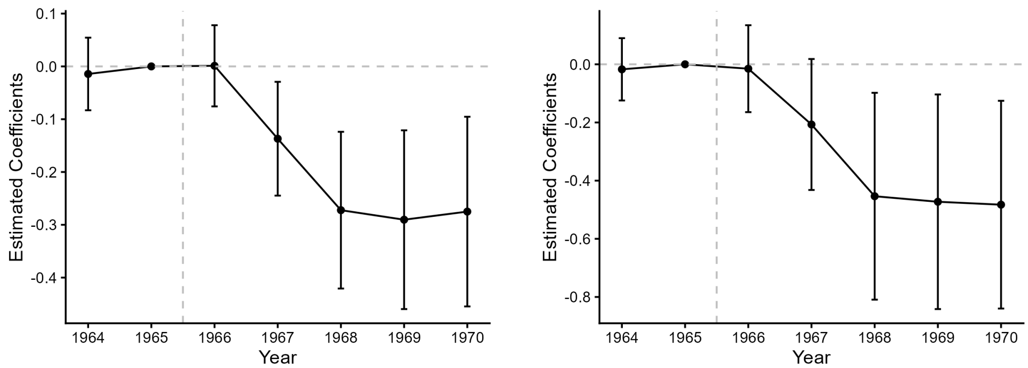
Note: The figure shows the relationship between private health insurance coverage for the elderly in 1963 (x-axis) and union membership in 1964 (y-axis). Data on state-level union density are from Hirsch et al. (2001); the insurance coverage is from Finkelstein (2007). Each circle corresponds to each state in the U.S. and the size of the circles represents the size of the state population in 1960.

Figure A.5: Union Density by Policy Exposure



Note: This figure displays the trend in union density in states with high policy exposure (red solid line) and the trend in union density in states with low policy exposure (blue dashed line). The high-exposure states are those with Blue Cross coverage above the median, and the remaining states are low-exposure states. Data on state-level union density are from Hirsch et al. (2001); the Medicare exposure measure is from Finkelstein (2007).

Figure A.6: Impact of Medicare Introduction on Union: Controlling Medicaid

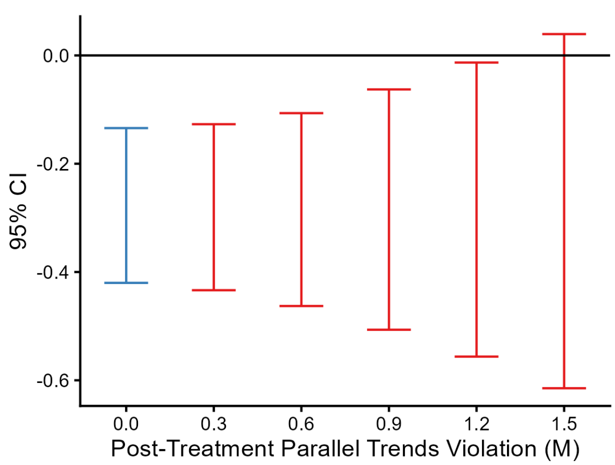


(a) Blue Cross Coverage

(b) Any Insurance Coverage

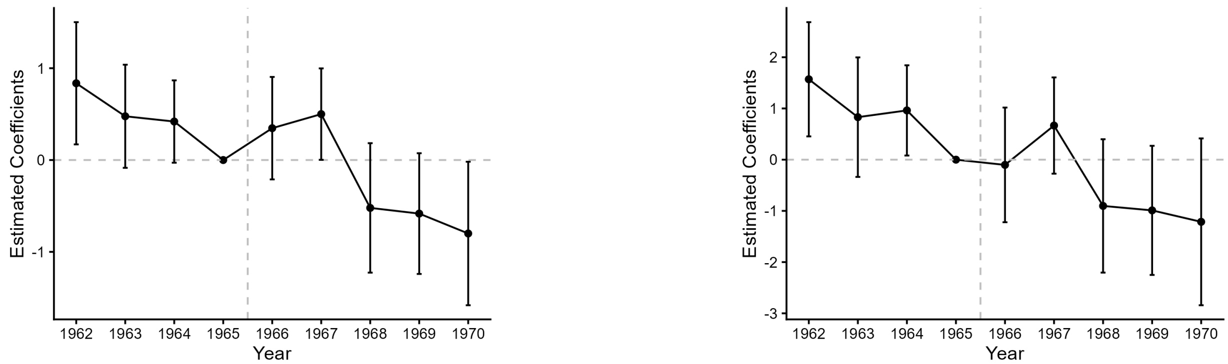
Note: This figure displays the estimated coefficients of equation (3) where we additionally control for four dummies for years before/after Medicaid implementation in each state. Data on state-level union density are from Hirsch et al. (2001); the Medicare exposure measure is from Finkelstein (2007). The error bars represent the 95% confidence intervals based on standard errors clustered at the state level.

Figure A.7: Robust Confidence Interval for $\hat{\beta}_{1968}$ based on Rambachan and Roth (2023)



Note: This figure reports the robust confidence set for $\hat{\beta}_{1968}$, based on the event-study specification in equation (3). Data on state-level union density are from Hirsch et al. (2001); the Medicare exposure measure is from Finkelstein (2007). For example, the confidence interval under $M = 1.2$ suggests that we restrict the post-treatment violations of parallel trends to be not 1.2 times as large as the maximal pretreatment violation of parallel trends, and we still find a statistically significant negative impact of the Medicare introduction.

Figure A.8: Impact of Medicare Introduction on Union Elections

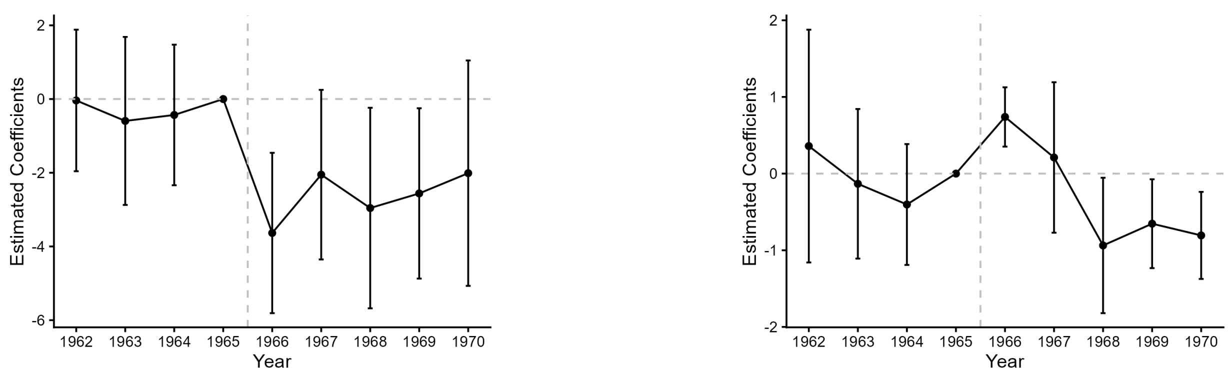


(a) Blue Cross Coverage

(b) Any Insurance Coverage

Note: This figure displays the estimated coefficients of equation (3) where the outcome is the log number of elections. Data on union elections are NLRB election data from Sojourner and Yang (2022); the Medicare exposure measure is from Finkelstein (2007). Panel (a): *Exposure* is Blue Cross insurance coverage in 1963. Panel (b): *Exposure* is any insurance coverage in 1963. The error bars represent the 95% confidence intervals based on standard errors clustered at the state level.

Figure A.9: Impact of Medicare Introduction on Union Elections

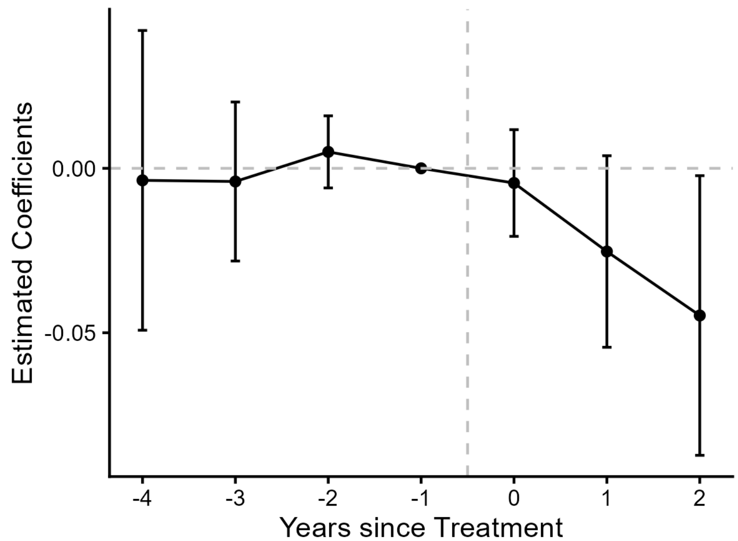


(a) Low Exposure

(b) High Exposure

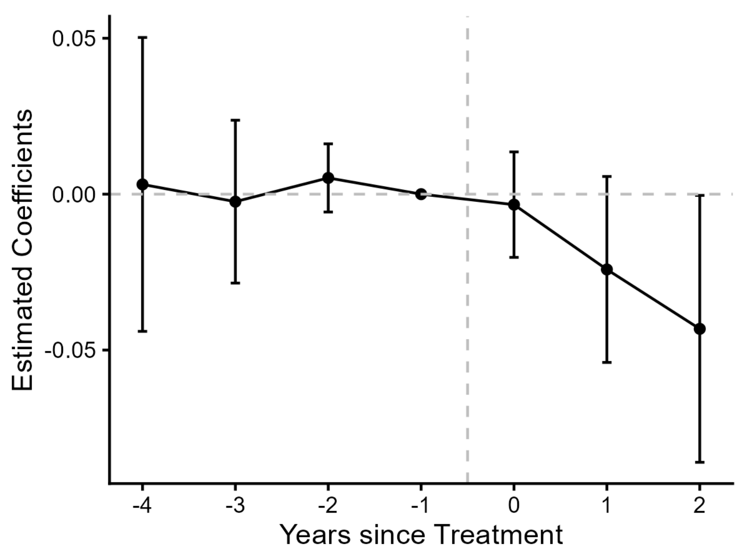
Note: This figure displays the estimated coefficients of equation (3) where the outcome is the log number of elections. Data on union elections are NLRB election data from Sojourner and Yang (2022); the Medicare exposure measure is from Finkelstein (2007). Low Exposure: *Exposure* is below median. High Exposure: *Exposure* is above median. *Exposure* is Blue Cross coverage in 1963. The error bars represent the 95% confidence intervals based on standard errors clustered at the state level.

Figure A.10: Estimated Impact of Medicaid Implementation on Union Density



Note: This figure displays the estimated coefficients of equation (A6). Data on state-level union density are from Hirsch et al. (2001); Medicaid adoption timing is from Gruber (2003). The error bars represent the 95% confidence intervals based on standard errors clustered at the state level.

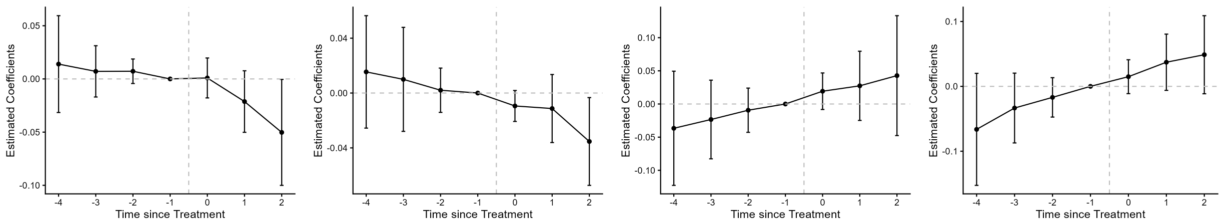
Figure A.11: Estimated Impact of Medicaid Implementation on Union Density: Controlling Medicare



Note: This figure displays the estimated coefficients of equation (A6), additionally controlling for Medicare exposure by including $\mathbb{1}\{t > 1965\} \times \textit{High exposure}_s$ where $\textit{High exposure}_s$ is an indicator for Blue Cross coverage higher than the median. Data on state-level union density are from Hirsch et al. (2001); Medicaid adoption timing is from Gruber (2003). The error bars represent the 95% confidence intervals based on standard errors clustered at the state level.

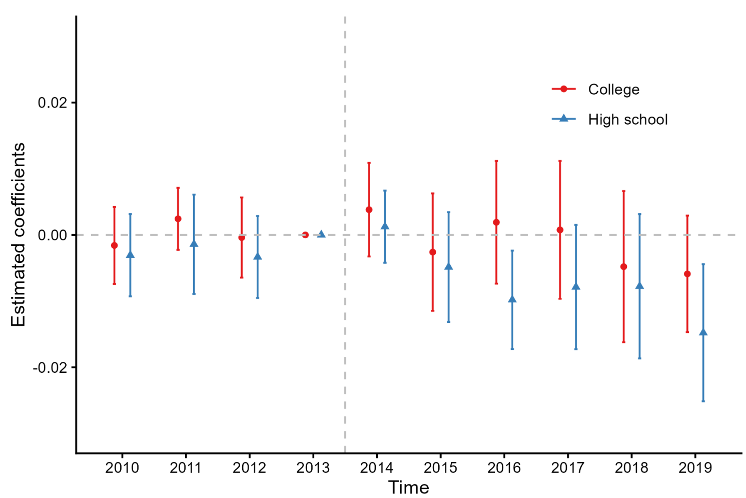
Figure A.12: Estimated Impact of Medicaid on Union Density: Standard TWFE Event Study and Placebo

(a) Actual ($k = 0$) (b) Placebo, $k = 1$ (c) Placebo, $k = 2$ (d) Placebo, $k = 3$



Note: Each panel reports estimates of equation (A7) for different values of the placebo shift k. Panel (a) uses the actual Medicaid adoption dates ($k = 0$); panels (b)–(d) shift each state's adoption date backward by k years for $k = 1, 2, 3$. Data on state-level union density are from Hirsch et al. (2001); Medicaid adoption timing is from Gruber (2003). The error bars represent 95% confidence intervals based on standard errors clustered at the state level.

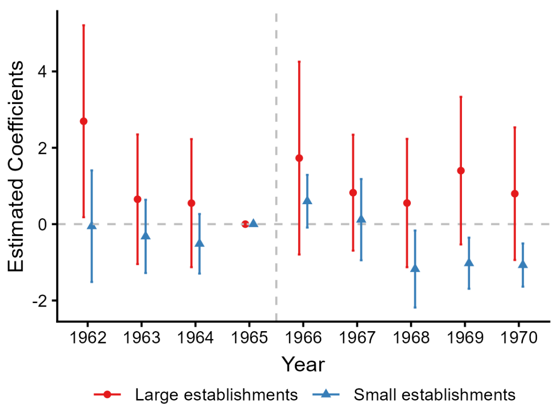
Figure A.13: ACA Medicaid Expansion Impact on Union Membership



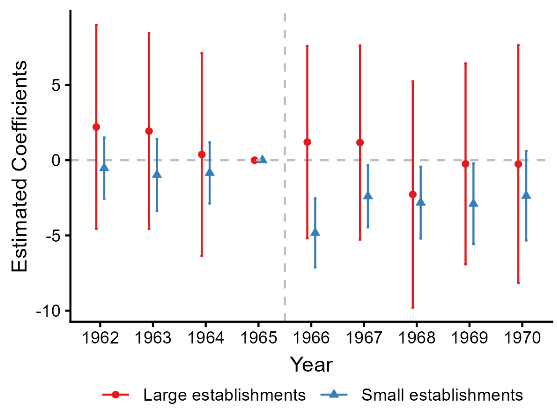
Note: This figure shows the estimated coefficients of equation $Union_{ist} = \sum_{\tau=2010,\tau\neq 2013}^{2019} \beta_\tau \times$ ACA Medicaid$_s$ $\times$ $\mathbb{1}[t = \tau] + x'_{ist}\gamma + \alpha_s + \mu_t + \varepsilon_{ist}$, where ACA Medicaid$_s$ is an indicator taking 1 if a state expanded Medicaid in January 2014. Data come from the CPS 2010-2019. States that expanded Medicaid in other periods during 2010-2019 are excluded. Other variables are the same as in equation (A8). Person-level weights are used. Standard errors are clustered at the state level. The error bars represent 95% confidence intervals.

Figure A.14: Testing the Fixed Cost Mechanism: Medicare Impact on Elections by Establishment Size

(a) High Exposure States

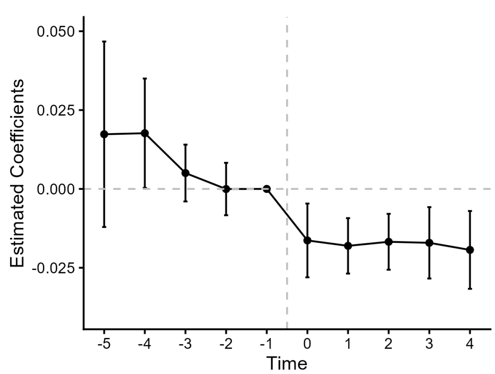


(b) Low Exposure States

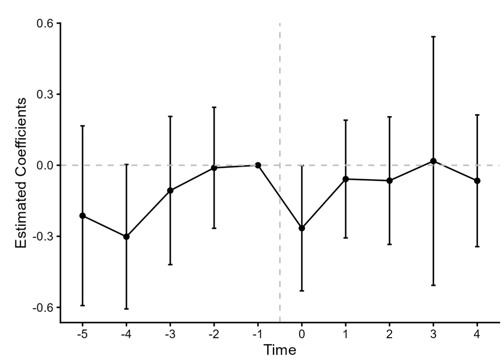


Note: The figure displays the event study plot of the regression of the inverse hyperbolic sine transformation of the number of union elections on the exposure to the Medicare introduction. Data on union elections are NLRB election data from Sojourner and Yang (2022); the Medicare exposure measure is from Finkelstein (2007). Panel (a) uses the sample of states with the exposure above the median and panel (b) uses the sample of states with the exposure below median. The exposure is based on Blue Cross coverage. We control for political variables, Medicaid timing dummies, state-level demographics, manufacturing employment share interacted with year dummies in addition to the state fixed effects and the year fixed effects. The standard errors are clustered at the state level. 95% confidence intervals are displayed.

Figure A.15: RTW Law Impact on Unionization



(a) Union membership



(b) Union elections

Note: Panel (a): The estimated coefficients of equation (A10) and their 95% confidence intervals are displayed. Data come from the CPS 2009-2019. Standard errors are clustered at the state level. Panel (b): The estimated coefficients of equation (A11) and the 95% confidence intervals are displayed. The sample comes from NLRB election data from Sojourner and Yang (2022). Standard errors are clustered at the state level.

Table A.1: Union Membership and Insurance Coverage

	ESHI (binary)			
	(1)	(2)	(3)	(4)
Union	0.215***	0.134***	0.058***	0.056***
	(0.010)	(0.010)	(0.018)	(0.018)
log(firm size)		0.043***	0.022***	0.021***
		(0.002)	(0.003)	(0.003)
log(earnings)		0.128***	0.039***	0.038***
		(0.005)	(0.005)	(0.005)
Mean outcome	0.72	0.72	0.72	0.72
Observations	32,787	32,787	30,192	30,192
Age & Individual FE			X	X
Industry & Occupation				X

Note: This table reports the estimation results of the regression of an indicator for ESHI on union membership based on equation (1) with various sets of covariates. Data come from the HRS 1992-2019. Year fixed effects and region fixed effects are controlled in all the specifications. Person-level analysis weights are used. Standard errors are clustered at the individual level. $^{*}p < 0.1$;$^{**}p < 0.05$;$^{***}p < 0.01$.

Table A.2: Unionization and Insurance Provision in Firm-Side Data

	ESHI (binary)	
	(1)	(2)
Union	0.264***	0.131***
	(0.028)	(0.028)
Mean outcome	0.50	0.50
Observations	21,292	21,292
Estab. size & Industry		X

Note: This table reports the establishment-level estimation results of the regressions of an indicator for insurance on an indicator for union and covariates. Data come from the 1997 Robert Wood Johnson Foundation Employer Health Insurance Survey. The second column controls for the log of establishment size and industry fixed effects. Establishment weights are used. Robust standard errors are reported. $^{*}p < 0.10$; $^{**}p < 0.05$; $^{***}p < 0.01$.

Table A.3: Estimated Impact of Medicare Introduction on Union Density: Interaction with Age Composition

	log(union density)	
	(1)	(2)
Exposure	0.230	0.079
	(0.300)	(0.222)
Exposure x Old	-0.665	-0.578**
	(0.471)	(0.248)
Blue Cross	X	
Any Coverage		X
Observations	343	343

Note: This table reports the estimated coefficients of the regression equation (A5). Data on state-level union density are from Hirsch et al. (2001); the Medicare exposure measure (the 1963 fraction of the elderly with retiree insurance coverage) is from Finkelstein (2007). *Old* is the share of workers aged 40 or older in 1964 computed in the CPS. The covariates are the same as in the baseline specification (2). Standard errors are clustered at the state level. $^{*}p < 0.10$; $^{**}p < 0.05$; $^{***}p < 0.01$.

Table A.4: Correlation between Union Membership and Demographic Variables Related to Insurance Demand

	Union membership	
	(1)	(2)
log(age)	0.085***	0.043***
	(0.003)	(0.002)
Married	0.009***	0.001
	(0.002)	(0.002)
Has child	0.016***	0.007***
	(0.002)	(0.001)
Bad health	-0.022***	-0.010
	(0.008)	(0.007)
log(earnings)		0.043***
		(0.001)
Demographics	X	X
Occ., Ind, Firm size		X
Observations	239,928	239,925

Note: This table reports the estimation results of the regression of union membership on various demographic variables. Data come from the CPS 2001–2019. Column (1) controls for demographic variables (sex, education, race), column (2) additionally controls for firm-related variables (firm size, 3-digit occupation, 3-digit industry, and log earnings). Year and state fixed effects are included in all columns. Robust standard errors are reported. $^{*}p < 0.1;^{**}p < 0.05;^{***}p < 0.01$.

Table A.5: RTW Law Impact on ESHI: 2SLS Regression Results

	ESHI	
	(1)	(2)
Union	0.346***	0.274***
	(0.083)	(0.077)
Mean outcome	0.64	0.83
KP F-stat	19.05	18.00
Observations	583,591	446,937

Note: The table reports the 2SLS regression results of the impact of RTW laws on ESHI, using the interaction between industry dummies and RTW laws as instruments for union membership. Data come from the CPS 2001–2019. The first- and second-stage specifications are given by (A3) and (A4), respectively. Column (1) includes the sample of employed workers while column (2) focuses on employed workers who are either covered by ESHI or not covered at all. Individual characteristics (age, education, race), 3-digit occupation, 3-digit industry, year fixed effects, and state fixed effects are controlled. Standard errors are clustered at the state $\times$ broad industry level. $^{*}p < 0.10$; $^{**}p < 0.05$; $^{***}p < 0.01$.

Table A.6: Union Membership and Job Losing

	Job Losing			
	Pooled		High school	College
	(1)	(2)	(3)	(4)
Union	-0.0020***	-0.0020***	-0.0028***	-0.0012***
	(0.0001)	(0.0001)	(0.0002)	(0.0002)
Mean outcome	0.007	0.007	0.008	0.006
Demographics		X	X	X
Observations	4,549,537	4,549,537	1,721,606	2,827,931

Note: This table reports the results from the regression of job loss on union membership based on equation (A1). Data come from the SIPP panels 1996, 2001, 2004, and 2008. Demographic controls include dummies for age, sex, race, and education. Person-level weights are used. State and year fixed effects are controlled in all specifications. Robust standard errors are reported. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

Table A.7: Union Membership and ESHI Coverage After Retirement

	ESHI after 65			
	(1)	(2)	(3)	(4)
Union	0.125***	0.100***	0.042	
	(0.014)	(0.014)	(0.030)	
Union to Nonunion				-0.013
				(0.034)
Nonunion to Union				0.079*
				(0.047)
Covariates		X	X	X
Individual FE			X	X
Mean outcome	0.15	0.15	0.15	0.15
Observations	11,675	11,675	9,861	9,861

Note: This table reports the estimation result of equation (1), with the outcome being post-retirement ESHI coverage. Data come from the HRS 2000-2019. Year and region fixed effects are controlled in all the specifications. In columns (2)-(4), we control for the quadratic polynomials for age, log earnings, log firm size, sex, education, and dummies for occupation and industry. Columns (3) and (4) additionally control for individual fixed effects. Person-level analysis weights are used. Standard errors are clustered at the individual level. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

Table A.8: Medicare Impact on Union Elections: DID Estimate

	log(elections)	
	(1)	(2)
Exposure × Post	-0.622**	-1.311**
	(0.308)	(0.491)
Coverage	Blue Cross	Any
Observations	432	432

Note: This table reports the estimation result of equation (2) with the outcome replaced by log union elections. Data on union elections are NLRB election data from Sojourner and Yang (2022); the Medicare exposure measure is from Finkelstein (2007). The treatment variable *Exposure* is based on Blue Cross insurance coverage in 1963 in column (1) and on any insurance coverage in column (2). State population in 1960 is used as weights. Standard errors are clustered at the state level. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

Table A.9: ACA Medicaid Expansion Impact on Union Membership

	Union Membership		
	All	High School	College
	(1)	(2)	(3)
ACA Medicaid	-0.003	-0.005**	-0.001
	(0.003)	(0.003)	(0.003)
Mean outcome	0.118	0.103	0.125
Observations	1,177,605	393,106	784,332

Note: This table reports the estimation result of equation (A8). Data come from the CPS 2010-2019. The first column uses the whole sample. The second column restricts the sample to individuals whose highest grade is not greater than the high-school graduate. The third column restricts the sample to individuals whose highest grade is greater than the high-school graduate. Person-level weights are used. Standard errors are clustered at the state level. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

Table A.10: Unemployment Insurance Impact on Union Membership

	Union Membership			
	(1)	(2)	(3)	(4)
Replacement Rate	-0.215***	-0.215***	-0.217***	-0.218***
	(0.020)	(0.020)	(0.020)	(0.021)
UI Duration FE		X	X	X
RTW Law			X	X
Political Control				X
Observations	2,680,492	2,680,492	2,680,492	2,598,607

Note: This table reports the estimation result of equation (A9). Data come from CPS 2000-2019. The information on UI generosity is obtained from "Significant Provisions of State Unemployment Insurance Laws" published by the BLS. Dummies for age, gender, education, occupation, industry, year fixed effects, and state fixed effects are controlled in all specifications. Standard errors are clustered at the state level. $^{*}p < 0.1;^{**} p < 0.05;^{***} p < 0.01$.

Table A.11: List of Externally Set Parameters

Parameter	Description	Value	Target
γ	Discount rate	0.988	5% annual interest rate
σ	Elasticity of substitution between skills	1.5	Johnson (1997)
μ	Match efficiency	1.0	Normalization
ζ	CRRA consumption term	1.0	
ϵ	Consumption floor ($1K)	1.0	
β_u	Bargaining power of union workers	0.5	
β_n	Bargaining power of nonunion workers	0.5	
N_z	Measure of workers of each type	0.41, 0.59	Fraction of each skill group
M	Measure of total firms	0.042	Average firm size
δ_u	Job destruction rate (union)	0.05, 0.03	See text
δ_n	Job destruction rate (nonunion)	0.06, 0.03	See text
b_x	Consumption during job loss ($1K)	5.73, 9.89	70% of average wages for each skill
$\mu_{H,z}$	Medical exp. distribution: location	-1.21, -1.08	Medical exp. distribution for each skill
$\sigma_{H,z}$	Medical exp. distribution: scale	1.73, 1.56	Medical exp. distribution for each skill
$p_{0,z}$	Medical exp. distribution: mass at zero	0.23, 0.11	Medical exp. distribution for each skill

Note: Monetary values are in $1,000 in year 2007. In columns that have two numbers, the first number corresponds to the value for low-skill workers $x = 1$, and the second for high-skill workers $x = 2$. For "See text", refer to the description of the job destruction rates in Section 5.1.

Table A.12: Moment Informativeness

Moment	$\bar{z}$	z slope	ϕ	FC^{union}	c_0	σ_{union}	FC^a	λ	σ_a	κ	A	Beta: a	Beta: b
Panel A: Unionization													
union density	0.212	0.487	0.096	0.287	0.324	0.491	0.004	0.312	0.000	0.005	0.000	0.054	0.041
emp. share 10 (union)	0.018	0.060	0.087	0.027	0.277	0.083	0.040	0.180	0.061	0.089	0.046	0.006	0.005
emp. share 100 (union)	0.504	0.183	0.241	0.499	0.928	1.214	0.010	1.045	0.126	0.047	0.087	0.002	0.042
emp. share 10 (non-union)	0.038	0.083	0.238	0.028	0.655	0.126	0.127	0.219	0.157	0.217	0.105	0.011	0.006
emp. share 100 (non-union)	0.683	0.235	0.478	0.825	1.810	1.748	0.033	1.483	0.197	0.141	0.129	0.005	0.074
Panel B: Wages													
wage (low skill)	6.604	0.017	1.659	0.563	0.310	0.236	3.133	0.032	9.176	26.627	15.944	0.394	0.315
wage (high skill)	0.748	0.002	0.062	0.003	0.000	0.183	0.025	1.208	0.544	0.009	7.342	0.061	0.003
union premium (low skill)	0.218	0.052	0.388	0.314	0.356	0.254	0.305	0.005	0.001	0.218	0.030	0.101	0.073
union premium (high skill)	0.003	0.000	0.019	0.008	0.023	0.004	0.011	0.000	0.000	0.012	0.000	0.003	0.002
Panel C: Insurance													
insured rate (non-union)	0.158	0.221	0.041	0.251	0.010	0.318	0.109	0.097	0.004	0.089	0.007	0.000	0.001
insured rate (union)	0.133	0.167	0.032	0.195	0.007	0.203	0.081	0.001	0.002	0.053	0.007	0.006	0.029
insured rate (low skill)	0.267	0.322	0.120	0.448	0.054	0.512	0.255	0.080	0.010	0.179	0.011	0.001	0.000
insured rate (high skill)	0.130	0.213	0.027	0.214	0.004	0.262	0.102	0.072	0.011	0.072	0.012	0.003	0.007
offer rate (small firms)	0.060	0.242	0.372	0.262	0.327	0.057	0.034	1.011	0.552	0.143	0.152	0.000	0.005
offer rate (large firms)	0.227	0.005	0.138	0.084	0.098	0.090	0.017	0.012	0.250	0.104	0.008	0.322	0.365
union ESHI effect	0.082	0.008	0.177	0.188	0.165	0.367	0.148	1.562	0.011	0.152	0.007	0.013	0.004
Panel D: Others													
unemployment rate	0.066	0.236	0.036	0.000	0.098	0.005	0.067	0.088	0.087	0.060	0.004	0.015	0.002
emp. share 10 (low skill)	0.021	0.073	0.248	0.034	0.659	0.143	0.106	0.230	0.109	0.218	0.080	0.001	0.008
emp. share 100 (low skill)	0.779	0.432	0.378	1.027	1.581	2.020	0.002	1.415	0.107	0.098	0.073	0.036	0.160
emp. share 10 (high skill)	0.022	0.039	0.249	0.020	0.616	0.108	0.106	0.209	0.088	0.238	0.063	0.000	0.006
emp. share 100 (high skill)	0.545	0.196	0.505	0.706	1.809	1.548	0.045	1.355	0.191	0.143	0.135	0.019	0.162

Note: This table reports the informativeness measure E_4 from Honoré et al. (2020). Each entry is the proportional increase in the asymptotic variance of the column parameter that results from excluding the row moment from the estimation, i.e., $I_k = \mathrm{diag}(\tilde{\Sigma}_k - \Sigma)/\mathrm{diag}(\Sigma)$, where Σ is the asymptotic variance-covariance matrix using all moments and $\tilde{\Sigma}_k$ is the corresponding matrix obtained when moment k is dropped.

Table A.13: List of Externally Set Parameters (RTW Economy, 2007)

Parameter	Description	Value	Target
γ	Discount rate	0.988	5% annual interest rate
σ	Elasticity of substitution between skills	1.5	Johnson (1997)
μ	Match efficiency	1.0	Normalization
ζ	CRRA consumption term	1.0	
ϵ	Consumption floor ($1K)	1.0	
β_u	Bargaining power of union workers	0.5	
β_n	Bargaining power of nonunion workers	0.5	
N_z	Measure of workers of each type	0.42, 0.58	Fraction of each skill group
M	Measure of total firms	0.043	Average firm size
δ_u	Job destruction rate (union)	0.04, 0.02	See text
δ_n	Job destruction rate (nonunion)	0.05, 0.03	See text
b_x	Consumption during job loss ($1K)	5.78, 10.03	70% of average wages for each skill
$\mu_{H,z}$	Medical exp. distribution: location	-1.21, -1.08	Medical exp. distribution for each skill
$\sigma_{H,z}$	Medical exp. distribution: scale	1.73, 1.56	Medical exp. distribution for each skill
$p_{0,z}$	Medical exp. distribution: mass at zero	0.23, 0.11	Medical exp. distribution for each skill

Note: This table reports externally set parameters for the RTW economy in 2007. Monetary values are in $1,000 in year 2007. For the "Value" column with two numbers, the first number corresponds to the value for low-skill workers $x = 1$, and the second for high-skill workers $x = 2$. For "See text", refer to the description of the job destruction rates in Section 5.1.

Table A.14: List of Externally Set Parameters (No RTW Economy, 2007)

Parameter	Description	Value	Target
γ	Discount rate	0.988	5% annual interest rate
σ	Elasticity of substitution between skills	1.5	Johnson (1997)
μ	Match efficiency	1.0	Normalization
ζ	CRRA consumption term	1.0	
ϵ	Consumption floor ($1K)	1.0	
β_u	Bargaining power of union workers	0.5	
β_n	Bargaining power of nonunion workers	0.5	
N_z	Measure of workers of each type	0.42, 0.58	Fraction of each skill group
M	Measure of total firms	0.042	Average firm size
δ_u	Job destruction rate (union)	0.05, 0.03	See text
δ_n	Job destruction rate (nonunion)	0.06, 0.03	See text
b_x	Consumption during job loss ($1K)	5.78, 10.03	70% of average wages for each skill
$\mu_{H,z}$	Medical exp. distribution: location	-1.21, -1.08	Medical exp. distribution for each skill
$\sigma_{H,z}$	Medical exp. distribution: scale	1.73, 1.56	Medical exp. distribution for each skill
$p_{0,z}$	Medical exp. distribution: mass at zero	0.23, 0.11	Medical exp. distribution for each skill

Note: This table reports externally set parameters for the no-RTW economy in 2007. Monetary values are in $1,000 in year 2007. For the "Value" column with two numbers, the first number corresponds to the value for low-skill workers $x = 1$, and the second for high-skill workers $x = 2$. For "See text", refer to the description of the job destruction rates in Section 5.1.

Table A.15: List of Internally Estimated Parameters (RTW Economy, 2007)

Parameters	Description	Estimate
A	TFP	40.0
$Beta(a,b):a$	Returns to scale distribution	0.72
$Beta(a,b):b$	Returns to scale distribution	0.70
$\bar{z}_1$	Low-skill worker relative skill: average	0.56
z_1: slope	Low-skill worker relative skill: slope	1.50
FC^a	Fixed cost of insurance provision	9.0
λ (1e-3)	Union ESHI cost advantage	1.19
σ_a	Std. dev. of insurance cost shock	9.00
FC^{union}	Fixed cost of unionization	21.8
σ_{union}	Std. dev. of union cost shock	4.00
c_0 (1e-2)	Marginal cost of union threat	2.07
κ	Vacancy posting cost	2.50
ϕ	Break-up cost	16.0

Note: This table reports the estimated model parameters for the RTW economy in 2007. Monetary values are 2007 USD.

Table A.16: List of Internally Estimated Parameters (No RTW Economy, 2007)

Parameters	Description	Estimate
A	TFP	42.3
$Beta(a,b):a$	Returns to scale distribution	0.72
$Beta(a,b):b$	Returns to scale distribution	0.70
$\bar{z}_1$	Low-skill worker relative skill: average	0.56
z_1: slope	Low-skill worker relative skill: slope	1.50
FC^a	Fixed cost of insurance provision	9.0
λ (1e-3)	Union ESHI cost advantage	1.19
σ_a	Std. dev. of insurance cost shock	9.00
FC^{union}	Fixed cost of unionization	21.8
σ_{union}	Std. dev. of union cost shock	4.00
c_0 (1e-2)	Marginal cost of union threat	2.68
κ	Vacancy posting cost	2.80
ϕ	Break-up cost	16.0

Note: This table reports the estimated model parameters for the no-RTW economy in 2007. Monetary values are 2007 USD.

Table A.17: Model Fit (RTW Economy, 2007)

Moment	Model	Data	Moment	Model	Data
(a) Cross-sectional moments					
ESHI coverage			*Firm size*		
ESHI coverage: union	0.86	0.91	Emp. share 10+: union	0.95	0.96
ESHI coverage: nonunion	0.74	0.69	Emp. share 10+: nonunion	0.83	0.85
ESHI coverage: low skill	0.67	0.58	Emp. share 100+: union	0.75	0.85
ESHI coverage: high skill	0.80	0.79	Emp. share 100+: nonunion	0.61	0.60
Wages and labor market			*Skill composition*		
Average wage: low skill	8.31	8.26	Emp. share 10+: low skill	0.69	0.83
Average wage: high skill	14.4	14.3	Emp. share 10+: high skill	0.94	0.88
Union density	0.04	0.05	Emp. share 100+: low skill	0.47	0.57
Unemployment rate	0.04	0.04	Emp. share 100+: high skill	0.72	0.64
(b) ESHI offer rates					
ESHI offer rate: small (<50)	0.37	0.48	ESHI offer rate: large (50+)	0.88	0.95
(c) Regression coefficients					
Union ESHI effect	0.05	0.06	Union wage premium: low skill	0.11	0.12
			Union wage premium: high skill	0.04	0.04

Note: This table reports the targeted data moments and their simulated counterparts for the RTW economy in 2007. "Emp. share $x+$: union" ("nonunion") is the share of union (nonunion) workers employed at firms whose total workforce is at least x where the numerator sums employment across union (nonunion) firms with $\geq x$ total employees and the denominator sums employment across all union (nonunion) firms. "Emp. share $x+$: low (high) skill" is defined analogously, restricting the numerator and denominator to low- (high-)skill employment, while the size cutoff x continues to refer to the firm's total employment across both skill groups.

Table A.18: Model Fit (No RTW Economy, 2007)

Moment	Model	Data	Moment	Model	Data
(a) Cross-sectional moments					
ESHI coverage			*Firm size*		
ESHI coverage: union	0.89	0.90	Emp. share 10+: union	0.98	0.95
ESHI coverage: nonunion	0.73	0.74	Emp. share 10+: nonunion	0.82	0.86
ESHI coverage: low skill	0.67	0.65	Emp. share 100+: union	0.91	0.79
ESHI coverage: high skill	0.80	0.84	Emp. share 100+: nonunion	0.58	0.59
Wages and labor market			*Skill composition*		
Average wage: low skill	8.44	8.50	Emp. share 10+: low skill	0.69	0.82
Average wage: high skill	15.1	15.1	Emp. share 10+: high skill	0.94	0.90
Union density	0.12	0.13	Emp. share 100+: low skill	0.48	0.54
Unemployment rate	0.05	0.05	Emp. share 100+: high skill	0.72	0.67
(b) ESHI offer rates					
ESHI offer rate: small (<50)	0.38	0.48	ESHI offer rate: large (50+)	0.89	0.95
(c) Regression coefficients					
Union ESHI effect	0.05	0.06	Union wage premium: low skill	0.13	0.12
			Union wage premium: high skill	0.04	0.04

Note: This table reports the targeted data moments and their simulated counterparts for the no-RTW economy in 2007. "Emp. share x+: union" ("nonunion") is the share of union (nonunion) workers employed at firms whose total workforce is at least x where the numerator sums employment across union (nonunion) firms with $\geq x$ total employees and the denominator sums employment across all union (nonunion) firms. "Emp. share x+: low (high) skill" is defined analogously, restricting the numerator and denominator to low- (high-)skill employment, while the size cutoff x continues to refer to the firm's total employment across both skill groups.

Table A.19: List of Externally Set Parameters (RTW Economy, 1955)

Parameter	Description	Value	Target
γ	Discount rate	0.988	5% annual interest rate
σ	Elasticity of substitution between skills	1.5	Johnson (1997)
μ	Match efficiency	1.0	Normalization
ζ	CRRA consumption term	1.0	
ϵ	Consumption floor ($1K)	0.1	
β_u	Bargaining power of union workers	0.5	
β_n	Bargaining power of nonunion workers	0.5	
N_z	Measure of workers of each type	0.87, 0.13	Fraction of each skill group
M	Measure of total firms	0.057	Average firm size
δ_u	Job destruction rate (union)	0.06, 0.03	See text
δ_n	Job destruction rate (nonunion)	0.07, 0.03	See text
b_x	Consumption during job loss ($1K)	0.51, 0.80	70% of average wages for each skill
$\mu_{H,z}$	Medical exp. distribution: location	-3.52, -3.84	Medical exp. distribution for each skill
$\sigma_{H,z}$	Medical exp. distribution: scale	1.02, 1.01	Medical exp. distribution for each skill
$p_{0,z}$	Medical exp. distribution: mass at zero	0.12, 0.09	Medical exp. distribution for each skill

Note: This table reports externally set parameters for the RTW economy in 1955. Monetary values are in $1,000 in year 1955. For the "Value" column with two numbers, the first number corresponds to the value for low-skill workers $x = 1$, and the second for high-skill workers $x = 2$. For "See text", refer to the description of the job destruction rates in Section 5.1.

Table A.20: List of Externally Set Parameters (No RTW Economy, 1955)

Parameter	Description	Value	Target
γ	Discount rate	0.988	5% annual interest rate
σ	Elasticity of substitution between skills	1.5	Johnson (1997)
μ	Match efficiency	1.0	Normalization
ζ	CRRA consumption term	1.0	
ϵ	Consumption floor ($1K)	0.1	
β_u	Bargaining power of union workers	0.5	
β_n	Bargaining power of nonunion workers	0.5	
N_z	Measure of workers of each type	0.84, 0.16	Fraction of each skill group
M	Measure of total firms	0.057	Average firm size
δ_u	Job destruction rate (union)	0.07, 0.04	See text
δ_n	Job destruction rate (nonunion)	0.08, 0.04	See text
b_x	Consumption during job loss ($1K)	0.66, 0.92	70% of average wages for each skill
$\mu_{H,z}$	Medical exp. distribution: location	-3.52, -3.84	Medical exp. distribution for each skill
$\sigma_{H,z}$	Medical exp. distribution: scale	1.02, 1.01	Medical exp. distribution for each skill
$p_{0,z}$	Medical exp. distribution: mass at zero	0.12, 0.09	Medical exp. distribution for each skill

Note: This table reports externally set parameters for the no-RTW economy in 1955. Monetary values are in $1,000 in year 1955. For the "Value" column with two numbers, the first number corresponds to the value for low-skill workers $x = 1$, and the second for high-skill workers $x = 2$. For "See text", refer to the description of the job destruction rates in Section 5.1.

Table A.21: List of Internally Estimated Parameters (RTW Economy, 1955)

Parameters	Description	Estimate
A	TFP	2.9
$Beta(a,b) : a$	Returns to scale distribution	0.22
$Beta(a,b) : b$	Returns to scale distribution	0.56
$\bar{z}_1$	Low-skill worker relative skill: average	0.78
z_1: slope	Low-skill worker relative skill: slope	0.40
FC^a	Fixed cost of insurance provision	0.4
λ (1e-6)	Union ESHI cost advantage	0.54
σ_a	Std. dev. of insurance cost shock	0.55
FC^{union}	Fixed cost of unionization	0.9
σ_{union}	Std. dev. of union cost shock	0.19
c_0 (1e-2)	Marginal cost of union threat	2.87
κ	Vacancy posting cost	0.05
ϕ	Break-up cost	1.5

Note: This table reports the estimated model parameters for the RTW economy in 1955. Monetary values are 1955 USD.

Table A.22: List of Internally Estimated Parameters (No RTW Economy, 1955)

Parameters	Description	Estimate
A	TFP	3.8
$Beta(a,b) : a$	Returns to scale distribution	0.20
$Beta(a,b) : b$	Returns to scale distribution	0.64
$\bar{z}_1$	Low-skill worker relative skill: average	0.78
z_1: slope	Low-skill worker relative skill: slope	0.40
FC^a	Fixed cost of insurance provision	0.3
λ (1e-6)	Union ESHI cost advantage	0.50
σ_a	Std. dev. of insurance cost shock	0.40
FC^{union}	Fixed cost of unionization	0.6
σ_{union}	Std. dev. of union cost shock	0.20
c_0 (1e-2)	Marginal cost of union threat	3.76
κ	Vacancy posting cost	0.06
ϕ	Break-up cost	1.5

Note: This table reports the estimated model parameters for the no-RTW economy in 1955. Monetary values are 1955 USD.

Table A.23: Model Fit (RTW Economy, 1955)

Moment	Model	Data	Moment	Model	Data
(a) Cross-sectional moments					
ESHI coverage			*Firm size*		
ESHI coverage: union	0.83	0.84	Emp. share 10+: union	0.97	0.99
ESHI coverage: nonunion	0.69	0.60	Emp. share 10+: nonunion	0.82	0.86
ESHI coverage: low skill	0.71	0.64	Emp. share 100+: union	0.85	0.95
ESHI coverage: high skill	0.75	0.71	Emp. share 100+: nonunion	0.56	0.59
Wages and labor market			*Skill composition*		
Average wage: low skill	0.72	0.73	Emp. share 10+: low skill	0.83	0.83
Average wage: high skill	1.25	1.14	Emp. share 10+: high skill	0.94	0.88
Union density	0.17	0.19	Emp. share 100+: low skill	0.59	0.57
Unemployment rate	0.05	0.05	Emp. share 100+: high skill	0.69	0.64
(b) ESHI offer rates					
ESHI offer rate: small (<50)	0.38	0.48	ESHI offer rate: large (50+)	0.77	0.95
(c) Regression coefficients					
Union ESHI effect	0.05	0.06	Union wage premium: low skill	0.15	0.12
			Union wage premium: high skill	0.05	0.04

Note: This table reports the targeted data moments and their simulated counterparts for the RTW economy in 1955. "Emp. share $x+$: union" ("nonunion") is the share of union (nonunion) workers employed at firms whose total workforce is at least x where the numerator sums employment across union (nonunion) firms with $\geq x$ total employees and the denominator sums employment across all union (nonunion) firms. "Emp. share $x+$: low (high) skill" is defined analogously, restricting the numerator and denominator to low- (high-)skill employment, while the size cutoff x continues to refer to the firm's total employment across both skill groups.

Table A.24: Model Fit (No RTW Economy, 1955)

Moment	Model	Data	Moment	Model	Data
(a) Cross-sectional moments					
ESHI coverage			*Firm size*		
ESHI coverage: union	0.86	0.75	Emp. share 10+: union	0.91	0.98
ESHI coverage: nonunion	0.72	0.56	Emp. share 10+: nonunion	0.79	0.83
ESHI coverage: low skill	0.76	0.62	Emp. share 100+: union	0.67	0.84
ESHI coverage: high skill	0.83	0.68	Emp. share 100+: nonunion	0.64	0.55
Wages and labor market			*Skill composition*		
Average wage: low skill	0.96	0.94	Emp. share 10+: low skill	0.81	0.82
Average wage: high skill	1.37	1.31	Emp. share 10+: high skill	0.94	0.90
Union density	0.36	0.39	Emp. share 100+: low skill	0.63	0.54
Unemployment rate	0.07	0.05	Emp. share 100+: high skill	0.75	0.67
(b) ESHI offer rates					
ESHI offer rate: small (<50)	0.36	0.48	ESHI offer rate: large (50+)	0.83	0.95
(c) Regression coefficients					
Union ESHI effect	0.06	0.06	Union wage premium: low skill	0.12	0.12
			Union wage premium: high skill	0.06	0.04

Note: This table reports the targeted data moments and their simulated counterparts for the no-RTW economy in 1955. "Emp. share $x+$: union" ("nonunion") is the share of union (nonunion) workers employed at firms whose total workforce is at least x where the numerator sums employment across union (nonunion) firms with $\geq x$ total employees and the denominator sums employment across all union (nonunion) firms. "Emp. share $x+$: low (high) skill" is defined analogously, restricting the numerator and denominator to low- (high-)skill employment, while the size cutoff x continues to refer to the firm's total employment across both skill groups.

Table A.25: Deunionization Analysis (Forward Simulation)

	Social Insurance	Tech Change	RTW Laws
Contribution to union decline (%)	15.2	33.2	6.1

Note: This table reports the fraction of the decline in union density between 1955 and 2007 explained by social insurance, skill-biased technological changes, and RTW laws.

Table A.26: Additional Policy Experiments

	(1) Baseline	(2) Mandate	(3) Job security
Union density (%)	10.30	9.14	8.82
ESHI coverage (%)			
Overall	72.9	100.0	72.9
Union	86.6	100.0	85.2
Nonunion	71.3	100.0	71.8
Low skill	66.2	100.0	66.3
High skill	77.1	100.0	77.2
Unemployment rate (%)			
Overall	4.7	4.2	4.4
Low skill	9.1	8.1	8.5
High skill	1.6	1.6	1.5
Output per capita (% change)	-	0.27	0.21
Labor productivity (% change)	-	-0.20	-0.13
Average wage (% change)	-	-1.77	-0.10
Skill wage gap (log points)	57.9	61.6	58.3

Note: This table reports the general equilibrium impacts of each policy change. Column (2) is the economy where all firms are forced to offer insurance to workers. In column (3), nonunion firms are forced to provide better job security than the baseline.

Table A.27: Social Insurance Expansion: Tax Adjustment

	Endogenous tax	Fixed tax
Contribution to union decline (%)	15.8	15.9
Endogenous union:		
ESHI rate (p.p. change)	-4.6	-4.3
Skill wage gap (log points change)	-2.0	-2.0
Welfare (% change)	0.9	0.9
Fixed unionization:		
ESHI rate (p.p. change)	0.6	0.6
Skill wage gap (log points change)	-2.1	-2.1
Welfare (% change)	0.7	0.7

Note: This table reports the social insurance expansion simulation with and without the tax adjustment.

Table A.28: Deunionization Analysis: Robustness

	Social Insurance	Tech Change	RTW Laws
High consumption floor	16.1	32.6	1.7
TFP and curvature correlation	20.6	34.8	1.9
Convex threat cost	17.8	30.2	2.5

Note: This table reports the robustness check of the deunionization analysis.