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The impact of unions on nonunion wage setting: threats and bargaining

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The Impact of Unions on Nonunion Wage Setting: Threats and Bargaining*

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Abstract

In this paper we provide new estimates of the impact of unions on nonunion wage setting. We allow the presence of unions to affect nonunion wages both through the typically discussed channel of nonunion firms emulating union wages in order to fend off the threat of unionisation and through a bargaining channel in which nonunion workers use the presence of union jobs as part of their outside option. We specify these channels in a search and bargaining model that includes union formation and, in our most complete model, the possibility of nonunion firm responses to the threat of unionisation. Our results indicate an important role played by union wage spillovers in lowering wages over the 1980-2010 period. We find de-unionisation can account for 38% of the decline in the mean hourly wage between 1980 and 2010, with two-thirds of that effect being due to spillovers. Both the traditional threat and bargaining channels are operational, with the bargaining channel being more important.

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1 Introduction

Private sector unionisation in the United States is very nearly dead. In 2019, only 6.2% of private sector workers belonged to a union (U.S. Bureau of Labor Statistics (2021)). Recently, however, there have been some glimmers of revival, including successful unionisation drives at Amazon and Starbucks that are raising questions about whether a resurrected union movement could help slow or reverse the trend toward increasing inequality. We potentially can learn something about that possibility by examining the impacts on inequality of the de-unionisation that has occurred over the last 50 years. De-unionisation can affect inequality through various channels: by shifting workers from the union sector, where both wages are higher and wage differentials among workers tend to be smaller, to the nonunion sector where those differentials are larger; through any impacts on employment (perhaps through a general lowering of labour costs); through any accompanying changes in inequality within the union sector; and through indirect effects on wage setting in nonunion firms. These latter, spillover, effects are important since their existence could imply that the impact of de-unionisation on inequality is larger than what is calculated based just on the shifting of workers from the union to the nonunion sector. In this paper, we build a model of the impact of unions on wage setting in the nonunion sector and use it in estimation based on Current Population Survey (CPS) data to re-assess the role of de-unionisation in movements in the wage structure in the US.

The idea that unions could impact nonunion wage setting goes back at least to Lewis (1963). The core idea raised in that paper, and discussed in subsequent papers such as Rosen (1969), is that nonunion firms raise their wages in response to the 'threat' that their workers will unionise, presumably imposing extra costs beyond direct wage increases.[1] Our model incorporates that 'threat' effect plus an added mechanism through which the union sector affects nonunion wages: a bargaining channel whereby the outside options of nonunion workers and, through that their bargained wage, are affected by their ability to find high paying union jobs. In a sense, both are threat channels, with one being the threat of workers to leave and find a union job and the other channel being the threat to unionise the nonunion workplace. In order to differentiate these two channels, we will call the effects due to concern about being unionised the traditional threat effect and the other channel the bargaining effect. Our model makes clear the difficulties inherent in separately identifying these two effects while controlling for selection into the union/nonunion sectors. Part of the contribution of this paper is to provide estimates of spillover effects through both of these channels. This has the potential to expand our notion of the extent of the impact of de-unionisation on the wage structure, with the goal of our empirical work being to estimate the size of any such expansion.

The existing literature on union wage spillovers implements a specification in which nonunion wages are regressed on the percent of organized workers in labour markets defined in various ways. Evidence based on variation in this proxy for union power is mixed and sensitive to the included control terms, with the preponderance of studies finding a small

[1]For instance, Starbucks recently offered wage increases to "company-operated stores" but not in "unionized stores, or to stores that may be in the process of unionizing." NLRB has accused the announcement as a threat, designed to have a chilling effect on impending union votes. (New York Times, May, 2022).

positive spillover effect, though Neumark and Wachter (1995) estimate a negative effect.[2] In an important analysis, Farber (2005) carefully considers the role played by omitted variables, in turn introducing industry and state fixed effects. His analysis finds great sensitivity in estimates to the source of variation used, providing some context for the variation in estimates across earlier studies. When controlling for a wide range of potential omitted variables his results indicate at most a small positive effect of union power on nonunion wages.

In a recent paper, Fortin et al. (2021) estimate the role played by union threat effects using variation in the unionisation rate at the industry×state level included as an additional covariate in their distribution regression approach. They find positive effects of the unionisation rate operating primarily at the part of the wage distribution just below the median and their counterfactual exercise indicates that spillovers double the measured impact of deunionisation in increasing wage inequality in the US. While this is very useful, this paper shares with all of the early analyses a lack of an identification strategy for addressing the potential endogeneity of the union proportion. This stems from the lack of an effective instrument for the union proportion.[3] Virtually none of the papers in the literature even mention the twin problem of selectivity that could bias estimates: as the proportion of workers who are unionized declines, the composition of nonunion workers and firms will change.

In contrast to the existing literature on union spillovers that largely relies on reduced form estimation, our approach formalises union spillovers in a search and bargaining framework, endogenising the process of union formation and incorporating wage effects arising through differences in the bargaining process. In making clear what is being identified in the model and the variation used, we overcome the problems inherent in early studies of likely biases due to omitted characteristics and selection into the union sector, and we estimate an effect with a clear theoretical basis and interpretation.

Our model is based on that of **?** (henceforth TD), whose work is informed by the contributions of Pissarides (1986), Açıkgöz and Kaymak (2014), and Krusell and Rudanko (2016), among others. The TD model is centred around union threat effects which happen through the hiring channel. In the model, more skilled workers tend to dislike unionisation and firms skew their hiring toward these workers in order to stack the unionisation vote. Though this effect is certainly interesting, we believe it is likely of second order importance relative to a more direct firm response through raising wages to lessen the gains from unionising and direct union busting actions which raise the costs of unionisation. Our model focuses on these latter effects instead of the hiring channel.

[2] Freeman and Medoff (1981) find a non-significant positive correlation between the proportion union and nonunion wages, a result confirmed by Donsimoni (1981). Conversely, Holzer (1982) finds a large positive effect, with a 10% increase in the proportion unionised associated with a 4% increase in nonunion wages. Kahn (1980) provides additional evidence in support of a positive spillover effect, while Hirsch and Neufeld (1987) find a positive spillover effect at the industry level, but no significant effect operating in the local labour market. Dickens and Katz (1986) also find a positive correlation between the union proportion and wages, although Podgursky (1986) finds this effect exists only for large establishments.

[3] Farber (2005) presents event studies of the enactment of Right to Work (RTW) laws in Idaho and Oklahoma, while noting that RTW's are more likely to be enacted in weak union states and, so, face endogeneity issues of their own. In an earlier version of the paper, (Fortin et al., 2019), the authors extend the analysis of Farber (2005) using variation in right-to-work laws in an event-study framework. They find evidence of reductions in nonunion wages with the introduction of right to work laws but the estimates are poorly defined because there are few states that switch RTW status in their time period.

Additionally, our framework is informed by Beaudry et al. (2012) (henceforth BGS), which formalises the impact of changing alternative job prospects (outside options) on wages due to changes in the industrial composition of work. Following BGS, we model local labour markets composed of industries and firms with workers able to transition between jobs in proportion to job prevalence.[4] As in BGS we will exploit cross-city, within-industry variation - in our case, to identify the effect of declining unionisation and changes in the composition of union work on nonunion wages over the period 1980-2020.

Combining these elements, we derive an empirical specification which incorporates spillover effects operating through both the bargaining and standard threat channels, formalises selectivity, and makes it straightforward to see barriers to identification. Specifically, changes in outside options associated with the union sector may be correlated with unobserved local productivity shocks. As in BGS and Beaudry et al. (2014), we overcome this problem using Bartik-style instruments related to worker outside options. For nonunion workers, outside options are related to the probability the worker could transit to a union job (which we allow to vary by industry and over time) times the expected wage the worker could get in that job. It also depends on expected wages in nonunion jobs in the local economy. The Bartik instruments use versions of these outside options based on start of period industry and union employment composition in a locality interacted with changes in industry growth, industry premia, and the probability of moving to a union job at the national level. It is this outside option for nonunion workers that identifies the bargaining channel for union effects. We get extra power to identify the bargaining effect because improvements in outside options have the same effect on bargained wages whether they stem from reduced probabilities of finding a union job or reductions in the number of high-rent nonunion jobs. That means we get identification from both changes in unionisation and shifts in industrial structure in both the nonunion and union sectors. We argue that the validity of our Bartik instrument depends on a random walk type assumption that we show implies an overidentifying restriction. We test that restriction and cannot reject it. We also show that within the context of the model, we identify the threat channel by the impact on nonunion wages of the interaction of the probability a firm in a given industry×city cell would face a union election (which shows the size of the direct threat) with the outside option value for union workers (which captures the size of what the firm needs to respond to in order to prevent unionisation). We construct and implement similar Bartik instruments related to this component.

The results from our estimation point to the importance of both spillover channels. Between 1980 and 2010, the mean real wage in the US fell 16% (holding composition in terms of education, experience, race and gender constant). Through a decomposition exercise based on our estimates, we find that de-unionisation accounts for 38% of the decline. A third of that impact arises from a standard shift share effect (because workers shifted away from higher paying union jobs) with the other two-thirds from spillover channels. Unions have spillover effects on nonunion wages and they are sizeable.

While both the traditional threat and bargaining effects show up significantly in our

[4] Although we specify the firm as the level at which workers become unionised in the model, our estimation exploits variation at the industry-city level. In that sense, firms can be thought of more accurately as establishments, or more generally, the level at which the unionisation vote takes place. Given the emphasis on the enterprise level in the Wagner Act, aggregation to the city level may ignore some complexity, particularly if there are notable transitions between different union/nonunion establishments in the same firm.

estimates, our decomposition exercise indicates that the spillover effects are almost entirely due to the latter. The threat probability was too low, even in 1980, to play a substantial role. This is potentially important for considerations about trying to raise wages through policy tools. The threat effect is unique – it can only be harnessed by increasing unionisation. The bargaining effect, in contrast, is more general. It is about more good paying jobs representing improved outside options for workers in all other jobs, something that Beaudry et al. (2012) and Caldwell and Danieli (2021) point out can have sizeable positive effects on wages in a location. Unions are one way to create such a higher wage option but other policies, such as eliminating non-compete arrangements could also have such an impact (Johnson et al. (2020)).

Although spillovers roughly tripled the standard shift share effect of unionisation over the long run, we find no evidence of bargaining spillovers in the 1980s – the decade of the largest declines in unionisation. This is because those declines were matched with increases in the union wage premium, increasing the value of the outside option of nonunion workers at the same time as declining probabilities of finding union jobs were reducing it. Our model provides an explanation for the increased wage premium in the 1980s that echoes an argument in Farber (2005). While both union and nonunion wages faced downward pressures from technological change, trade, etc.., the substantial reduction in the risk of being unionised in the decade meant that, in addition, nonunion firms no longer had to pay higher wages in order to stave off unionisation. As a result, nonunion wages fell faster than union wages. After 1990, the threat of unionisation stabilized at a low level, causing the union wage premium to decline, and the outside option effect of unions began to reflect the falling unionisation rate alone. The potential lesson for any re-unionisation efforts is that spillover effects onto nonunion wages may arise through the traditional threat channel but the implied increase in nonunion wages will dampen the bargaining channel. Union jobs would be more plentiful but not pay as high a premium over nonunion jobs as before re-unionisation. Eventually, as the unionisation threat stabilized, the extent of spillover onto nonunion wages would increase, but that could take time to be fully realized.

Our work is also related to the substantial literature that investigates patterns in declining unionisation, estimating both movements in the union wage premium and the role of declining unionisation in driving increasing wage inequality. Card et al. (2004) and Card et al. (2018) provide comprehensive summaries of the research in this area following the early contribution of Freeman (1980). Farber et al. (2021) provide the most comprehensive account of the relationship between union density and inequality in the U.S., introducing new survey data that allows them to push their analysis back to the 1930s. They find that increasing unionisation had a substantial impact on decreasing inequality after WWII while the reversal in the unionisation trend had a smaller effect on increasing inequality in the last 50 years. Their estimates allow for spillover effects onto nonunion wages but they do not study spillovers directly. Our results imply that spillovers may have played an important role in their estimated inequality impacts from unions and provide an explanation for why those impacts were less evident at the time of the big union decline in the 1980s.[5]

[5] Other papers in this literature include an important contribution by DiNardo et al. (1996), who introduce a semi-parametric re-weighting technique building on the work of Oaxaca (1973) which attributes 14% of the increase in wage inequality over 1979-1988 (for men) to declining unionisation. Extensions of this work are found in DiNardo and Lemieux (1997) and Fortin et al. (2019). Further studies by Card (2001), Card

The remainder of the paper is organised as follows: In section 2, we present our versions of our model both with and without incorporated firm responses to unionisation threats. In section 3, we describe the data and the construction of our key outside option variables and provide a discussion of our instrumental variables. Section 4 contains descriptive patterns and our estimation results. In section 5, we present a counterfactual exercise designed to demonstrate the impact of spillovers on movements in the wage structure and the role played by our two channels. Section 6 contains conclusions.

2 The Model

2.1 Model Set-up

Our model is based on that of **?** (TD) which places union formation and wage setting in a search and bargaining model. Unions are able to bargain a higher wage because they can threaten to take the whole workforce out of production, while an individual, nonunion worker can only threaten to withdraw her own labour. In the TD model, firms employ workers of different skill levels who have different preferences about unions. In particular, since unions compress skill differentials, more skilled workers would vote against a union and less skilled workers would vote in favour. Firms facing a union threat can resist unionisation by taking advantage of these preferences: hiring a larger proportion of skilled workers in order to construct a workforce that will vote against a union. Our interest is in the potential spill-over effects of unions on nonunion wages and, partly for that reason, we focus instead on the concept of nonunion firms resisting unionisation through paying higher wages. We also view this type of wage emulation channel as well as direct legal and illegal campaigns by firms as likely more immediate responses to unionisation threats than altering the skill composition of the workforce. In addition, we will allow for the possibility of both direct and wage emulation effects of unions on wages in a given industry affecting wage setting in other industries in a local economy. Following Beaudry et al. (2012)(BGS), this can happen because having higher rent job options in a city increases the value of the outside option for workers in all industries as they bargain with their employers. Through the rest of the paper, we will refer to effects of unions on nonunion wage setting in order to resist unionisation as standard threat effects (to reflect that these are what have been discussed in the previous literature) and indirect effects through impacts on bargaining options as bargaining spillover effects. To focus attention on whether these channels are sizeable, we will alter the TD model by having only one skill level but multiple industries.

With that in mind, we start with an aggregate production function of the form:

$$Q = \left[\sum_i a_i Z_i^{\chi}\right]^{\frac{1}{\chi}}, \quad \chi < 1,$$

et al. (2004) and Gosling and Lemieux (2001) provide estimates of the union contribution to wage inequality, comparing patterns across male/female workers, private/public sector workers, and countries. In a recent contribution, Card et al. (2018) compare the experience of Canada and the United States, focusing in particular in the increasing prominence of the public sector as a source of unionised employment. See also recent studies by **?** and Firpo et al. (2018).

where Q is total output for the national economy and is equal to a CES aggregate of I intermediate goods indexed by i. The intermediate goods (Z_i's) are produced in local economies, or cities, of which there are C. Thus, $Z_i = \sum_c Y_{ic}$, where Y_{ic} is the amount of intermediate good i produced in city c. The price for Q is normalized to 1 and the price for intermediate good i is given by r_i. For simplicity, and to focus attention on wage setting, we will ignore firm entry and exit and assume that there is a fixed number of firms operating in each industry by city cell. Thus, $Y_{ic} = \sum_f y_{icf}$, where y_{icf} is the output of firm f operating in industry i in city c.

Firms and workers operate in a labour market that includes frictions. Unemployed workers and vacancies posted by firms meet according to a matching function, $M(U_c, V_c)$, where U_c is the number of unemployed workers in city c and V_c is the number of unfilled vacancies. As is standard in the search literature, we will assume that the matching function is constant returns to scale. BGS show that in steady state this implies that the probability that a vacancy in city c meets an unemployed worker, q_{vc}, and the probability that an unemployed worker in city c meets a vacancy, q_{uc}, can be written as functions of the employment rate for the city, ER_c. We will assume that workers are not mobile across cities and that there is no on-the-job search.[6] Our model is partial equilibrium in that we treat q_{vc}, q_{uc}, and ER_c as given rather than solving for them from the model. We make this simplifying assumption in order to maintain our focus on wage outcomes.

Once workers and firms meet, they bargain a wage to divide the match surplus. Following TD, in a nonunion firm, this bargaining is between the individual worker and the firm while in a union firm it is between the set of employees and the firm. Employees at nonunion firms have the opportunity each period to vote on whether to form a union. Once unionised, a firm stays unionised – there is no decertification. Both union and nonunion firms choose the optimal number of workers to hire, taking account of the bargained wage.

2.2 Workers

We begin by characterizing the choices and environment faced by workers. We will work in discrete time and assume that all workers have the same skill level. At any moment, a worker can be unemployed and searching for work or employed at one of three types of firms in a particular industry. The first type of firm is a simple nonunion firm in which the firm and worker bargain a wage and the firm chooses an optimal number of employees without direct regard to the threat of unionization. As we will see, these firms exist in situations where the costs and benefits of unionization are such that workers at the firm do not want to form a union. Workers in these firms have a value function given by:

$$W^n_{icf}(w^n_{icf}) = w^n_{icf} + \rho(\delta U^u_{ic} + (1-\delta)W^n_{icf}(w')), \tag{1}$$

where ρ is the discount rate, w^n_{icf} is the nonunion wage that would be paid at firm f in the given $i-c$ cell, δ is an exogenous probability that the worker's job is terminated, and U^n_{ic} is the value of unemployment in city c for a worker whose previous job was a nonunion job in industry i. In this specification, w' corresponds to the wage that will be paid in the next period if the job is not terminated. Following TD, we assume that workers and firms believe

[6] Beaudry et al. (2014) provide a model with mobility.

that next period's wage will be set optimally and that they cannot affect it through actions this period. This assumption rules out, for example, the relevance of reputations.

Alternatively, workers may vote to unionise, in which case their value function is:

$$W^u_{icf}(w^u_{icf}) = w^u_{icf} + \psi_{icf} + \rho(\delta U^u_{ic} + (1-\delta)W^u_{icf}(w')), \tag{2}$$

where u superscripts correspond to unionised values, and ψ_{icf} is the non-wage benefit to workers from being in a union in this particular firm, which could be related to the work environment in the firm. We assume that a value for ψ_{icf} is drawn separately for each firm from a common distribution, $f(\psi)$, and that the union amenities are created by the union itself rather than the firm and do not enter the cost function for the firm.

Firms can respond to a union threat by paying workers a wage that is high enough to make them indifferent between unionising and remaining non-union (what we will call an emulation wage). Given this, there is a third possible type of firm which we will call a union emulating firm in which workers are paid a wage, $w^{n*}_{icf} > w^n_{icf}$ but do not get union benefits. As a result, the value function associated with working at this type of firm is the same as for the simple non-union firm but with a wage of w^{n*}_{icf}.

The other potential state for workers is being an unemployed searcher. For reasons that will become apparent momentarily, the value of unemployment differs by previous industry and by whether the person was a union or nonunion worker in their previous job. The value function for a person who was formerly in a nonunion job in industry j is given by the follow equation (an analogous expression exists for those formerly in a union job in industry j):

$$\begin{aligned} U^n_{jc} = b + \rho\Bigg(& q_{uc}\left[T^n_{jc}\sum_i \eta^u_{ic}W^u_{ic} + (1-T^n_{jc})\sum_i \eta^n_{ic}\left(p^n_{ic}W^n_{ic} + (1-p^n_{ic})W^{n*}_{ic}\right)\right] \\ & + (1-q_{uc})U^n_{jc}\Bigg). \end{aligned} \tag{3}$$

where b is the flow value of being unemployed; η^u_{ic} and η^n_{ic}, respectively, are the proportions of employment in the union and nonunion sectors in city c that are in industry i. For workers in city c and previously employed in a nonunion job in industry j, the probability of finding a job in the union sector is T^n_{jc}. With probability, $(1-T^n_{jc})$, these workers match with a nonunion firm. We allow the probability of finding a union job to differ by whether the worker was previously in a union job and by their industry of previous employment.[7] This is the reason for indexing the value of unemployment by previous industry and union status.

We assume that search is random – that, for example, workers are unable to differentiate between nonunion firms paying regular nonunion wages and nonunion firms paying emulation wages.[8] As a result, they find employment in proportion to the relative abundance of these

[7] For union job searchers the probability is T^u_{jc}.

[8] Our model therefore abstracts away from issues related to workers queueing for union jobs (see Abowd and Farber (1982) for a theoretical treatment of queuing and supportive empirical evidence). This queuing mechanism could imply an additional spillover channel whereby the existence of union firms drives down vacancy filling rates in the nonunion sector, pushing up wages. The prevalence of queueing is likely driven by union wage premia and the relative likelihood of finding union work such that queuing effects are likely to enter through the outside option channel in our model.

firms as measured by p_{ic}, which is the fraction of jobs in industry i in city c that are in non-emulation firms. $W^u_{ic} = E(W^u_{icf})$ is the expected value of employment in a union firm in the $i-c$ cell with the expectation taken across the distribution of union firms in that cell (the other expectations are defined analogously).

Taken together, (3) says that an unemployed worker gets a flow value, b, and with probability q_{uc} meets a vacancy. Workers have different propensities for finding employment in the union and nonunion sectors, but otherwise find jobs in proportion to their share of local employment. The expression implies that the value of search is higher when the local employment structure has more high value industries and more unionised firms since that means searchers are more likely to find those high value jobs. To simplify other expressions in the model we will often summarize the expected value of a job as:

$$\tilde{W}^n_{jc} = T^n_{jc} \sum_i \eta^u_{ic} W^u_{ic} + (1 - T^n_{jc}) \sum_i \eta^n_{ic} \left(p^n_{ic} W^n_{ic} + (1 - p^n_{ic}) W^{n*}_{ic} \right). \tag{4}$$

2.3 Firms

All firms in a given industry-city cell have a common production function given by:

$$y_{icf}(n) = \epsilon_{ic} n_f - \frac{1}{2} \sigma_i n_f^2,$$

where ϵ_{ic} is a local productivity shock, n_f is the number of employees, and $\sigma_i > 0$ is a parameter reflecting potential span of control issues.[9] This specification implies that technology is common across cities within an industry but that there is comparative advantage in producing each intermediate good by city. We assume that the technology is common to all three types of firms (unionized, non-union, and union emulators). The literature on union effects on productivity seems to us to be inconclusive and so we adopt an agnostic take in which unions affect firm activity by affecting wages but not through technological adaptations.[10] We assume that the σ_i's are sufficiently smaller than 1 such that, combined with the assumption of a fixed number of firms in each $i-c$ cell, they imply that production of any good is spread across cities.[11]

Firms choose a number of vacancies to post in order to attain a profit maximizing number of employees given the bargained wage. The cost of hiring is linear in the number of vacancies posted. As a result, the value function for a nonunion firm is given by:

$$J^n_{icf}(n) = \max_v r_i y_{icf}(n') - w^n_{icf}(n') n' - \kappa v + \rho J^n_{icf}(n'),$$

subject to the equation of motion:

$$n' = n(1 - \delta) + q_{vc} v,$$

[9] For notational simplicity, we will drop the f subscript on n_f.

[10] Hirsch and Link (1984) and Addison and Hirsch (1989) summarise the early research in this area which finds largely inconclusive and mixed evidence on the effect of unionisation on productivity.

[11] We work with a quadratic production function to permit tractability in deriving our wage expressions. It is worth noting that TD showed that using a Cobb-Douglas type form for production has the unfortunate implication that general productivity shifts such as ϵ_{ic} do not determine wages. In that sense, our results are not perfectly generalizeable but a quadratic function captures the main points we want to emphasize.

where $J^n_{icf}(n)$ is the value of non-union firm f when it ends the previous period with n employees. Of those, a fraction, δ, leave the firm for exogenous reasons. The firm posts v vacancies at a cost, κ, per vacancy and fills them with probability, q_{vc}. Following TD, we will assume that the firm can only post positive vacancies and that the number of vacancies is sufficiently large that we can treat $q_{vc}v$ as a deterministic number according to a law of large numbers. This allows us to rewrite the value function as:

$$J^n_{icf}(n) = \max_{n'} r_i y_{icf}(n') - w^n_{icf}(n')n' - \kappa \frac{n' - n(1-\delta)}{q_{vc}} + \rho J^n_{icf}(n').$$

Note that we have written the value function as if the firm exists in a stationary environment in which it assigns a probability of zero to its workers trying to unionise in the future. We will return to that assumption below. The value functions for firms when they are unionised or acting as union emulators are identical in structure to the nonunion firm value function, with the only change being the substitution of the relevant wage ($w^{n*}_{icf}(n')$ or $w^u_{icf}(n')$) in the expression. We next turn to using these value functions in combination with those for the workers to characterize wage bargaining solutions. Once we have those solutions, we will be in a position to discuss union formation.

2.4 Wage Determination

2.4.1 Collective Bargaining

Once the firm hires workers, the workforce will vote on unionisation. If a union is formed then the union bargains collectively on behalf of workers. Following TD, wages are bargained according to Nash bargaining over the entire surplus to production from hiring n workers. By bargaining collectively, the union is able to effectively threaten that the entire workforce will quit this period. Such an action would impose two costs on firms: first, the firm would produce zero units of output this period; second, as the firm employs n workers every period in steady state, the number of vacancies required to achieve this optimal workforce will be much larger, increasing vacancy filling costs. These costs are taken into account when determining firm and worker surplus.

Wages (in a unionised firm) are set according to the Nash Bargaining condition:

$$\beta S^u = (1-\beta)n(W^u_{icf}(w) - U^u_{ic}),$$

where S^u represents the firm's surplus. On the right hand side is the sum of workers' surplus, which is given by the gain to employment for all workers hired by the firm. Since the workers are identical, we use a specification that focuses on the total surplus and assume that the union members will all get an equal share of the part of the surplus captured by the union. This ignores issues related to seniority, for example.[12]

[12]See Abraham and Medoff (1984, 1985) who present evidence of the importance of seniority for layoffs and promotions, and see Abraham and Farber (1988) for evidence that the seniority wage profile is steeper under collective bargaining.

We will focus on a steady state in which the wage and optimal number of workers for a firm are constant across periods. Given this, the workers' surplus can be re-written as:

$$
\begin{aligned}
W^u_{icf}(w) - U^u_{ic} =& \frac{1}{1-\rho(1-\delta)}(w^u_{icf}(n^u_{icf}) + \psi_{icf}) - \frac{(\rho-1)b}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))} \\
&- \frac{(\rho-1)\rho q_{uc}}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))}\tilde{W}^u_{ic}.
\end{aligned}
$$

On the firm side, the surplus from a successful bargain with a union is given by the difference between producing this period with n_u workers (the optimal number of workers with a bargained union wage) and not producing this period along with the cost of rehiring the entire workforce the next period. The firm surplus can be expressed as follows (with a detailed derivation provided in Appendix A):

$$
S^u = \left(r_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf})\right) + \rho(1-\delta)\frac{\kappa n^u_{icf}}{q_{vc}}.
$$

The surplus equals the current period profit plus the cost of replacing the portion of the work force that would not normally turn over, which is the relative cost of being in the bargaining break-down option.

Solving the bargaining expression for steady state wages yields:

$$
\begin{aligned}
w^u_{icf}(n^u_{icf}) =& \frac{\beta(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))}\frac{r_i y_{icf}(n^u_{icf})}{n^u_{icf}} + \frac{\beta\rho(1-\delta)(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))}\frac{\kappa}{q_{vc}} \\
&+ \frac{(1-\beta)(1-\rho)}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}b - \frac{(1-\beta)}{1-\beta\rho(1-\delta)}\psi_{icf} \\
&+ \frac{(1-\rho)\rho(1-\beta)q_{uc}}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc})}\tilde{W}^u_{ic}.
\end{aligned}
$$

2.4.2 Individual Bargaining

Non-union wages are also set through Nash bargaining. However, the value for the firm of the option corresponding to a break down in negotiations relates only to the loss of an individual worker (and the indirect effects the removal of one worker has on the others) rather than to the complete stoppage of production that occurs under collective bargaining. The Nash bargaining condition is the same as before, and we derive the following expression for the firm surplus (see Appendix A.1 for details):

$$
S^n = r_i\frac{\partial y_{icf}(n^n_{icf})}{\partial n} - w^n_{icf}(n^n_{icf}) - n^n_{icf}\frac{\partial w^n_{icf}(n^n_{icf})}{\partial n} + \frac{\rho(1-\delta)\kappa}{q_{vc}}.
$$

Substituting this expression into the Nash bargaining condition and (again, following TD) solving the differential equation in wages yields the following expression for non-union wages:

$$
\begin{aligned}
w^n_{icf}(n^n_{icf}) =& \frac{1-\rho(1-\delta)}{1-\beta\rho(1-\delta)}\frac{\beta r_i}{1+\beta}\left(\frac{\partial y_{icf}(n^n_{icf})}{\partial n} + \beta\epsilon_{ic}\right) + \frac{\beta\rho(1-\delta)(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))}\frac{\kappa}{q_{vc}} \\
&+ \frac{(1-\beta)(1-\rho)}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}b + \frac{(1-\rho)\rho(1-\beta)q_{uc}}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}\tilde{W}^n_{ic}.
\end{aligned}
$$

This wage differs from the union wage in two ways. First, under collective bargaining the union negotiates over the surplus generated from total output while for non-union workers what matters is the marginal surplus. As a result, individual wages are determined by average output for union workers but marginal product for nonunion workers. Second, union wages include a compensating differential component with wages declining when union amenities are greater. On the other hand, outside options – which are determined by both the value of non-work time and the expected value of finding a new job after unemployment – have the same effects on the union and nonunion wages.

2.4.3 Wage Equations: Empirical Specifications

The wage equations derived to this point include firm size as an argument of either the average or marginal product of workers. In Appendix A.3, we derive the optimal firm size expressions for union and nonunion firms and then substitute those into the wage equation. Thus, we implement quasi-reduced form wage specifications that allow us to avoid dealing with firm size endogeneity issues and with problems with the CPS data where there is categorical data on the size of the whole firm rather than on the establishment or workplace. It is the latter that is most relevant for our analyses.[13] It is worth noting that nonunion firms have a larger optimal size, though this is mitigated to some extent by the fact the union amenities lower the wage differential between union and nonunion firms.

In our empirical work, we will focus on linearizations of wages around the point where all cities have the same proportions in each industry.[14] This is done partly to make explicit the role of the employment rate in the city, ER_c. In steady state, the worker job finding rate, q_{uc} and the firm worker finding rate, q_{vc} can be written as functions of ER_c. The other main driving force is the outside option for workers, which is a function of the expected value of a job ($\tilde{W}^n_{ic}$ for a nonunion worker currently in sector i in city c). Working from the fact that the expected value of a job is a function of wages, one can show that the outside option for a nonunion worker can be represented as a weighted average over wages in different sectors in the local economy. We will write that weighted average as $E^n_{ic}(w)$ for a nonunion worker and $E^u_{ic}(w)$ for a union worker.

Thus, wages can be written as:

$$w^n_{icf} = \gamma^n_{0i} + \gamma^n_1 E^n_{ic} + \gamma^n_2 ER_c + \gamma^n_4 \epsilon_{ic} \tag{5}$$

and

$$w^u_{icf} = \gamma^u_{0i} + \gamma^u_1 E^u_{ic} + \gamma^u_2 ER_c - \gamma^u_3 \psi_{icf} + \gamma^u_4 \epsilon_{ic} \tag{6}$$

for nonunion and union workers, respectively, where

$$E^u_{ic} = T^u_{ic} \sum_j \eta^u_{jc} w^u_{jc} + (1 - T^u_{ic}) \sum_j \eta^n_{jc} \left(p^n_{jc} w^n_{jc} + (1 - p^n_{jc}) w^{n*}_{jc} \right) \tag{7}$$

and

$$E^n_{ic} = T^n_{ic} \sum_j \eta^u_{jc} w^u_{jc} + (1 - T^n_{ic}) \sum_j \eta^n_{jc} \left(p^n_{jc} w^n_{jc} + (1 - p^n_{jc}) w^{n*}_{jc} \right), \tag{8}$$

[13] Although we have used 'firms' throughout this paper as our terminology of choice, one can think of union certification at the level of the workplace or establishment without making any modification to the model.

[14] See Appendix A.4 for details.

where $p^n_{jc}w^n_{jc} + (1 - p^n_{jc})w^{n*}_{jc}$ is the observed mean nonunion wage. Note that the union and nonunion wage equations have different error terms, with the nonunion error term consisting of the productivity shock, ϵ_{ic}, while the union error term includes the productivity shock but also the unobserved (to the econometrician) value of union amenities, ψ_{icf}. Given that we have assumed, so far, that workers are identical, higher wages in an industry correspond to rents – differences in pay over and above what is required for the marginal worker to want to join that industry. Those rents are maintained because of the frictions in the labour market. It is important that we are considering rents since wage differences across industries that correspond to compensating differentials (say, for dangerous work) cannot be the basis of bargaining a higher wage with your current employer. If a higher wage corresponds to a compensating differential then there is no net gain to moving to the dangerous job and, so, no basis on which to threaten your current employer during bargaining.

2.5 Union Determination

Our theoretical analysis occurs at the level of the firm and, therefore, we are interested in the question of whether firms become unionised. We assume there is no decertification, so unionisation is an absorbing state for a firm.[15] But all firms – union and nonunion – die with probability δ_d in a period (which is a component of the separation rate, δ, seen earlier). The dying firms are replaced with new firms started by new entrepreneurs. In steady state, the number of dying firms equals the number of newly born firms. All firms are born nonunion and then workers at the firm decide whether to unionise. Recall that firms are born with a draw of a level of amenities that would be provided at that firm if it were unionised, ψ_{icf}.

2.5.1 Union Determination Without Firm Responses

We begin with a simple model in which firms are passive players in unionisation. That is, unionisation is determined entirely by the workers and firms do not try to respond by, for example, emulating union wages. In terms of the model derived so far, this implies that the probability of meeting an emulating firm $(1 - p^n_{ic})$ is zero and the outside option wage terms are adjusted accordingly.

In the simple model, the workers at a firm compare the value of the job continuing as a nonunion job to the value of the job being union minus the cost of unionising. The wages they use in this exercise are the ones we arrived at in the previous section that reflect the optimal hiring decisions by firms. We will assume that the decision is made according to a median voter model with the median voter not at risk of losing her job when employment is reduced after unionisation. We also assume that workers do not care about the employment outcomes of those who do lose their jobs, implying that we can focus exclusively on wages. Given this, a firm becomes unionised if:

$$W^u_{icf}(w^u_{icf}) - \lambda^*_{ct} > W^n_{icf}(w^n_{icf}),$$

[15]Decertification is much less common than certification. Using election data from the NLRB (discussed below) we find that certification elections outnumber decertification elections over 1980-2010 by at least 5 to 1. The same is true if we consider the number of workers involved in elections. Fortin et al. (2019) present a figure showing the ratio of eligible workers certified over the time horizon 1978-2017.

where λ^*_{ct} is the fixed cost to workers of unionising a firm in a city c. Substituting in steady state expressions for the value functions based on (2) and (1), we can define an index function:

$$I_{icf} = (w^u_{icf} - w^n_{icf}) + \psi_{icf} - (1 - \rho(1-\delta))\lambda^*_{ct}, \tag{9}$$

such that a firm is unionised if $I_{icf} > 0$ and remains nonunion otherwise. Here, we have assumed that all workers in the firm share the costs of unionisation equally. The term multiplying the fixed cost per worker puts the one-time fixed cost of unionising in the same present value terms as the flow of wages and union amenities.

We can substitute in the union and nonunion linearized wage expressions, (6) and (5), into (9) to obtain:

$$I_{icf} = \alpha_{0i} + \gamma^u_1 E^u_{ic} - \gamma^n_1 E^n_{ic} + \alpha_2 ER_c + (1 - \gamma^u_3)\psi_{icf} - \lambda_{ct} + \alpha_4 \epsilon_{ic}, \tag{10}$$

where $\lambda_{ct} = (1 - \rho(1-\delta))\lambda^*_{ct}$ and $\alpha_4 \geq 0$.[16]

This is a very standard selection set-up and implies:

$$E(w^n_{icf}|I_{icf} \leq 0) = \gamma^n_{0i} + \gamma^n_1 E^n_{ic} + \gamma^n_2 ER_c + \gamma^n_4 E(\epsilon_{ic}|I_{icf} \leq 0). \tag{11}$$

Notice that the error-mean term is a function of λ_{ct} – the cost of unionising – but the unconditional nonunion wage equation is not. Thus, measures of the cost of unionising are available as exclusion restrictions that identify selection effects separately from direct determinants of the nonunion wage. If we consider two cities, c and c', that are identical except that c' has higher costs of unionisation then unionisation will be lower in c'. Moreover, because the coefficient on the productivity term, ϵ_{ic}, is positive in the index function and recalling that ψ_{icf} and ϵ_{ic} are assumed to be independent, union firms tend to have higher productivity. Thus, the marginal firms that would be unionised in c but non-union in c' will be at the low end of the productivity range for union firms but the high end for nonunion firms. This has implications for a simple specification using the proportion union to capture spillover effects, as the estimated coefficient on the union proportion would be biased downward. Industry-city cells with higher unionisation rates would be ones with lower productivity among nonunion firms.

In what follows, we first estimate the regression specification given by (11). This allows us to investigate the presence and size of spillovers of union power on nonunion wages. We do not attempt to estimate a specification for local union wages because low unionisation rates, especially in the later years of our sample, imply sample sizes in industry $\times$ city cells that are too small to work with.

2.5.2 Incorporating Firm Responses

Next, we consider a more realistic setting in which firms at risk of being unionised can respond to forestall unionisation. In TD, firms respond by hiring more skilled workers who they know will vote against unionisation. While this response is possible, it seems to us to likely be of second order importance relative to more direct responses. In real world

[16] This condition on α_4 is shown in Appendix A.4 and arises because nonunion wages are determined by the marginal product of one worker while union wages relate to average product of all workers.

descriptions of firm reactions to the threat of unionisation,[17] firms adopt some combination of three possible responses: 1) paying higher wages to reduce the net benefit of unionising; 2) providing better working conditions to match what workers would get in a union setting (in the context of our model where only union workers get the amenities, ψ_{icf}, nonunion firms would provide similar amenities to their workers); and 3) intimidation (which, in our model, would correspond to increasing the cost of unionisation, λ_c). We begin by focusing on the wage emulation response, returning to the other potential responses later.

Before discussing the responses, note that firms do not need to consider employing any response if their workers are happily nonunion. That is, if the costs of unionising, the amenities that would be available to the workers at this firm if they unionise, the union wage they would get if they organise, and the nonunion wage they get if they don't are such that $I_{icf} < 0$ then there is no reason for the firm to bear costs to incentivize its workers not to form a union. We are interested in the set of firms that are unionised (i.e., for which $I_{icf} > 0$) but near the margin of being so in our first, non-responsive firm model. Recall that the fixed cost of forming a union depends on the legal climate in each state while the non-wage benefits of unionising are firm specific. As a result, not all firms in the same industry and city will have the same worker preferences about unionisation.

Wage Response First, consider the possibility that firms respond to their workers' desire to unionise by offering the workers a wage that just offsets the benefit of unionising. In particular, based on the discussion of worker preferences on whether to unionise underlying the index function (9), the wage the firm would need to offer to make workers indifferent about forming a union is:

$$w^*_{icf} = w^u_{icf} - \lambda_c + \psi_{icf}.$$

Now consider worker preferences about unionising when they take account of the non-emulation nonunion wage we derived earlier, w^n_{icf}, the cost of unionising and the union wage. In this situation, we can define a $\psi^*_{icf} = \lambda_c - (w^u_{icf} - w^n_{icf})$: the value for amenities such that workers at firm f are indifferent between whether they organise or not. For $\psi \leq \psi^*_{icf}$, workers will not organise, and the firm will pay the regular nonunion wage, w^n_{icf}. For $\psi > \psi^*$, workers will prefer to unionise if offered the regular nonunion wage and firms will consider the value of offering the emulation wage, w^*_{icf}, instead and forestalling unionisation. We show in Appendix A.5 that the firm will be willing to pay the higher, emulation wage until the point where the costs and benefits to unionisation overlap, which happens when $\psi_{icf} = \lambda_c$. At that point, the emulation wage equals the union wage. Beyond it, the wage required to prevent union formation exceeds the union wage, and to fight the union would lower firm profits. For all $\psi_{icf} > \lambda_c$, then, the firm will become unionised. Thus, we can characterize the firm union status as follows:

- $\psi_{icf} < \psi^*_{icf}$: f is nonunion and pays w^n_{icf},
- $\psi^*_{icf} \leq \psi_{icf} < \lambda_c$: f is nonunion but emulates unionised firms and pays w^*_{icf},

[17] See The Guardian: 'Pay a living wage': Bernie Sanders accuses Disney of dodging fair pay, The Guardian: Why Target's anti-union video is no joke, and The Guardian: Delta workers seeking to unionize say they are 'under siege' by management.

- $\psi_{icf} \geq \lambda_c$: f is union and pays w^u_{icf}.

Note that both cut-offs rise with λ_c so that states that implement policies that raise the cost of unionising will have more nonunion firms. As the unionising costs rise, the actual nonunion wage will not change. However, the observed nonunion wage will decline because the fraction of nonunion workers who are paid the higher, emulation wage, will decline.[18] Moreover, the emulation wage w^*_{icf}, will also decline. Thus, we would observe a decline in unionisation combined with an increase in the observed union wage differential. This pattern might seem to imply that de-unionisation is arising because of union rigidity on wages. But this can happen even without unions being rigid (i.e., even though union wages will decline with declines in ϵ). The pattern of declining unionisation with an increasing union wage differential is what we observe during the 1980s.

To obtain an empirical specification related to the model including union emulation, first note that the linearized version of the union emulation wage is given by:

$$w^*_{icf} = \gamma^u_0 + \gamma^u_1 \tilde{E}^u_{ict} + \gamma^u_2 ER_c + (1-\gamma^u_3)\psi_{icf} - \lambda_{ct} + \gamma^u_4 \epsilon_{ic}/ \tag{12}$$

There is one additional complication that we have yet to address, which is that for certain values of ϵ there may be no range over which workers wish to unionise. In terms of the thresholds defined above this occurs when $\psi^* = \lambda_c - (w^u - w^n) \geq \lambda_c$, which occurs when the nonunion wage exceeds the union wage. In this instance, for all values of ψ, workers will prefer to remain nonunionised. We define the productivity draw that equalises the costs and benefits of unionisation as ϵ^*. For productivity draws above this threshold the union wage will exceed the nonunion wage and there will be some set of ψ values over which unionisation is preferred.

The observed mean nonunion wage in an *ic* cell is given by a weighted average of the actual nonunion wage, and the emulation wage. We define this observed wage as $\bar{w}^n_{icf}$:

$$\begin{aligned} E(\bar{w}^n_{ic}) =& \frac{Pr(I_{icf} < 0)}{Pr(I_{icf} < 0) + Pr(I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*))} E(w^n_{icf} | I_{icf} < 0) \\ &+ \frac{Pr(I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*)}{Pr(I_{icf} < 0) + Pr(I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*))} E(w^*_{icf} | I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*) \end{aligned} \tag{13}$$

The weights are the probability of firms being of each type conditional on being observed as a nonunion firm. To form this regression equation, we first need to specify the probabilities that make up the weights:

$$Pr(I_{icf} < 0) = \int_0^\infty \int_{-\infty}^{\frac{\lambda_c - \Delta - \alpha_4 \epsilon}{(1-\gamma_3)}} f(\psi) g(\epsilon) d\psi d\epsilon \tag{14}$$

and

$$Pr(I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*) = \int_{\epsilon *}^\infty \int_{\frac{\lambda_c - \Delta_{ic} - \alpha_4 \epsilon}{(1-\gamma_3)}}^{\lambda_c} f(\psi) g(\epsilon) d\psi d\epsilon, \tag{15}$$

[18] To see this note that the cut-off for the lower end of the emulation range, $\psi_u = \lambda_{ct} - (w^u - w^n)$, will rise faster than the upper end (which is λ_{ct}) as λ increases. This can be shown by noting that $\frac{\partial \psi_u}{\partial \lambda_{ct}} = 1 - \frac{\partial w^u}{\partial \lambda_{ct}}$ and that $\frac{\partial w^u}{\partial \lambda_{ct}} < 0$ since union wages are lower when union related amenities are higher.

where $\Delta = \alpha_{0i} + \gamma_1^u E_{ic}^u(w) - \gamma_1^n E_{ic}^n(w) + \alpha_2 ER_c$, the union benefit threshold is $\psi^* = \frac{\lambda_c - \Delta - \alpha_4 \epsilon}{(1-\gamma_3)}$, and recalling that the range for the productivity shock, ϵ_{ic} is $[0, \infty]$. The first probability is the probability that workers would choose to be nonunion and the second is the probability that firms are nonunion but emulate union wages.

Substituting equation (11) for $E(w_{icf}^n | I_{icf} < 0)$ and the relevant expectation over (12) for $E(w_{icf}^* | I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*)$ in equation (13), we arrive at an expression for the expected nonunion wage that incorporates firm emulation responses:

$$E(\bar{w}_{ic}^n) = \gamma_{0i}^{n*} + \gamma_1^n (1 - P(\psi_{icf}, \lambda_c)) E_{ic}^n + \gamma_1^u P(\psi_{icf}, \lambda_c) E_{ic}^u + \gamma_2^* ER_c + \mu(\epsilon_{ic}, \psi_{icf}, \lambda_c), \quad (16)$$

where $P(\psi_{icf}, \lambda_c)$ is the probability of a nonunion firm being a wage emulator given in (15).[19]

We implement both the specification assuming no firm responses, equation (11), and the one incorporating wage emulation, equation (16), in the empirical work that follows. Note that the emulation specification differs from the no-response specification, in part, because of the inclusion of the value of outside options for union workers. The latter is based on the probability of a union worker finding a union job and should not determine nonunion wages in the absence of the standard threat effect. In a separate estimation not reported here, we estimated a linearized version of (16) with E_{ic}^n, E_{ic}^u, $P(\psi_{icf}, \lambda_c)$, and ER_c as regressors. The fact that E_{ic}^u entered significantly in that estimation is strong evidence in favour of the standard threat effect being real.

Amenity and Intimidation Responses The other ways firms can respond to a unionisation threat are to provide amenities that match what unions provide and to raise the costs of unionising. We discuss both in more detail in Appendix A.5. For amenities, we describe a situation in which providing amenities has an increasing and convex marginal cost function with the marginal cost of the initial units provided being below a dollar for one dollar worth of amenities as valued by the worker. In that case, nonunion firms would want to use amenities as a response to a union threat until the point where the marginal cost of a dollar's worth of amenities rises to one dollar. After that, they would respond through wage emulation. However, if it is cost effective for a nonunion firm to use amenities to respond to a union threat, it would also be cost effective for it to pay in amenities instead of wages even in the absence of such a threat. Thus, if a threat emerges, the nonunion firm will already be providing amenities up to the point where their marginal cost equals a dollar. In that case, there is no room for the firm to respond to a union threat using amenities. Instead, it will respond through wage emulation.

Firms could also respond to a unionisation threat by increasing the cost of unionising, λ_c, at a cost to themselves. For example, they could lock out the workers and either not produce or hire scabs who are less productive than the actual workers. In the Appendix A.5, we set out the value function for a firm that employs intimidation and compare it to the value function when the firm chooses the wage emulation response. We show that, intimidation is a more cost effective response when workers are close to indifferent about unionising but becomes less effective when the level of intimidation required is large (i.e., as the value of ψ rises). We show that under reasonable assumptions there is a new cut-off value, ψ_{icf}^b

[19] $\mu(\epsilon_{ic}, \psi_{icf}, \lambda_c)$ is an expected error mean term equal to $\gamma_4^n E(\epsilon_{ic} | I_{icf} < 0) - \lambda_c + E(\gamma_4^u \epsilon_{ic} + (1 - \gamma_3^u)\psi_{icf} | I_{icf} > 0, \psi_{icf} < \lambda_c, \epsilon > \epsilon^*)$.

such that, $\psi^*_{icf} < \psi^b_{icf} < \lambda_c$. Firms with $\psi^*_{icf} < \psi_{icf} \leq \psi^b_{icf}$ will use intimidation and, as a result, will remain nonunion and pay the nonunion wage. Firms with $\psi^b_{icf} < \psi_{icf} \leq \lambda_c$ will be nonunion but pay the emulation wage. As before, firms with ψ_{icf} below ψ^*_{icf} will be nonunion, paying the nonunion wage, and firms with ψ_{icf} above λ_c will be unionised. Thus, introducing intimidation serves to expand the region over which unions are nonunion and pay the simple nonunion wage to include the range over which firms use intimidation. We do not have a way to separately identify ψ^b_{icf} from ψ^*_{icf} and so cannot distinguish between a model in which there is no intimidation and the relevant cut-off for the nonunion region is ψ^*_{icf} and one in which there is intimidation and the relevant cut-off is ψ^b_{icf}. Since both cut-offs are functions of the same variables – λ_c, the expected rents, and the employment rate – there is essentially no impact on our empirical specifications of including or not including intimidation. We will proceed as if there is no intimidation in order to simplify the exposition.

3 Implementation of the Wage Specifications

In this section we present details and results regarding estimation of the nonunion wage equations (11) and (16) derived from the model presented in Section 2.5.

3.1 Data

In our analysis we use data from the Current Population Survey Merged Outgoing Rotation Groups for 1983-2020 and the CPS May extracts for 1978-1982. We are interested in comparisons across steady states over a medium-long time horizon and, as such, we consider variation over 10 year periods. For each time period, we pool observations across 3 years to reduce statistical noise. We consider variation across 1980, 1990, 2000, 2010, and 2020 using the years 1978-1980, 1988-1990, 1998-2000, 2008-2010, and 2018-2020.

From this data, we keep all workers between the ages of 25-65 who do not report being in school either full-time or part-time. We follow Lemieux (2006) in the construction of our wage data, working with weekly wages. We use an aggregated grouping of industry codes based on the 1980 industrial classification from the Census Bureau. We obtain a consistent industry classification using crosswalks provided by IPUMS and the Census Bureau that map the 1970, 1990, and 2000 industry codes to the 1980 classification. The result is a consistent classification system with 51 industries. Appendix B contains additional processing details.

We construct a set of cities with as consistent geographic boundaries as possible given data limitations in the CPS. We are constrained by the number of SMSA's available in the May extract data and end up with 43 cities. Making use of the limited number of counties identified in the CPS, we are able to create a set of cities which are reasonably, though not always perfectly, consistent over time.[20] The final geographic definition we use pools data for these 43 cities and the remaining population. Specifically, we create additional regions made up of the remaining state population absent the population living in these 43 cities.

[20]The metropolitan area definition used by the IPUMS identifies a general pattern of expanding metropolitan area definitions over time that we overcome to some extent, but not perfectly: `https://usa.ipums.org/usa/volii/county_comp2b.shtml`. Estimation using states as the geographic unit yields very similar results, suggesting that issues related to geographic definitions are not driving our results.

In the end, our core geographic measure is composed of 93 areas that are fairly consistently defined over the course of the sample period.

Additionally, we use data on union elections to proxy for the costs of unionisation, λ_{ct}, in our model. The idea is that locations where the proportion of union certification elections that result in a certification is high are more union friendly. To obtain these proportions we use National Labor Relations Review Board (NLRB) case data for the three year periods for which we use CPS data.[21] We focus on certification elections and cases where a conclusive decision on certification was reached.[22] We use the county of the unit involved in the election to construct our geographic measures, aggregating counties to our city definition discussed above. We also use this data to construct expected probabilities of a firm facing a unionisation election. Unfortunately, the elections data ends before 2020 and so we estimate the full model over the years 1980, 1990, 2000 and 2010.

3.2 Empirical Implementation with Worker Heterogeneity

In order to take the model to the data, we must confront the fact that, while workers are homogeneous in our model, they are not in our data. Our approach is to treat individuals as representing different bundles of efficiency units of work and to assume those bundles are perfect substitutes in production. We then interpret w^n_{ict} in equation (5) as the cost per effective labour unit. Let effective labour units be $\exp(H_l'\beta_t + a_l)$, where H_l and a_l capture observable and unobservable skills of worker l, respectively. Adding industry, city and time subscripts, workers' observed nonuion log wages, $\ln w^{on}_{lict}$, are given by:

$$\ln w^{on}_{lict} = H_{lt}'\beta_t + \ln w^n_{ict} + a_{lt}. \tag{17}$$

The values of $\ln w^n_{ict}$ are our object of interest. To obtain a measure of these, we estimate (17) capturing $\ln w^n_{ict}$ as the coefficients on a complete set of industry×city fixed effects. Our specification of H_l includes a complete interaction of dummies for educational attainment, a quadratic in potential experience, and gender and race dummy variables. We estimate (17) using only nonunion workers, separately by year. This allows for flexible changes over time in the returns to education, etc.. The estimated vector of coefficients on the industry-city fixed-effects are regression-adjusted, average local industry wages, and we use these coefficients as our dependent variable in the regressions (11) and (16).[23] This procedure removes skill and demographic variation from our wage measure. We also form an analogous regression adjusted union wage measure, $\ln w^u_{ict}$, using only data on unionised workers.

3.3 The Outside Option Terms

Central to our empirical work are the outside option terms characterising alternative job prospects in either the union or nonunion sectors. As defined above, these terms are composed of the rents a worker would get in expectation when searching for a new job ($\sum_i \eta^u_{ic} w^u_{ic}$

[21]Our thanks to Hank Farber for providing this data.

[22]As opposed to the case being dismissed or withdrawn.

[23]As described in section 3.5, we work with an adjusted version of our wage specification in order to move to working with mean log wages.

for union jobs and $\sum_i \eta_{ic}^n w_{ic}^n$ for nonunion jobs) and the probability of finding work in the union sector (T_{jc}^n, for a worker formerly in a nonunion job in sector j). We use our regression adjusted wages in order get as close as possible to rents rather than skill differentials since wage differences that reflect skill differentials cannot be used as an outside option in bargaining (a janitor cannot use the opening of new jobs for lawyers in town to bargain a better wage).

In our model we assume that the relative likelihood of finding work in the union sector differs by city, previous industry of work, and whether the worker was previously unionised. Using matched CPS data, sample sizes are not sufficient to estimate a fully flexible characterisation of transition paths in this manner. However, working at the national level, we can construct T_{it}^n – the probability that a worker observed in a nonunion job in industry, i, in year, t, who separates from that job in the following year is observed in a union job (in any industry) in year $t+1$. We construct the same probabilities for workers starting in union jobs, creating all of these transition rates separately for each of our sample periods.[24] We then combine these rates with the proportion of jobs in the city that are unionised, P_c, in order to capture local variability in finding unionised jobs.

More specifically, we construct our transition rate measures as:

$$T_{ic}^u = \frac{T_i^u P_c}{T_i^u P_c + (1 - T_i^u)(1 - P_c)} \text{ and } T_{ic}^n = \frac{T_i^n P_c}{T_i^n P_c + (1 - T_i^n)(1 - P_c)}, \tag{18}$$

This is a simplified version of similar measures constructed by Tschopp (2017) who uses rich data to calculate transitions between industries. As in Tschopp (2017), we interpret this as a measure of relative mobility into the union sector.[25] Thus, for a construction worker in Detroit, for example, it measures the relative likelihood of matching with a union job (in any industry) compared to moving into either a union or nonunion job (in any industry). If there are many union jobs in Detroit, and construction workers tend to transition into unionised employment, then these workers will have a higher relative mobility measure of finding unionised employment compared to workers in cities with lower rates of unionisation and who work in industries with relative low transition rates into unionised jobs.

Working with the wage rent variables and the transition rates, we form our measure of the outside option value as:

$$E_{jc}^n = (1 - T_{jc}^n)\bar{w}_c^n + T_{jc}^n \bar{w}_c^u, \tag{19}$$

where $\bar{w}_c^n = \sum_i \eta_{ic}^n w_{ic}^n$ is the mean, residualized nonunion wage in the city and $\bar{w}_c^u$ is the mean, residualized union wage. Thus, the outside option for a nonunion worker in industry j in city c is the sum of the probability the person gets a nonunion job times the expected rent from a nonunion job and the probability they get a union job times the expected rent for those jobs.

[24] Our framework assumes that bargaining effects operate only through the unemployment channel, that is, workers must first transition through unemployment to access other jobs. However, due to data limitations, our transition measures use transitions between sectors which may, or may not, have included an intervening unemployment spell. Thus, the union outside option term may reflect on-the-job search dynamics. Formally modelling on-the-job search, or job laddering is beyond the scope of this paper. As noted by Beaudry et al. (2012) it is not straightforward and is sensitive to modelling of the search process and its relationship to wage determination.

[25] Tschopp's specification would take the same form if the economy was composed of two sectors.

Changes in the outside option are driven by five factors. The first is changes in the local composition of nonunion jobs as captured in the employment shares, i.e., the η^n_{ic}'s. In essence, if high-paying jobs such as those in the steel industry are replaced with lower paying service sector jobs then the outside option for all workers in the city is reduced. The second factor is the wage rents in the nonunion sector: even if there is no shift in the industrial composition of nonunion jobs, if the steel industry stops paying higher wages then it no longer offers an attractive outside option for workers in other industries. The third factor is the probability the nonunion worker can get a union job. If union jobs pay, on average, higher rents then a decrease in the probability of getting a union job means lower access to those rents and, therefore, a less valuable outside option. The value of the union option is also altered if there is a shift in the composition of union jobs (the η^u_{ic}'s) or the wages in the union sector, which are the fourth and fifth factors.

3.4 Constructing Emulation Probabilities

Since the emulation probabilities are unobserved in our data, we construct a proxy. To do so, we use information from the NLRB and the County Business Patters (CBP). The idea is to proxy the threat of emulation, $P(\psi_{icf}, \lambda_{ct})$, by the likelihood of that a firm in a given ic cell faces a union election, taking into account the number and composition of establishments at the city-industry level. To this end, our proxy is constructed by first calculating the ratio of the predicted number of elections to the number of establishments in an ic cell. We use this ratio to calculate the probability that an establishment had least one election in the past four years. The predicted number of elections come from a negative binomial regression, where the observed number of elections are regressed on a number of variables at the ic and city level. These variables include polynomials in the number of workers and the size of establishments in ic along with the unemployment rate, the participation rate, and the average age at the c level plus year- and city-fixed effects. The predicted number of elections come from the fitted values of the model. The details of our procedure are contained in Appendix B.1.

3.5 Dealing with Endogeneity

We will estimate our derived specification in first differences in order to eliminate any industry×city-time invariant characteristics. Ignoring selection issues for the moment, this means in the case without firm responses we consider the regression:

$$\Delta \ln w^n_{ict} = \gamma^n_{0it} + \gamma^n_1 \Delta E^n_{ict} + \gamma^n_2 \Delta ER_{ct} + \gamma^n_4 \Delta \epsilon_{ict}. \tag{20}$$

where we work with an adjusted version of equation (5) in which we average across firms, add a time subscript, and divide both sides by a base wage, w_0, in order to have a dependent variable in log differences.[26] Given that this specification includes a complete set of industry×time-period fixed effects, the relevant identifying variation for the estimated coefficients comes from across-city within-industry variation. Intuitively, this means that we

[26] As described in section 3.2, we create our dependent variable as residualized mean wages in an ic cell from log wage regressions. We work with log wages in order to match the human capital literature and to make our estimated coefficients easier to interpret.

identify the impact of outside options by comparing wage changes in the same industry in two different cities that are experiencing different changes in the quality of outside employment prospects, holding the employment rate constant. Given the inherent reflection problem, due to the fact that E^n_{ict} includes local, mean nonunion wages, we would not expect OLS estimation to yield consistent estimates of γ^n_1. To eliminate this mechanical correlation and to address other potential identification threats, we utilize an instrumental variables strategy as follows.

We begin the construction of our instruments by replacing local wages in E^n_{ict} with what we refer to as industry-level rents. In particular, we estimate separate wage regressions for each of our set of sample years at the national level, working with pooled union and nonunion workers. The regressions include the same set of skill and demographic variables used when forming our residualized wages for the dependent variable plus a complete set of industry dummy variables interacted with a union dummy. We interpret the coefficients on the industry dummies as rents that are allowed to differ in the union and nonunion sectors.[27] Using them, we form:

$$\tilde{E}^n_{jct} = (1 - T^n_{jct}) \sum_i \eta^n_{ict} \nu^n_{it} + T^n_{jct} \sum_i \eta^u_{ict} \nu^u_{it}, \tag{21}$$

where ν^n_{it} is the wage premium for nonunion workers in industry i. Because the premia are at the national level, they break the direct reflection link. The result is an outside option variable expressed such that workers in cities with a concentration in industries that pay high rents at the national level are able to bargain high wages.

In order to use this version of outside option values as the basis for an instrumental variables strategy, we consider changes in each of the right-hand-side terms. The change in the first term equals:

$$\Delta \left((1 - T^n_{jct}) \sum_i \eta^n_{ict} \nu^n_{it} \right) = \left((1 - T^n_{jct}) \sum_i \eta^n_{ict} \nu^n_{it} \right) - \left((1 - T^n_{jct-1}) \sum_i \eta^n_{ict-1} \nu^n_{it-1} \right).$$

Our concern is that movements in the η_{ict}'s and T^n_{jct} could be correlated with the productivity changes in the error term since sectors with greater increases in productivity will increase their share of employment. Following the standard Bartik approach, we construct predicted values of each that depend on the start of period industrial composition of the city combined with national level changes. That is, we define an instrument:

$$IV1^n_{jct} = \left((1 - \hat{T}^n_{jct}) \sum_i \hat{\eta}^n_{ict} \nu^n_{it} \right) - \left((1 - T^n_{jct-1}) \sum_i \eta^n_{ict-1} \nu^n_{it-1} \right),$$

where $\hat{\eta}^n_{ict}$ corresponds to predicted values of end of period industrial shares in city c for nonunion jobs, and $\hat{T}^n_{jct}$ is the predicted end of period probability of a nonunion worker finding a union job. We define a second instrument, IV^u_{jct}, corresponding to the second term in (21) analogously. We construct predicted employment levels using start-of-period levels

[27] We define the industry dummy variables such that the coefficient values are defined relative to the overall average wage.

at the industry-city level combined with national level growth rates for the relevant industry. Likewise, we combine national level predictions in employment growth with local start-of-period industrial-union composition to form $\hat{T}^n_{jct}$, which captures the idea that if there are declines in union employment in sectors with high local employment, then this will predict a decline in the local union proportion. We spell out the details in the construction of $\hat{\eta}^n_{ict}$ and $\hat{T}^n_{jct}$ in Appendix B.3

To understand the conditions under which $IV1^n_{jct}$ and $IV1^u_{jct}$ are valid instruments, recall that we include a complete set of time×industry effects and, so, we are working with within-industry, cross-city variation. Note also that the variation in the instrument comes from start of period, cross-city differences in the industrial proportions (the η_{ict}'s) and the start of period, cross-city differences in the relative shares of union employment in industries (as reflected in $\hat{P}_{ct}$). Validity of the instrument requires that these are uncorrelated with the relevant variation in the error term: cross-city variation in productivity growth. That is, we require an assumptions that the productivity process follows a random walk (since, as BGS show, the η_{ict}s can be written as functions of the ϵ_{ict}s). We can assess this assumption using an over-identification test which we discuss when we present our results.[28]

For reasons that will become apparent when we discuss addressing selectivity, we are also interested in instruments that do not rely on the predicted change in P_{ct}. Thus, we form a second set of instruments given by:

$$IV2^n_{jct} = \left((1 - T^n_{jct-1}) \sum_i \hat{\eta}^n_{ict} \nu^n_{it} \right) - \left((1 - T^n_{jct-1}) \sum_i \eta^n_{ict-1} \nu^n_{it-1} \right),$$

with an analogous expression for $IV2^u_{jct}$. Notice that these instruments rely on the $\hat{\eta}_{ict}$'s but not on $\hat{P}_{ct}$. Because the outside option that a worker uses in bargaining depends on the expected value of wages in other jobs, not on the component parts of that expectation, the two instrument sets should yield the same estimates. That is, it doesn't matter whether a worker's outside option worsens because a high rent paying industry leaves town (as emphasized in the difference between η^n_{ict-1} and $\hat{\eta}^n_{ict}$) or because of de-unionisation that implies a loss of access to union rents (as captured in the difference between T_{jct-1} and T_{jct}). $IV2^n_{jct}$ uses only the industry composition change variation while $IV^n_{jct}1$ uses both and, theoretically, they should both provide the same answer even though part of the basis of their variation is quite different.

The validity of our instruments relies on the exogeneity of the start of period industry shares and union shares. Our theory implies over-identifying restrictions that allow us to test that exogeneity which we present in the Results section. But we can also follow the advice in **?** about checking patterns and correlations for further, suggestive evidence that the exogeneity requirements are met in our case. These checks are weakened to some extent by the fact that our situation differs from the classic Bartik case because our key endogenous variables (the outside option terms) vary at the ic level rather than just the c level. They also include different national level variables in the same expression, requiring us to make

[28] As **?** and BGS point out, Bartik instruments are functions of the start of period values for the η_{ict}'s – the local industrial composition – and any combination of those values can be used as an instrument. BGS argue that in our case one can find specific combinations within the theory by examining decompositions of the outside option variables that both have intuitive appeal and imply testable over-identifying restrictions.

restrictive assumptions (such as that the industry wage premia are the same in the union and nonunion sectors) in order to fit into the **?** framework. When we do that, the Rotemberg weights for each of our four instruments – weights showing which industries are the main drivers behind the variation in our instruments – point to the top five weighted industries including: mining; motor vehicles and equipment; retail trade; construction; and lumber and wood products. Apart from retail trade, this list is reassuring because it consists of sectors with high wage premia that, at least at one time, were highly unionised. Thus, they seem like a good set of industries for identifying the impacts of variation in access to high rent jobs. **?** also suggest looking for correlates of the baseline industry proportions on which our instruments are built to see if those suggest possible issues. In our case, we would be worried about correlations with variables that might predict growth in city-level productivity. Given that we control for education, age, gender and the employment rate, we do not have candidates for other variables that could fit this bill.

4 Results

4.1 Descriptive Patterns

Before turning to estimation we first present key patterns in unionisation over our sample period. As is well known, the decline of unionisation in the United States over 1980-2019 (and for other rich world nations over a similar time frame, see Schmitt and Mitukiewicz (2012) and Lesch (2004)), has been remarkable.

In Figure 1 we plot the fraction of workers unionised at the city level over 1980-2019 for each city, highlighting a subset of cities with particularly large or small declines in unionisation. We also highlight the national average (the thick black, dashed line) in the figure. On average, about 25% of jobs were unionised at the city level in 1980, but this number declines to 17% by 1990 and then to 13% by 2019. In cities, like Detroit, Gary, and Pittsburgh, where the union sector played a much larger role in the 1980 economy, the declines are substantial: respectively 21, 29, and 22 percentage points by 2019. Smaller declines (under 10%) are observed in cities with low initial rates of unionisation, such as Dallas, Washington, and Rochester. Thus, there is a considerable range in the changes in unionisation across cities and, importantly, there is variation in the decade in which the declines occur. This will allow us to separate the effects of union declines from general trends.

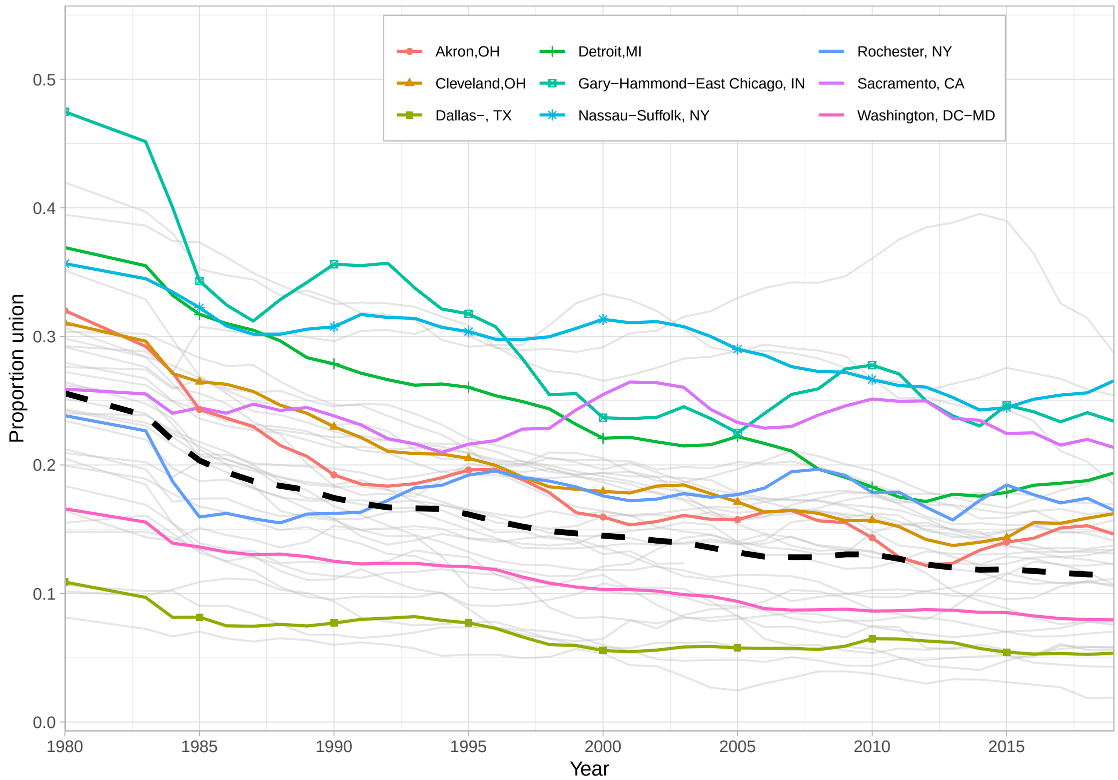


Figure 1: Union percentage over time

Notes: Data comes from the CPS. The figure denotes the fraction of unionised workers at the the city level over the 1980-2019 time period calculated as a three-year moving average. The dashed black line refers to the level of unionisation at the national level.

Observed declines in rates of unionisation are large, which naturally has implications for our outside option terms. In particular, as discussed in Section 3.3, outside options depend on the probability of a nonunion worker getting a union job, T^n_{ict}, and the average wage in the union sector, $\bar{w}^u_{ct}$. Similarly, as union rates fall, so does the threat of unionisation, through $P(\psi_{icf}, \lambda_{ct})$. The theoretical framework we present above suggests that shifts in these terms will impact nonunion wages through both bargaining and threat channels.

As a first step in establishing the relevance of these channels, Figure 2 plots the 2010-1980, city-level change in log nonunion wages against: (a) changes in nonunion-to-union transition rates, T^n_{ct}; (b) changes in average wages in the union sector, $\bar{w}^u_c$; and (c) changes in the probability of a firm facing a union certification election, $P(\psi_{icf}, \lambda_{ct})$. Each of the three panels shows a strong, positive association between changes in the nonunion wage and changes in the outside option and threat components.

In our empirical work, we attempt to separately identify spillover effects from bargaining and threat channels. To do so, we require that cities experience different declines in nonunion-to-union transition rates (that help identify bargaining effects) and union election rates (that help in identifying threat effects). Figure 2 highlights several cities to emphasise this variation in the data. Cleveland and Pittsburgh, for instance, both faced relatively large declines in transition rates (panel A), but very different change in election probabilities (panel C). Likewise, Sacramento experienced large declines in transition rates compared to Dallas, but have similar changes in election probabilities. Overall, the variability across cities in movements in outside option and threat components will help separately identify the bargaining effect from the threat effect.

4.2 Estimation Results

We now turn to estimation results based on the models presented in section 2, beginning with the simple specification (11) and then moving to the specification that includes emulation effects (16). As described in Section 3.2, we work in ten year differences and use the residualized wage as the dependent variable.

4.2.1 Simple Specification

Our simple specification relates changes in nonunion wages to changes in the values of outside options for nonunion workers and the employment rate. It is useful, however, to begin with a more standard specification in order to see how our results relate to the existing literature. In papers including Freeman and Medoff (1981), Holzer (1982) and Hirsch and Neufeld (1987), the main regression takes the form:[29]

$$\ln w^n_{icf} = a_o + a_1 P_{ic} + a_2 x_{icf} + u_{icf} \tag{22}$$

where, P_{ic} is the proportion of workers in the i-c cell who are unionised[30], x_{icf} is a vector of other controls, a_2 is a parameter vector of the same length as x_{icf}, and u_{icf} is an error term. In earlier papers, the x_{icf} vector typically includes local demand and supply shifters, such as the proportion of teenagers in the region, the unemployment rate, local lagged per capita income growth, and average firm size, amongst others. Additionally, more recent papers by Neumark and Wachter (1995) and Farber (2005) control for full industry-year and/or city-year fixed effects in order to better account for relevant omitted variables.

There are strong similarities between the specification given by equation (22) and our specification (given in equation (11)). In particular, previous studies have included controls similar to the employment rate and the industry-year effects that are in equation (11). The main difference is that union effects are represented by the simple proportion union variable in equation (22) whereas in our specification, theory indicates that they should be part of the outside option term, interacting with changes in union wage premia. The specification does not contain the union proportion on its own.

In the first column of Table 1, we present the results from estimating the standard specification, (22), using our data. Given the way we created the dependent variable, we are controlling for education, age, gender and race in a flexible way. We do not include industry controls, in order to match the older part of the literature. We include changes in the union proportion in an industry-city cell as our core explanatory variable, finding a positive and statistically significant relationship between growth in nonunion wages and growth in unionisation that is broadly similar to previous estimates. Hirsch and Neufeld (1987) find

[29] Or is pooled for union and nonunion workers, with the proportion union interacted with a union dummy.
[30] Typically this metric is calculated either at the national-industry level, or at the city/state level.

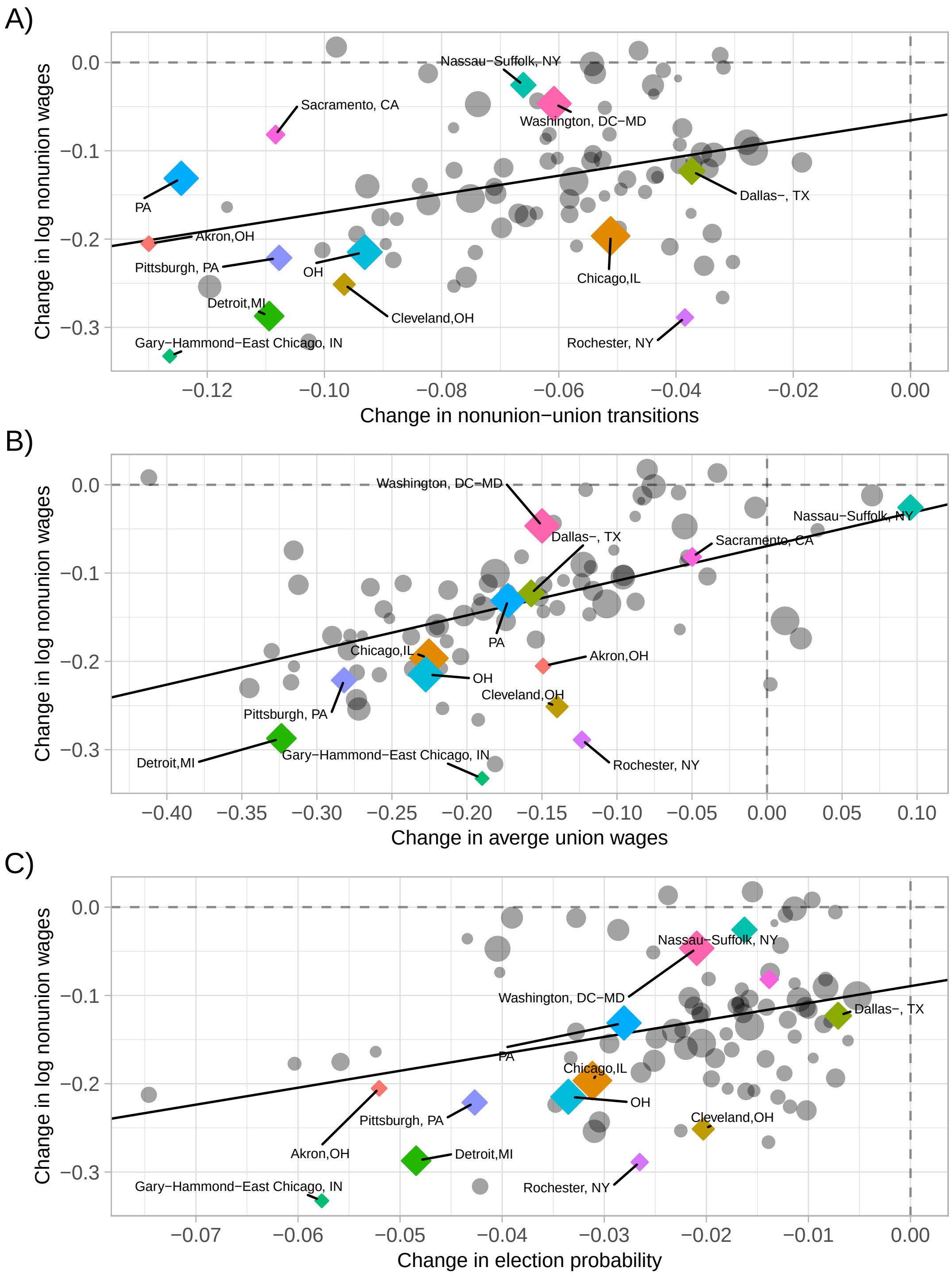


Figure 2: Wages and de-unionisation

Notes: Data from the CPS and the NLRB. In all three panels the y-axis denotes the change in regression adjusted nonunion wages at the city level between 2010 and 1980, and the marker size is relative to the size of the city in 1980. In panel A the x-axis variable is the change in our measure of nonunion-to-union transition rates, T^{n}_{jct}, aggregated to the city level. In panel B, the x-axis variable is the change in the average, regression adjusted union wage at the city level. In panel C, the x-axis denotes the change in the probability of a firm facing a union certification election calculated from the NLRB data and aggregated to the city level. Appendix B contains more information on our data construction.

a coefficient estimate on the union proportion in the region of .25-.58 for nonunion workers when exploiting industrial variation, although they find little evidence of a spillover effect operating at the local level. Holzer (1982) finds a positive spillover effect using the rate of unionisation at the SMSA for white males, although his results are sensitive to the inclusion of supply and demand shifters and the sample time frame.

Table 1: Non-Union Wages and Outside Options: OLS and 2SLS Estimates

	OLS			2SLS			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔP_{ic}	0.26***	0.030*	0.013	0.014			
	(0.024)	(0.017)	(0.015)	(0.015)			
$\Delta((1-T^N_{ict})\cdot \bar{w}^N_{ct})$			0.92***	0.69***	0.69***	0.61***	
			(0.016)	(0.074)	(0.074)	(0.10)	
$\Delta(T^N_{ict}\cdot \bar{w}^U_{ct})$			0.80***	0.64***	0.64***	0.69***	
			(0.029)	(0.063)	(0.062)	(0.088)	
ΔE^N_{ict}							0.65***
							(0.067)
ΔER_c			-0.091*	0.14	0.14	0.23*	0.18*
			(0.051)	(0.10)	(0.10)	(0.13)	(0.096)
Observations	9024	9024	9024	9024	9024	9024	9024
R^2	0.026	0.387	0.531	0.222	0.223	0.211	0.218
Year × Ind.	No	Yes	Yes	Yes	Yes	Yes	Yes
Instrument set:				$IV1^N_{jct}$	$IV1^N_{jct}$	$IV2^N_{jct}$	$IV2^N_{jct}$
				$IV1^U_{jct}$	$IV1^U_{jct}$	$IV2^U_{jct}$	$IV2^U_{jct}$
First-Stage p-Stat.:							
$\Delta((1-T^N_{ict})\cdot \bar{w}^N_{ct})$				0.000	0.000	0.000	
$\Delta(T^N_{ict}\cdot \bar{w}^U_{ct})$				0.000	0.000	0.000	
ΔE^N_{ict}							0.000
Over-id. p-val				.	.	.	0.457

Notes: This table displays results from the estimation of equations (22) (columns 1 and 2) and (5) (columns 3 - 7) via OLS (columns 1 - 3) and 2SLS (columns 4 - 7). The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

In column (2), we control for industry-year fixed effects which leads to a large decrease in the magnitude of the coefficient on the proportion union from .26 to .03. In comparison, Neumark and Wachter (1995) employ separate industry and year effects, working with national level variation. They estimate a negative effect of the union proportion on nonunion wages. Farber (2005) exploits variation in the probability of unionisation across states and industries in the cross-section, and in turn controls for state and industry fixed effects. He finds a coefficient estimate on the probability of unionisation around .18 which declines significantly over time. This result is more similar to ours than that of Neumark and Wachter (1995), but in both cases the source of variation is not directly comparable. In taking first differences we are controlling for any fixed city-industry effects, and we additionally control for common industry trends in wage growth.

In the third column of Table 1, we present OLS estimates of our simple specification but (contrary to that derivation) include the proportion union in the ic cell. We break the outside option term into its two components: the probability of finding a nonunion job times the expected wage in nonunion jobs ($(1 - T_{jc}^n) \sum_i \eta_{ic}^n w_{ic}^n$); and the probability of finding a union job times the expected wage in those jobs ($T_{jc}^n \sum_i \eta_{ic}^u w_{ic}^u$). Note that the coefficients on these variables should be equal since what matters is the overall outside option value and these are just components of that value. In other words, it should not matter to a firm whether a worker's outside option loses value because a high rent unionised firm leaves town or a nonunionised firm in an industry that also pays high rents shuts down.

The results in column (3) indicate that the option values associated with union and nonunion jobs have positive and significant effects on nonunion wages and are similar, though certainly not identical, in size. In addition, the employment rate enters significantly but with the theoretically incorrect sign. In column (4), we present results of IV estimation using the $IV1^n$ and $IV1^u$ instruments. We do not instrument for ΔER_{ct} even though there are clear reasons to assume it is correlated with our error term.[31] We follow Stock and Watson (2011) in interpreting the employment rate as a control variable - a variable that is not of direct interest in its own right but is useful for picking up its own effect and those of correlated omitted variables. In our case, we view the employment rate as capturing its own effect plus the impact of general, local demand shifts. This allows us to isolate the outside option effects we care about from demand effects. We present the p-values for the Sanderson-Windmeijer test statistics for weak instruments (Sanderson and Windmeijer, 2016) at the bottom of the table. These take values of 0.001 or less in all cases, indicating that we do not face weak instrument problems.

The results in column (4) again show positive and significant outside option effects, with the estimated coefficients for the two components being very similar in size. This and the fact that the employment rate now takes the correct sign are strongly supportive of the model since there is no mechanical reason why the two outside option terms should have similar sized effects. In the context of the model, in which the composition of local employment does not determine productivity within a specific industry, the significance of these effects implies that wages are partly driven by bargaining responses to rents in the local economy. Following BGS, we view the fact that we get these estimates while controlling for industry specific trends and controlling for the local employment rate as reinforcing this interpretation. Note, further, that the coefficient on the change in the union proportion is now 0.014 and smaller than its standard error. This suggests that any threat effects captured by the union

[31]We did implement a specification in which we instrumented for ΔER_{ct} using the classic Bartik, labour demand instrument. We find that this instrument is highly correlated with our other instruments and results in large estimates of the impact of the employment rate that are out of line with those in BGS. The estimates for the proportion union and the outside option value coefficients are very similar regardless of whether we instrument for the employment rate. The results in which we instrument for the employment rate can be found in Appendix C.

proportion variable occurs through the bargaining channel reflected in our outside options variables. In column (5), we present our specification, dropping the proportion union, with no impact on our key estimated coefficients.

To understand the magnitude of our estimated outside option coefficients, recall from equation (11) that γ_1^n is the impact of a 1 dollar increase in the expected value of the outside option for a nonunion worker in industry i in city c on that worker's wage. But this is only the immediate impact of a shift in, say, the unionisation rate. The resulting increase in the wage in i is an increase in the outside option for workers in other industries, inducing further increases in their wages that then imply an increase in the value of the outside option and the wage in industry i, which implies a further increase in the outside option for other industries, and so on. In the end, the total impact of a 1 dollar increase in the outside option value for the mean wage in industry i is $\frac{\gamma_1^n}{1-\gamma_1^n}$.[32] Thus, our estimated initial impact of a one unit change in the value of the outside option of 0.66 (the average of the two estimates in column (5)) becomes 1.78 once we include feedback loops of the spillovers.

Following BGS and Green (2015), we can discuss this total effect in relation to a standard shift share estimate of the impact of a change in the industrial composition of nonunion jobs in a city on the mean nonunion wage, $\bar{w}_{ct}^n$. In particular, we can decompose the change in $\bar{w}_{ct}^n$ as:

$$\Delta\bar{w}_{ct}^n = \sum_i \Delta\eta_{ict}^n w_{ict}^n + \sum_i \eta_{ict+1}^n \Delta w_{ict}^n \tag{23}$$

The first term on the right hand side is the 'between' component, showing the effect of changes in the industrial composition. The second term is the 'within' component, showing the effect of changes in wages within sectors. For a clean decomposition, these two terms are assumed to be independent. But in our model, a change in composition has both its direct 'between' effect and an effect through inducing changes in wages within industries because of the alteration in the outside option value. Thus, if the industrial composition changes by enough to increase the 'between' component by 1 unit, in a standard setting, the average nonunion wage in the city will also increase by 1 unit. With bargaining induced spillovers, we also need to add in the effects on wages within industries, noting from (19) that a 1 unit increase in the mean nonunion wage would increase the outside option value, E_{jc}^n by $(1-T_{jc}^n)$ units. Thus, the total effect on wages would be $1+\frac{\gamma_1^n}{1-\gamma_1^n}\cdot(1-T_{jc}^n)$ units. Using the average of the two outside option coefficients in column (5) and an average value of T_{jc}^n of 0.04, this implies that to get the total impact of a change in industrial composition on the mean nonunion wage we would multiply the standard 'between' component measure of that impact by 2.86. This is very similar in size to what BGS found not taking account of the union sector.

In column (6), we estimate a specification in which use the $IV2$ instruments. Recall that these instruments do not use variation in the local proportion of workers who are unionised. The fact that we get very similar results with these instruments and the $IV1$ instruments means that changes in outside options induced by either changes in the probability of accessing union jobs or the industrial composition of jobs have the same effect. This is implied by the theory (since firms don't care why the outside option value changes) and serves as

[32] Here, we have assumed that γ_1^u=γ_1^n so that the spillover effects on union and nonunion wages are the same.

an over-identifying test of the model. It also implies that, when it comes to bargaining, de-unionisation is not special – it is a reason that worker bargaining power is reduced just like the decline in any other high wage option they might lose. As we will discuss in the next section, the fact that we get the same results when we don't use variation in the proportion unionised is also useful when we investigate potential selection issues.

In column (7), we impose the restriction that the coefficients on the two components of the outside option take the same value in an estimation using the $IV2$ instruments. Not surprisingly, given the similarity of the coefficients in column (6), we cannot reject the restriction at any standard level of significance (the p-value associated with testing the restriction is given at the bottom of column (5)). We view this as another over-identifying test of our model since, again, there is no mechanical reason why the two terms should have the same effect (and they are based on quite different variation – in nonunion versus union industrial proportions and wage premia) but theoretically their effects should be identical.

4.2.2 Controlling for Selectivity

Thus far we have presented estimates using IVs to break the linkage between local productivity shocks and growth in outside options. Section 2.5, however, makes clear that there is likely selectivity into the union sector based on productivity draws. This arises because productivity shocks are differently weighted in the union and nonunion sector due to alternative methods of wage bargaining.

We address selection through a generalized Heckman two-step approach (see Heckman (1979), Dahl (2002), Snoddy (2019)). The idea in this approach is that the error mean term in (11), $E(\epsilon_{ic}|I_{icf} \leq 0)$, creates an omitted variables bias that can be addressed by including a fitted error mean term as a covariate in the regression. Further, the error mean can be expressed as a non-linear function of the probability of selection (the probability of being nonunion in our case) or of exogenous variables that drive that probability. The fact that that the error mean term is a function of P_{ict} implies that instrumental variables approaches to estimating the standard specification are invalid in the presence of selection. In that situation the variation in P_{ict} induced by an instrument identifies a coefficient that is a combination of the causal effect of the union proportion on wages and the fact that when P_{ict} changes, the composition of nonunion firms (and the mean of the error term) necessarily changes. However, our instrument set includes instruments that are not a function of the union proportion and we have seen that we obtain very similar results whether or not we include the union proportion and union transition rate variables as instruments. Thus, we can take an approach in which we use the restricted instrument set that does not include the union proportion to identify the effects of our rent variables, using functions of the union proportion or related variables to absorb the selectivity effect.

Given these arguments, we examine potential selection effects using two sets of variables. First, we include a quartic in the change in the proportion of workers in the industry×city×time cell who are unionised. In doing this, we are taking the model very seriously. In particular, we are taking advantage of the difference between the transition probability of workers (used in the construction of the outside options), and the proportion of workers who are unionised at a point in time. The latter is the theoretically correct variable for dealing with selection since differences in the proportion who are unionised is directly related to differences in se-

Table 2: Non-Union Wages and Outside Options: Controlling for Selectivity

	OLS	2SLS		OLS	2SLS	
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta((1-T^N_{ict})\cdot\bar{w}^N_{ct})$	0.92***	0.61***		0.92***	0.61***	
	(0.016)	(0.100)		(0.020)	(0.11)	
$\Delta(T^N_{ict}\cdot\bar{w}^U_{ct})$	0.80***	0.69***		0.82***	0.72***	
	(0.030)	(0.089)		(0.032)	(0.080)	
ΔE^N_{ict}			0.65***			0.68***
			(0.067)			(0.065)
ΔER_c	-0.090*	0.23*	0.19*	-0.028	0.51**	0.39***
	(0.051)	(0.13)	(0.097)	(0.072)	(0.21)	(0.14)
Observations	9024	9024	9024	6860	6860	6860
R^2	0.53	0.21	0.22	0.54	0.25	0.25
Year × Ind.	Yes	Yes	Yes	Yes	Yes	Yes
Instrument set:		$IV2^N_{jct}$	$IV2^N_{jct}$		$IV2^N_{jct}$	$IV2^N_{jct}$
		$IV2^U_{jct}$	$IV2^U_{jct}$		$IV2^U_{jct}$	$IV2^U_{jct}$
First-Stage F-Stat.:						
$\Delta((1-T^N_{ict})\cdot\bar{w}^N_{ct})$		0.000			0.000	
$\Delta(T^N_{ict}\cdot\bar{w}^U_{ct})$		0.000			0.000	
ΔE^N_{ict}			0.000			0.000
Over-id. p-val		.	0.474		.	0.326
Selection Controls						
P_{ic} Quartic	Yes	Yes	Yes	No	No	No
Election Vars.	No	No	No	Yes	Yes	Yes
Joint Tests:						
p-val	.855	.867	.8214	.17	.306	.1732
F-Stat	.334	.316	.3821	1.54	1.2008	1.5148

Notes: This table displays results from the estimation of equation (5) via OLS (columns 1 and 4) and 2SLS (columns 2,3,5, and 6). The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

lectivity. We present both OLS and IV estimates of our main specification including the quartic in the change in the union proportion. For the IV estimates, we use the $IV2$ set – the instruments that do not include any changes in the unionisation rate.

Following Fortin et al. (2019), we also estimate a specification in which we proxy for costs of unionisation using National Labor Relations Board (NLRB) data on certification elections. In particular, we calculate the number of workers involved in certification elections in each city over three year windows around the years 1980-1990-2000-2010, divided by the number of nonunion workers in the city. We also construct a second measure: the fraction of certification elections won over the same three year window. We see low values of both variables as reflecting higher costs to successful unionisation in a location. We use this to address selection by including quadratics in both variables and their interaction.

We present results from the specifications using the quartic in unionisation proportions to control for selection in columns (1)-(3) of Table 2, with column (1) containing OLS estimates and the IV estimates in columns (2) and (3). A test of the hypothesis that the parameters in the quartic equal zero is not rejected at any standard significance level, and the estimates for the key covariates change very little from Table 1. Columns (4)-(6) contain the results when we use the NLRB variables to address selection. Because the elections data series does not extend to 2020, we estimate this specification using the 1980 through 2010 data. Here, we can again not reject the null hypothesis of no selectivity effects, though the associated p-values are much lower than with the union proportion quartic. The main estimated effects are more erratic, though the outside option effect is very similar to the estimates in Table 1 if we impose the restriction that the union and nonunion outside options have the same effect. We conclude that selectivity is not a central issue driving our results.

4.2.3 Results Allowing for Firm Response

We now turn to our full specification, (16), in which we allow for nonunion firms to respond to the threat of unionisation by increasing their wages. As we saw in section 2, we do this by multiplying the outside option variable for a nonunion worker, E^n_{ic}, by the probability the firm is not a union emulator and introducing a new variable equal to the outside option value for a union worker, E^u_{ic}, multiplied by the probability the firm is an emulator. We construct the outside option variables as before, with E^u_{ic} and E^n_{ic} differing in the transition probabilities to entering union jobs for union and nonunion workers (with the estimated probabilities from the national data being substantially larger for the union workers). The probability of a nonunion firm being a union emulator is not directly observed and we proxy for it using our constructed probability that firms in the ic cell faced a unionisation drive in the previous four years, described in section 3.4. As we saw earlier, the probability of a union election moves differently from the probability a worker can transfer to a union job. We view variation in the probability of an election times the outside option value for union workers as reflecting the traditional threat effect and variation in the probability of a nonunion worker obtaining a union job times the expected rents in those jobs as the bargaining effect. These correspond to the two main right hand side variables in our full specification.

We present the results for the full, firm response model in Table 3. The first column contains IV estimates, using the $IV1$ instrument set along with $\hat{P}(\psi_{icf}, \lambda_c)$ and interactions between it and the $IV1$ instruments as instruments (calling this instrument set $IV3$). Recall from equation (16) that the coefficient on the first term (the probability of not being a union emulator times the expected value of wages for a nonunion worker) is the same as on the outside option value in the simple specification. And, in fact, the estimated coefficient for that first term is very similar to what we observed in Table 1. For the second term (the union worker outside option value times the probability that the firm is an emulator), the coefficient is the effect of an increase in the outside option value on a union worker's wage and turns out to be very similar in size to the effect for nonunion workers. In column (2), we replace the $IV1$ instruments with $IV2$ instruments and find even greater similarity in the outside option effects for union and nonunion workers at similar size effects to column (1)[33] . In column (3), we include the quartic in changes in the proportion union to address selection. As in Table 2, the main estimated effects change very little when we control for selectivity.

The key result from these estimates is that both the bargaining channel (as captured in the coefficient in the first row) and the standard threat channel (represented by the coefficient in the second row) matter. In fact, they have similar sized coefficients, though there is no theoretical reason why they would be equal. If $\gamma_1^n = \gamma_1^u$ (as seems to be a reasonable assumption given the estimates) and $T_{ic}^n = T_{ic}^u$ then the impact of a one dollar increase in the mean nonunion wage due to a shift in industrial composition would have the same total effect as in the simple regression (i.e., $1 + \frac{\gamma_1^n}{1-\gamma_1^n} \cdot (1 - T_{jc}^n)$ or about 2.9 times the standard 'between' effect computation of the industrial composition shift).[34] This occurs because the mean nonunion wage is part of both E_c^n and E_c^u and the former is multiplied by $(1 - P(\psi_{icf}, \lambda_c))$ while the latter is multiplied by $P(\psi_{icf}, \lambda_c)$. The significance of the union worker outside option is, in some ways, remarkable. Its effect is identified relative to the nonunion worker outside option only because the transition probabilities to union jobs are larger for union workers.

In column (4), we interact the outside option variables with an indicator variable for whether the city is located in a state with Right to Work (RTW) laws at the start of the decade. Given this specification and the fact that only 6 states adopt RTW laws in our time period (Fortin et al. (2019)), this indicator is not about the effects of becoming a RTW state but about whether nonunion wage setting is different in states with ongoing RTW environments. The impact of the nonunion option value is slightly larger in RTW states. But more importantly, adding together the union outside option coefficient with the coefficient on its interaction with the RTW indicator yields a statistically insignificant effect of only 0.24 in RTW states. That is, we cannot reject the hypothesis that there is no traditional threat effect in states where RTW laws eliminate the threat of being unionised. This is both interesting in its own right and a further piece of evidence for our claim that we are identifying emulation effects.

Taken together, our results indicate that de-unionisation affected nonunion wage setting through both bargaining and standard threat channels, though the latter does not operate in RTW environments. The effects appear to be sizeable relative to what one would compute with simple shift-share calculations but we will confirm their overall impact in a

[33] We refer to the set of instruments composed of $IV2$ and interactions with $\widehat{P(\psi_{icf}, \lambda_c)}$ as $IV4$.

[34] If, as we observe, $T_{ic}^u > T_{ic}^n$ then the impact of a one dollar change in the mean non-union wage is smaller by $\gamma_1^u P(\psi_{icf}, \lambda_c)(T_{ic}^u - T_{ic}^n)$ than it would be if $T_{ic}^u = T_{ic}^n$, which is a small adjustment.

counterfactual exercise after investigating heterogeneity across different demographic and skill groups.

Table 3: Non-Union Wages and Outside Options: Including Wage Emulation Effects

	2SLS				
	(1)	(2)	(3)	(4)	(5)
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)$	0.70***	0.69***	0.70***	0.70***	0.68***
	(0.071)	(0.072)	(0.073)	(0.071)	(0.075)
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)$	0.64***	0.70***	0.63***	0.54***	0.52
	(0.094)	(0.088)	(0.093)	(0.13)	(0.88)
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)\cdot RTW$				0.12***	
				(0.035)	
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)\cdot RTW$				-0.30*	
				(0.18)	
ΔER_c	0.33**	0.36**	0.34**	0.25*	0.46***
	(0.15)	(0.15)	(0.15)	(0.15)	(0.15)
Observations	6860	6860	6860	6860	6860
R^2	0.256	0.254	0.256	0.256	0.251
Year × Ind.	Yes	Yes	Yes	Yes	Yes
$P(\psi_{icf},\lambda_c)\times$ Ind.					Yes
Instrument set:	$IV3^N_{jct}$ $IV3^U_{jct}$	$IV4^N_{jct}, IV5^N_{jct}$ $IV4^U_{jct}, IV5^U_{jct}$	$IV3^N_{jct}$ $IV3^U_{jct}$	$IV3^N_{jct}, IV3^N_{jct}\times RTW$ $IV3^U_{jct}, IV3^U_{jct}\times RTW$	$IV3^N_{jct}$ $IV3^U_{jct}$
Select controls					
ΔP_{ic} Quartic			Yes		
First-Stage p-Stat.:					
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)$	0.000	0.000	0.000	0.000	0.000
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)$	0.000	0.000	0.000	0.000	0.000
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)\cdot RTW$				0.000	
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)\cdot RTW$				0.000	
Over-id. p-val	.	0.021	.	.	.

Notes: This table displays results from the estimation of equation (16) via 2SLS. The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

4.2.4 Heterogeneity in Spillover Effects

There is considerable heterogeneity in the unionisation experience. Farber et al. (2021) demonstrate that union density was higher among lower educated workers in the US during the highest unionisation period (particularly between 1940 and 1960). Other evidence indicates that unionisation was particularly prevalent among low-skilled men (although as highlighted by Card et al. (2018), there has been a remarkable rise in the share of unionised jobs held by women). The decline in unionisation might then be expected to have had a bigger impact on these groups, and so it would be useful to know if they have a bigger or smaller reaction to outside options in their wage setting.

In Table 4, we present estimates of the bargaining and standard threat effect coefficients (γ_1^n and γ_1^u in equation (16)) for a set of sub-populations defined by gender, age, and education. Each row corresponds to estimates for a different subsample. To construct industry by location cells of sufficient size, we shift to using states rather than cities as our geographic dimension. Estimating our specification with the full sample with states generates very similar effects (though slightly larger in absolute value) to those presented in the tables up to this point, which use cities. The 3rd and 4th columns of the table also show the p-values from SW weak instrument tests for the instruments corresponding to the two outside option terms. In all cases, these p-values are 0.02 or less, implying no weak instrument problems. The sample size varies across the sub-populations because we eliminate ic cells with fewer than 20 observations in each case.

The first two rows contain separate results for men and women. These indicate that both the bargaining and standard threat effects are larger for women, implying that declines in unionisation would have a more negative effect on nonunion wages for women than men. On the other hand, differences by age group are not generally substantial. The other noticeable pattern is the relative importance of the traditional threat effect compared to the bargaining effect for the less educated – and, particularly, for less educated males. While all the effects are statistically significant and of a similar order of magnitude to our whole sample estimates, the traditional threat effect estimate is twice as large as the bargaining effect for the low (high school graduates or less) educated while the opposite is true for those with a post-secondary education. That is the relatively more salient threat for the lower educated is to unionise the current workplace rather than to leave and find a better job, which fits with characterizations of unionisation among less-educated workers as raising wages through organizing relatively homogeneous workers (in terms of skill) who do not have much individual power because their specific outside options are not strong.

We have also examined potential heterogeneity between public and private sector unions. Thus far we have included the public sector, both in the construction of our outside option terms, and as an observation on the left hand side of the wage equation. Card et al. (2018) however outline the marked difference in unionisation between the private and public sectors since 1980 such that unionisation is now 5 times higher in the public sector. To the extent that wage setting is different in the public sector, these shifts in composition could be driving some of our results. In Appendix C, we present results excluding the public sector both in the construction of the dependent variable and in the construction of our outside option variables and associated instruments. Our results are robust to these changes, though our estimated spillover effects are slightly larger in the simple specification.

5 Counterfactual Exercise

Our results thus far indicate a significant relationship between quality job opportunities in both the nonunion and union sectors and nonunion wage setting. However, the exact

Table 4: Subsample Analysis - Coefficient Estimates on Outside Options

	(1)	(2)	(3)	(4)	(5)
	Coefficient		First-stage p-value		
Sample	γ_1^n	γ_2^n	(1)	(2)	N
Men	0.55**	0.44**	0.00	0.00	5073
	(0.13)	(0.18)			
Women	0.73**	0.96**	0.00	0.00	4164
	(0.07)	(0.17)			
Age 20–26	0.69**	0.40**	0.00	0.00	4371
	(0.08)	(0.19)			
Age 36–55	0.69**	0.67**	0.00	0.00	4618
	(0.14)	(0.19)			
$\leq$ HS	0.50**	1.00**	0.00	0.00	4486
	(0.11)	(0.25)			
$>$ HS	0.59**	0.41*	0.00	0.00	4835
	(0.12)	(0.23)			
Men: Age 20–26	0.69**	0.74**	0.00	0.00	2486
	(0.15)	(0.25)			
Men: Age 36–55	0.53	1.12	0.02	0.02	2834
	(0.37)	(0.68)			
Men: $\leq$ HS	0.52**	1.10*	0.00	0.00	2258
	(0.14)	(0.62)			
Men: $>$ HS	0.85**	0.55*	0.00	0.00	3358
	(0.16)	(0.33)			

Notes: This table displays results from the estimation of equation (16) via 2SLS using 2SLS on separate subsamples.

magnitude of the estimated effects remain unclear. In this section, we pursue a counterfactual exercise, asking what path mean wages in a typical city would have followed if unionisation rates and union wage premia had remained at their 1980 levels. This both provides a way of characterizing the size of our estimated effects and some insight into whether deunionisation played an important role in wage changes over the last four decades.

5.1 Loss of Union Power and Movements in the Average Wage

Our focus is on changes in total mean wages at the city level, expressed as the weighted average of nonunion and union mean wages, where the weight is the proportion unionised at the city level, P_{ct}:

$$w_{ct} = P_{ct} \times w^u_{ct} + (1 - P_{ct}) \times w^n_{ct} \tag{24}$$

In order to focus attention on trends related to unions, undistracted by changes in education, age, etc., we use the residualized industry-city wages that formed the dependent variable in our regressions combined with industrial shares at the local level to create city level wages (i.e., $w^n_{ct} = \sum_i \eta^n_{ict} w^n_{ict}$ and $w^u_{ct} = \sum_i \eta^u_{ict} w^u_{ict}$).

Changes in the strength of unions affect movements in the overall average city wage through four channels. The first is changes in the union proportions, P_{ct}. This is the most direct effect of deunionisation: the shift in workers out of higher paid (union) jobs to lower paid (nonunion) jobs, holding constant the wages in each sector. This is the 'between' component in standard decompositions and is often presented as the effect of unions on the wage structure. The second channel is the effect of de-unionisation on nonunion wages through the classic threat of unionisation route. In our full wage specification, (16), this is captured by changes in the probability that nonunion firms face union election drives ($P(\psi_{icf}, \lambda_c)$). The loss of union power also affects nonunion wages through reducing outside options (reductions in bargaining power). This happens, in part, because of reductions in the probability a nonunion worker can find a union job. In our specification, this shows up in two places. The first is a decline in T^n_{jc} , which refers to a fall in the probability of getting a union job in any sector. But the impact on worker outside options will obviously be greater if it is mostly union jobs in, say, the high-paying manufacturing industry that are lost. At the same time, we don't want to assign all industrial changes as union effects. Instead, we assume that shifts in the industrial distribution for nonunion workers (the η^n_{ic}'s) capture changes in the overall economy while a change in the industrial distribution for union workers relative to what happens for nonunion workers (η^u_{ic} - η^n_{ic}) is a union decline effect. These shifts in T^n_{jc} and the (η^u_{ic} - η^n_{ic})'s form the third union impact channel.

The value of the outside option for nonunion workers is also affected by declines in the union wage premia (the fourth channel). If unions become less effective at unifying worker resistance during bargaining or afraid to threaten the withdrawal of the whole workforce in a new policy environment then the union wage premium will decline. In that case, the value of the outside option of finding a union job for a nonunion worker also declines. Here, too, we focus on the relative decline of union versus nonunion wages (the union wage premium).[35]

[35] A fifth channel is through selection, as shifting firms from being union to nonunion changes the productivity composition of nonunion firms. In our model, this effect implies an increase in the observed nonunion wage, offsetting the negative effects of declining union power operating through the first two channels. How-

Versions of the third and fourth channels also exist for union wage determination: as unions are weakened other union workers face declines in their own outside options.

In Figure 3, we plot percentage changes in these key drivers (P_c, $P(\psi_{icf}, \lambda_c)$, T^n_{jc}, and $(w^u - w^n)$) relative to their 1980 values. For each of the probabilities, we work with values at the industry-city cell level, aggregating them to a single value using the proportion of all nonunion workers in each cell in 1980. The union wage premium comes from our estimates at the nation level. Our model indicates that the different types of unionisation probabilities are relevant for different parts of the wage setting process. P_c does not appear directly in our model but is part of the first step of our decomposition, as we discussed earlier, and is what is used in other papers. T^n_{jc} is relevant for outside options since it shows whether workers are actually able to move into union jobs. Its movements are obviously related to the decline in the proportion of workers who are unionised, though one could imagine it declining either faster than that proportion (if older union workers keep their jobs but new job searchers have difficulty getting into a union job) or slower than that proportion (if the proportion declines quickly because union workers suddenly start taking early retirement). In fact, the figure shows that the two proportions move similarly in the 1980s but the probability of entering a union job declines faster after 1990. $P(\psi_{icf}, \lambda_c)$ is relevant for the traditional threat effect since it shows changes in whether a firm can expect to face a unionisation drive. It falls the fastest of any of the unionisation measures, particularly in the 1980s when the policy environment was shifting strongly against unionisation. It is worth noting, though, that the probability of a firm facing a union election was small even in 1980 (on the order of 4%).

Perhaps the most interesting line in Figure 3 is the one corresponding to the union wage premium. The premium actually increases in the 1980s before showing a sizeable decline in the 1990s and a smaller one thereafter.[36] Our model provides an explanation for the increase in the 1980s based on the emulation channel that echoes a discussion in Farber (2005). Recall that the observed mean non-union wage equals a weighted average of the pure nonunion wage (w^n), paid by firms under no threat of unionisation, and the emulation wage (w^{n*}), paid by firms seeking to stave off unionisation. The weights are $P(\psi_{icf}, \lambda_c)$ and $(1 - P(\psi_{icf}, \lambda_c)$, respectively. Suppose that larger forces (trade, technological change, etc) drive down both w^n and the union wage, w^u, to the same extent. If the threat of unionisation declines at the same time then the observed nonunion wage will fall farther because there will be fewer emulating firms and the emulation wage they have to pay won't be as high. This pattern of faster decline in mean observed wages in the nonunion sector is what we observe in the 1980s. It is striking that this is the decade in which the union threat (the probability of facing a union election) fell fastest relative to other unionisation probabilities.

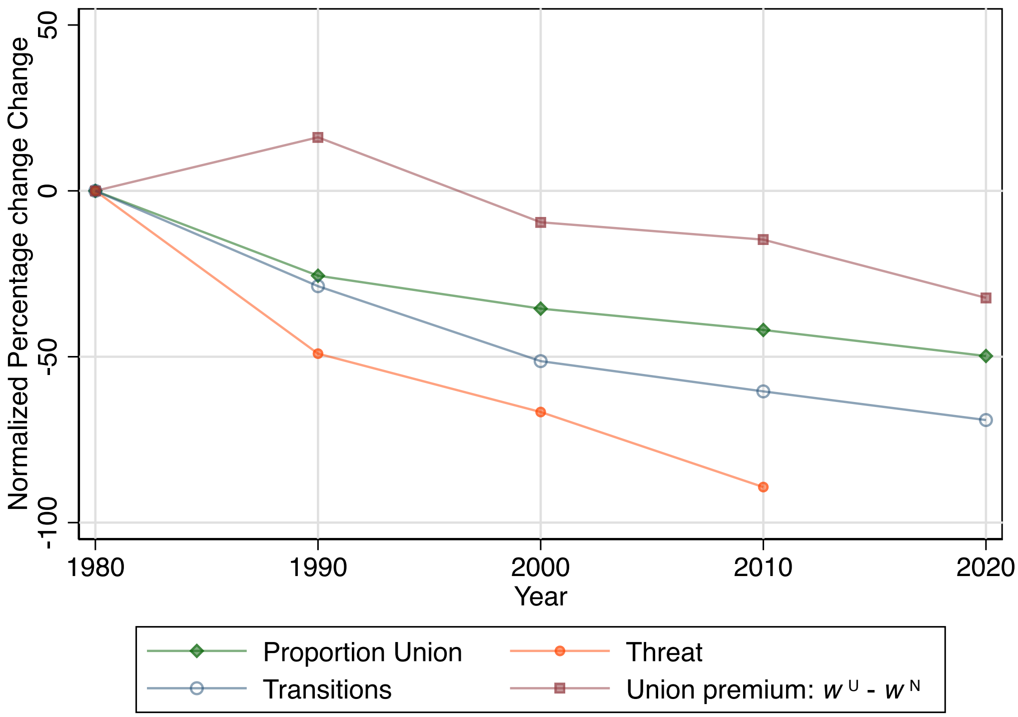
5.2 Overall Decomposition

We present our decomposition of the overall trend in average city wages in Figure 4. Recall that we are working with residualized wages after taking out education, race, age, and gender

ever, we find no evidence that selection had substantial effects on observed mean wages and, so, do not include this channel in our decompositions.

[36] Farber et al. (2021) plot union wage premiums over an extended time period. Their plot differs from ours in showing a flat premium over the 1980s but is similar in showing a decline after 1990. Their estimates are based on family income and do not include controls for education that are part of our estimation.

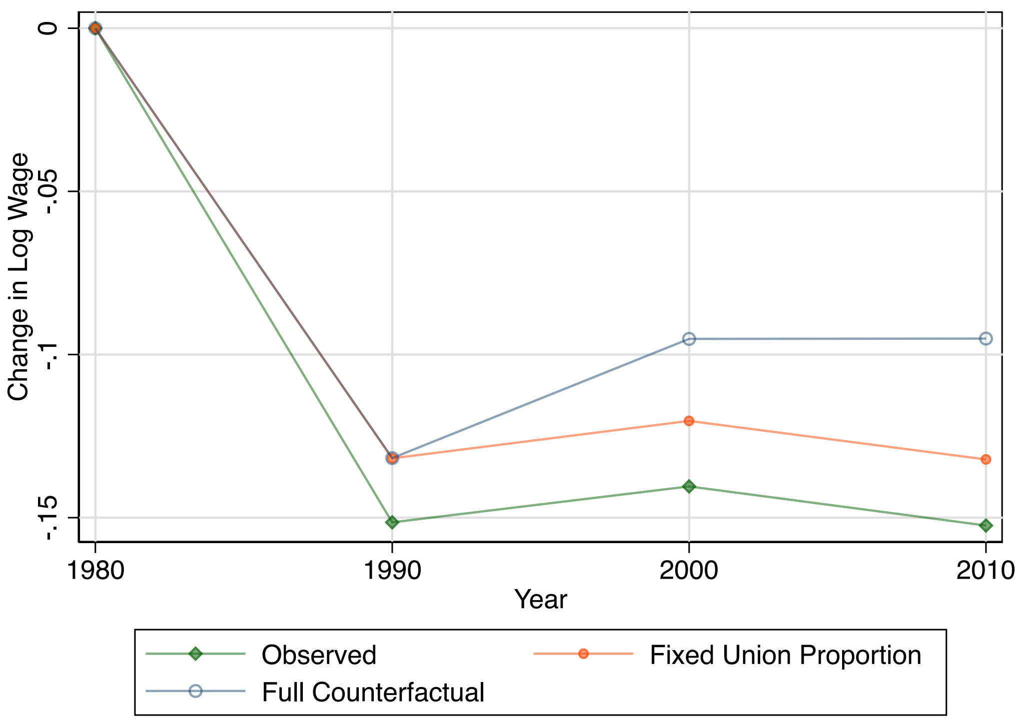
Figure 3: Components of Decomposition



Notes: Data from the CPS and the NLRB. Each series is represented as a percentage change from the corresponding 1980 level. Proportion union, union premium and transitions are constructed from the CBS data, discussed in Appendix B. The treat of union election comes from NLRB data, and is described in detail in Appendix B.1.

effects. We set the wage level to correspond to the mean wage across all worker types.[37] To estimate time trends we average across cities each period, using city populations as weights. Thus, we are showing the components of mean wage movements for an average city, not the components of the trend in mean wages at the national level. The bottom line in Figure 4 is the actual trend in the (residualized) mean wage for an average city. It depicts an overall real wage trend that is strongly decreasing between 1980 and 1990 – falling by approximately 15% in that decade – followed by a see-saw pattern of mild increases in the 1990s and declines in the 2000's.[38]

Figure 4: Average Wage Decomposition



Notes: Data from the CPS and the NLRB. Each series is represented as a log change from the corresponding 1980 level. Wage data is from both union and nonunion workers and is adjusted for worker characteristics. The 'Observed' wage series represents the national average of city-industry wages using the size of the city-industry in 1980 as fixed weights. 'Fixed Union Proportion' holds the proportion of union workers fixed at the 1980 levels. 'Full Counterfactual', in addition, holds the threat, union premium and union transitions at 1980 levels. Details of series construction described in main text.

In our decomposition, we hold the driving forces related to the de-unionisation impact channels constant at their 1980 levels while allowing all other factors determining wages to vary. Our first component corresponds to the first channel: holding the the union proportions, P_{ct}, in (24) constant at their 1980 values. It is given by the 'Fixed Union Proportion'

[37] In particular, mean wages correspond to the wages of white workers, holding the proportion of education×gender groups at their 1980 levels.

[38] We end our figure in 2010 because we only have data on one element of our decomposition – the part related to union elections - up to that year.

line in the figure. According to that line, this channel for the impact of the unionisation decline generated a 2.0% drop in the mean wage in the 1980s, accounting for about 13% of the overall drop in the mean wage in that decade, with about the same effect on the drop between 1980 and 2010.

In the second step in our decomposition, we construct counterfactual outside option values for both nonunion and union workers in each ic cell by setting $P(\psi_{icft}, \lambda_{ct})$, T^n_{jct}, T^u_{jct}, $(\eta^u_{ict} - \eta^n_{ict})$and $(w^u_{ict} - w^n_{ict})$ equal to their 1980 values.[39] Notice that we hold constant differences in industrial structure and wages between union and nonunion sectors, allowing the outside option values to move with movements in the η^n_{ict}'s and w^n_{ict}. We use the counterfactual outside option values to predict mean nonunion and union wages in each ic cell and year using the estimated coefficients from our preferred specification in column (3) of Table 3.[40] These initial estimated wages, however, are only first round effects of de-unionisation. If these counterfactual wages in a particular ic cell are higher than what was actually observed then the outside options for other workers would also be higher and, so, we create a second round of counterfactual values of the outside options in which we use the first round counterfactual wages and then form a second round of counterfactual wages using the updated outside options. This new set of wages, in turn, implies a new set of outside options and, therefore, new wages. We continue to iterate on this process until the predicted wages change by less than 0.1 of a percent. This provides estimates of the effect of reductions in the threat of unionisation that reflects the complete feedback loop inherent in bargaining schemes. It also means that the union premia used in the outside option terms are consistent with the premia we would calculate from the set of counterfactual wages.[41]

The top line in the figure shows the combination of the two steps in the decomposition – the full counterfactual effect of holding union related factors at their 1980 values. For the first decade – 1980 to 1990 – there is no spillover effect of de-unionisation. Instead, the effects are completely captured by the standard 'between' component in a shift-share analysis. But in the ensuing decades, spillovers begin to emerge and for the full period from 1980 to 2010, amount to almost twice the simple 'between' component (accounting for a 0.037 decline in the overall wage compared to 0.02 from the union proportion component). Over this full period, the two components together imply that de-unionisation can account for 37.6% of the total decline in the mean wage over that period. This is similar to Fortin et al. (2021)'s finding that taking account of spillovers roughly doubles the estimated 'shift-share' impact of de-unionisation on wage inequality over the 1979-2017 period (though, in our case, it much more than doubles it). Our results focus on wage levels rather than inequality and provide

[39]For union wages, we use equation (6), the fact that the coefficient on the E^u_{ic} term in the full specification equals γ^u_1 according the theory, and an assumption that $\gamma^u_2 = \gamma^n_2$.

[40]In particular, we construct an initial counterfactual wage for each nonunion worker in 1990 by subtracting the bargaining effect variable $((1 - P(\psi_{icf,1990}, \lambda_c) \cdot E^n_{ic,1990})$ times its estimated coefficient (0.70) from their 1990 wage then adding back $0.7 \cdot (1 - P(\psi_{icf,1980}, \lambda_c) \cdot E^c_{ic,1980}$, where $P(\psi_{icf,1980})$ is the relevant probability of facing a union election in 1980 and $E^c_{ic,1980}$ is the outside option value in 1980. We do the same for union wages, using our estimate of γ^u_1 from Table 3, column 3 (0.63).

[41]Amazon touted its recent move to a \$15 minimum wage for its workers as a reason for workers not to unionise at one of its plants. Source: Amazon to hike wages for over 500,000 workers. Our procedure would capture that initial wage increase as a standard threat effect but would also take account of the possibility that the increased wage at Amazon represents an improved outside option for other nonunion workers, leading to increases in their wages.

an identification strategy for supporting a claim that we have estimated causal effects.

We will examine why there was no spillover effect from 1980 to 1990 in the next section when we examine nonunion wage movements on their own. First, though, we show our counterfactual exercise for various sub-groups in the top panel of Table 5.2. Men experienced a decline in the mean real wage between 1980 and 2010 that was over double that experienced by women, but with identical sized spillover effects, deunionisation plays a larger role in explaining the decline for women than men (43% for women versus 27% for men). Older aged workers (age 36 to 55) experienced larger real wage declines than for younger workers (age 20 to 26) but the union effects are similar in levels for both groups, implying they are proportionally more important for younger workers. The combination of high school graduates and drop-outs faced real wage declines that were almost triple those with post-secondary education. Perhaps not surprisingly, the effect of de-unionisation was also much larger for the lower educated group, accounting for 43% of their real wage decline. For the post-secondary educated, the spillover effects actually imply increases in mean wages. This arises because union jobs for this education group became more concentrated in higher paying (public sector) jobs, implying increased average union wages that more than offset declines in the probability of getting a union job in the calculation of their outside option term.

Table 5: Decomposition results: 1980-2010

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	All	Men	Women	Age 36-55	Age 20-26	$\leq$ HS	$>$ HS
	1980-2010						
(1) Observed	–0.152	–0.204	–0.088	–0.271	–0.189	–0.226	–0.084
(2) Union Prop.	0.020	0.027	0.011	0.019	0.026	0.043	0.010
(3) Full Counterfactual	0.037	0.027	0.027	0.031	0.026	0.054	–0.019
Non-union:							
(4) Full Counterfactual	0.018	0.015	0.015	0.014	0.016	0.029	–0.004
(4a) Fixed Transitions	0.021	0.019	0.012	0.021	0.020	0.037	0.003
(4b) Fixed Union Prem.	–0.002	–0.003	0.004	–0.005	–0.003	–0.007	–0.006
(4c) Threat Effect	–0.001	–0.001	–0.001	–0.002	–0.001	–0.001	–0.002
Union:							
(5) Full Counterfactual	0.019	0.013	0.012	0.017	0.010	0.025	–0.015
(6) Total	0.057	0.055	0.038	0.050	0.052	0.097	–0.009
(7) Total/Observed	–0.376	–0.268	–0.429	–0.183	–0.276	–0.432	0.112

Notes: This table displays results from the decomposition for union and nonunion workers from 1980-2010. Each column contains the decomposition results for a different subsample. All figures are log changes from 1980 levels. Details described in main text.

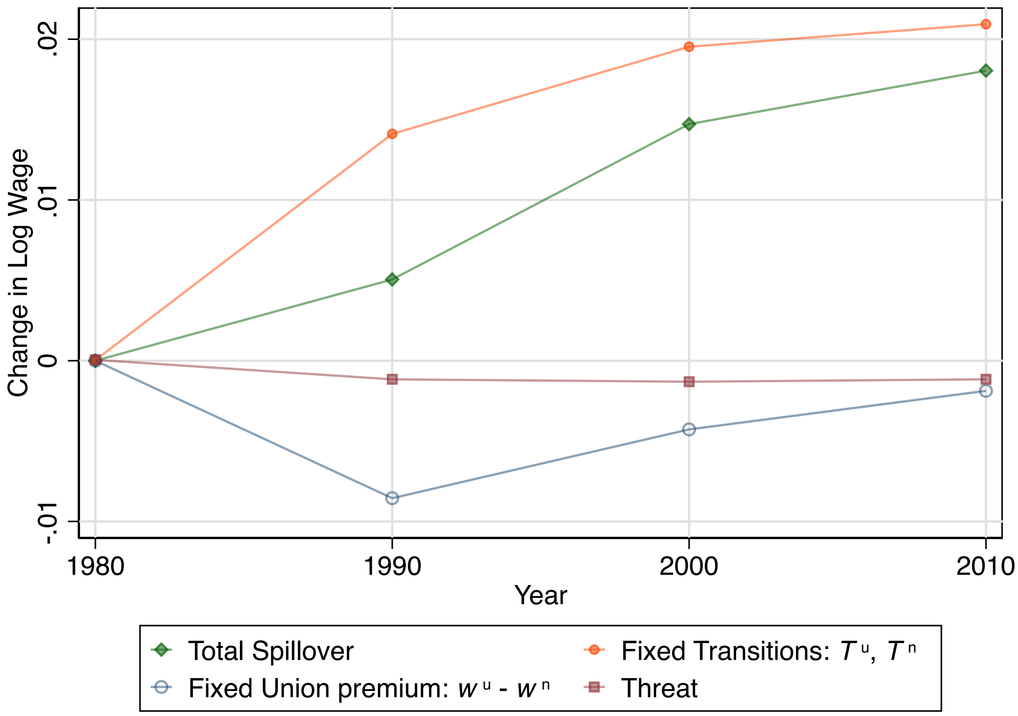
6 Decomposing Nonunion Wages

We next turn to decomposing movements in mean nonunion wages, which, of course, are the focus of our estimation. We do this with a similar approach, holding constant at their 1980 values (in turn) $P(\psi_{icf}, \lambda_c)$ (the threat effect), T^n_{jc} plus $(\eta^u_{ict} - \eta^n_{ict})$ (the effect of reduced probabilities of finding union jobs), and $(w^u - w^n)$ (the effect of changes in the union wage premium). As with the full decomposition, we incorporate subsequent changes in outside option values through an iterative process. Unlike the full decomposition, the result is not completely consistent. We do not update union wages in the iterative process, keeping them at the values in the specific year. Nonetheless, we view the result as useful in showing the broad strokes of which elements of de-unionisation had the most salient spillover effects onto nonunion wages.

In figure 5 we present the movements in the different union-related components of the nonunion mean wage decomposition. The line with square symbols shows the impact of holding the probability of a firm facing a unionisation drive constant at 1980 values. That impact is small, which arises because the threat probability was quite small even in 1980. Thus, while our estimates show clear evidence of the standard emulation threat effect, their actual impact on nonunion wage movements was small. That means that the sizeable spillover effect that emerges by 2010 in figure 4 is almost completely accounted for by the bargaining channel. This has potentially important implications for policy making aimed at raising wages since the threat effect can only be harness by increasing unionisation. But the bargaining channel is not unique to unions - any policy that pushes up the outside option value for workers (such as eliminating non-compete clauses (Johnson et al. (2020) or expanding commuting options Hafner (2022)) can have this effect and our results imply that this channel can be powerful. This is reminiscent of the results in Caldwell and Danieli (2021), who show that wages are increasing in their index of the value of outside options. Their index increases when workers have greater probabilities of transferring to other occupations and job opportunities. Our result is driven by decreases in the probability a worker can transfer to a union job.

But why was this channel so silent during the 1980s – the decade of US unionisation's biggest collapse? The answer can be seen in two offsetting impacts on outside option values revealed in Figure 5. On one side is the predicted impact from holding the probability a nonunion worker can move to a union job constant (the combination of T^n_{jc} and $(\eta^u_{ict} - \eta^n_{ict)})$, which is reflected in the 'Fixed Transitions' line in the figure. This line indicates that nonunion wages would have been 0.023 log points higher in 1990 and 0.016 log points higher in 2010 if these transition probabilities had not declined, accounting for between 10 and 15 percent of the declines in nonunion wages over this period. These effects are similar in magnitude to the standard 'between' effect due to the changes in P_c in the first step of the overall mean wage decomposition. But over the 1980s, they were nearly offset by the impact of the increase in the union wage premium that we saw in figure 3. Thus, the lack of a spillover effect in that decade in the overall decomposition appears to arise from the changes in the union environment having little impact on outside option values as declining probabilities of getting a union job were balanced with higher wage premia if a worker did manage to get one.[42] As we described earlier, our model provides an explanation for why union premia would increase exactly when union power is being most substantially reduced. It stems from the reduction in the need for some nonunion firms to emulate union wages since they no longer fear their shop being unionised.

Figure 5: Decomposition components: Nonunion workers



Notes: Data from the CPS and the NLRB. Each series is represented as a log change from the corresponding 1980 level. Wage data is for nonunion workers and is adjusted for worker characteristics. Each series corresponds to a component of the decomposition, described in main text.

The second panel in Table 5.2 shows the decomposition for nonunion wages for the full 1980-2010 sample period for different sub-groups. As with the decomposition of the overall wage, the union counterfactual effect is similar in size between men and women and across age groups. Also similarly to the overall wage decomposition, the implied deunionisation effect is much larger for high school or less educated workers than for workers with a post-secondary education. The nonunion wage decomposition indicates that this effect is due to the reduced probability of the lower educated nonunion workers moving to union jobs.

[42]Note that the components in the nonunion wage decomposition will not add up to what is observed in the full decomposition in figure 4 because they do not include impacts on union wages and union wages are not updated in the iteration process in the nonunion decomposition.

7 Conclusion

In this paper we provide new estimates of the impact of unions on nonunion wage setting. We allow the presence of unions to affect nonunion wages both through the typically discussed channel of nonunion firms emulating union wages in order to fend off the threat of unionisation and through a bargaining channel in which nonunion workers use the presence of union jobs as part of their outside option. We specify these channels in a search and bargaining model that includes union formation and, in our most complete model, the possibility of nonunion firm responses to the threat of unionisation. By formalising wage setting and union formation we derive a specification grounded in theory that provides guidance on what to control for, how to interpret our coefficients and what is in the error term. Based on that, we derive a set of instruments and a model-based over-identification test, the values for which imply that our identification strategy is appropriate for this data.

Our estimates indicate that deunionisation in the US after 1980 had a substantial effect on nonunion wages, in particular, and the wage structure in general. In a decomposition exercise, holding the probability a worker can find a union job, the probability a firm faces a unionisation drive, and union wage premia constant at their 1980 levels would have undone 38% of the 16% decline in the mean (composition constant) real wage in a typical city in the US between 1980 and 2010. While we find evidence for spillover effects of unions onto nonunion wage setting through both the traditional threat channel and the bargaining channel, it is the latter that dominates. That is important for policy makers looking for tools to help in raising wages. The union threat channel can only be implemented by increasing union power. But the bargaining channel is not specific to unions. Any policy that raises worker outside option values will raise wages for a wide set of workers (Beaudry et al. (2012), Caldwell and Danieli (2021)). Unions are just one mechanism for doing that – though our estimates indicate a powerful and direct one.

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Estimating Union Wage Spillovers: The Role of Bargaining and Emulation Effects On-line Appendices

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A Mathematical Appendix

A.1 Derivation of the Firm Surplus

Here we provided details regarding the derivation of the firm surplus under collective bargaining, and individual bargaining.

A.1.1 Collective Bargaining

As noted in text, the surplus from a successful bargain with a union is given by the difference between producing this period with n_u workers (the optimal number of workers with a bargained union wage) and not producing this period along with the cost of rehiring the entire workforce the next period. Noting that the firm has already hired its replacements for workers lost due to normal turnover at the time of the bargaining, this is given by

$$S^u = \big(p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf}) + \rho J^u_{icf}(n^u_{icf})\big) - \big(\pi(0) + \rho J^u_{icf}(0)\big) \tag{1}$$

Where $\pi(0) = 0$ corresponds to earning zero profits and $J^u_{icf}(0)$ is the value of a union firm starting the period with no workers. Due to the linear hiring costs, firms will hire back their optimal number of workers,[1] n^u_{ic}, every period, and, as a result, the expression for the value with no workers is:

$$J^u_{icf}(0) = p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf}) - \kappa \frac{n^u_{icf}}{q_{vc}} + \rho J^u_{icf}(n^u_{icf}) \tag{2}$$

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[1] Acemoglu and Hawkins (2014) develop a search model with quadratic hiring costs that leads to implications for the firm size distribution and dynamics that we do not consider here.

Substituting this into S^u yields,

$$\begin{aligned} S^u &= \left(p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf}) + \rho J^u_{icf}(n^u_{icf})\right) \\ &- \rho\left(p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf}) - \kappa\frac{n^u_{icf}}{q_{vc}} + \rho J^u_{icf}(n^u_{icf})\right) \\ &= (1-\rho)p_i y_{icf}(n^u_{icf}) - (1-\rho)n^u_{icf} w^u_{icf}(n^u_{icf}) + \rho\kappa\frac{n^u_{icf}}{q_{vc}} + \rho(1-\rho)J^u_{icf}(n^u_{icf}) \end{aligned} \tag{3}$$

Using the definition for $J^u_{icf}(n_u)$, this can be written,

$$\begin{aligned} S^u = (1-\rho)p_i y_{icf}(n^u_{icf}) - (1-\rho)n^u_{icf} w^u_{icf}(n^u_{icf}) + \rho\kappa\frac{n^u_{icf}}{q_{vc}} \\ + \rho(p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf}) - \kappa\frac{n^u_{icf}\delta}{q_{vc}}) \end{aligned} \tag{4}$$

Simple algebra then yields:

$$S^u = \left(p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf}(n^u_{icf})\right) + \rho(1-\delta)\frac{\kappa n^u_{icf}}{q_{vc}} \tag{5}$$

A.1.2 Individual Bargaining

Following Taschereau-Dumouchel (2020) (hereafter, TD), we approach the problem of solving for non-union wages in this bargaining environment by first calculating the effect of losing h marginal units of labour and then sending h to zero in order to get the marginal contribution of a single worker. In doing this, we make use of the expression $n-h$ to refer to the removal of h workers from the number of hires n and write the firms surplus from having versus removing h workers as

$$S^n(h) = p_i y_{icf}(n^n_{icf}) - n^n_{icf} w^n_{icf}(n^n_{icf}) + \rho J^n_{icf}(n^n_{icf}) - [\pi(n^n_{icf}-h) + \rho J^n_{icf}(n^n_{icf}-h)] \tag{6}$$

with,

$$\pi(n_n - h) = p_i y_{icf}(n^n_{icf}-h) - (n^n_{icf}-h) w^n_{icf}(n^n_{icf}-h) \tag{7}$$

$$J^n_{icf}(n^n_{icf}-h) = p_i y_{icf}(n^n_{icf}) - n^n_{icf} w^n_{icf}(n^n_{icf}) - \kappa\frac{(n^n_{icf}-h)\delta}{q_{vc}} + \rho J^n_{icf}(n^n_{icf}) - \kappa\frac{h}{q_{vc}} \tag{8}$$

Substituting and rearranging yields,

$$\begin{aligned} S^n(h) = \left(p_i y_{icf}(n^n_{icf}) - n^n_{icf} w^n_{icf}(n^n_{icf})\right) - (p_i y_{icf}(n^n_{icf}-h) - (n^n_{icf}-h) w^n_{icf}(n^n_{icf}-h)) \\ + \frac{h\rho(1-\delta)\kappa}{q_{vc}} \end{aligned} \tag{9}$$

Dividing by h and taking the limit $\lim_{h\to 0}$ yields the following expression for the firm surplus

$$\lim_{h\to 0}\frac{S^n}{h} = S^n = p_i\frac{\partial y_{icf}(n^n_{icf})}{\partial n} - w^n_{icf}(n^n_{icf}) - n^n_{icf}\frac{\partial w^n_{icf}(n^n_{icf})}{\partial n} + \frac{\rho(1-\delta)\kappa}{q_{vc}} \tag{10}$$

A.2 Nonunion Wage Equation Derivation

Plugging the worker and firm surplus into the Nash bargaining condition gives the following:

$$\beta\left(p_i\frac{\partial y_{icf}(n^n_{icf})}{\partial n} - w^n_{icf}(n^n_{icf}) - n^n_{icf}\frac{\partial w^n_{icf}(n^n_{icf})}{\partial n} + \frac{\rho(1-\delta)\kappa}{q_{vc}}\right) = (1-\beta)\left(w^n_{icf}(n^n_{icf}) + \frac{\rho(1-\delta)w'}{1-\rho(1-\delta)} + \frac{(\rho-1)b}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))} + \frac{(\rho-1)\rho q_{uc}}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))}E^n_{ic}\right) \tag{11}$$

which yields a simple differential equation in wages

$$\begin{aligned} w^n_{icf}(n^n_{icf}) + \beta n^n_{icf}\frac{\partial w^n_{icf}(n^n_{icf})}{\partial n} &= \beta p_i\frac{\partial y_{icf}(n^n_{icf})}{\partial n} + \frac{\beta\rho(1-\delta)\kappa}{q_{vc}} \\ &- \frac{(1-\beta)\rho(1-\delta)w'_{icf}}{1-\rho(1-\delta)} + \frac{(1-\beta)(1-\rho)b}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))} \\ &+ \frac{(1-\beta)(1-\rho)\rho q_{uc}}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))}E^n_{ic} \end{aligned} \tag{12}$$

Solving this expression yields the following wage equation for non-union wages:

$$\begin{aligned} w^n_{icf}(n^n_{icf}) = \frac{\beta p_i}{1+\beta}\frac{\partial y_{icf}(n^n_{icf})}{\partial n} + \frac{\beta^2}{1+\beta}p_i\epsilon_{ic} + \frac{\beta\rho(1-\delta)\kappa}{q_{vc}} - \frac{(1-\beta)\rho(1-\delta)w'_{icf}}{1-\rho(1-\delta)} \\ + \frac{(1-\beta)(1-\rho)b}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))} + \frac{(1-\beta)(1-\rho)\rho q_{uc}}{(1-\rho(1-\delta))(1-\rho(1-q_{uc}))}E^n_{ic} \end{aligned} \tag{13}$$

In steady state this becomes

$$\begin{aligned} w^n_{icf}(n^n_{icf}) = \frac{1-\rho(1-\delta)}{1-\beta\rho(1-\delta)}\frac{\beta p_i}{1+\beta}\left(\frac{\partial y_{icf}(n^n_{icf})}{\partial n} + \beta\epsilon_{ic}\right) + \frac{\beta\rho(1-\delta)(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))}\frac{\kappa}{q_{vc}} \\ + \frac{(1-\beta)(1-\rho)}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}b + \frac{(1-\rho)\rho(1-\beta)q_{uc}}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}E^n_{ic} \end{aligned} \tag{14}$$

A.3 Firm Size Derivation

For union firms, the first order condition is:

$$\frac{\partial J^u_{icf}(n^u_{icf})}{\partial n} = p_i\frac{\partial y_{icf}(n^u_{icf})}{\partial n} - w^u_{icf} - n^u_{icf}\frac{\partial w^u_{icf}(n^u_{icf})}{\partial n} - \frac{\kappa\delta}{q_{vc}} = 0 \tag{15}$$

Using the quadratic production function and the expression for the union wage, this becomes:

$$\begin{aligned} p_i\left(\epsilon_{ic} - \sigma_i n^u_{icf}\right) - \frac{\beta p_i(1-\rho(1-\delta))}{1-\beta\rho(1-\delta)}(\epsilon_{icf} - \frac{1}{2}\sigma_i n^u_{icf}) \\ - D^u_{icf} + \frac{\beta p_i(1-\rho(1-\delta))}{1-\beta\rho(1-\delta)}\frac{1}{2}\sigma_i n^u_{icf} - \frac{\kappa\delta}{q_{vc}} = 0 \end{aligned} \tag{16}$$

where D^u_{icf} contains the elements of the union wage expression that do not vary with n. Rearranging this expression, we arrive at:

$$\begin{aligned} n^u_{icf} =& \frac{1}{\sigma_i p_i}\Bigg[p_i\epsilon_{ic} + \psi_{icf} - \frac{1-\beta\rho^2(1-\delta)^2}{1-\beta}\frac{\kappa}{q_{vc}} - \frac{1-\rho}{1-\rho(1-q_{uc})}b \\ & - \frac{(1-\rho)\rho q_{uc}}{1-\rho(1-q_{uc})}E^u_{ic}\Bigg] \end{aligned} \tag{17}$$

Similarly for nonunion firm size:

$$\frac{\partial J^n_{icf}(n^n_{icf})}{\partial n} = p_i\frac{\partial y_{icf}(n^n_{icf})}{\partial n} - w^n_{icf} - n^n_{icf}\frac{\partial w^n_{icf}}{\partial n} - \frac{\kappa\delta}{q_{vc}} = 0 \tag{18}$$

Using the production function and the nonunion wage expression, this becomes:

$$\begin{aligned} & p_i\left(\epsilon_{ic} - \sigma_i n^n_{icf}\right) - \frac{(1-\rho(1-\delta))}{1-\beta\rho(1-\delta)}\frac{\beta p_i}{1+\beta}(\epsilon_{icf} - \sigma_i n^n_{icf}) - D^n_{icf} \\ & \qquad + \frac{(1-\rho(1-\delta))}{1-\beta\rho(1-\delta)}\frac{\beta p_i}{1+\beta}\sigma_i n^n_{icf} - \frac{\kappa\delta}{q_{vc}} = 0 \end{aligned} \tag{19}$$

where D^n_{icf} contains the elements of the nonunion wage expression that do not vary with n. Rearranging this expression, we arrive at:

$$\begin{aligned} n^n_{icf} = & \frac{1+\beta}{(1+\beta\rho(1-\delta))}\cdot\frac{1}{\sigma_i p_i}\Bigg[p_i\epsilon_{ic} - \frac{(1-\beta\rho^2(1-\delta)^2)}{1-\beta}\frac{\kappa}{q_{vc}} - \frac{1-\rho}{1-\rho(1-q_{uc})}b \\ & - \frac{(1-\rho)\rho q_{uc}}{1-\rho(1-q_{uc})}E^n_{ic}\Bigg] \end{aligned} \tag{20}$$

A.4 Wage Equation Linearisation

First note that the contribution of firm size to the nonunion and union wage equations is given by:

$$\tilde{w}^n_{icf}(n) = -\frac{\beta}{1+\beta}\frac{1-\rho(1-\delta)}{1-\beta\rho(1-\delta)}\sigma p_i n \tag{21}$$

$$\tilde{w}^u_{icf}(n) = -\frac{\beta}{2}\frac{1-\rho(1-\delta)}{1-\beta\rho(1-\delta)}\sigma p_i n \tag{22}$$

where, for the firm size contribution to wages is smaller for union wages as $\beta \in (0,1)$. Plugging firm size into the respective nonunion and union wage equations gives:

$$\begin{aligned} w^n_{icf} = & \frac{\beta(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))}\cdot\frac{\beta\rho(1-\delta)}{1+\beta\rho(1-\delta)}p_i\epsilon_{ic} \\ & + \frac{1-\rho}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}\frac{1-\beta^2\rho(1-\delta)}{1+\beta\rho(1-\delta)}b \\ & + \frac{(1-\rho)\rho q_{uc}}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))}\frac{1-\beta^2\rho(1-\delta)}{1+\beta\rho(1-\delta)}E^n_{ic} \\ & + \frac{\beta(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))(1-\beta)}\frac{1+\rho(1-\delta)(1-\beta-\beta^2\rho(1-\delta))}{1-\beta\rho(1-\delta)}\frac{\kappa}{q_{uc}} \end{aligned} \tag{23}$$

$$
\begin{aligned}
w^u_{icf} = & \frac{\beta(1-\rho(1-\delta))}{2(1-\beta\rho(1-\delta))} p_i \epsilon_{ic} - -\frac{2-\beta-\rho(1-\delta)}{2(1-\beta\rho(1-\delta)} \psi_{icf} \\
& + \frac{1-\rho}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))} \frac{2-\beta-\rho(1-\delta)}{2} b \\
& + \frac{(1-\rho)\rho q_{uc}}{(1-\beta\rho(1-\delta))(1-\rho(1-q_{uc}))} \frac{2-\beta-\rho(1-\delta)}{2} E^u_{ic} \\
& + \frac{\beta(1-\rho(1-\delta))}{(1-\beta\rho(1-\delta))(1-\beta)} \frac{1+\rho(1-\delta)(2(1-\beta)-\beta\rho(1-\delta))}{2} \frac{\kappa}{q_{uc}}
\end{aligned}
\tag{24}
$$

We can conclude that for values of $\beta \in (0, 1)$, an increase in industrial prices, or a sectoral productivity shock will have larger increase on union wages. Following BGS, the rates of arrival can be expressed as functions of the city employment rate. We re-write the wage equations making explicit that the coefficients on the key variables are nonlinear functions of the employment rate:

$$
w^n_{icf} = \tilde{\beta}^n_1 p_i \epsilon_{ic} + \tilde{\beta}^n_{2c}(ER_c) b + \tilde{\beta}^n_{2c}(ER_c) \tilde{E}^n_{ic} + \tilde{\beta}^n_{3c}(ER_c) \tag{25}
$$

$$
w^u_{icf} = \tilde{\beta}^u_1 p_i \epsilon_{ic} + \tilde{\beta}^u_{2c}(ER_c) b + \tilde{\beta}^u_{2c}(ER_c) \tilde{E}^u_{ic} + \tilde{\beta}^u_{3c}(ER_c) + \tilde{\beta}^u_4 \psi_{icf} \tag{26}
$$

where $\tilde{E}^n_{ic}$ and $\tilde{E}^u_{ic}$ are outside options, following BGS expressed in terms of weighted averages over wages.

We take a linear approximation of the wage equations above with respect to the vector $[p_i, ER_c, \epsilon_{ic}, \tilde{E}^n_{ic}]$. We expand around the point where cities have a common industrial structure: $[p, ER, \epsilon, 0]$. Our final linearised wage equations are:

$$
w^n_{ic} = \gamma^n_{0i} + \gamma^n_1 \tilde{E}^n_{ic} + \gamma^n_2 ER_c + \gamma^n_4 \epsilon_{ic} \tag{27}
$$

and,

$$
w^u_{ic} = \gamma^u_{0i} + \gamma^u_1 \tilde{E}^u_{ic} + \gamma^u_2 ER_c - \gamma^u_3 \psi_{icf} + \gamma^u_4 \epsilon_{ic} \tag{28}
$$

where $\frac{\gamma^u_{0i}}{\gamma^n_{0i}} = \frac{\gamma^u_4}{\gamma^n_4} = K \geq 1$ such that $\gamma^u_{0i} = \gamma^n_{0i} K$ and $\gamma^u_4 = K \gamma^n_4$.

A.5 Firm Responses to a Unionisation Threat

A.5.1 Wage Responses

We can characterize the firm decision on whether to pay an emulation wage and, so, stay nonunion by examining the value of the firm if it pays this wage versus if it pays the union wage.

The value (in steady state) of the firm if it is unionised is:

$$
J^u_{icf} = \frac{1}{1-\rho} \left[p_i y_{icf}(n^u_{icf}) - n^u_{icf} w^u_{icf} - \frac{n^u_{icf} \delta \kappa}{q_{vc}} \right] \tag{29}
$$

While the value of the firm if it pays the wage to prevent unionisation is:

$$
J^*_{icf} = \frac{1}{1-\rho} \left[p_i y_{icf}(n^*_{icf}) - n^*(w^u_{icf} - \lambda_c + \psi_{icf}) - \frac{n^*_{icf} \delta \kappa}{q_{vc}} \right] \tag{30}
$$

where, n^*_{icf} is the optimal firm size when a nonunion firm pays a wage of w^*_{icf}.

Now, take the derivative of both value functions with respect to ψ_{icf}:

$$\frac{\partial J^u_{icf}}{\partial \psi} = \frac{1}{1-\rho}\left[\frac{\partial n}{\partial \psi}\left[p_i \frac{\partial y_{icf}}{\partial n} - w^u_{icf} - n^u_{icf}\frac{\partial w^u_{icf}}{\partial n} - \frac{\delta \kappa}{q_{vc}}\right] - n^u_{icf}\frac{\partial w^u_{icf}}{\partial \psi}\right] \tag{31}$$

and,

$$\frac{\partial J^*_{icf}}{\partial \psi} = \frac{1}{1-\rho}\left[\frac{\partial n}{\partial \psi}\left[p_i \frac{\partial y_{icf}}{\partial n} - w^*_{icf} - \frac{\delta \kappa}{q_{vc}}\right] - n^*_{icf}\frac{\partial w^u_{icf}}{\partial \psi} - n^*_{icf}\right] \tag{32}$$

Note that in both cases, the term in brackets multiplying $\frac{\partial n}{\partial \psi}$ is the first order condition associated with choosing n and, so, equals zero. As a result:

$$\frac{\partial J^u_{icf}}{\partial \psi} = \frac{1}{1-\rho}\left[-n^u_{icf}\frac{\partial w^u_{icf}}{\partial \psi}\right] \tag{33}$$

and,

$$\frac{\partial J^*_{icf}}{\partial \psi} = \frac{1}{1-\rho}\left[-n^*_{icf}\frac{\partial w^u_{icf}}{\partial \psi} - n^*_{icf}\right] \tag{34}$$

$\frac{\partial J^u_{icf}}{\partial \psi}$ is positive since $\frac{\partial w^u_{icf}}{\partial \psi}$ is negative. An increase in ψ reduces the wage that union firms have to pay and their profits increase as a result. But for emulation firms, an increase in ψ requires a one for one increase in the wage they have to paid (partially offset by the fact that the union wage they are trying to emulate has dropped). That is, $\frac{\partial J^*_{icf}}{\partial \psi} < 0$.

Note that at $\psi^u = \lambda_c - (w^u - w^n)$, workers are just indifferent between whether they organize or not. For $\psi_{icf} < \psi^u$, workers will not organize and the firm will be nonunion and will pay the nonunion wage derived earlier, w^n_{icf}. At $\psi_{icf} = \psi^u$, the wage a firm needs to pay to prevent unionisation is just w^n_{icf} and at that wage and associated optimal firm size, the value of the firm is greater than its value unionised. Therefore, at $\psi_{icf} = \psi^u$, $J^*_{icf} > J^u_{icf}$. With J^u_{icf} rising and J^*_{icf} declining with increases in ψ, there will be a point, $\tilde{\psi}$, at which $J^*_{icf} = J^u_{icf}$. This arises when $\psi = \lambda_c$ and, therefore, $w^*_{icf} = w^u_{icf}$.

A.5.2 Amenity Response

Firms could respond to the threat of unionisation through means other than raising wages. The first possibility is that they respond by increasing amenities for the workers. Since this is one of the things workers get out of a union, a direct response of this type seems possible. In our specification, workers utility on a union job is a linear function of the value of union amenities available to him if the firm is unionized with the value being expressed in dollar equivalents. In order for firms to want to provide amenities rather than wages as a means of resisting unionisation, the cost of providing the amenity must be less than or equal to the dollar valuation that the worker gets from the amenity. Otherwise, wages (increasing which costs the firm a dollar and gives the worker a dollar in value) will be a more cost-effective response. Assume, in particular, that providing amenities has an increasing and convex marginal cost function with the marginal cost of the initial units provided being below a dollar for one dollar worth of amenities as valued by the worker. In that case, nonunion firms would want to use amenities as a response to a union threat until the point where the marginal cost of a dollar's worth of amenities rises to one dollar. After that, they would respond through wage emulation.

However, if it is cost effective for a nonunion firm to use amenities to respond to a union threat, it would also be cost effective for it to pay in amenities instead of wages even in the absence of such a threat. Thus, if a threat emerges, the nonunion firm will already by providing amenities up to the point where their marginal cost equals a dollar. In that case, there is no room for the firm to respond to a union threat using amenities. Instead, it will respond through wage emulation.

A.5.3 Intimidation response

As a third potential response we allow firms to increase the costs to unionise for workers. Recall that workers in state s face a fixed cost of unionising, λ_s. Firms can increase that cost at a cost to themselves. For example, they could lock out the workers and either not produce or hire scabs who are less productive than the actual workers. The firm could also take legal actions to delay the union vote, imposing more costs on the workers.

The firm's threat, of course, needs to be credible. Recall that workers will choose to unionize if,

$$W^u_{icf}(w^u_{icf}) - W^n_{icf}(w^n_{icf}) - \lambda^*_s > 0 \tag{35}$$

or,

$$\frac{1}{1-\rho(1-\delta)}[w^u_{icf} + \psi_{icf} - w^n_{icf}] - \lambda_s > 0 \tag{36}$$

To halt unionization, the firm would have to pay a cost per worker of,

$$\lambda_f = \frac{1}{1-\rho(1-\delta)}[w^u_{icf} + \psi_{icf} - w^n_{icf}] - \lambda_s \tag{37}$$

At the moment that workers threaten to unionise, they know that firms will compare the value of the firm continuing as non-union minus the cost of carrying through on their intimidation to the value of the firm as a unionised firm. That is, they compare J^u_{icf} to:

$$J^b_{icf} = J^n_{icf} - n^n_{icf}[\frac{1}{1-\rho(1-\delta)}[w^u_{icf} + \psi_{icf} - w^n_{icf}] - \lambda^*_s] \tag{38}$$

$$= \frac{1}{1-\rho}[p_i y(n^n_{icf} - w^n_{icf}n^n_{icf} - \frac{\kappa\delta}{q_{vc}}n^n_{icf} - -n^n_{icf}[\frac{1}{1-\rho(1-\delta)}[w^u_{icf} + \psi_{icf} - w^n_{icf}] - \lambda^*_s]$$

If δ is small this is approximately equal to:[2]

$$J^b_{icf} \approx \frac{1}{1-\rho}[p_i y(n^n_{icf} - (w^u_{icf} + \psi_{icf} - \lambda_s)n^n_{icf} - \frac{\kappa\delta}{q_{vc}}n^n_{icf}] \tag{39}$$

$$= \frac{1}{1-\rho}[p_i y(n^n_{icf} - w^*_{icf}n^n_{icf} - \frac{\kappa\delta}{q_{vc}}n^n_{icf}]$$

[2]The equation when we do not set δ to zero is messier but the main logic continues to hold: the value function is equivalent to one where the firm is hiring workers as if its paying the nonunion wage but the actual wage is something else. With a non-zero δ, it is essentially a weighted average of w^*_{icf} and w^n_{icf}.

Several points follow from this equation. First, $J^b_{icf} < J^*_{icf}$ since J^b_{icf} is equivalent to a value function in which firms pay a wage of w^*_{icf} but do not hire optimally for that wage (instead, they hire n^n_{icf} workers). Second, at ψ_u, firms only need to pay a small cost to deter unionisation since workers are almost indifferent between being union and nonunion. That means that the value from implementing the intimidation is close to the value of being nonunion and, therefore, above the value of being a union firm. Third,

$$\frac{\partial J^b_{icf}}{\partial \psi} = \frac{1}{1-\rho}[\frac{\partial n^n}{\partial \psi}\left[p_i \frac{\partial y_{icf}}{\partial n} - w^*_{icf} - \frac{\delta \kappa}{q_{vc}}\right] - n^n_{icf}\frac{\partial w^u_{icf}}{\partial \psi} - n^n_{icf}] \tag{40}$$

Note that $\frac{\partial n^n}{\partial \psi} = 0$ and, so,

$$\frac{\partial J^b_{icf}}{\partial \psi} = \frac{1}{1-\rho}[-n^n_{icf}\frac{\partial w^u_{icf}}{\partial \psi} - n^n_{icf}] \tag{41}$$

With $w^*_{icf} \geq w^n_{icf}$, $n^*_{icf} \leq n^n_{icf}$ and so $\frac{\partial J^b_{icf}}{\partial \psi}$ is greater in absolute value than $\frac{\partial J^*_{icf}}{\partial \psi}$. Fourth, given these results, there is a cut-off value ψ_b below which firms would credibly intimidate and above which they would not resist unionisation. This cut-off is lower than the cut-off separating union emulation firms from unionised firms (which equals λ_s.

Together, these results imply that firms with $\psi_{icf} < \psi_u$ would be nonunion and pay the nonunion wage, w^n_{icf}. Firms with $\psi_u < \psi_{icf} \leq \psi_b$ would credibly threaten retaliation if workers tried to unionise and so remain non-union, paying the true nonunion wage, w^n_{icf}. Firms with $\psi_b < \psi_{icf} \leq \lambda_s$ would pursue wage emulation, paying w^*_{icf}. And firms with $\psi_{icf} > \lambda_s$ would become unionised and pay w^u_{icf}. But workers could recognize that $J^*_{icf} > J^b_{icf}$. In terms of estimation, there would still be a set of firms paying the nonunion wage, a set paying the union emulation wage, and a set paying the union wage. The model implies that the first set could be divided into two groups but it is not clear there is any advantage to doing so since we are just studying the overall nonunion wage and both of these subgroups pay the same wage. Moreover, the relevant cut-offs are:

$$\psi^u = \lambda_s - (w^u - w^n) = \frac{\lambda_s - (\alpha_{0i} + \gamma^u_0 E_u(w) - \gamma_1 E_n(w) + \alpha_2 ER_c) - \alpha_4 \epsilon}{(1-\gamma_3)} \tag{42}$$

and,

$$\psi^b = \frac{1}{n^n_{icf}}[p_i(y(n^n_{icf}) - y(n^u_{icf})) - w^u_{icf}(n^n_{icf} - n^u_{icf}) - \frac{\kappa \delta}{q_{vc}}(n^n_{icf} - n^u_{icf}) + \lambda_s n^n_{icf}] \tag{43}$$

The latter expression only implicitly identifies ψ^b because n^u_{icf} and w^u_{icf} are both functions of ψ. But the key point is that both are functions of the same variables - λ_s, the expected rents, and the employment rate. Given the data we have, there is no way for us to identify one of these thresholds from the other. To do so, we would need data that allows us to determine which firms are under threat of intimidation if workers tried to unionise and which are not.

There is an assumption under which the firm intimidation option becomes irrelevant. Suppose that workers know that $J^*_{icf} > J^b_{icf}$. (Note that this does not imply that firms will just automatically choose to use emulation because J^n_{icf} is bigger than both of them and under a credible intimidation approach, firms get that value since they never have to actually spend the money to force unionisation costs higher given that workers won't try to unionise.) Workers then could threaten to unionise in each period, getting firms to choose emulation over intimidation. If the cost of threatening is low enough then we would not see intimidation, only emulation. We will proceed as if that is the situation and only emulation exists but note that our key conclusions do not change if intimidation is being used. In particular, the direction of change in response to increases in λ_s is the same, i.e., wage emulation will decline and the observed nonunion wage will decline for that reason alone.

B Data Appendix

Our CPS data is downloaded from the National Bureau of Economic Research (NBER).

We construct potential experience as max(min(age-years of schooling-6, age-16),0), dropping those with negative potential experience. We use the approach in Jaeger and Page (1996) to convert the years of completed schooling recorded in the MORG prior to 1992 to the post-1992 education categories. Because of limitations in the union coverage question, we define union workers as workers reporting belonging to a labour union.

We follow Lemieux (2006) closely in the construction of our wage data, working with weekly wages. Specifically, wages are based on individuals reporting employment in the reference week as wage and salary workers. We drop observations with allocated wages, and for workers paid hourly we use hourly earnings multiplied by usual weekly hours worked. For workers not paid hourly, we use edited weekly earnings, multiplying the weekly earnings topcode by 1.4 for topcoded observations. Wages are converted to 2000 dollars using a CPI deflator. We drop observations with an hourly wage below 1 or greater than 100 in 1979 dollars. All calculations use the earnings weights provided in the data. We aggregate highest degree obtained into four categories (less than high school, high school graduate, some post secondary, and university degree). For years before 1992, we use Table 5 from Park (1994) to construct education categories from the number of completed years of education.

We define industry using an aggregated grouping of industry codes based on the 1980 industrial classification from the Census Bureau. We obtain a consistent industry classification using crosswalks provided by IPUMS and the Census Bureau that map the 1970, 1990, and 2000 industry codes to the 1980 classification.[3] The result is a consistent classification system with 51 industries.[4]

We construct a set of cities with as consistent geographic boundaries as possible given data limitations in the CPS. We are constrained by the number of SMSA's available in the May extract data and end up with 43 cities. Making use of the limited number of counties identified in the CPS, we are able to create a set of cities which are reasonably, though not always perfectly, consistent over time.[5] The final geographic definition we use pools data

[3] Available at `https://www.census.gov/topics/employment/industry-occupation/guidance/code-lists.html` and `https://usa.ipums.org/usa/volii/occ_ind.shtml`

[4] Appendix table Table 6 shows the relationship between this detailed industry definition and the 1990 industrial classification system used by the Census Bureau.

[5] The metropolitan area definition used by the IPUMS identifies a general pattern of expanding metropolitan area definitions over time that we overcome to some extent, but not perfectly: `https://usa.ipums.`

for these 43 cities and the remaining population. Specifically, we create additional regions made up of the remaining state population absent the population living in these 43 cities. In the end, our core geographic measure is composed of 93 areas that are fairly consistently defined over the course of the sample period.

Additionally, we use data on union elections to proxy for the costs of unionisation, λ_{ct} in our model. The idea is that locations where the proportion of union certification elections that result in a certification is high are more union friendly. To obtain these proportions we use National Labor Relations Review Board (NLRB) case data for the three year periods for which we use CPS data.[6] We focus on certification elections and cases where a conclusive decision on certification was reached.[7] We use the county of the unit involved in the election to construct our geographic measures, aggregating counties to our city definition discussed above.

B.1 Emulation Probabilities

The procedure to construct $P(\psi_{icf}, \lambda_c)$

1. Using NLRB data, count the number of elections in each ic cell from 1977-2010, NE_{ic}.
2. Using the CBP data, count the number of establishments in each ic cell from 1977-2010, $Estab_{ic}$
3. Using the CPB data, construct a vector of predictor variables $\mathbf{X}_{ic}$ that contains the number of workers in each ic, its square, and variables capturing the number of establishments in Employment Class Size 1-4, 5-9,10-19,20-49,50-99,100-249,250- 499, 1000+, 1000-1499,1500-2499,2500-4999,5000+. Employment counts are taken from Eckert et. al.'s web page.
4. Using CPS data, we construct a vector of predictor variables $\mathbf{W}_c$ containing the unemployment rate, non-participation rate, average age, fraction or workers who are unionised and all of their squares,
5. We fit a negative binomial regression where NE_{ic} is the dependent variable and predictor variables include $\mathbf{W}_c$, $\mathbf{X}_{ic}$, and industry, year, and city fixed effects using $Estab_{ic}$ as the exposure variable.
6. Using the predicted values from above regression, we calculate the likelihood of an election occurred in an ic cell over 4 year periods prior estimation years.

Our proxy is constructed by first calculating the ratio of the predicted number of elections to the number of establishments in an ic cell. Next, we use this ratio to calculate the probability that an establishment had least one election in the past four years. The predicted number of elections come from a negative binomial regression, where the observed number

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`org/usa/volii/county_comp2b.shtml`. Estimation using states as the geographic unit yields very similar results, suggesting that issues related to geographic definitions are not driving our results.

[6] Our thanks to Hank Farber for providing this data.

[7] As opposed to the case being dismissed or withdrawn.

of elections are regressed on a number of variables at the ic level. These variables include polynomials in the number of workers in ic, unemployment rate, participation rate, and average age at the c level, variables capturing the size of establishments, and industry, year and city fixed effects. The predicted number of elections come from the fitted values of the model.

B.2 Union Transition rates

To create the T_i^u and T_i^n variables, we construct transitions using additional data from IPUMS-CPS. For years 1990, 2000, 2010, and 2016 all transitions are constructed using IPUMS data, which contains a necessary unique identification variable. For 1980, we match IPUMS identification data to the May extracts, as union data is not contained in IPUMS for these years. We perform the match using household identifiers and personal characteristics. It is not possible to track individuals for most of 1981 and for all of 1982 in the May extracts. To overcome this limitation we extend the range of years used to calculate transitions. Using the May extracts we match individuals from 1977 to 1981, and we match individuals from 1983 to 1984 using the MORG data.

B.3 Endogeneity and Instrument Construction

In order to construct our instruments,$IV1_{jct}^n$ and $IV1_{jct}^u$, we need (1) estimates of the national industrial premia, (2) to predict local union and nonunion employment composition (since the η's are potentially correlated to the error terms in (11) and (16), and (3) to predict local union transition rates, since these terms depend on local union proportions.

Estimating national wage premia. In particular, we estimate separate log wage regressions for each of our set of sample years at the national level, working with pooled union and nonunion workers. The regressions include the same set of skill and demographic variables used when forming our residualized wages for the dependent variable plus a complete set of industry dummy variables interacted with a union dummy. We interpret the coefficients on the industry dummies as rents that are allowed to differ in the union and nonunion sectors. We define the industry dummy variables such that the coefficient values are defined relative to the overall average wage. We then replace the wages, w_{ic}^u and w_{ic}^n, with the industry premia, which we call ν_i^u and ν_i^n, in the outside option expressions.

Prediction local industrial composition. There is clear reason to be concerned about such a correlation for the employment rate variable, which is at the city level of aggregation. In addition, BGS show that local industrial composition captured in the η_{ic}^u and η_{ic}^n terms in the outside options value expression, (19), can be written as functions of ϵ_{ict}. Whether this implies an endogeneity problem depends on the time series processes of the productivity shocks. If they follow a random walk specification in which the changes in ϵ_{ict} are independent of their levels, then there is no endogeneity issue with this variable. Otherwise, there is reason to treat it as potentially endogenous.

We construct the predicted nonunion industrial shares in steps. We first construct predicted employment levels using start-of-period levels at the city level combined with national level growth rates for the relevant industry:

$$\hat{N}^n_{ict} = N^n_{ict-1} \cdot \left(\frac{N^n_{it}}{N^n_{it-1}} \right)$$

We then form predicted city level employment as $\hat{N}^n_{ct} = \sum_i \hat{N}^n_{ict}$ and, from that, we construct predicted employment shares as $\hat{\eta}^n_{ict} = \frac{\hat{N}^n_{ict}}{\hat{N}^n_{ct}}$.

Predicting union transition rates. As shown in (18), T^n_{jct} is constructed from a combination of national-level probabilities of nonunion workers from a given industry finding a union job and the local proportion of workers who are unionised. The first of these varies at the national-industrial level and so, in our regressions including industry fixed effects, it does not represent a problematic source of variation. This is not the case for the locally defined union proportion term. To address this term, it is useful to make clear how the union proportion term is tied to the local industrial structure:

$$P_{ct} = \frac{\sum_i N^u_{ict}}{\sum_{s=u,n} \sum_i N^s_{ict}}$$

Given this, it is possible to predict changes in local unionisation using national changes in the composition of work. That is, we use national level predictions in employment in jobs to predict local employment growth. If there are declines in union employment in sectors with high local employment, then this will predict a decline in the local union proportion. Our predicted union proportion term is calculated as:

$$\hat{P}_{ct} = \frac{\sum_i \hat{N}^u_{ict}}{\sum_{s=u,n} \sum_i \hat{N}^s_{ict}}$$

where, $\hat{N}^u_{ict}$ is constructed in the same way as $\hat{N}^n_{ict}$, as described above. Replacing P_{ct} with $\hat{P}_{ct}$ yields $\hat{T}^n_{jct}$.

C Additional Results

In this Appendix, we present alternative estimates of our main specifications. In the first part, we re-form both our dependent variable and our outside option variables and instruments, dropping public sector industries. In the second part, we show our main specification results when we define location by state rather than by the SMSA x state approach we use in our main data.

C.1 Results without Public Sector

Table 1: Non-Union Wages and Outside Options: OLS and 2SLS Estimates

	OLS			2SLS			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔP_{ic}	0.29***	0.033*	0.020	0.020			
	(0.026)	(0.018)	(0.016)	(0.016)			
$\Delta((1-T^N_{ict})\cdot\bar{w}^N_{ct})$			0.91***	0.70***	0.70***	0.58***	
			(0.016)	(0.055)	(0.056)	(0.087)	
$\Delta(T^N_{ict}\cdot\bar{w}^U_{ct})$			0.78***	0.65***	0.65***	0.80***	
			(0.031)	(0.055)	(0.055)	(0.099)	
ΔE^N_{ict}							0.69***
							(0.053)
ΔER_c			-0.11**	0.11	0.11	0.26**	0.13
			(0.051)	(0.087)	(0.088)	(0.12)	(0.086)
Observations	8273	8273	8273	8273	8273	8273	8273
R^2	0.029	0.404	0.551	0.235	0.235	0.215	0.233
Year × Ind.	No	Yes	Yes	Yes	Yes	Yes	Yes
Instrument set:				$IV1^N_{jct}$	$IV1^N_{jct}$	$IV2^N_{jct}$	$IV2^N_{jct}$
				$IV1^U_{jct}$	$IV1^U_{jct}$	$IV2^U_{jct}$	$IV2^U_{jct}$
First-Stage F-Stat.:							
$\Delta((1-T^N_{ict})\cdot\bar{w}^N_{ct})$				44.51	44.53	27.31	
$\Delta(T^N_{ict}\cdot\bar{w}^U_{ct})$				164.62	160.19	66.17	
ΔE^N_{ict}							25.81
Over-id. p-val				.	.	.	0.06

Notes: This table displays results from the estimation of equations (22) (columns 1 and 2) and (5) (columns 3 - 7) via OLS (columns 1 - 3) and 2SLS (columns 4 - 7). The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

Table 2: Non-Union Wages and Outside Options: Including Wage Emulation Effects

	2SLS				
	(1)	(2)	(3)	(4)	(5)
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)$	0.71***	0.69***	0.71***	0.70***	0.70***
	(0.058)	(0.060)	(0.059)	(0.058)	(0.059)
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)$	0.66***	0.72***	0.66***	0.58***	0.53
	(0.083)	(0.085)	(0.083)	(0.10)	(0.79)
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)\cdot RTW$				0.10***	
				(0.034)	
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)\cdot RTW$				-0.31*	
				(0.16)	
ΔER_c	0.31**	0.34***	0.31**	0.26**	0.42***
	(0.13)	(0.13)	(0.13)	(0.13)	(0.13)
Observations	6284	6284	6284	6284	6284
R^2	0.270	0.267	0.270	0.270	0.267
Year × Ind.	Yes	Yes	Yes	Yes	Yes
$P(\psi_{icf},\lambda_c)\times$ Ind.					Yes
Instrument set:	$IV3^N_{jct}$	$IV4^N_{jct}, IV5^N_{jct}$	$IV3^N_{jct}$	$IV3^N_{jct}, IV3^N_{jct}\times RTW$	$IV3^N_{jct}$
	$IV3^U_{jct}$	$IV4^U_{jct}, IV5^U_{jct}$	$IV3^U_{jct}$	$IV3^U_{jct}, IV3^U_{jct}\times RTW$	$IV3^U_{jct}$
Select controls					
ΔP_{ic} Quartic			Yes		
First-Stage F-Stat.:					
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)$	50.85	20.65	50.51	50.47	46.62
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)$	25.71	10.97	25.80	21.11	65.12
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E^N_{ic}\right)\cdot RTW$				266.48	
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E^U_{ic}\right)\cdot RTW$				45.07	
Over-id. p-val	.	0.01	.	.	.

Notes: This table displays results from the estimation of equation (16) via 2SLS. The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

C.2 Results using States

Table 3: Non-Union Wages and Outside Options: OLS and 2SLS Estimates

	OLS			2SLS			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔP_{ic}	0.43***	0.044**	0.024	0.025			
	(0.037)	(0.020)	(0.016)	(0.016)			
$\Delta((1 - T^N_{ict}) \cdot \bar{w}^N_{ct})$			0.92***	0.68***	0.69***	0.65***	
			(0.019)	(0.087)	(0.086)	(0.10)	
$\Delta(T^N_{ict} \cdot \bar{w}^U_{ct})$			0.77***	0.64***	0.65***	0.66***	
			(0.032)	(0.061)	(0.060)	(0.073)	
ΔE^N_{ict}							0.65***
							(0.072)
ΔER_c			-0.12*	0.13	0.12	0.16	0.16
			(0.059)	(0.12)	(0.12)	(0.13)	(0.11)
Observations	6425	6425	6425	6425	6425	6425	6425
R^2	0.055	0.590	0.694	0.238	0.239	0.233	0.234
Year × Ind.	No	Yes	Yes	Yes	Yes	Yes	Yes
Instrument set:				$IV1^N_{jct}$	$IV1^N_{jct}$	$IV2^N_{jct}$	$IV2^N_{jct}$
				$IV1^U_{jct}$	$IV1^U_{jct}$	$IV2^U_{jct}$	$IV2^U_{jct}$
First-Stage F-Stat.:							
$\Delta((1 - T^N_{ict}) \cdot \bar{w}^N_{ct})$				18.57	18.46	15.91	
$\Delta(T^N_{ict} \cdot \bar{w}^U_{ct})$				84.41	82.76	74.51	
ΔE^N_{ict}							13.40
Over-id. p-val				.	.	.	0.92

Notes: This table displays results from the estimation of equations (22) (columns 1 and 2) and (5) (columns 3 - 7) via OLS (columns 1 - 3) and 2SLS (columns 4 - 7). The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

Table 4: Non-Union Wages and Outside Options: Including Wage Emulation Effects

	2SLS				
	(1)	(2)	(3)	(4)	(5)
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E_{ic}^N\right)$	0.67***	0.68***	0.66***	0.67***	0.61***
	(0.074)	(0.076)	(0.076)	(0.072)	(0.100)
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E_{ic}^U\right)$	0.76***	0.77***	0.75***	0.63***	1.48
	(0.11)	(0.11)	(0.12)	(0.14)	(1.34)
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E_{ic}^N\right)\cdot RTW$				0.070**	
				(0.029)	
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E_{ic}^U\right)\cdot RTW$				-0.32**	
				(0.16)	
ΔER_c	0.36**	0.33**	0.37**	0.30*	0.55**
	(0.16)	(0.16)	(0.16)	(0.16)	(0.21)
Observations	4791	4791	4791	4791	4791
R^2	0.267	0.269	0.268	0.269	0.254
Year × Ind.	Yes	Yes	Yes	Yes	Yes
$P(\psi_{icf},\lambda_c)\times$ Ind.					Yes
Instrument set:	$IV3_{jct}^N$	$IV4_{jct}^N, IV5_{jct}^N$	$IV3_{jct}^N$	$IV3_{jct}^N, IV3_{jct}^N\times RTW$	$IV3_{jct}^N$
	$IV3_{jct}^U$	$IV4_{jct}^U, IV5_{jct}^U$	$IV3_{jct}^U$	$IV3_{jct}^U, IV3_{jct}^U\times RTW$	$IV3_{jct}^U$
Select controls					
ΔP_{ic} Quartic			Yes		
First-Stage F-Stat.:					
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E_{ic}^N\right)$	25.91	8.99	25.81	22.81	20.20
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E_{ic}^U\right)$	34.43	17.34	34.36	19.32	32.48
$\Delta\left((1-P(\psi_{icf},\lambda_c))\cdot E_{ic}^N\right)\cdot RTW$				132.98	
$\Delta\left(P(\psi_{icf},\lambda_c)\cdot E_{ic}^U\right)\cdot RTW$				21.84	
Over-id. p-val	.	0.24	.	.	.

Notes: This table displays results from the estimation of equation (**??**) via 2SLS. The dependent variable is the decadal change in the regression adjusted average hourly wage of nonunion workers in an industry-city cell, using CPS data from 1980-2019 across 50 industries and 93 cities. Standard errors, in parentheses, are clustered at the city-year level.

C.3 Decomposition of Non-union wages

Table 5: Outside Options Contribution to Changing Wages - Subsample Analysis

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	All	Men	Women	Age 36-55	Age 20-26	$\leq$ HS	$>$ HS
				1980-2010			
Observed	–0.126	–0.179	–0.066	–0.246	–0.161	–0.193	–0.075
Fixed Threat	–0.008	–0.008	–0.008	–0.010	–0.006	–0.006	–0.010
Fixed Transitions	0.022	0.025	0.018	0.021	0.025	0.038	0.014
Fixed η^U	0.006	0.002	–0.001	0.008	0.001	0.009	–0.005
Fixed w^U	–0.002	–0.004	0.004	–0.006	–0.003	–0.008	–0.006
Total	0.018	0.016	0.013	0.013	0.017	0.033	–0.007
Total/Observed	–0.142	–0.087	–0.197	–0.053	–0.108	–0.171	0.095
				1980-1990			
Observed	–0.140	–0.181	–0.094	–0.362	–0.180	–0.174	–0.103
Fixed Threat	–0.004	–0.005	–0.004	–0.005	–0.003	–0.003	–0.005
Fixed Transitions	0.016	0.016	0.015	0.014	0.017	0.024	0.008
Fixed η s	0.002	0.001	0.000	0.003	0.003	0.002	–0.001
Fixed w^U	–0.009	–0.008	–0.011	–0.010	–0.009	–0.014	–0.009
Total	0.005	0.005	0.001	0.002	0.008	0.009	–0.008
Total/Observed	–0.036	–0.027	–0.006	–0.005	–0.047	–0.051	0.077

Notes: This table displays results from the decomposition for non-union workers.

D Online Appendix: City and Industry Construction

Table 7: Changes to SMSA Definitions 1973-2010

	1973-1980		1981-1989	1993-2003	2004-2020
Chicago	Cook	Lake	Kendall Added		
	Du Page	McHenry	Grundy Added	Dekalb Added	
	Kane	Will			
Philadelphia	Burlington	Chester		Salem Added	
	Camden	Delaware			
	Gloucester	Montgomery			
	Bucks	Philadelphia			
Detroit	Lapeer	Oakland	Monroe Added	Lenawee Added	

	Livingston	St.Clair			Washtenaw Added	
	Macomb	Wayne				
Washington	District of Columbia	Arlington	Calvert Added		Fauquier Added	King George Dropped
	Montgomery	Fairfax	Charles Added		Clarke & Warren Added	Rappahannock Added
	Prince George's	Fairfax city	Frederick Added		Culpeper Added	
	Alexandria	Falls Church	Loudoun Added		King George Added	
			Prince William Added		Spotsylvania Added	
			Masassas Added		Jefferson Added	
			Masassas Park Added		Fredericksburg Added	
			Stafford Added		Berkeley Added	
Boston	Essex	Plymouth	Bristol Added			Bristol Dropped
	MiddleSex	Suffolk				Essex Dropped
	Norfolk		Worchester Added			
Pittsburgh	Allegheny	Washington	Fayette Added		Butler Added	Armstrong Added
	Beaver	Westmoreland				
St Louis	Clinton	Jefferson	Jersey Added		Lincoln Added	Macoupin Added
	Madison	St. Charles			Warren Added	Bond Added
	Monroe	St. Louis				Calhoun Added
	St. Clair Franklin	St. Louis city				
Baltimore	Anne Arundel	Carroll	Queen Anne's Added			
	Baltimore city	Harford				
	Baltimore	Howard				
Cleveland	Cuyahoga	Lake			Added Ashtabula	
	Geauga	Medina			Added Lorain	
Houston	Brazoria	Liberty			Added Chambers	Added Austin
	Fort Bend	Montgomery				Added Galveston
	Harris	Waller				
Newark	Essex	Sussex				Union Dropped
	Morris	Union				
Minneapolis-	Anoka	Ramsey	Isanti Added		Sherburne Added	
St Paul	Carver	Scott				
	Chisago	Washington				
	Dakota	Wright				
	Hennepin					
Dallas-	Collin	Wise	Wise Dropped		Henderson Added	Wise Added
Fort Worth	Dallas	Hood	Hood Dropped		Hunt Added	Somerwell Added
	Denton	Johnson			Hood Added	

	Ellis	Tarrant				
	Kaufman	Parker				
	Rockwall					
Seattle-Everett	King	Snohomish			Island Added	Pike Added
Atlanta	Cherokee	Gwinnett	Barrow Added		Butts dropped	Butts Added
	Clayton	Henry	Coweta Added		Carroll Added	Dawson Added
	Cobb	Newton	Spalding Added		Bartow Added	Haralson Added
	De Kalb	Paulding				Heard Added
	Douglas	Rockdale				Jasper Added
	Fayette	Walton				Lamar Added
	Forsyth	Butts				Meriwether Added
	Fulton					Morgan Added
Cincinnati	Dearborn	Clermont			Ohio Added	Union Added
	Boone	Hamilton			Gallatin Added	Bracken Added
	Campbell	Warren			Grant & Brown Added	Butler Added
	Kenton				Pendelton Added	
Kansas City	Johnson	Jackson	Lafayette Added		Clinton Added	Linn Added
	Wyandotte	Platte	Leavenworth Added			Bates Added
	Cass	Ray	Miami Added			Caldwell Added
	Clay					
Denver	Adams	Denver				Adams Dropped
	Arapahoe	Douglas				Broomfield Added
	Boulder	Jefferson				Clear Creek Added
						Elbert & Park Added
						Gilpin Added
Indianapolis	Boone	Johnson				Brown Added
	Hamilton	Marion				Putnam Added
	Hancock	Morgan				
	Hendricks	Shelby				
New Orleans	Jefferson	St. Bernard	St Charles Added		St James Added	
	Orleans	St.Tammany	St John the Bap. Added		Plaquemines Added	
Tampa-	Hillsborough	Pinellas	Hernando Added			
St Petersburg	Pasco					
Portland	Clackamas	Washington			Clark Added	
	Multnomah	Yamhill			Columbia Added	
Columbus	Delaware	Madison	Licking Added		Licking Dropped	Licking Added
	Fairfield	Pickaway	Union Added			Hocking Added
	Franklin					Morrow Added
Rochester	Livingston	Orleans			Genesee Added	

	Monroe	Wayne				
	Ontario					
Sacramento	Placer	Yolo			El Dorado Added	
	Sacramento					
Birmingham	Jefferson	Walker	Blount Added		Walker Dropped	Walker Added
	Shelby	St. Clair				Bibb & Chilton Added
Albany-	Albany	Schenectady	Greene Added		Greene Dropped	
Schenectady-Troy	Rensselaer	Montgomery			Schoharie Added	
	Saratoga					
Norfolk-	Currituck	Portsmouth	Currituck Dropped		Currituck Added	
Portsmouth	Chesapeake	Virginia Beach	Gloucester Added		Isle of Wight Added	
	Norfolk		Hampton & Suffolk Added James & York Added Newport News Added Poquoson Added Williamsburg Added		Mathews Added	Gloucester Added
Greensboro-	Forsyth	Yadkin	Davie Added		Alamance Added	Alamance Dropped
Winston-Salem-	Guilford	Stokes				
High point	Randolph	Davidson				
Gary-Hammond	Lake	Porter				Jasper Added
East Chicago						Newton Added
Portland	Clackamas	Multnomah			Columbia Added	
	Washington	Yamhill				

Notes: Changes to the counties/cities/parishes, included in the SMSA definitions over the sample period. There are no county changes for New York, Patterson, Nassau-Suffolk, Los Angeles, San Francisco, Anaheim, Milwaukee, San Diego, Buffalo, Miami, San-Bernadino, San Jose, Akron.

Table 6: SMSA Rankings

1980 Rank	SMSA	1980 Rank	SMSA
1	New York, NY	23	Patterson-Clifton-Passaic, NJ
2	Los Angeles-Long Beach, CA	24	San Diego, CA
3	Chicago, IL	25	Buffalo, NY
4	Philadelphia, PA	26	Miami, FL
5	Detroit, MI	27	Kansas City, MO, KS
6	San Francisco-Oakland, CA	28	Denver, CO
7	Washington, DC, MD, VA	29	San Bernardno-Riverside-Ontario, CA
8	Boston, MA	30	Indianapolis, IN
9	Nassau-Suffolk, NY	31	San Jose, CA
10	Pittsburgh, PA	32	New Orleans, LA
11	St Louis, MO, IL	33	Tampa- St Petersburg, FL
12	Baltimore, MD	34	Portland, OR
13	Cleveland, OH	35	Columbus, OH
14	Houston, TX	36	Rochester, NY
15	Newark, NJ	37	Sacramento, CA
16	Minneapolis-St Paul, MN	38	Birmingham, AL
17	Dallas-Fort Worth, TX	39	Albany-Schenectady-Troy, NY
18	Seattle-Everett, WA	40	Norfolk-Portsmouth, VA
19	Anaheim-Santa Ana-, Garden Grove, CA	41	Akron, OH
		42	Gary-Hammond-East Chicago, IN
20	Milwaukee, WI	43	Greensboro-Winston-Salem- High Point, NC
21	Atlanta, GA		
22	Cincinnati, OH		

Notes: SMSAs consistently available from 1978-2010, ranked by population size in 1980.

Table 8: Aggregated Industry Definitions

Category	Code	1990 Industry Codes
Agriculture Service	1	12, 20, 21 , 30
Other Agriculture	2	10 - 11
Mining	3	40 - 50
Construction	4	60
Lumber and Wood Products, except Furniture	5	230 - 241
Furniture and Fixtures	6	242
Stone Clay, Glass, and Concrete Product	7	250 - 262
Primary Metals	8	270 - 280
Fabricated Metal	9	281 - 300
Not Specified Metal Industries	10	301
Machinery, except Electrical	11	310 - 332
Electrical Machinery, Equipment, and Supplies	12	340 - 350
Motor Vehicles and Equipment	13	351
Aircraft and Parts	14	352
Other Transportation Equipment	15	360 - 370
Professional and Photographic Equipment, and Watches	16	371 - 382
Toys, Amusements, and Sporting Goods	17	390
Miscellaneous and Not Specified Manufacturing Industries	18	391 - 392
Food and Kindred Products	19	100 - 122
Tobacco Manufactures	20	130
Textile Mill Products	21	132 - 150
Apparel and Other Finished Textile Products	22	151 - 152
Paper and Allied Products	23	160 - 162
Printing, Publishing and Allied Industries	24	171 - 172
Chemicals and Allied Products	25	180 - 192
Petroleum and Coal Products	26	200 - 201
Rubber and Miscellaneous Plastics Products	27	210 - 212
Leather and Leather Products	28	220 - 222
Transportation	29	400 - 432
Communications	30	440 - 442
Utilities and Sanitary Services	31	450 - 452, 460 - 472
Wholesale Trade	32	500 - 571
Retail Trade	33	580 - 691
Banking and Other Finance	34	700 - 710
Insurance and Real Estate	35	711 - 712
Private Household Services	36	761
Business Services	37	721, 722, 731 - 750, 892
Repair Services	38	751 - 760
Personal Services, except Private Household	39	762 - 791
Entertainment and Recreation Services	40	800 - 802, 810
Hospitals	41	831
Health Services, except Hospitals	42	812 - 830, 832 - 840
Educational Services	43	842 - 860
Social Services	44	861 - 871
Other Professional Services	45	730, 841, 872 - 891, 893
Forestry and Fisheries	46	31 - 32
Justice, Public Order and Safety	47	910
Administration Of Human Resource Programs	48	922
National Security and Internal Affairs	49	932
Other Public Administration	50	900, 901, 921, 930, 931

Notes: List of aggregated industries and corresponding 1990 codes used by the US Census Bureau.

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