A Macroeconomic Analysis of the Rising Skill Premium and Deunionization in the United States*

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Abstract

During the past 50 years, the US economy has been characterized by a rapid decline of labor unions and a substantial rise in the wage inequality. This paper proposes that the rise in the skill premium in the non-union sector, for instance, due to technical change, can potentially explain these trends. Based on the premise that labor unions compress wages between skilled and unskilled workers, a larger skill premium encourages skilled workers to withdraw from the union. If this is accompanied by a fall in the productivity of unskilled workers, firms become more reluctant to hire the relatively expensive union workers, reinforcing the decline in the unionization rate. To evaluate our hypothesis, we develop a macroeconomic model of endogenous union membership with heterogenous agents, where union members are selected from the middle of the skill distribution and have significant wage gains that are decreasing in skill, consistent with the US evidence. The model predicts that the rise in skill prices in the non-union sector explains 30-60% of the decline in the unionization rate. It was argued that the declining union activity contributed to the rise in wage inequality by changing the labor force composition. We find this effect to be much smaller due to selection into union jobs.

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The US economy witnessed significant changes in the earnings distribution throughout the 20th century. The Gini coefficient of earnings inequality was 0.48 in 2004, up from a century low of 0.32 in the late-1950s (Figure 1). The trends in wage inequality were closely mirrored by the rate of unionization, which rose sharply after the Great Depression, peaking at around 28% in the 1950s, and decreasing ever since. These trends are not peculiar to the US. Australia, the Netherlands and the UK experienced similar hikes in wage inequality accompanied by declines in the unionization rate between 1980 and 1992, while labor unions were gaining strength in Italy and Belgium, where the wage inequality declined (Wallerstein, 1999; Wallerstein and Western, 2000).[1]

The negative correlation between wage inequality and unionization is commonly interpreted as the outcome of efforts by labor unions to achieve a more equal distribution of wages among their members. Although earlier work found that unions increased the overall wage inequality by negotiating higher wages relative to non-union workers (Lewis, 1963), more recent studies, based on extensive micro-level census data, disagree (Hirsch, 1982; Freeman, 1980; Card, 1996). Egalitarian union practices, such as limited inclusion of performance-based pay in union contracts or negotiation of across-the-board wage raises for all workers, result in a compressed wage structure in union establishments. Freeman and Medoff (1984) particularly emphasize the strong negative impact of unions on plant level wage inequality.

The abundance of evidence on wage compression has led to claims that the decline of the unions was a major factor explaining the rise in wage inequality in the US (Freeman, 1993; DiNardo, Fortin, and Lemieux, 1996; Card, 1998; Fortin and Lemieux, 1997; Koeniger, Leonardi, and Nunziata, 2007). These studies, aimed at measuring the rise in wage inequality due to deunionization, take unions as exogenous institutions. Nevertheless, when the prevalence of unions is considered as an outcome of economic incentives of firms and workers, one could also argue that a third factor caused the observed trends, given the egalitarian practices by labor unions.

In a labor market with free mobility, the extent of unionization depends crucially on the productivity distribution, workers' choices among union and non-union jobs, and firms' hiring decisions. If unions establish transfers among their members to achieve equality, then workers weigh the rents extracted by negotiating collectively against the transfer payments within the union. These payments may discourage skilled workers from joining the union, while encouraging the unskilled workers. On the other hand, if unionized firms have discretion over whom to hire in response to wages set by the union, they would do so selectively, admitting only those who are productive enough to make up for the high union wage. In such an economy, a change in the productivity differences among workers, for instance, due to technical change, alters the checks and balances within the unionized establishments, leading to a change in the wage distribution as well as the unionization rate. This is the direction of causality we wish to examine in this paper.

We hypothesize that a rise in the skill premium in the non-union sector, for instance, led by a skill-biased technical change, could both raise the wage inequality and reduce the rate of union-

[1] Wallerstein (1999) documents a significant negative link between the degree of centralization of wage-setting institutions and wage inequality, in his study of 16 OECD countries.

ization. Consider the particular case of a technical change that increases the productivity of skilled workers while deflating the value of skills possessed by low-productivity workers.[2] The higher skill premium raises the opportunity cost of being a union member for skilled workers, and encourages them to get jobs in the non-union sector, where they are better rewarded for their skills. Meanwhile, as the productivity of the unskilled workers declines, firms become reluctant to hire them at the relatively higher union wage. This could potentially generate a large decline in the rate of unionization. We argue below that this scenario is a reasonable description of the US deunionization experience.

Our main goal is to provide a theory of deunionization that is consistent with the observed relationship between skills, union participation, and wage distribution. To this end, we develop a macroeconomic model of an economy with search frictions, heterogeneous workers, and endogenous union participation. Our model generates positive and significant wage gains to union workers that are decreasing in skill, and predicts that union workers are selected from the middle of the productivity distribution, consistent with the evidence in the literature (Johnson, 1975; Lewis, 1986; Card, 1996; Lemieux, 1998; Hirsch and Schumacher, 1998).

We use data on union participation and wage rates by skill, as measured by education and experience in 1978, to identify the parameters of our model.[3] Then, we use our model to evaluate the change in the rate of unionization in response to the observed rise in skill prices in the union and in the non-union sectors. The rate of unionization in our sample declines from 32.1% to 10.2% between 1978 and 2007. The model predicts a decline of 7 to 13 percentage points in the unionization rate, corresponding to a third to over a half of the observed decline since 1978.

We also consider the compression effect of labor unions on the wage distribution. Once we account for non-random selection into unions using our model, we find that unions have little effect on the equilibrium wage distribution, unlike previous studies (Freeman, 1993; DiNardo, Fortin, and Lemieux, 1996; Fortin and Lemieux, 1997). In particular, the selection of the workers from the middle of the skill distribution leads to an overestimation of the effect of unions on wage dispersion.[4] Studies that attempted to control for selection, such as Card (2001) and Lemieux (1998) are closer to our findings.[5]

In the next section, we briefly discuss the US deunionization experience and provide an evaluation of the existing explanations in the literature. Among these, our work is closer to Acemoglu, Aghion, and Violante (2001), who provide a theoretical assessment of the incentives of skilled workers in face of a skill-biased technical change. While our results support some of the conjectures in their paper, our approach differs considerably from theirs as we discuss in the next section.

[2]Although we use technical change as an example, an alternative explanation of lower productivity, for instance, based on the emergence of international competition for low-skill jobs, would be equivalent.

[3]Our approach requires micro-level data on union coverage in the US, which has only been available since 1978.

[4]With a similar reasoning, Boyle (2008) recently showed that the discipline-independent pay schedules in New Zealand universities fostered the research performance of disciplines with otherwise low outside options, while severely curbing the quality of those with high market values.

[5]Card (2001) only controls for selection, based on observables, and therefore finds a somewhat larger effect than our model would predict. Lemieux (1998) attempted to control for selection with respect to unobservables using panel data from Canada.

We describe our model in section 2 and present the conditions under which the rising skill premium leads to a decline in the unionization rate. In section 3, we estimate some of the model parameters and present our calibration. We discuss our results in section 4. Following a robustness analysis in section 5, we conclude in section 6.

1 The Deunionization Experience in the United States

We begin by discussing the timing of the changes in the unionization rate, wage inequality, and skill premium, to put the alternative explanations in perspective. Based on Figure Figure 1, we deduce that the rise in the wage inequality began around 1960. The rate of unionization, on the other hand, begins to decline in the early-1950s when almost a third of the workers belonged to unions. This decline was limited to three percent until 1970, and thereafter, the unionization rate decreased by 13 percent. The antecedence of the rise in the wage inequality suggests that deunionization was more likely the outcome than the cause.

To place the change in the skill premium in context, Figure 2 depicts the unionization rate against the wage ratios of college graduates to high school graduates, and of high school graduates to workers with eight years of education or less. Both measures declined during the first half of the 20th century, and have been increasing since 1950, indicating a strong negative correlation between unionization and skill premium.

An exception to this relationship is the downswing in the college wage premium in the 1970s, when unionization was declining rapidly. However, this does not challenge our argument since the decline in the skill premium in the 1970s is attributed to an increase in the supply of college workers, and not to a drop in the demand for skill (Welch, 1979; Katz and Murphy, 1992; Goldin and Katz, 2008; Kaymak, 2009). As college workers have relatively low unionization rates, a larger supply of college workers adds to deunionization while suppressing the college premium. Furthermore, since the larger supply of college graduates is likely endogenous to the higher return to college, the behavior of the college premium and union density in the 1970s is not at odds with our hypothesis. Therefore, we conclude that the timing of events is *prima facie* consistent with our argument that the higher skill premium led to deunionization.[6]

Since our analysis requires micro-level data on wages and union coverage, we limit our study to the 1978-2007 period. We focus on male workers in the private sector to maintain a relatively homogenous sample of workers. Although the unionization rate among female workers is generally lower than that of male workers, a mildly declining pattern can also be observed among female workers. In contrast, the unionization rate in the public sector differs considerably from the general trends in Figure 1. The unions rose significantly in the public sector during the 1970s and have remained fairly stable since the mid-1980s (Farber, 2005).

The nature of the contracts and the limited room for performance-based pay in the government

[6]The unavailability of micro-level data on wages and union status for most of the 20th century prevents us from examining these patterns separately for the union and non-union sectors. Clearly, any reasoning based only on these figures is suggestive, not conclusive.

sector are similar to unionized establishments. Therefore, the choice between a union job and a non-union job, is akin to the choice between a public sector job and private sector job. In fact, Katz and Krueger (1991) found that the wage differentials by skill in the public sector were unresponsive to the rising skill premium in the private sector, during the 1970s and the 1980s. Furthermore, when faced by the demand by unions for an increase in compensation, public sector employers can partially meet the higher costs through an increase in taxes, whereas, the private sector employers face stiff budget constraints. Therefore, the employers' incentives are also different in the public sector. Thus, we consider the public sector as being outside the scope of our analysis.[7]

Figure 3 shows the increase in the wage inequality and the skill premium in our sample. The wage ratio of the 90th percentile to the 10th percentile of the wage distribution (90-10 ratio) increased considerably, consistent with the general trends in wage inequality in Figure 1. Our sample is representative of the general trends, discussed above; however, the decline in private sector unionization is more pronounced. Figure 4 shows the changes in the 10th, 50th, and 90th percentiles of the wage distribution, relative to their levels in 1973. The rise in the wage inequality, in large part, is due to the decline in wages below the median, and the rise in wages of the top 10 percent. This peculiar feature of the rise in wage inequality in the US is a crucial aspect of our results in section 4. The corresponding decline in the unionization rate is shown in Figure 5. The union coverage among male workers in the private sector declined from 32% in 1978 to 10% in 2007. The union membership rate is close to the union coverage rate, suggesting that the membership rates in Figure 1 are an accurate description of the trends in the union coverage rate.

To understand the role of selection into the union sector, Figure 6 displays the union participation rate by skill[8] for 1978 and 2007. The union participation rate has an inverse-U shape, consistent with the argument that the skilled workers prefer the non-union sector for better pay, while union jobs are "hard to get" for low-skill workers (Farber, 1983; Abowd and Farber, 1982; Card, 1996). The selection of the workers from the middle of the skill distribution affects how the overall unionization rate responds to the rise in the skill premium. For instance, if the withdrawal of the skilled workers from the union is matched by the inclusion of less skilled workers, the unionization rate would stay constant in response to technical change.

The unionization rates in Figure 6 decline for all skill groups, between 1978 and 2007. Furthermore, the hump-shaped profile of union participation by skill flattens over time. We think that a coherent theory of deunionization must be consistent with these facts. Next, we provide a review of the existing explanations for deunionization, in light of the empirical facts above.

1.1 Alternative Explanations

Industrial Composition of Employment. The composition of the economy in the US has shifted away from sectors where unions have been strong, traditionally, such as the manufacturing sector, to

[7] Farber (2005) also points out the differences between public and private sector in terms of legislation that allows for collective bargaining, and attributes the rise in public sector unionization in the 1970s to legalization of collective bargaining for public sector employees at the state-level.

[8] Skill is defined by the predicted wages in the non-union sector, based on the returns to education and experience.

sectors where unions are harder to organize, such as services.[9] Could this explain the observed decline in the unions? To evaluate this proposition, Table 7 shows union membership and coverage rates according to major industries, between 1973 and 2007. The unions appear to be strongest in transportation, and communication and utilities, followed by manufacturing, construction, and mining. In 1973, these industries had unionization rates of over 40%, which decreased drastically over the next three decades, to about 10-20% in 2007. The strong decline in unionization rates within each industrial sector indicates that deunionization in the US involves more than just composition effects.

We estimated the industrial composition effects, at the two-digit level,[10] keeping the industrial shares of employment constant at their 1978 level and updating only the sectoral union coverage rates. The changes in the union density within sectors amount to a 13.8% decline in unionization between 1978 and 1999, leaving a decline of only 3.4% for the between industry effects.[11] We conclude that the composition effects constitute a fifth of the decline in the unionization rate in the US. Our findings are in line with those of Farber (1990) and Farber and Krueger (1992), who estimated similar composition effects for the 1980s.

Globalization, Trade, and Competition. Farber and Western (2000) and Baldwin (2003) argue that the possibility of outsourcing less-skill-intensive goods to developing countries not only eliminated the surpluses from production bargained with unions but was also used as a threat by firms to persuade union workers to vote against unionization in their plants. The increased competition in the labor market, either directly through outsourcing, or indirectly through import-substitution of labor intensive products, could have been responsible for the decline of union activity; however, we find this argument to be in line with our model to the extent that it explains the decline in wages of low-skill workers. Our results do not depend on reasons for the widening productivity differences, and therefore, carry over to more general cases where skill prices are altered by, *inter alia*, increased competition in the labor market due to expansion of US trade or outsourcing. Nevertheless, the decline in the unionization rate was not confined to the tradable goods sector, and has been almost uniform across industries. Furthermore, the union wage premium in our sample has been stable for the last 30 years. Thus, any theory based on depletion of economic rents available to unions, must adhere to a union strategy that trades size for larger wage premiums.

Anti-Union Politics. One could also attribute deunionization to the anti-union efforts of the Reagan government in the US and the Thatcher government in the UK. (Howell, 1995; Farber and Western, 2001) Although we recognize that the political agenda can put an effective pressure on labor unions, the timing of events suggests that the economic forces underlying deunionization

[9] The cost of organizing a union could vary by industry, due to such factors as geographical dispersion of production units, capital intensity, the size distribution of establishments, or local legislation.

[10] We mapped the 3-digit census coding to the 2-digit classification, as defined by the NBER. Due to the change in the census industrial classification system in 2000, we confined the compositional analysis to the years 1978-1999. Since the rate of unionization declined by 4% after 1999, we think that our composition analysis would not be altered much by including the period after 1999.

[11] We also repeated the analysis using the union membership rate between 1973 and 1999, and found that the change in industrial composition accounts for a 4.3% decline, relative to a total decline of 20.2%.

were already underway before Reagan's presidency. Farber and Western (2001) provide empirical evidence that the fall in the annual number of union elections, in fact, preceded the appointment of the Reagan Labor Board in 1983. Therefore, the political factors, at best, may have amplified the existing deunionization trend.

Technical Change. Technical change, as a possible cause of deunionization, has not received any attention in the literature, with the exception of Acemoglu, Aghion, and Violante (2001). They provide a theoretical assessment of the incentives of skilled workers in response to skill-biased technical change, and argue that skilled workers would leave the union for better wages. Although this element of their paper is common with ours,[12] our approach differs considerably. First, their mechanism abstracts from the firms' role in the unionization process, thereby, producing unions that consist mostly of unskilled workers, and a competitive sector that contains only skilled workers. This is inconsistent with the hump-shaped selection pattern we document above. Furthermore, the decline in unions, in their study, is led entirely by skilled workers, contradicting the US experience, as shown in Figure 6. Most importantly, we empirically evaluate our model using US data, and make quantitative predictions; whereas, their focus is completely theoretical.

In the next section, we present our model, which explains a decline in the unionization rate consistent with the empirical evidence we document above.

2 A Model of Union Participation

For our analysis, we employ a Mortensen-Pissarides (MP) search model (Mortensen and Pissarides, 1994; Pissarides, 2000). The search and matching frictions in this framework lead to rents to be divided between the employer and the worker, and allows for collective bargaining as an alternative to bilateral bargaining between the firm and individual workers. The rents come from the possibility of being unemployed at any point in time, from the perspective of the worker and the contingency of staying vacant, from the firm's point of view.

The explicit matching process also allows us to introduce strategic incentives at the worker-firm level to model the selection into unions. Union membership is endogenous in our model. The union is an outcome of the interaction of firms' and workers' incentives, and it is the composition of the workers in the union that determines union wages. To the best of our knowledge, this is the first model of endogenous union selection in the literature[13].

[12] Although the prediction is similar, the means with which they generate this prediction is different. The unskilled workers in their model are in the union sector by construction, and some skilled workers are randomly assigned to the union sector due to information frictions, and choose to remain, thanks to switching costs; whereas, workers in our model become union members by choice. Therefore the skilled workers in their model have negative wage gains, while all workers in our model have a positive union wage premium, even though some seem to have negative gains due to selection bias.

[13] Another attempt to endogenize unions is by Ebell and Haefke (2006). However, we find that the absence of heterogeneity and selection, and lack of predictive power on aggregate union rate makes it unsuitable as a model of deunionization.

2.1 Environment

There is a continuum of workers with a unit measure and a large measure $N >> 1$ of firms. We assume that each worker is endowed with a time-invariant productivity $(s, x) \in \mathcal{S} \times \mathcal{X}$ with joint distribution $(s, x) \sim \mu(.,.)$. *Skill* s represents the productivity level observed by the firms, union, and the econometrician alike. *Ability* x of the worker is independent of s and is evaluated by the firms but not observed or priced by the union. The ability component is also not observed by the econometrician. We assume a representative sector and a representative union.

2.2 Preferences and Technology

We assume that workers are risk-neutral and that they maximize

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \delta^t c_t$$

subject to relevant constraints. Here, $0 < \delta < 1$ denotes the subjective discount factor of the worker and c_t is the consumption at time t. There are no savings; therefore, all earnings are consumed in each period.

At any point in time, each worker i is either employed or unemployed. When employed, the worker earns wage w_{it}, which depends on the worker's productivity and union status. When unemployed, the worker receives the unemployment benefit b.

All firms are *ex-ante* identical and they maximize the discounted profits,

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \delta^t \pi_t$$

subject to the relevant constraints. There is free entry to the market. Should a firm decide to post a vacancy, it has to incur a fixed cost of κ. When a vacant position is filled, the firm-worker pair produces

$$f(s_i, x_i) = \exp(\psi_t(s_i + x_i)),$$

where $s_i + x_i$ is the productive capacity of the worker i, and ψ_t captures the marginal productivity of skill at time t. The profit from a filled position is

$$\pi_t = f(s_i, x_i) - w_{it}$$

2.3 Matching and Recursive Formulation

Every period, vacant positions, and unemployed workers are randomly matched. Given the measure of vacant positions, v_t, and the unemployed, u_t, the measure of matched pairs is represented by the following constant-returns-to-scale (CRS) matching technology:

$$m(v_t, u_t) = \eta u_t^{1-\alpha} v_t^{\alpha}$$

We define $\theta_t \equiv v_t/u_t$ as the labor market tightness. The CRS assumption allows us to express the relevant matching variables in terms of θ. In particular,

$$p(\theta_t) \equiv m(v_t, u_t)/u_t = m(\theta_t, 1) \qquad \text{and} \qquad q(\theta_t) \equiv m(v_t, u_t)/u_t = m(1, \theta_t^{-1}) = p(\theta_t)/\theta_t.$$

Under a law of large numbers, $p(\theta_t)$ denotes the probability of an unemployed worker being matched with a firm and $q(\theta_t)$ is the probability of a vacant position being matched with a worker.[14]

2.3.1 Workers' Value Functions

We are now ready to define the value functions of the agents in the economy. We drop the time subscript, since we are interested in the steady-state of the economy. For what is to follow, let $M^w(s, x)$ be the value of a match for the worker and $M^f(s, x)$ the value of a match for the firm. These values are determined by a strategic interaction between the firm and the worker when they are matched.

Value of a non-union job to a worker is given by:

$$W^n(s, x) = w^n(s, x) + \delta\big[\lambda U(s, x) + (1 - \lambda)M^w(s, x)\big] \tag{1}$$

Each worker receives the competitive wage $w^n(s, x)$ and the match breaks next period with an exogenous probability of λ, in which case the worker gets $U(s, x)$. With probability $(1 - \lambda)$, the match is retained.

Similarly, the value of a union job is,

$$W^u(s, x) = w^u(s) + \delta\big[\lambda U(s, x) + (1 - \lambda)M^w(s, x)\big] \tag{2}$$

Note that the union wage depends only on the skill level $s \in \mathcal{S}$. Since union wages are determined collectively at a larger scale, we think it plausible that individual firms are better than unions in evaluating their employees individually. Given our assumption of independence between s and x, we allow any productivity trait that can be projected on easily observable credentials, such as education, experience or seniority, to be priced into the union wage.

The value of unemployment for a worker with skills (s, x) is represented by:

$$U(s, x) = b + \delta\big[p(\theta)M^w(s, x) + (1 - p(\theta))U(s, x)\big],$$

where b is the flow benefit of being unemployed, which includes the value of leisure as well as

[14]We assume that the firm cannot direct its search toward a specific skill group. Market segmentation by skill does not change our results, even when we calibrate our parameters to vary by skill groups.

unemployment benefits received by the worker. With probability $p(\theta)$, the worker is matched with a firm and gets $M^w(s,x)$. The worker remains unemployed if either not matched with a firm (with probability $1-p(\theta)$), or if matched, but the match breaks without an agreement.

2.3.2 Firms' Value Functions

The firm's benefit from a filled position depends on the union status of the worker. The values are,

$$J^n(s,x) = f(s,x) - w^n(s,x) + \delta\big[\lambda V + (1-\lambda)M^f(s,x)\big] \tag{3}$$

$$J^u(s,x) = f(s,x) - w^u(s) + \delta\big[\lambda V + (1-\lambda)M^f(s,x)\big] \tag{4}$$

for a non-union worker and a union worker, respectively.

From the firm's perspective, the only difference between the two positions is the wage rate paid to the worker. The value of a vacancy is given by:

$$V = -\kappa + \delta\left(q(\theta)\int_{\mathcal{S}\times\mathcal{X}} M^f(s,x)d\mu + (1-q(\theta))V\right) \tag{5}$$

A vacancy is matched with a worker at the rate $q(\theta)$ every period. With probability $(1-q(\theta))$, the position stays vacant.

2.3.3 Value of a Match and Two-Sided Selection into Unions

We can now define the match values $M^f(s,x)$ and $M^w(s,x)$. When a firm is matched to a worker, the firm observes the worker's productivity $(s,x)\in\mathcal{S}\times\mathcal{X}$ and both parties observe union wages $w^u(s)$. Then, they play a non-cooperative game, in which the worker moves first, decides whether or not to join the union, and the firm decides whether or not to hire the worker or stay vacant, conditional on the worker's decision. Let $\mathcal{G}(s,x)$ represent the extensive-form non-cooperative game between a firm and a worker of skill (s,x).[15] Figure 8 represents the game tree, decision nodes, and payoffs. The game specified above has a sub-game perfect equilibrium in pure strategies,

[15]Formally, we define the extensive-form game with perfect information for skill pair (s,x) as,

$$\mathcal{G}(s,x) = \langle \mathcal{T}, \{d_i\}_{i\in\{w,f\}}, a, \{\pi_i^{s,x}\}_{i\in\{w,f\}}\rangle$$

where

- Game tree $\mathcal{T}$ is a tuple $\langle N, D, p\rangle$ where N is the set of all nodes $N = \{0, U, N, (U,H), (U,NH), (N,H), (N,NH)\}$, $D \subset N$ is the set of decision nodes $D = \{0,U,N\}$ and $p : N\setminus\{0\}\to N$ is the node function that maps each node to the first node that precedes it.
- d_i for each $i\in\{w,f\}$ is a set of decision nodes where agent i gets to play. ($d_w=\{0\}$,$d_f=\{U,N\}$)
- a is a function that maps each decision node to a set of actions. ($a(U)=a(N)=${Hire , Not Hire}, $a(0)=${Union, Non-union}))
- $\pi_i^{s,x}$ for each $i\in\{w,f\}$ and each $(s,x)\in\mathcal{S}\times\mathcal{X}$ is a function that maps each terminal node to the payoff received by player i, i.e. $\pi_f^{s,x}(U,H)=J^u(s,x)$, $\pi_f^{s,x}(U,NH)=V$,$\pi_f^{s,x}(N,H)=J^n(s,x)$,$\pi_f^{s,x}(N,NH)=V$ and $\pi_w^{s,x}(U,H)=W^u(s,x)$, $\pi_w^{s,x}(U,NH)=U(s,x)$, $\pi_w^{s,x}(N,H)=W^n(s,x)$, $\pi_w^{s,x}(N,NH)=U(s,x)$.

since we can always find a pair of strategies that survive backward induction.

Let $M^f(s,x)$ and $M^w(s,x)$ be the sub-game perfect equilibrium payoffs[16] of the above game for the firm and the worker, respectively. Given that the union wages have a compressed structure, the equilibrium of this game would be such that union workers would be selected from the middle of the productivity distribution. Clearly, a worker with high productivity would opt for the non-union sector, where the worker is better rewarded for skills. On the other extreme, the firm would be very selective when it matches with a low productivity worker, due to relatively high union wages. In particular, it would not hire any worker whose productivity does not make-up for the high union wage. Such a worker would strategically choose the non-union job anticipating the firm's hiring decision. In this sense, this game formalizes the observation by Farber (1983): Both firm and worker incentives play a role in determining the union status of a worker.

The timing of the game is crucial for the equilibrium selection. A similar game, where the firm plays first and the worker responds, leads to an inefficient equilibrium. More specifically, a worker with productivity lower than the union wage would be unemployed in this alternative setting, since such a worker would always choose the union job conditional on being hired. An alternative arrangement, where the worker is employed, albeit at a lower (non-union) wage, is a clear improvement for both parties. Leaving aside this normative criterion, this timing also fails in a descriptive sense. In particular, it would not capture the inverse U-shaped union membership profile, as described earlier.

The strategic interaction we adopt seems non-standard, since the firm may refuse to hire a worker if the worker chooses to join the union. We think that this assumption is reasonable for two reasons: First, the firm never exercises this option at the equilibrium, since the firm never gets to play at this decision node with a low-productive worker. Second, at the steady-state equilibrium, a worker optimally looks for either a union job[17], or a non-union job. The equilibrium outcome mimics the dynamics of a directed search model, though such elements are not made explicit.

2.4 Wage Determination

All wages are determined by bargaining over the surplus. If a non-union worker is hired, the wage is determined by Nash bargaining, with worker bargaining power $\beta \in (0,1)$:

$$w^n(s,x) \in \arg\max \beta \log[W^n(s,x) - U(s,x)] + (1-\beta)\log[J^n(s,x) - V]$$

For the specification above, $w^n(s,x)$ satisfies,

$$\frac{W^n(s,x) - U(s,x)}{\beta} = \frac{J^n(s,x) - V}{1-\beta}.$$

[16] Given the equilibrium wage structure and an absolutely continuous distribution for (s,x), these payoffs and the equilibrium strategies are unique almost everywhere on $\mathcal{S} \times \mathcal{X}$ with respect to measure μ.

[17] We can make an alternative interpretation: If a worker has low productivity, the worker would not apply for a union job since the firm would hire the best workers in the queue of applicants for the position. A high enough waiting cost would induce the worker to apply for a non-union job in the first place.

It can be shown, by straightforward algebra that,

$$w^n(s,x) = \beta f(s,x) - \beta(1-\delta)V + (1-\beta)(1-\delta)U(s,x). \tag{6}$$

We let $b \equiv \rho w^n(s,x)$, for agents who bargain competitively, where $\rho \in (0,1)$. Importantly, the first-order condition on the Nash bargaining does *not* involve an indirect effect through the unemployment benefit. The bargaining parties take it as given,[18] as opposed to being a constant fraction of market wage, while evaluating the outside option of the worker. This allows us to simplify the wage functions and obtain a functional form that can be used for empirical analysis.

Unionized workers, and firms hiring them, bargain collectively. We model this as a two-stage process. We can think of the first stage as the firm leader and union leader bargaining over the total surplus generated by union jobs. We assume that the union has a higher bargaining power in wage negotiations than an individual worker. The second stage involves the distribution of accrued union rents among the union members.

Let $\mathcal{U} \subset \mathcal{S} \times \mathcal{X}$ denote the set of values for $(s,x) \in \mathcal{S} \times \mathcal{X}$, for which the worker would choose[19] to be a union member. Also, let $\gamma > \beta$ represent the union bargaining power. The union share of the total surplus, $\int_{\mathcal{U}} w^u d\mu$, solves the following problem:

$$\max \gamma \log \left(\int_{\mathcal{U}} [W^u(s,x) - U(s,x)] d\mu \right) + (1-\gamma) \log \left(\int_{\mathcal{U}} [J^u(s,x) - V] d\mu \right) \tag{7}$$

Similarly to the workers bargaining individually, we let $b \equiv \rho w^u(s)$ for the union workers.

Due to the linear nature of the union bargaining problem, we need to impose an additional parametric structure on the union wages. The difficulty arises from the fact that the collective bargaining problem gives us one restriction; however, we need a continuum of such restrictions to pin down all wages in the union[20]. For the second stage, we choose a particular way in which the unions distribute the aggregate surplus under the premise that unions compress wages: the return to skills in the union is lower than that in the non-union sector. Although we do not explicitly model the wage determination within the union, we account for the union's impact on rent distribution over time in our empirical analysis. In this sense, this abstraction does not alter our quantitative results.

[18] This is not an unreasonable assumption when we consider that a worker-firm match has measure zero.

[19] In equilibrium, this set is induced by the equilibrium strategies for $\mathcal{G}(s,x)$.

[20] Suppose $\bar{w}^u$ is the wage level when we make the extreme assumption that every union member is paid the same wage. We can show that any mean-preserving spread $\tilde{w}^u(s,x)$ around $\bar{w}^u$, i.e., $\int_{\mathcal{U}} \tilde{w}^u(s,x) d\mu = \bar{w}^u$, also satisfies the bargaining first-order-condition.

2.5 Equilibrium

The law of motion[21] for unemployed workers implies a steady-state unemployment of

$$u = \frac{\lambda}{p(\theta) + \lambda} \tag{8}$$

At the steady-state, the number of vacancies equals

$$v = \frac{\lambda \bar{N}}{q(\theta) + \lambda} \tag{9}$$

where $\bar{N}$ is the measure of firms operating at the steady-state.

The equilibrium in this economy is defined as follows:

Definition 1 *A* ***steady-state equilibrium with unions*** *consists of the wages $w^u : \mathcal{S} \to \mathbb{R}_+$, $w : \mathcal{S} \times \mathcal{X} \to \mathbb{R}_+$, extensive-form games $\mathcal{G}(s,x)$, value functions $M^f, M^w, J^u, J, W^u, W^n, U : \mathcal{S} \times \mathcal{X} \to \mathbb{R}$, $V \in \mathbb{R}$, and labor market tightness $\theta \in \mathbb{R}_+$, such that*
Given w^u and w^n,

1. *$M^f(s,x)$ and $M^w(s,x)$ constitute the sub-game perfect equilibrium payoffs of the game $\mathcal{G}(s,x)$ for each $(s,x) \in \mathcal{S} \times \mathcal{X}$.*

Given equilibrium strategies for $\mathcal{G}$,

2. *Union Wages: w^u is determined by collective bargaining (7).*

3. *Non-union wages: $w^n(s,x)$ is determined by individual bargaining (6) for each $(s,x) \in \mathcal{S} \times \mathcal{X}$.*

4. *Optimal vacancy posting: $V \geq 0$.*

5. *Free entry: $V \leq 0$.*

6. *Labor market tightness: $\theta = v/u$ where u satisfies (8) and v satisfies (9).*

Consistency

7. *Wages w^u, w^n and the equilibrium strategies for $\mathcal{G}$ are consistent with each other.*

2.6 Analysis and Empirical Implications

We start our analysis of the equilibrium by characterizing the unionized workers. Essentially, four important conditions apply to the productivity of a worker. The first two are employability margins, dictated by the firm. For a worker to be hired, the firm requires that the value of a position filled by a worker must be greater than the value of keeping the position vacant for another period. Since the value to the firm of a filled position depends on the negotiated wage, this requires that

[21] The evolution of the measure of unemployed agents is represented by $u_{t+1} - u_t = \lambda(1 - u_t) - p(\theta_t) u_t$.

$J^n(s,x) \geq V$ must hold for all non-unionized workers, and $J^u(s,x) \geq V$ for all union workers. The first condition determines the hiring margin to be a function of s and x. The latter condition is more demanding at the equilibrium, since the union workers enjoy higher wages, in general. As we will demonstrate shortly, this condition will set a lower bound for the productivity of union workers.

The worker's decision similarly introduces two margins. First, for the worker to accept a position, the value of having a job should be higher than staying unemployed for another period. This requires that $W^n(s,x) \geq U(s,x)$. The second condition is induced by the worker's decision to join the union, and requires that $W^u(s,x) \geq W^n(s,x)$.

The conditions above define two critical selection thresholds. To derive these thresholds, note that using the equilibrium condition $V = 0$ along with equations (3), (1) and (6), the value of unemployment for a non-union worker can be expressed as a fraction of productivity:[22]

$$U(s,x) = \frac{\beta\rho(1-\delta(1-\lambda)) + \beta\delta p(\theta)}{(1-\delta)\big[(1-\rho(1-\beta))(1-\delta(1-\lambda)) + \beta\delta p(\theta)\big]} f(s,x) \tag{10}$$

Substituting the expression above in (6), wage function in the non-union sector is:

$$\begin{aligned} w^n(s,x) &= \frac{\beta(1-\delta(1-\lambda)) + \beta\delta p(\theta)}{(1-\rho(1-\beta))(1-\delta(1-\lambda)) + \beta\delta p(\theta)} f(s,x) \qquad (11) \\ &= C(\beta)\exp(\psi(s+x)) \leq f(s,x), \qquad (12) \end{aligned}$$

where $0 < C(\beta) < 1$ is the fraction of production retained by the worker in wages and is increasing in β. This implies that, for any filled position, the firm claims the remainder $1 - C(\beta)$ fraction of the production. Using equation (5), the value to a filled non-union position is simply:

$$J^n(s,x) = \frac{f(s,x) - w^n(s,x)}{(1-\delta(1-\lambda))} = (1 - C(\beta))f(s,x).$$

However, this expression is non-negative for all values of s and x, implying that, at the equilibrium, a filled non-union position is always preferable to keeping the position vacant for another period, from the firm's perspective.[23] This is a direct consequence of the zero-profit condition for posted vacancies at the equilibrium and of our assumption that the flow value of unemployment is proportional to the worker's productivity. It is straightforward to show, in a similar manner, that the optimality of a non-union match is mutual, i.e., $W^n(s,x) \geq U(s,x)$ holds for all values of s and x; therefore, it is mutually beneficial not to break the match for any worker-firm pair.

We now turn to the selection conditions on the unionized workers. We impose the following union wage scheme as a means of distributing the surplus:

$$w^u(s) = \phi_0 \exp(\phi_1 s) \tag{13}$$

[22]At the equilibrium, the unionization decision is time-consistent for a given worker. This is implicit in the calculation of the unemployment value for the union worker.

[23]Note that the equality is only possible at the limit when $s + x \to -\infty$.

where $\phi_1 \leq \psi$ is the exogenously fixed return to skill in the union. Given ϕ_1, the union bargaining problem can be recast as an optimal choice of a single parameter ϕ_0:

$$\phi_0 \in \arg\max \gamma \log \left(\int_{\mathcal{U}} [W^u(s,x) - U(s,x)] d\mu \right) + (1-\gamma) \log \left(\int_{\mathcal{U}} [J^u(s,x) - V] d\mu \right)$$

A log-linear distributive rule appears to accurately describe the union wages in the US data. In addition, since the non-union wage function (11) has a similar form, wage compression can be measured simply by ϕ_1/ψ, which makes the qualitative and empirical analysis convenient.

Conditional on being hired, a worker prefers to join the union if $W^u(s,x) \geq W^n(s,x)$, or equivalently, when the union wage is higher than the competitive alternative $w^u(s) \geq w^n(s,x)$. Given a union wage policy (ϕ_0, ϕ_1), the union wage is higher if and only if $(\psi - \phi_1)s + \psi x \leq \ln \phi_0 - \ln C(\beta)$. When union wage policy has a compressed structure, $\phi_1 < \psi$, this condition is satisfied only for the lower values of s and/or x. Therefore, the first implication of our model is that workers with higher productivity choose not to be a part of the union, since their skill is better rewarded in the competitive market. Second, conditional on having a high value of skill s, the workers who prefer to join the union have lower unobserved abilities, x.

On the other hand, the firm agrees to offer a union position if and only if $J^u(s,x) \geq V$, or equivalently, if the worker's productivity is higher than the union wage: $f(s,x) \geq w^u(s)$. Any worker who does not meet this requirement is not desirable by the firm, should the worker choose to join the union. Since the worker is better off employed than not, the worker would choose to bargain individually, in our game-theoretic framework. Given equation (13), the firm has a positive surplus from hiring a union worker if $(\psi - \phi_1)s + \psi x \geq \ln \phi_0$. Once again, with $\phi_1 < \psi$, this condition defines a minimum requirement for skill s (or x), conditional on x (or s), to be a union member. The following proposition summarizes the skill composition of union workers.

Proposition 1 *Given the union wage policy (ϕ_0, ϕ_1), a worker is a union member if and only if* $\ln \phi_0 \leq (\psi - \phi_1)s + \psi x \leq \ln \phi_0 - \ln C(\beta)$.

For a given level of ability, x_o, the selection conditions above define the following two thresholds,

$$\underline{s}(x_o) = \frac{\ln \phi_0 - \psi x_o}{\psi - \phi_1} \tag{14}$$

$$\bar{s}(x_o) = \frac{\ln \phi_0 - \ln C(\beta) - \psi x_o}{\psi - \phi_1} \geq \underline{s}(x_o), \tag{15}$$

which constitute the skill composition of the union workers at the equilibrium.

The above characterization of unionized workers is partial, since we have taken the union wage as given. At the equilibrium, union wages implied by the skill composition of union workers are consistent with the union wage function used to determine the skill composition of unions in Proposition 1. To characterize the union wage function, recall that the union bargains for all workers that join the union and distributes the surplus according to the wage policy, defined by equation (13). The following equality shows the budget constraint of the union.

$$\int_{\mathcal{U}} C(\gamma)\exp(\psi(s+x))dF(s,x) = \int_{\mathcal{U}} \phi_0 \exp(\phi_1 s)dF(s,x). \tag{16}$$

The left side of this equation is the part of the collective marginal product retained by the union at the solution to (7). With linear utility, this is equivalent to the aggregate of all individual bargaining outcomes for union workers, assuming that they each have a bargaining weight of γ. Since $C(\gamma) > C(\beta)$ when $\gamma > \beta$, union workers keep a larger portion of their product than if they were to bargain individually and pool their wages. The second part of the equation is the total union wage bill. Given the skill price in the union, ϕ_1, ϕ_0 can be determined by the equation above. Equilibrium is reached when union participation can be rationalized by the union wages, as calculated above.[24]

2.6.1 Union Wage Premium

The union wage premium for a worker with productivity (s,x) is $w^n(s,x) - w^u(s) = \ln\phi_0 - (\psi - \phi_1)s - \ln C(\beta) - \psi x$. For a given value of ability x_o, the union premium attains its maximum at the lower skill threshold $\underline{s}(x_o)$ and equals $-\ln C(\beta)$. The lowest observed union premium is zero for the worker at the upper margin. This is consistent with the findings in the literature that show that the union wage premium is decreasing in skill. The observed union premium, conditional on s; however, transcends both of these bounds. Due to selection, the average value of x, conditional on being a union member, will be lower than the unconditional average for low enough values of observed skill s. This will generate an artificially large wage gain for low values of observed skill and artificially low and potentially negative wage gains for high-skill workers.

2.6.2 Rising Skill Premium

To see the implications of a rise in the skill premium for the rate of unionization, we consider two different scenarios. In the first scenario, the rise in the productivity of skilled labor is accompanied by a fall in the productivity of the less skilled. This is shown in Figure 9a as a rotation of $\ln f(s,x_o)$ for a fixed value of x_o. Since wages in the non-union sector are proportional to productivity, the wage curve for this sector also rotates. Suppose that the skill price within the union, ϕ_1, is fixed. Then, unionized workers with the highest productivity will choose to opt out of the union, since the cost of wage transfers within the union increases relative to gains from collective bargaining. This results in an inward shift of $\bar{s}(x_o)$. Similarly, the union workers with the lowest productivity cannot stay in the union since their productivity now less than compensates for the high union

[24] Essentially, this is the computational method for finding the equilibrium. We find the fixed point of the mapping $\zeta : \mathbb{R}_{++} \to \mathbb{R}_{++}$

$$\zeta(\tilde{\phi}) \equiv \frac{\int_{\mathcal{U}(\tilde{\phi})} C(\gamma)\exp(\psi(s+x))d\mu}{\int_{\mathcal{U}(\tilde{\phi})} \exp(\phi_1 s)d\mu}$$

induced by (16), by updating the value of $\tilde{\phi}$.

wage. Consequently, $\underline{s}(x_o)$ increases, further reducing the size of the union. This effect is *ceteris paribus*; in particular, we keep the union wage function fixed. Nevertheless, in this scenario, the average productivity of unionized workers changes only slightly, since the thresholds are moving inward.

In the second scenario, we consider an across-the-board rise in the productivity, where the productivity of the skilled workers increases disproportionately, raising the skill premium (Figure 9b). Once again, productivity and the competitive wage function shift up, which results in a downward shift of $\bar{s}(x_o)$. Since productivity is higher for all workers; however, low-skill workers are now more profitable to the firm relative to their union wage. Firms hire additional low-skill union workers, which decreases the threshold productivity to join the union to $\bar{s}'(x_o)$. The relative skill composition of unionized workers declines, and the rate of unionization may increase or decrease, depending on the distribution of skill.

A relaxation of wage compression within the union has effects similar to those of the decreasing skill premium. An increase in the return to skill ϕ_1, for union workers without an accompanying increase in productivity, makes the union more attractive at the upper margin. The effect on low-skill workers' union participation depends crucially on whether or not this increase comes about with rising wages for all, or at the expense of low-skill workers. Two scenarios are again possible: one, where the unionization rate is not much affected, and the other, where it substantially increases.

The impact of combining these two effects calls for a quantitative analysis. Both effects, in isolation, can increase or decrease the unionization rate, depending on the initial composition of the union workers and the overall distribution of productivity. In the next section, we present the empirical analysis and quantify these effects.

3 Empirical Analysis

We evaluate our model economy to reproduce the labor market conditions in 1978 and then use our model to predict the unionization rate in 2007, based on the estimated changes in the skill prices and the productivity distribution in 2007. We estimate the skill prices in the union and the non-union sectors, as well as the distribution of skills in the labor force, and calibrate the remaining parameters related to search and matching frictions. In the next subsection, we outline the estimation of our parameters, based on an auxiliary model that assumes that the union status is random. Then, we describe how we uncover the true parameters using our modeling restrictions.

3.1 Estimation of an Auxiliary Model with Random Union Status

We identify the parameters of our model related to the productivity distribution and skill prices by first estimating an auxiliary model under the assumption that union membership is random. Some of these estimates are biased when sorting of workers into union jobs is non-random. We then use the selection, implied by our model, to gauge the bias and identify the underlying true parameters. If one were to estimate these parameters using model-simulated data, ignoring selection, the estimates would coincide with those obtained in the actual data. Essentially, this is an indirect inference approach.

3.1.1 Skill Prices

Let $j \in \{n, u\}$ denote the non-unionized and the unionized categories, respectively. Marginal productivity of worker i at time t is given by,

$$\log y_{it} = X_{1it}\Gamma_{jt} + X_{2it}\Psi_{jt} + \varepsilon_{ijt}, \tag{17}$$

where, X_1 is a vector of skill characteristics and X_2 is a vector of controls for non-skill-related characteristics. Note that skill prices vary across the two sectors. Under random selection into union jobs,

$$E[\varepsilon_{iut}|X_{1it}, X_{2it}] - E[\varepsilon_{int}|X_{1it}, X_{2it}] = 0,$$

i.e., conditional on observable skills, the union workers are perfectly substitutable with non-union workers. This would allow for a consistent estimation of skill prices Γ_{jt} in the two sectors using standard methods. When selection into union jobs is non-random, the expression above is positive for low values of observed skill, since, among these workers, only those with high unobservable skills join the union. Similarly, it is negative for high values of observed skill. When this sorting behavior is ignored, the skill premium is overestimated in the non-union sector and underestimated in the union sector.

3.1.2 Skill Distribution

We begin by defining an observable skill composite for workers, corresponding to s in our model, based on the observed prices in the non-union sector. A skill composite for a worker, priced at time t is obtained by estimating the equation (17) above for non-unionized workers and calculating the prediction,

$$\hat{s}_{it} = X_{1it}\hat{\Gamma}_{nt}$$

for all workers, including those in the union sector. Note that the time t price of the composite skill calculated at time t is one by definition. We normalize ψ accordingly to be one for 1978. The standard deviation of $\hat{s}_{it}$ gives an estimate of σ_s when union status is random. When selection is present, it overestimates the dispersion in skills.

3.1.3 Wage Compression in the Union Sector

To estimate the degree of wage compression, we take the predicted composite skill index for unionized workers and project their actual wage on their predicted skills and other control variables as follows.

$$\log w_{itu} = \phi_t \hat{s}_{it} + X_{2it}\Psi_{ut} + \epsilon_{itu} \tag{18}$$

The coefficient $\phi_t < 1$ measures the degree of compression in the return to skill in the unionized sector. Under non-random selection, this procedure overestimates the wage compression by the union (underestimates ϕ), since the skill prices in the non-union sector are overestimated in equation (17).

3.1.4 *Estimation of the Change in the Skill Distribution*

The change in the skill distribution between years t_0 and t can be captured by calculating the skill composite of person i, based on t_0 prices as:

$$\tilde{s}_{it,t_0} = X_{1it}\hat{\Gamma}_{nt_0},$$

and comparing the the cross-sectional distribution of $\tilde{s}_{it,t_0}$ relative to $\hat{s}_{it_0}$.

3.1.5 *Estimation of the Change in the Skill Premium*

The change in the price of skill between t_0 and t is identified by the difference between the distributions of $\tilde{s}_{it_0,t}$ and $\hat{s}_{it}$. It can be measured by regressing actual wages of non-union workers at time t on their predicted skills, based on t_0 prices:

$$\log w_{itn} = \psi_t \log \tilde{s}_{it} + X_{2itn}\Psi_{tn} + \epsilon_{itn} \tag{19}$$

The estimate of ψ_t gives the price of skill in year t, relative to ψ_{t_0}. Since the estimated skill prices in both years are biased upwards, the direction of the bias in their ratio depends on the size of the bias in each year. In our case, we expect a smaller bias in 2007 due to lower unionization rate.

3.2 Introducing Selection Using the Model of Union Participation

Three of our parameters cannot be estimated directly using the standard methods in the previous section: the dispersion of unobserved ability, σ_x, and the bargaining powers of the union and of the individual workers. To identify these parameters, and to correct the auxiliary estimates in the previous section for selection, we rely on our model's predictions for union participation by skill.

3.2.1 *Ability Distribution*

Given the estimate of skill dispersion, σ_s, dispersion of unobserved ability, σ_x determines the curvature of the unionization profile. For low values of σ_x, union participation by skill approaches a step function, with 100% unionization for mid-skill groups, and 0 elsewhere. For large values of σ_x, the union profile becomes flat. Therefore, the curvature identifies a unique value of σ_x. In section 5, we check the robustness of our results to this identification strategy, by using the component of residual wage variance that can be attributed to worker-specific effects, to identify σ_x instead.

Similarly, we calibrate the σ_x in 2007 to the rise in the wage variance that can be attributed to worker heterogeneity. We allude to the findings in the literature to define bounds on how much the worker-specific variation in wages has changed and report our results for both cases.

3.2.2 Individual and Collective Bargaining Powers in Wage Negotiations

The literature provides little guidance on the bargaining powers of unions and individual workers, relative to that of the employers. In our benchmark estimation, we target the unionization rate by skill to identify the bargaining parameters. The aggregate unionization rate and the shape of the unionization profile by skill determines the values of bargaining parameters as follows: Given β, low values of γ produce relatively small wage gains to bargaining collectively. Consequently, skilled workers prefer the competitive sector due to the transfers within the union sector. On the other hand, since the union wage is close to the competitive wage, unskilled union workers are affordable for firms, creating a union that consists mostly of workers with low skills. The unionization rate in this case is mostly decreasing in skill. In addition, the overall union density is low, since the measure of workers with low skills is small. For high values of γ, the skilled workers find it optimal to join the union, and the firms find the unskilled union workers too costly. As a result, the union contains workers with high skills. In this case, the rate of unionization is increasing in skill. Therefore a hump-shaped unionization by skill, as observed in the data, is obtained for a non-extreme value of γ. Once the value of γ is determined, relative to β, the aggregate rate of unionization pins down the value of β. Evidently, high values of individual bargaining power, β, limit the potential gains to bargaining collectively, and yield low overall unionization rates.

For our benchmark economy, we simultaneously choose the collective and the individual bargaining powers, the skill price in the union sector, along with the standard deviations of the distributions of skill and ability, to match the observed rate of unionization by deciles of the predicted skill distribution and the auxiliary estimates of the standard deviation of the skill distribution and the estimated wage compression within the union. Because the moments outnumber the parameters, we weight each data moment by its standard deviation. The next section presents the results.

4 Results

4.1 Calibration of the Search Environment

We set the time period for our analysis to a month and use a discount rate of $\delta = 0.99^{1/3}$. We choose the vacancy cost, κ, to normalize the equilibrium labor market tightness θ to unity. When $\theta = 1$, the probability of being matched with an employer is η. We set $\eta = 0.32$, to target the average duration of unemployment in 1978, which is 3.09 months. Given the job finding rate, we set the separation rate to $\lambda = 0.02$, to match the average unemployment rate of 6.03%, for years

1978-1980 (see equation (8)).[25] We set $\alpha = 0.5$.[26] We set $\rho = 0.4$ as reported by Shimer (2005) to be the fraction of earnings enjoyed when unemployed. We assume that s and x have independent normal distributions with mean zero. This implies that when the skill premium, ψ, rises, non-union wages decrease for half of the workers, comparable to the 57% observed in the data. The first panel in Table 1 summarizes the calibrated parameters.

4.2 The Productivity Distribution, Wage Compression, and Unionization by Skill

The data we use for our estimations is a combination of the May Supplements to the Current Population Surveys for years 1978-1980, and the Merged Outgoing Rotation Group files for the years 2004-2007. We use education and potential experience as measures of skill. We categorize education into five levels: less than high school, high school, some college (less than 16 but more than 12 years of education), college, and more than college. We then estimate (17) for the non-union workers including a full set of indicators for each year of potential experience (age - years of education - 6), each education category and a full set of interactions in X_1. We also include indicators for race, marital status, and survey year in X_2.[27]

Based on the estimated coefficients, we construct the skill composite for all workers. The standard deviation of the skill index is 0.28. A regression of actual wages on the skill composite for union workers yields 46% wage compression ($\hat{\phi}_{1978}/\hat{\psi}_{1978} = 0.54$) in 1978. We use these auxiliary estimates, along with the union participation by deciles of predicted skill composite in 1978, as shown in Figure 6, to estimate the underlying parameters that are corrected for selection. Since the model is overidentified (5 parameters and 13 moments), we weight each moment by its standard deviation.

The second panel in Table 1 summarizes the parameters and the corresponding moments. The corrected estimate of skill dispersion is 0.26, slightly lower than the auxiliary estimate, as expected. The actual wage compression by the union in 1978 is 39%, implying a 7% selection bias when non-random selection into the union is ignored.

The remaining three parameters, namely the standard deviation of ability, σ_x, and the bargaining parameters, α and β are identified by the union participation profile. Figure 10 compares the skill distribution of unionization in the model with the data. The model closely replicates the observed hump-shaped pattern in the data. The ability of the model to match the entire profile with a few parameters confirms our assumptions regarding the selection process, and the distributions of skill and ability.

The implied bargaining powers for individual workers and the union are 0.24 and 0.39. Estimates of workers' bargaining power are hard to come by in the literature. Cahuc, Postel-Vinay, and

[25] We combine years 1978-1980 (and years 2005-2007) to obtain larger samples, and more robust estimates of time-variant moments, such as the unemployment rate. For the rest of the paper, all estimations for 1978 actually use years 1978-1980, and all estimates for 2007 are based on years 2005-2007.

[26] The standard Hosios efficiency condition does not hold, but, a *version* of the Hosios condition holds in this economy that maps the bargaining powers β and γ to the elasticity parameter α. We confirmed that a change in the parameter α has almost no effect on our results.

[27] Since we combine years 1978, 1979, and 1980, we include two survey year indicators for 1979 and 1980.

Robin (2006) provide estimates for France that vary between 0 and 0.3. This is not, however, the share of rent accrued by workers in their model. Due to on-the-job search, workers get 10 to 60% of total surplus, depending on their skill. The average bargaining power in our benchmark model is 0.29 in 1978 when we combine the unionized and non-unionized workers. This is within the range of values provided by Cahuc, Postel-Vinay, and Robin (2006).

The estimate of the standard deviation of unobserved ability, σ_x, in 1978 is 0.11. Combined with the estimate of, σ_s, the variance of log-wages implied by the model is 0.08. This indicates that the heterogeneity in individual skill characteristics constitutes about a quarter of the total variance of log-wages in the data.

4.2.1 Union Wage Premium

To evaluate the cross-sectional predictions of our model, we examine the union wage premium by skill. The first column in Table 2 shows the log wage differences between union and non-union workers for each quintile of the predicted skill distribution in the data. Workers with lower skills gain 0.40 additional log points, substantially more than high-skill workers who seem to gain -0.04. The average union premium is 0.20 log points. The model's predictions with the benchmark parameters is shown in the second column. The model successfully captures the declining union premium pattern by skill, but undershoots the average premium by 0.15. In section 5, we examine the sensitivity of our results to alternative parameter values that improve the fit of the model to the observed wage gaps.

To illustrate the implications of selection for the union premiums, the third column shows the hypothetical return to being a union member for each worker. Since the actual return varies by worker, we report the average return for different groups. The third column shows the average wage gain for all workers in each quintile. Since the union workers are selected positively at the lower tail of the productivity distribution, the average return to being a union member is lower than the raw wage difference. Similarly, the actual union wage premium is higher for skilled workers, relative to the second column, since they are selected negatively to the union sector. The average return is 0.06, almost unchanged compared to the second column due to the two-sided nature of selection. The positive bias in the wage premium of the unskilled is offset by the negative bias in the estimated union premium for the skilled. These are consistent with Card (1996), who finds a strong selection bias in union premiums, yet, a small bias on average. Hirsch and Schumacher (1998) also finds evidence, using NLSY data, that union workers with high measured skills have relatively low unmeasured skills.

The fourth column displays the return to being a union member for the union workers only, also referred to as the treatment effect on the treated. The return to union membership varies little across quintiles, because higher observed skills in upper quintiles are compensated by lower unobserved ability, which affect the return to union membership in opposite directions (See section 2.6.1).

4.3 Technical Change, Deunionization, and Wage Inequality

We now turn to the estimation of the changes in the skill prices and the productivity distribution in 2007. Using the distribution of workers by education and experience, we predict a composite skill for these workers based on the prices estimated for 1978. The standard deviation of the skill index is 0.29, suggesting that the compositional changes in education and experience has added little to the increase in wage variance. The estimation of equation (19) for non-union workers in 2007, indicates a rise of 11% in the price of composite skill since 1978. The auxiliary estimate of the wage compression in 2007 is 35%. Next, we correct the auxiliary estimates for selection.

We rely on the increase in the total wage variance to identify the ability dispersion in 2007. The total wage variance in our sample increases 41% from 0.26 to 0.36. If the share of individual heterogeneity in wage variance was constant over time, this would correspond to a 41% rise in the model variance. We consider this case as a lower bound for how much worker-specific variation increased over time. In fact, the estimates reported in Haider (2001) show a rise in the share of individual factors in the total wage variance from 55% in 1978 to 70% in 1991. We are not aware of a comparable study for more recent years. At the other extreme, we assume that the rise in wage variation is all due to changing composition and prices of skills in the labor force. Since the variance in the data increases by 0.10, this implies a 75% (=$0.10/(0.55\times0.26)$) increase in the model wage variance. We consider this as an upper bound on how much the change in worker characteristics could have contributed to the rise in wage inequality. We refer to the two cases as A and B, respectively. Note that the rise in the variance of wages comes from *(i)* rising skill prices, *(ii)* changing skill composition and *(iii)* changing union composition. The model takes all of these into account when matching the variance figures reported above.

To correct the selection bias in the estimates of the changes in skill prices and the productivity distribution between 1978 and 2007, we keep the calibrated parameters constant, and evaluate our model to reproduce the auxiliary estimates in the data. The four remaining parameters, the skill prices in union and non-union sectors, and the standard deviations of skill and ability are identified exactly by the four corresponding moments: the variance of log-wages in 2007, the estimated values of the rise in the skill premium, the wage compression in 2007, and the standard deviation of observed skill in 2007.

Table 1 shows the results. We obtain two sets of estimates, depending on the variance of wages we adopt for 2007. In both cases, the estimate of skill dispersion is 0.26, which is very close to the estimate in 1978, and the skill price in the union sector is 0.75. The non-union skill price is up 12% in Case A, and 14$ in Case B, indicating 33% and 34% wage compression, respectively. The wage compression, with respect to observed skill, has eased slightly, which is incorporated in our model. The main difference appears in the estimate of ability dispersion, σ_x, which increases to 0.14 in Case A, and to 0.19 in Case B.

4.3.1 Rising Skill Premium and Deunionization

Overall, we think that the model provides a good description of the cross-sectional distribution of union participation and wages in 1978. Next, we calculate the unionization rate predicted by our model based on the estimated skill prices and the skill composition in 2005-2007. We keep all of the parameters related to the search and matching framework including the bargaining powers constant. Although one might suspect a decline in the union wage premium, as unions decline, the estimates in the data are stable during the period we consider. We therefore think that the relevant exercise, for our purposes, is to hold constant the relative bargaining powers of unions and individual workers.

In our conservative case for deunionization, the aggregate union coverage declines by 7.2 percentage points, from 32.1% to 24.9%. This is a third of the total drop in the unionization rate from 1978 to 2007. In Case B, which we take to constitute an upper bound on the contribution of changing skill prices to deunionization, the union coverage declines to 18.8%. Thus, we conclude from our benchmark results that the rising skill premium explains 30-60 % of the deunionization in the US since the late 1970s.

To evaluate the skill distribution of deunionization, Figure 11 compares the union participation rates by skill, before and after the rise in the skill premium. The unionization rates fall for most skill groups, and the union participation profile becomes flatter. Two factors help to explain this pattern. First, keeping all else constant, the nature of the change in the skill premium pushes both the lowest and the highest skilled workers out of the unions. Firms' hiring decisions play a crucial role for the decline in the unionization of low-skill workers. In particular, the decline in wages of low-skill workers in the non-union sector renders the low-skill union workers relatively more expensive. On the other hand, as the wages of skilled workers in the non-union sector rise, high-skill workers leave the union for better wages in the competitive sector. Second, our estimates indicate that the variation in wages, as explained by our skill measures, education, and experience, has risen relatively less than the variation in x. Since x is not observed, an increase in σ_x/σ_s makes the observed union profile flatter, raising the percent unionized for the highest and the lowest skill groups. This feature is particularly pronounced in Case B, where the estimated σ_x is much higher. In addition, since ability is not priced in the union, higher dispersion of ability increases the effective wage compression within the union, relative to the non-union sector, where both types of skills are valued equally. This results in a more homogeneous set of union workers, partially offsetting the effect of the rising skill price within the union on the wage inequality among unionized workers.

4.3.2 Decomposing the Decline of Labor Unions

The estimated unionization rate in 2007 is a result of the change in the skill premium and the skill distribution. The changes in the skill distribution are associated with shifts in the experience composition of the labor force, as well as changes in educational attainment. These alone could decrease the unionization rate when the skill distribution gets more dispersed, for instance, if more workers get college degrees and the union participation rate for college graduates is particularly low. To isolate the price effects from compositional changes, we evaluate a hypothetical economy where we raise only the skill price in the competitive sector to its estimated value. The skill composition, as well as the skill prices in the union sector, are kept constant at their 1978 levels. Table 4 shows that this by itself decreases the unionization rate to 25.9% in Case A, and to 25.2% in Case B. This effect is *ceteris paribus*, because it ignores the potential changes in composition of skills induced by the higher return to skill.

Next, we change the composition of skills along with the skill price in the non-union sector. This creates an additional 4.0% decline in the unionization rate in the conservative case, and 8.3% in case B. Since the change in the skill composition is, at least partly, endogenous to skill prices, we consider the additional decline as part of the effect of skill prices on the unionization rate.

The benchmark calibration also allows the skill price in the union to rise. This could be interpreted as the union's response to the rising skill prices in the competitive sector. This attempt to retain union workers clearly counteracts the deunionization, implied by rising skill premium in the competitive sector only. The results in Table 4 indicate that the ease of wage compression in the union sector has curbed the effects of the rise in the skill premium on unionization by 2-3 percentage points.

4.3.3 The Composition Effect of Unionization on Wage Inequality

In this subsection, we use the estimated wage functions to gauge the effect of unions on wage inequality. In particular, we ask the following question: What would the overall wage inequality be if all union workers were paid according to the non-union wage structure in the 1978 economy? Since the log-wage function in the competitive sector is $\ln C(\beta)+\psi(s+x)$, the hypothetical variance without unions would simply equal: $\psi^2(\sigma_s^2+\sigma_x^2) = 0.080$. As the actual variance of log-wages in the model is 0.0794 in 1978, the effect of unions on the variance of wages is -0.0006, which is less than 1% of the total model variance. This implies that the wage compression within the union sector is almost offset by the wage gap created between the union and non-union sector. Considering the increase in the model between 1978 and 2007, for the two scenarios, the composition effect of deunionization constitutes 1-2% of the total rise in wage inequality. Based on these figures, we conclude that the contribution of deunionization to the rise in wage inequality has been limited, unlike earlier findings (DiNardo, Fortin, and Lemieux (1996); Freeman (1993) among others).

To gauge the effect of selection, we replicate our calculations assuming that union status is random. We first use a simple two sector framework (Freeman, 1993). Let $\Delta_w = E[\ln w|U] - E[\ln w|N]$ be the average union wage premium and $\Delta_v = Var[\ln w|U] - Var[\ln w|N]$ be the causal effect of unions on the variance of wages. Denote the aggregate rate of unionization by ρ_u. Simple algebra implies that the total impact of unions on wage inequality is

$$\rho_u\Delta_v + \rho_u(1-\rho_u)\Delta_w^2,$$

where the first term is the decline in the variance due to wage compression among unionized workers, and the second term is the rise in the variance of wages between union and non-union sectors due to the union wage premium. Whether unions decrease the total variance depends on the size of the union wage premium relative to the wage compression.

The above calculation for our model implies that the impact of unions on variance is -0.030, which is 50-times the true effect! With selection of union workers from the middle of the skill distribution, $Var[\ln w|U]$ underestimates the hypothetical variance when union jobs are distributed randomly. The observed $Var[\ln w|N]$ is similarly high, not only because of the high skill prices in the non-union sector, but also because the non-union workers are selected from the tails of the skill distribution. This results in estimation of a large variance effect when selection is ignored. Accounting for the variation in union participation and union wage premium by observed skills[28] reduces the simple two-sector estimate by two thirds to -0.01, which is still 17-times the true effect. Correcting also for the selection bias in the union wage premium by observed skill reduces the estimate to -.006.[29]. While this brings the estimate closer to the true effect, it is still 10-times larger than the true effect.

To align the estimates with the true effect, the selection with respect to ability must also be taken into account. Since this component of productivity is not directly observed, correcting for non-random distribution of ability by union status is not trivial. We get around this problem by explicitly modeling the selection process under structural assumptions.

Most studies that estimate the compositional effect of deunionization to rising wage inequality compute the difference between the estimates of wage compression effects in two years. Therefore, even though the statistics are biased in each year, the bias in the difference is likely to be smaller.

4.3.4 Deunionization and Unemployment

The model also predicts a decline in the steady-state rate of unemployment for 2007. The decrease in the proportion of workers with higher bargaining power increases the average return to a filled vacancy from the firm's point of view. This leads to the creation of additional vacancies, raising the vacancy-unemployment ratio, θ, by 9-18%, and decreasing the unemployment rate by 3 basis points to approximately 5.7%.

5 Alternative Specifications and Sensitivity Analysis

The benchmark specification relied on the cross-sectional distribution of union participation by skill to identify the individual and collective bargaining powers, and the variance of unobserved ability.

[28] We use a variant of the formula above, extended to multiple skill levels where skill is measured by the deciles of workers' predicted wages in the non-union sector. We do not correct the estimated union premium figures for selection.

[29] A similar procedure is applied by Card (2001) to point out the effect of selection and heterogeneity on the estimates of wage compression by unions. The formula for the effect of unions on the variance of wages is $Var[\rho_u(s)\Delta_w(s)] + 2Cov[E[\ln w|s,N], \rho_u(s)\Delta_w(s)] + E[\rho_u(s)\Delta_v(s)] + E[\rho_u(s)(1-\rho_u(s))(\Delta_w(s)+\delta_x(s))^2 - \delta_x(s))^2]$, where $\delta_x(s) = \psi E[x|s,U] - E[x|s,N]$ is the bias in the union wage premium for workers of skill s. Note that this formula assumes that the unobserved productivity is equally valuable in both non-union and union sectors.

Since there is little guidance in the literature for the actual values of these parameters, we examine the sensitivity of our benchmark results to alternative values.

The bargaining parameters and the dispersion of skill also have implications for the wage distribution in each sector. Therefore, in this section, we identify these parameters using the variance of log-wages and the union wage premium by observed skill. Although some of our estimates change slightly as a result, the implied decline in the rate of unionization over time remains robust to alternative specifications.

The variance of wages in our benchmark case is 0.08, which is less than the estimates of worker heterogeneity found in the literature. The estimates provided in Haider (2001) indicate a corresponding variance of 0.14. Since his estimates do not control for firm-specific, or job-specific variation in wages, we consider this to be an upper bound. Combined with the estimates of such effects in Woodcock (2008), we deduce that the implied bias in Haider's estimate for 1978 is around 15%.[30] Therefore, for our robustness analysis, we target 55% and 40% of the total wage variance in 1978, which is equal to 0.14 and 0.10, respectively.

We maintain the estimates of wage compression and skill dispersion in our alternative specifications. We do not target the union participation profile, but we keep the aggregate rate of unionization among our targets. Additional moments that we use are the average union premium, or the distribution of union premium by skill, and the two alternative values for wage variance. This yields four specifications in total. Table 5 summarizes these specifications and reports the estimated parameter values in each specification. The estimates of the dispersion of skill and the relative skill price in the union are similar across different specifications. The dispersion of ability is larger when the target wage variance is higher. The estimates of bargaining parameters vary somewhat, with an average bargaining weight ranging between 0.20 and 0.34.

Table 6 shows the union wage premium as predicted by each of our four specifications. When we target the average union premium, the predicted distribution of the union premium by observed skill becomes very close to the observed distribution in the data. The moments, not surprisingly, become arbitrarily close to the data when we add the union premium profile in our target moments.

We now use our model to predict the decline in the unionization rate in response to changing skill distribution and skill prices. For each specification we re-estimate these changes for 2007. As in the benchmark case, we assume that the rise in the variance of wages was between 41% to 75%, and we report the unionization rate in both scenarios as bounds on the actual deunionization, as implied by the model. Table 7 shows the results. The conservative estimates place the unionization rate in 2007 at around 26-27%, whereas, the more permissive estimates are at around 21% to 22%. These figures imply that the rising skill premium explains 25-52% of the deunionization since 1978. Even though some of the parameter estimates were somewhat altered by alternative specifications,

[30] Table 2 in Woodcock (2008) reports the contribution of worker heterogeneity to total wage variance as 52% in the 1990s, after controlling for firm-specific and job-specific factors, compared to 70% ,reported in Haider (2001) for 1990-1991. This implies that the bias is about a quarter of the estimated value (or 17% of the total wage variance). Assuming that the relative size of the bias is constant over time, and given that the estimate in Haider (2001) for 1978-1980 is 57%, the actual contribution of worker heterogeneity must be 41% (or 39%) of the total variance.

the model's prediction for deunionization remains robust to these changes.

6 Conclusion

In this paper, we investigated the role of the rising skill premium on the rate of unionization and found that the rising skill premium explains at least a third of the deunionization in the US during the last three decades.

The decline in the unionization rate has been suggested to be one of the important determinants of the rise in wage inequality. The concurrence of the two events and cross-country evidence on the limited rise in wage inequality in more unionized economies is in line with this conjecture. We emphasized the unions as dynamic organizations composed of workers who respond optimally to market conditions, and showed that the causality could run the other way.

These predictions led us to investigate the role of deunionization, not as a main cause of the rise in wage inequality, but possibly as an amplification mechanism. However, we found that this effect is fairly limited quantitatively, in contradiction to the findings in the empirical literature, as well as to the predictions of the model by Acemoglu, Aghion, and Violante (2001). This is mainly due to selection biases in both the union wage premium and to the wage compression within the union. In particular, since union workers are selected from the middle of the skill distribution, the wage compression effect is exaggerated.

Although our model captures the main characteristics of unions, we have abstracted from a few potentially interesting extensions toward a more general theory of unions. First, the relatively stable union wage structure for the time period we analyze led us to refrain from modeling the political process that pertains to wage determination within the union. Explaining why skill prices are compressed for union workers would contribute to our understanding of unions. Second, we did not consider the association between union density and the presence of oligopolistic competition in the product market. In particular, larger mark-ups foster rent-sharing incentives and pave the way to union organization. As our results indicate that an important fraction of deunionization in the US remains unexplained, we think that investigating the connection between increased competition in the US and deunionization would be a promising next step.

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Data Appendix

Information on union status was first collected in 1973 in the Current Population Survey May supplements (May CPS). The survey question on whether or not the respondents' wage was covered under a collective agreement was added beginning in 1978. No union questions were used in the 1982 CPS. For later years, the survey questions on union status were included in the CPS Merged Outgoing Rotation Group (MORG) questionnaires. The data used in this paper is constructed from the May CPS extracts for 1973-1981 and from MORG files for 1983-2007 as provided in the National Bureau of Economic Research CPS collection.

We restrict the sample to male wage and salary workers in the private sector between the ages of 16 and 65, and with 0 to 39 years of potential experience. In all estimations, we use the hourly wage rate and CPS sampling weights. Hourly wages are reported as hourly earnings for those paid by the hour or as usual weekly earnings divided by usual weekly hours for other workers. The top-coded earnings observations are multiplied by 1.5 and the observations with hourly earnings below half the federal minimum wage in 2007 ($2.93=$5.85/2 in 2007 dollars) and those over $100 are dropped. Earnings are deflated by the CPI for personal consumption expenditures, as provided by the Bureau of Labor Statistics. Allocated earnings in the MORG data are excluded in all years, except for 1994 and 1995, when the census did not provide allocation flags. The May CPS does not include allocated earnings. We measure education by the completed years of schooling and categorize it into 5 groups; those with: less than 12, exactly 12, more than 12 but less than 16, exactly 16, and more than 16 years of education. Potential experience is calculated as age - years of education - 6.

Industrial Classification: The data from the May CPS (1973-1981) reports industry affiliation of workers based on the 1970 census classification. The MORG data uses the 1980 classification system of the census, for 1983-1999, with some changes beginning in 1992, and uses the 2002 classification for 2000-2007. To construct a time-consistent classification, we harmonized the industrial classification system in the MORG data using the classification mapping provided by Hirsch and MacPherson (2003)[31]. We then converted both the 1970 classification for 1972-1981 and the 1980 classification for 1983-1999 to a 2-digit NBER Industrial Classification. For our 2-digit industrial composition analysis, we excluded those in the armed forces or in public administration.The following industries were excluded for lack of sufficient observations throughout the sample period: toys, amusement, and sporting goods (NBER code 17), fabricated metal (NBER code 10) and other professional services (NBER code 45).

[31] A computer program is available online at http://unionstats.gsu.edu

Table 1: Calibrated and Estimated Parameter Values

Parameters	Value	Target Moments	
Calibrated Parameters (1978-1980)			
Unemployment Benefit (ρ)	0.400	Shimer(2005)	
Matching Coefficient (η)	0.324	Avg. unemp. duration	3.09 months
Matching Exponent (α)	0.500	Normalization	
Separation rate (λ)	0.020	Unemployment rate	6.03
Vacancy Cost (κ)	1.392	$\theta = 1$	
Discount rate (δ)	$0.99^{1/3}$	Annual interest rate	4%
Price of Skill (1978 - 1980) ($\psi_{1978-80}$)	1.00	Normalization	
Estimated Parameters			
1978 - 1980 Values			
Union Bargaining Power (γ)	0.387	Union Profile	See Figure 10
Competitive Bargaining Power (β)	0.240	Union Profile	See Figure 10
$\hat{\sigma}_{s,1978-80}$	0.261	Estimated Standard Deviation	0.284 (0.001)
$\hat{\sigma}_{x,1978-80}$	0.108	Union Profile	See Figure 10
$\hat{\phi}_{1978-80}$	0.605	Estimated Wage Compression	0.5359 (0.018)
2005 - 2007 Values: Case A			
$\hat{\psi}_{2005-07}$	1.120	Estimated Skill Price	1.11 (0.005)
$\hat{\phi}_{2005-07}$	0.751	Estimated Wage Compression	0.619 (0.015)
$\hat{\sigma}_{s,2005-07}$	0.261	Estimated Standard Deviation	0.285 (0.005)
$\hat{\sigma}_{x,2005-07}$	0.144	Percent Rise in Total Wage Variance	41%
2005 - 2007 Values: Case B			
$\hat{\psi}_{2005-07}$	1.136	Estimated Skill Price	1.11 (0.005)
$\hat{\phi}_{2005-07}$	0.748	Estimated Wage Compression	0.619 (0.015)
$\hat{\sigma}_{s,2005-07}$	0.263	Estimated Standard Deviation	0.285 (0.005)
$\hat{\sigma}_{x,2005-07}$	0.199	Percent Rise in Total Wage Variance	75%

Note. — The parameters were estimated simultaneously for each year and scenario. Numbers in parentheses are the standard errors that are used to weight the moments. See text for details.

Table 2: Benchmark Results: Union Wage Premium - 1978

Skill Quintile	Data[a]	Model	ATE[b]	TOT[c]
1	0.40	0.27	0.20	0.07
2	0.34	0.14	0.11	0.07
3	0.26	0.05	0.05	0.06
4	0.16	-0.04	0.00	0.06
5	-0.04	-0.16	-0.09	0.05
Average	0.20	0.05	0.06	0.06

Note. — The table shows the log wage differences between union and non-union workers by quintiles of predicted wages in the non-union sector. *a*. Estimates corrected for race, marital status, and survey year effects. The sample consists of male, private wage, and salary workers over the age of 16, from the CPS May supplements (1973-1981) and the monthly ORG files (1983-2007). *b*. The average wage gain for all workers in the quintile corrected for selection into unions. *c*. Actual real wage gains for union workers in the quintile corrected for selection into unions.

Table 3: Benchmark Results: Deunionization

		2005 - 2007	
	1978 - 1980	Case A	Case B
Data	32.1%	10.3%	10.3%
Model	32.1%	25.0%	18.5%

Note. — The table shows the extent of deunionization predicted by the model in response to rising skill prices in the non-union sector. Case A and Case B are lower and upper bounds for the effect of the rise in skill premium on deunionization.

Table 4: Decomposition of the Model Deunionization

Experiment	Case A	Case B
1978 - 1980 Economy	32.1%	32.1%
2005 - 2007 Economy with the new		
Non-union Skill Price only	26.2%	25.5%
Non-union Skill Price and Skill Composition	22.2%	17.2%
Union and Non-union Skill Prices and Skill Composition	25.0%	18.5%

Note. — The table shows the decomposition of the US deunionization into skill price and skill composition effects. Case A and Case B are lower and upper bounds for the effect of the rise in skill premium on deunionization.

Table 5: Sensitivity Analysis: Parameters

Target Moments	Parameter Estimates				
	$\sigma_{x,1978}$	$\sigma_{s,1978}$	β	γ	$\phi_{1,1978}$
Benchmark Specification	0.11	0.26	0.24	0.39	0.61
Average Union Premium with					
Low wage dispersion	0.18	0.25	0.18	0.31	0.61
High wage dispersion	0.27	0.25	0.13	0.24	0.62
Union Premium by Skill with					
Low wage dispersion	0.19	0.25	0.16	0.29	0.56
High wage dispersion	0.27	0.25	0.13	0.24	0.58

Note. — The table shows the sensitivity of the estimated parameters to selection of target moments. Each specification also targets the dispersion of wages explained by observed skills at non-union sector prices and the estimate of the relative price of skill in the union sector. The benchmark specification additionally targets the union participation profile by skill. Low(high) wage dispersion assumes that individual worker heterogeneity constitutes 40% (55%) of the total wage variance.

Table 6: Sensitivity Analysis: Union Wage Premium

		Skill Quintile				
	Mean	1	2	3	4	5
Data	0.20	0.40	0.34	0.26	0.16	-0.04
Model Targets						
Benchmark Specification	0.05	0.27	0.14	0.05	-0.04	-0.16
Average Union Premium with						
Low wage dispersion	0.20	0.36	0.25	0.17	0.09	-0.05
High wage dispersion	0.20	0.37	0.26	0.19	0.11	-0.01
Union Premium by Skill with						
Low wage dispersion	0.23	0.40	0.28	0.21	0.12	-0.04
High wage dispersion	0.22	0.41	0.29	0.21	0.13	0.01

Note. — The table shows the log wage differences between union and non-union workers by quintiles of predicted wages in the non-union sector. To maintain comparability with the data, all model statistics are calculated assuming that union status is random. All model specifications also target the dispersion of wages explained by observed skills at non-union sector prices and the estimate of the relative price of skill in the union sector. The benchmark specification additionally targets the union participation profile by skill. Low(high) wage dispersion assumes that individual worker heterogeneity constitutes 40% (55%) of the total wage variance.

Table 7: Sensitivity Analysis: Results

Target Moments	Rate of Unionization		
	1978 - 1980	2007 A	2007 B
Benchmark Specification	32.1%	24.9%	18.8%
Average Union Premium with			
Low wage dispersion	32.1%	26.0%	21.0%
High wage dispersion	32.1%	26.3%	22.1%
Union Premium by Skill with			
Low wage dispersion	32.3%	26.5%	21.2%
High wage dispersion	32.2%	26.6%	22.2%

Note. — The table shows the sensitivity of predicted deunionization to selection of target moments. Each specification also targets the dispersion of wages explained by observed skills at non-union sector prices and the estimate of the relative price of skill in the union sector. The benchmark specification additionally targets the union participation profile by skill. Low (high) wage dispersion assumes that individual worker heterogeneity constitutes 40% (55%) of the total wage variance.

Figure 1: Deunionization and Wage Inequality in US – Unionization is measured as a fraction of all employed workers. Wage inequality is measured by the Gini coefficient. Source: Kopczuk, Saez, and Song (2010) and Mayer (2004).

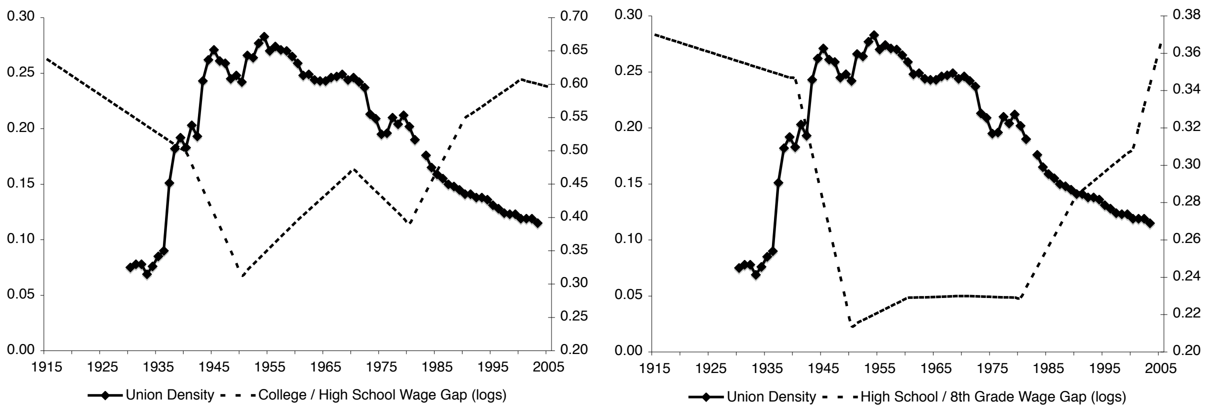


Figure 2: Union Membership and Education Premium in the U.S. (1973-2007) – Source: Goldin and Katz (2008) and Mayer (2004).

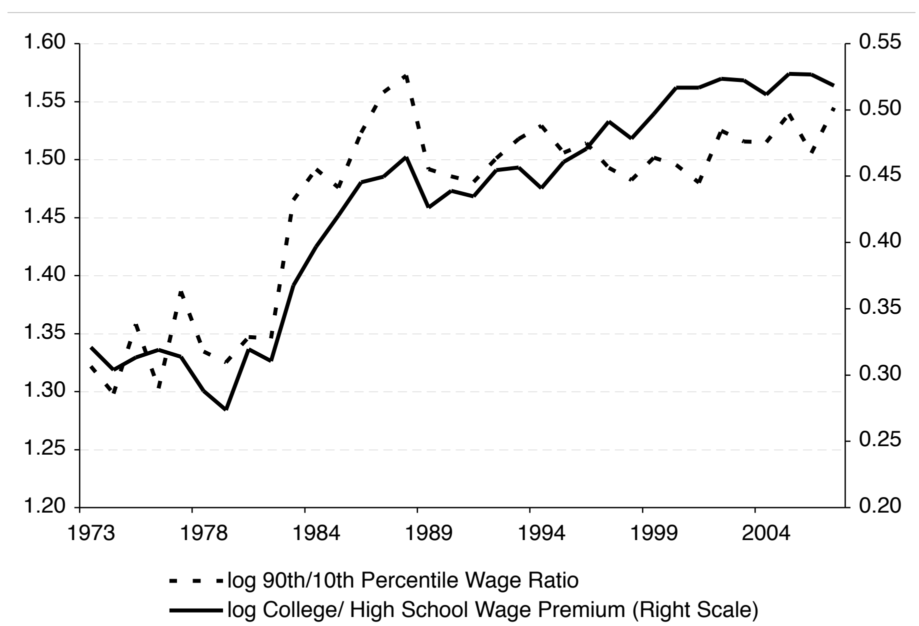


Figure 3: Rising Wage Inequality in the US – Hourly wages of male, private wage, and salary workers over the age of 16 from the CPS May supplements (1973-1981) and the monthly ORG files (1983-2007).

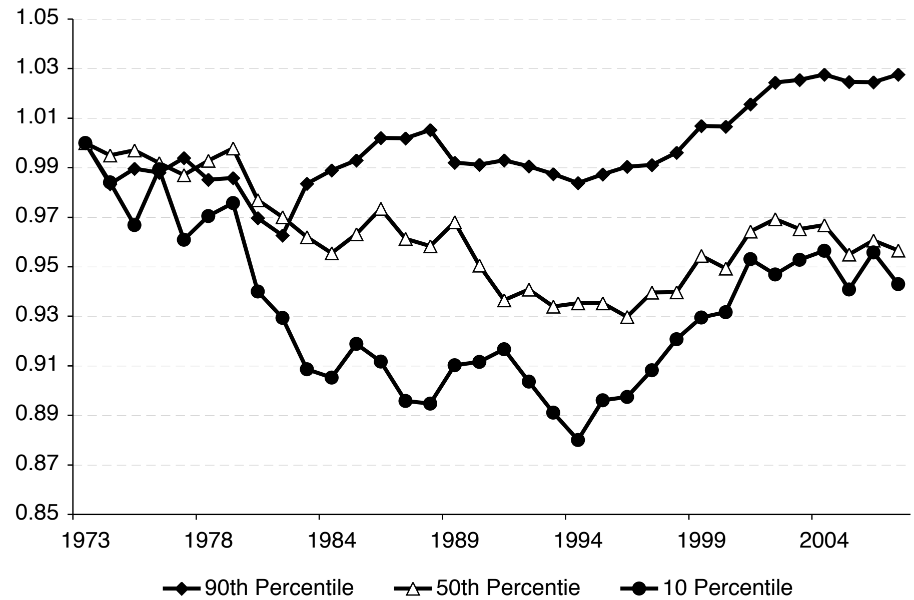


Figure 4: Changes in the Wage Distribution – Hourly wages deflated by the CPI index. The sample consists of male, private wage, and salary workers over the age of 16 from the CPS May supplements (1973-1981) and the monthly ORG files (1983-2007).

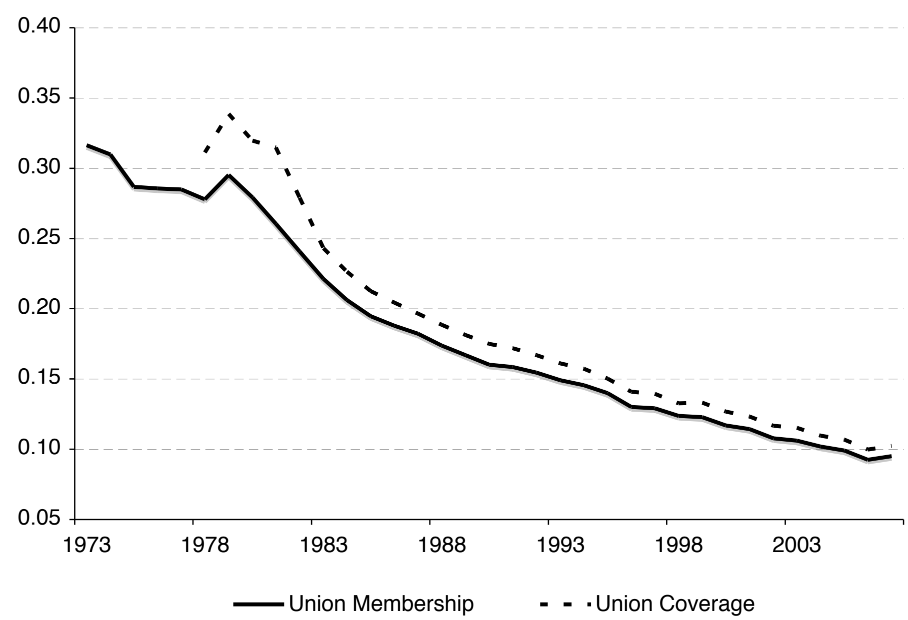
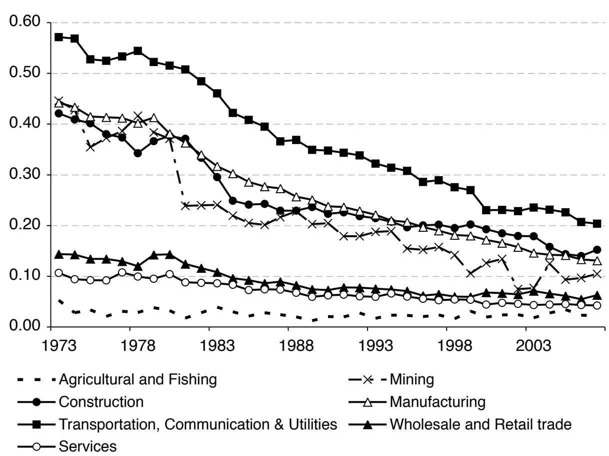
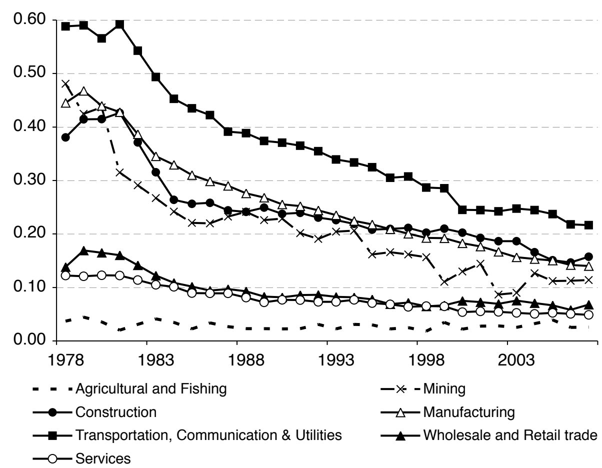


Figure 5: Rate of Unionization in the US – The sample consists of male, private wage, and salary workers over the age of 16 from the CPS May supplements (1973-1981) and the monthly ORG files (1983-2007). Membership is the percent of workers who belong to a union. Coverage is the percent of workers whose wages are determined by a collective agreement.

Figure 6: Union Coverage Density by Predicted Skill Deciles – The sample consists of male, private wage, and salary workers over the age of 16 from the CPS May supplements (1973-1981) and the monthly ORG files (1983-2007). Observations are stratified into deciles based on predicted wages in the non-union sector. See section 3.1 for details.



(a) Union Membership



(b) Union Coverage

Figure 7: Union Membership and Coverage by Industry in the US (1973-2007) – Percent of workers. The sample consists of male, private wage, and salary workers over the age of 16 from in the CPS May supplements (1973-1981) and the monthly ORG files (1983-2007). See the appendix for detailed industrial classifications.

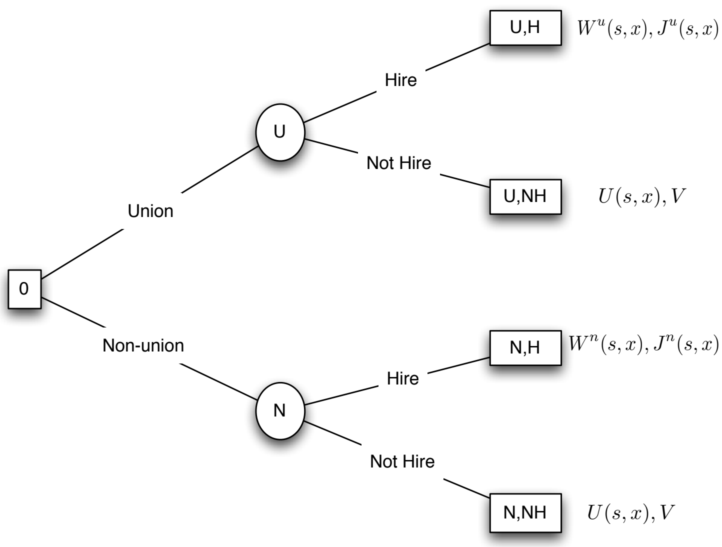
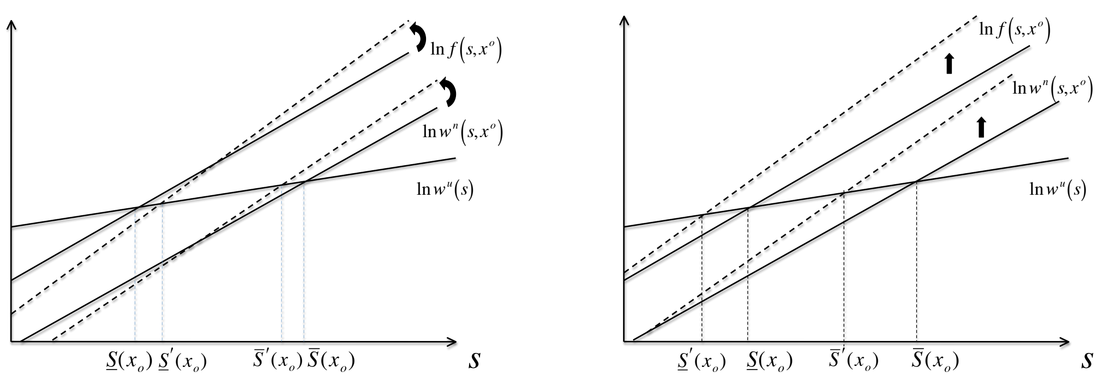


Figure 8: Game Tree and Payoffs for $\mathcal{G}(s, x)$



(a) Scenario I

(b) Scenario II

Figure 9: Rising Skill Premium and the Unionization Rate

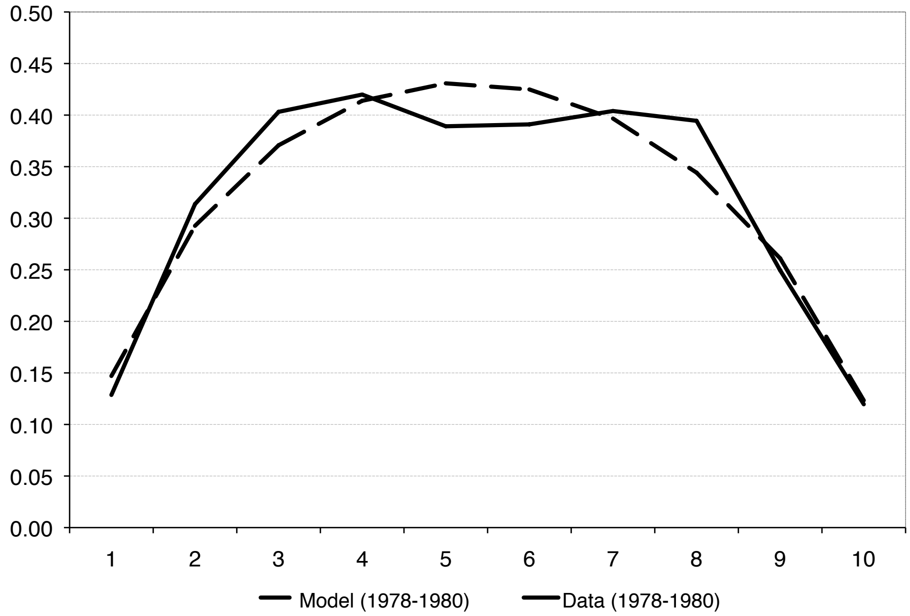


Figure 10: Model Unionization Rates by Skill Deciles: 1978 - 1980

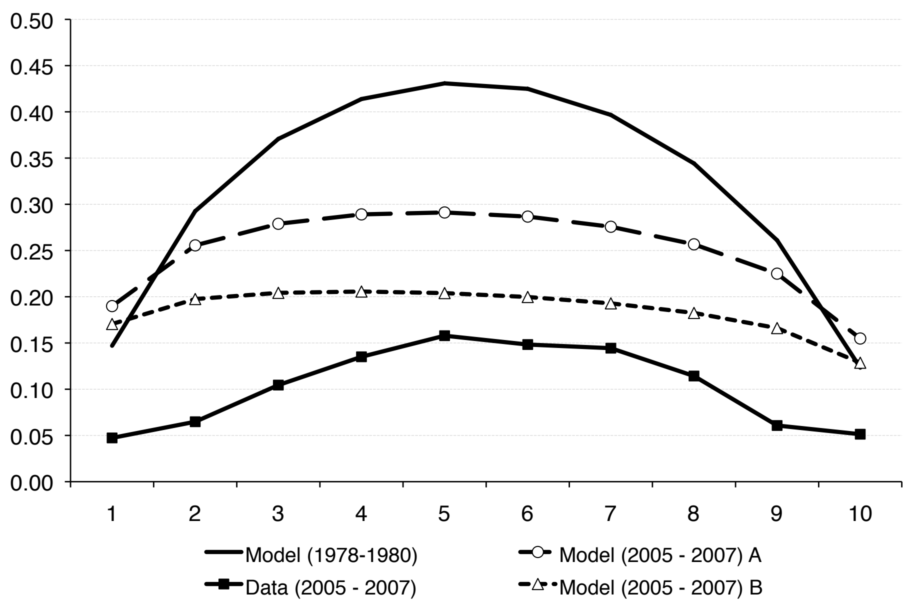


Figure 11: Model Deunionization by Skill Deciles: 1978 - 2007