Unions, Norms, and the Rise in American Wage Inequality

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Abstract

From 1973 to 2007, private sector union membership in the United States declined from 34 to 8 percent for men and from 16 to 6 percent among women. Inequality in hourly wages increased by over 40 percent in this period. We report a decomposition, relating rising inequality to the shrinking weight of the union wage distribution. We also argue that unions helped institutionalize norms of equity reducing the dispersion of nonunion wages in highly unionized regions and industries. Accounting for the effect of unions on union and nonunion wages suggests that the decline of organized labor explains a fifth to a third of the growth in inequality—an effect comparable to the growing stratification of wages by education.

The decline of organized labor in the United States coincided with a large increase in wage inequality. From 1973 to 2007 union membership in the private sector declined from 34 to 8 percent for men and from 16 to 6 percent for women. During this time, wage inequality in the private sector increased by over 40 percent. Union decline forms part of an institutional account of rising inequality that is often contrasted with a market explanation. In the market explanation, technological change, immigration, and foreign trade increased demand for highly-skilled workers, raising the premium paid to college graduates (for reviews, see Gottschalk and Danziger 2005; Autor, Katz, and Kearney 2008; Lemieux 2008).

Compared to market forces, union decline is often seen as a modest source of rising inequality (Autor, Katz, and Kearney 2008, 311). The effects of unions were viewed as indirect, mediating the influence of technological change (Acemoglu 2002), secondary, to other institutions like the minimum wage (Card and DiNardo 2002; DiNardo, Fortin, and Lemieux 1996), and limited, accounting for only a small fraction of rising inequality and only among men (Card, Lemieux, and Riddell 2004).

We revisit the effects of union decline on inequality offering two extensions to earlier research. First, we study the effects of union decline while controlling for education and other factors. Analyzing education alongside unions allows a comparison of market and institutional effects on rising inequality. Second, we examine union effects on nonunion wages, considering whether wage inequality is lower among nonunion workers in regions and industries that are highly unionized.

Union effects on unions on nonunion workers can be motivated in several ways. Nonunion employers may raise wages to avert the threat of union organization (Leicht 1989). We argue that unions also contribute to a moral economy that institutionalizes norms for fair pay, even for nonunion workers. In the early 1970s, when 1 in 3 male workers were organized, unions were often prominent voices for equity, not just for their members, but for all workers. Union decline marks an erosion of the moral economy and its underlying distributional norms. Wage inequality in the nonunion sector increased as a result.

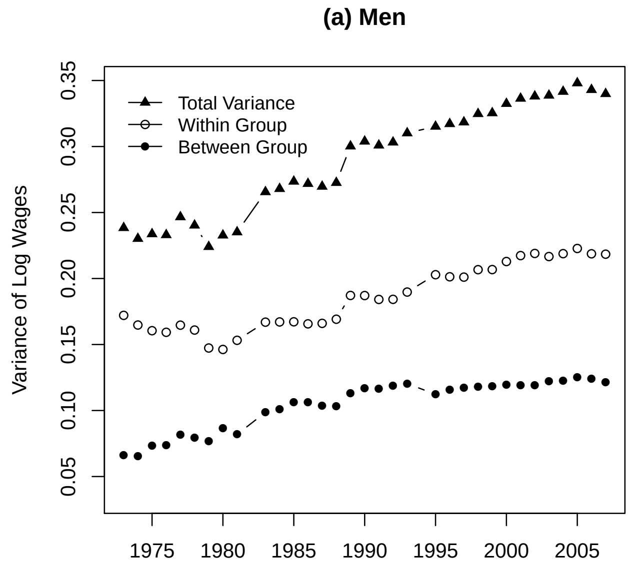
Our analysis estimates union effects on wage inequality by decomposing the growth in hourly wage inequality for full-time workers in the private sector. Analysis of the Current Population Survey (CPS) shows that union decline explains a fifth of the increase in inequality among men and none of the increase among women if only union wages are considered. The effect of union decline grows when we account for the link between unionization and nonunion wages. In this case, deunionization explains a fifth of the inequality increase for women, and a third for men. The decline of organized labor among men contributes as much to rising wage inequality as the growing stratification of pay by education.

Trends in Wage Inequality and Unionization

We analyze trends in the private sector—about 85 percent of nonfarm employment (US Census Bureau 2007, 403)—where the growth in inequality and the decline in union density was largest. Using data from the May and Merged Outgoing Rotation Group files of the CPS, inequality is measured by the variance in log hourly wages for men and women working full-time in private sector jobs. From 1973 to 2007, men's wage inequality increased by 40 percent, from .25 to .35, with most of the rise unfolding from 1978 to 2000 (Figure 1). Women's wage inequality increased by more, rising from .20 in 1973 to .30 in 2007. Overall trends were driven by movements at the top and the bottom of the wage distribution. Increasing inequality in the late 1970s and 1980s reflected falling wages at the bottom and rising wages at the top of the distribution. Since the late 1980s, the growth in wage inequality was propelled by wage increases for the highest-paid workers (Lemieux 2008, 26).

Figure 1 divides the total variance in log wages into between-group and within-group components. These components were obtained from a regression on log wages with age, race, ethnicity, education, region, union membership and industry-region unionization rates as predictors. Between-group inequality, measured by the variance of predicted wages, describes the dispersion of average wages across the groups defined by the predictors. Within-group inequality, measured by the residual variance, describes the spread of wages among workers in each of these groups. Rising inequality between and within groups increased total wage inequality in the 1980s. Since the 1990s, rising inequality was mostly due to increased within-group inequality.

Falling unionization accompanied rising wage inequality. CPS tabulations indicate union decline was especially large among men, falling from 35 percent in 1973 to under 10 percent by 2007. Employer surveys show that CPS survey respondents mistakenly report their union status 2 to 3 percent of the time (Card 1996). When unionization is very low, a relatively large number of nonunion respondents incorrectly report being union members. Figure 2 reports the observed proportion of unionized workers and an adjusted series that corrects for error in reported union status in the CPS. Adjusting for the reporting error, only 8 percent of private sector men and less than 4 percent of private sector women were union members by 2007.



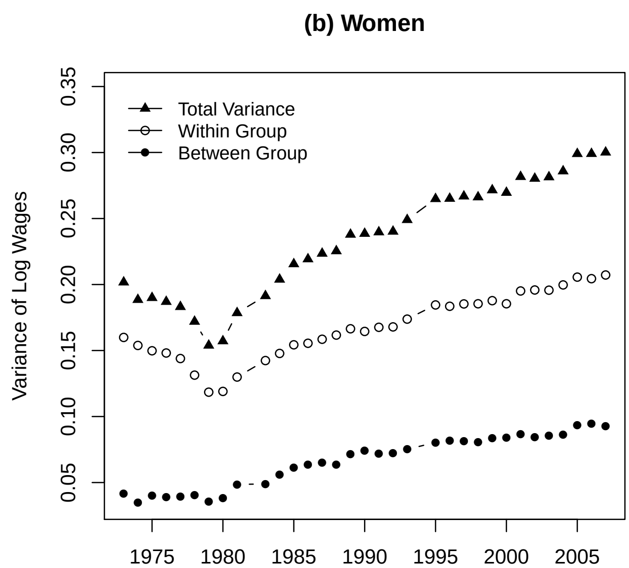


Figure 1. Inequality in hourly wages among full-time, private sector, men and women, 1973–2007. Figures are calculated from the May and Merged Outgoing Rotation Group files of the CPS.

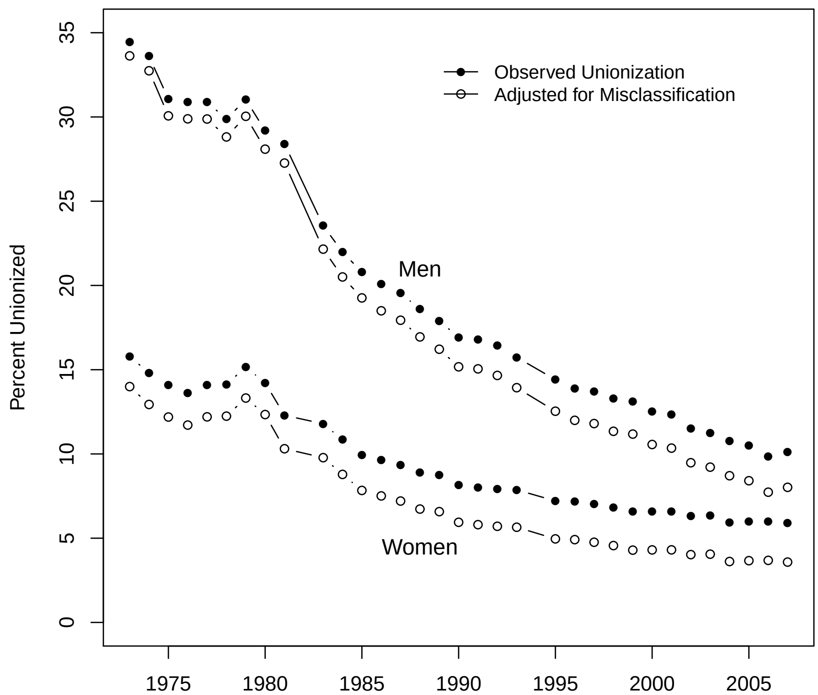


Figure 2. Union membership rates among full-time, private sector, men and women, 1973–2007, unadjusted and adjusted for reporting error in union status in the CPS.

US union decline is often explained by changes in the economy and intensified political conflict in the workplace (Farber and Western 2001; Freeman and Medoff 1984; Goldfield 1987). In this account, union firms could not respond to 1970s stagflation, industry deregulation, and economic globalization. The biggest driver of decline in the percentage unionized was employment growth outside of the traditional union strongholds of manufacturing, construction, and transportation, utilities, and communications. Faced with a newly competitive economic environment, employers in unionized industries also intensified their opposition and union employment and new organizing—at least through union elections—plunged through the 1980s (Tope and Jacobs 2009; Hirsch 2008).

Employer opposition unfolded in an increasingly adverse political context for labor. Hacker and Pierson (2010, chap. 5) report that an influx of corporate donations influenced policymakers to oppose pro-union reform of labor law in the 1970s. Union reform efforts were defeated, and the rightward shift of the National Labor Relations Board under Republican administrations made organizing more difficult in the 1980s. Political defeats in the 1970s and 1980s yielded an "enervated" labor law that enabled employers to block organizing campaigns and weaken existing unions (Cowie 2010, 288).

Unions and Inequality

Union decline and rising inequality motivated research on the share and the shape of the union wage distribution. We extend the research by linking unionization to wage inequality among nonunion workers.

Wages in the Union Sector

Research on unions and inequality focused on two effects. First, unions raise wages among less-educated and blue-collar workers. This *between-group* effect of unions reduces educational and occupational inequality. Second, collective bargaining standardizes wages within firms and industries. This *within-group* effect of unions on inequality reduces the spread of wages among union members with similar characteristics. Comparing these effects, Freeman (1980) found that unions reduce wage inequality more within educational and occupational groups, than between them. Because nonunion women were concentrated in occupations with low and relatively equal wages, the effect of unions on within-group inequality is largest among men (MacPherson and Stewart 1987).

The within-group and between-group effects of unions suggested union decline was associated with rising wage inequality. Declining unionization was estimated to explain 10 to 20 percent of the growth in men's wage inequality from the late 1970s through the late 1980s, (Freeman 1993; Card 1992). DiNardo, Fortin, and Lemieux (1996) studied the top and bottom tails of the wage distribution separately, finding that deunionization is chiefly associated with top-tail inequality, explaining nearly a third of the growth in the gap between the median wage and the ninetieth percentile. In that analysis, deunionization accompanied the declining middle of the pay distribution.

Unions and Nonunion Wages

Union wages were the main focus of research on inequality, but organized labor also affects nonunion workers. Economists contrast the effects of "spillover" and "threat." When unions raise wages for their members, employers may cut union employment forcing unemployed workers to find jobs in the nonunion sector. The spillover of workers into the nonunion labor market causes wages to fall. The threat effect results from nonunion employers raising wages to the union level to avert the threat of unionization. The two theories yield opposing predictions: unions reduce nonunion wages with spillover, but increase nonunion wages with threat. Empirical studies tend to support the threat effect, showing that nonunion wages are higher in highly unionized industries, localities, and firms (Neumark and Wachter 1995; Farber 2005; Leicht 1989).

The theory of union threat has distributional implications. If unions threaten to organize low-wage workers, employers may raise their wages thereby equalizing the wage distribution. Testing this theory, Kahn and Curme (1987) estimated the effect of industry unionization on the variance of nonunion wages for detailed industries and occupations for just one year, in the 1979 CPS. Consistent with the equalizing effect of union threat, they find less inequality among nonunion workers in highly-unionized industries (cf. Belman and Heywood 1990).

Unions and the Moral Economy

The theory of union threat takes a rationalistic view of employers and a minimalist view of labor market institutions. Employers minimize labor costs and only raise wages when threatened with even greater pay increases through unionization. Institutions are conceived minimally in the sense that unions are the key distortion in an otherwise competitive labor market.

We relax these assumptions, arguing that the labor market is embedded in a moral economy in which norms of equity reduce inequality in pay. The moral economy consists of norms prescribing fair distribution that are institutionalized in the formal rules and customs of the market. In a robust moral economy, violation of distributional norms inspires condemnation and charges of injustice. We often think of the moral economy historically—determining, for example, fair prices for bread and flour under the British Corn Laws (Thompson 1971) or the relative rank and standing of English workers in the nineteenth century (Polanyi 1957, 153).[1]

Unions are pillars of the moral economy in modern labor markets. Across countries and over time, unions widely promoted norms of equity that claimed the fairness of a standard rate for low-pay workers and the injustice of unchecked earnings of managers and owners (Webb and Webb 1911, 279–282; Hyman and Brough 1976). Comparative researchers emphasize the role of distributional norms governing European industrial relations (Elster 1989b; Swenson 1989). American labor never exerted the broad influence of the European unions, but US unions often supported norms of equity that extended beyond their own membership. In our theory of the moral economy, unions help materialize labor market norms of equity (1) culturally, through public speech about economic inequality, (2) politically, by influencing social policy, and (3) institutionally, through the rules governing the labor market.

Culturally, industrial unions in particular often used a language of social solidarity in public discourse and within firms. Walter Reuther's postwar leadership of the United Auto Workers (UAW) provides a key example. Labor historian, Nelson Lichtenstein (1995, 300), writes that Reuther

[1] The term moral economy was coined by Thompson (1971). Elster (1989a) argued that distributional norms are departures from strict rationality, eliciting strong emotions in response to violation.

aimed to "reshape the consciousness of millions of industrial workers, making them disciplined trade unionists, militant social democrats, and racial egalitarians." The UAW sought to develop a network of union-based community organizations, published hundreds of newspapers, and pressed Presidents Kennedy and Johnson on civil rights legislation (Lichtenstein 1995, 302–303; Boyle 1995). Within firms, unions have been common voices for equity, protesting the pay of upper management. Consistent with the egalitarian effect of union advocacy, studies of data from the mid-1970s to the early 2000s find that compensation is lower in unionized firms, and managerial employment is lower in highly unionized industries (DiNardo, Hallock, and Pischke 1997; Gomez and Tzioumis 2006).

Politically, American unions have been frequent advocates for redistributive public policy. Highly unionized states have higher minimum wages and their congressional representatives are more likely to support minimum wage increases (Cox and Oaxaca 1982; Kau and Rubin 1978).[2] Union political pressure also reached beyond wages to social legislation. For example, major unions regularly backed proposals for universal healthcare and supported the creation of Medicare in the mid-1960s (Derickson 1994). In the 1970s, unions joined with states in litigation opposing cuts to the federal food stamp program, and threatened to sue again in 1980 to keep the program solvent.

Institutionally, US industrial relations often extended union conditions to nonunion workers. When a third of the male labor force was organized, in the early postwar period, unions were national economic actors who shaped centralized wage policy. During World War Two, government

[2]Union self-interest may drive political support, however, where minimum wage increases reduces competition from nonunion workers.

boards with business and labor leaders helped set wage standards to control inflation and assist wartime production. Centralized wage policy continued during the Korean War when the tripartite Wage Standardization Board monitored wage increases among key firms and aimed to narrow inter-firm wage differentials, reducing wage inequality in the wider economy (Ross and Rothbaum 1954). Unions influenced national pay policy in the 1960s as the Kennedy and Johnson Administrations set wage and price targets to stabilize prices and promote the "distributional equity" of wages (Ulman 1998, 170). The Nixon Administration also adopted a tripartite wage policy. Concluding that "No program works without labor cooperation," (Matusow 1998, 160), Nixon's national pay board urged wage restraint in major contract negotiations, but also examined executive pay levels, supported raises for low-wage workers, and monitored merit pay increases (Mitchell and Weber 1998). In the 1970s, President Carter pursued a national wage policy with union representation in a Pay Advisory Committee that set industry wage and price guidelines.

Unions also helped to establish pay norms in local labor markets. In some industries, union influence was amplified by law. The federal Davis-Bacon Act and its state-level variants require public construction projects to pay at least the locally prevailing wages and fringe benefits. Studies from the 1970s show that Davis-Bacon raised construction wages in nonunion firms and reduced the difference between union and nonunion wages (Goldfarb and Morrall 1981).

Beyond federal contractors, major union agreements also set the pattern for industry-wide wage increases. Employers of the 1960s and 1970s used industry wage surveys to reduce wage differences between union and nonunion firms (Dunlop 1977, 15; Foulkes 1980; see also Jacoby 1997, 45–46). Of course, the adoption of union standards in nonunion firms was often intended to prevent unionization. Still, nonunion companies in the 1970s closely monitored union contracts even in lightly unionized industries where the threat of unionization was remote (Foulkes 1980, 156). Predominantly nonunion firms with small union workforces abandoned merit pay, and norms of equal treatment governed the distribution of fringe benefits. Summarizing his survey of pay practices in large nonunion firms, Foulkes (1980, 153) writes:

> In many environments, providing and demonstrating equity generally means that a company's pay rates favorably compare with those of unionized companies. It would be acceptable to say that the activities of many unions in the United States are benefiting many nonmembers; in other words, unions are doing much good for people who do not pay them any dues.

The slow postwar decline of unionization rates accompanied the erosion of the labor market's moral economy. The cultural, political, and institutional indicators of equitable pay norms declined with union membership. Many researchers see 1981 as a watershed year, when the Reagan administration defeated the air traffic controller's strike by hiring permanent replacements. Voss and Sherman (2000, 311) characterize the turning point as a change in norms when "corporate leaders stopped playing by the rules." Levy and Temin (2010) similarly divide the postwar history of the American labor market into the eras of the "Treaty of Detroit" until 1980, and the "Washington Consensus" which followed. The era of the Treaty of Detroit was named for the landmark wage agreement of 1948 between General Motors and the UAW that provided an annual wage increase of two percent plus cost-of-living. Wages across manufacturing industries moved broadly according to this formula until 1980. The Treaty of Detroit was succeeded by the Washington Consensus, an era of deregulation and eroded pay norms in which earnings inequality increased with the compensation of managers and professionals.

Because union decline varies across regions and industries, we view the transformation of the moral economy not as a discrete turning point but as a gradual process unfolding unevenly across the labor market through the 1980s and 1990s. Our research design exploits this variation to study how deunionization is associated with rising wage inequality among nonunion workers in regional labor markets.

Limitations and Rival Explanations

The contribution of the union wage distribution to wage inequality is directly observed, but union threat and norms of equity are unobserved mechanisms. In our approach, union threat and egalitarianism in the moral economy are indicated by industry-level unionization in a given region. This is an indirect measure compared to direct observation of the cultural, political, and institutional channels of union influence on nonunion wages. Still, the industry-region unionization rate captures in a single index the salience of organized labor in the surrounding labor market. The index can be measured across the national labor market over decades, and it usefully captures the uneven decline of American labor as a voice for equity in the economy as a whole.

Despite these virtues, industry-region unionization is likely correlated with economic and social conditions that are also associated with rising inequality. In particular, increased demand for highly skilled workers may have raised inequality and reduced unionization. The rising demand for skill, indicated by the increasing college wage premium, is regularly interpreted to result from technological change in which computerization has replaced routine work tasks (Autor, Levy, and Murnane 2003; Autor and Dorn 2009). In some accounts, firms adapted to technological change by introducing performance pay and other workplace incentives, weakening internal labor markets and increasing inequality within firms (Bloom and Reenen 2010; Cappelli 2001). Though skill-biased technological change has been the central focus, economists also argue that immigration and trade have shifted demand away from low-skill workers, further contributing to the wage premium of college graduates (Card 2009; Cline 1997).

Union decline may have enabled or influenced technological and organizational change, and economic globalization. So, for example, merit pay is less common in union firms (Lemieux, MacLeod, and Parent 2009, 22). Evidence from a plant restructuring indicates that unions shape the introduction of new technologies to moderate pay inequality and limit layoffs (Fernandez 2001, 316–17). Thus, deunionization may have indirect or mediated effects on inequality beyond the effects of union threat or normative influence.

Alternatively, unions flourished in concentrated, protected, smokestack industries employing relatively homogeneous workforces. Technological change, human resource management, and economic globalization may have undermined workplace solidarity, fueling deunionization and wage inequality. Declining unionization and rising inequality may share common roots in shifting market conditions. Effects we attribute to union decline may really be due to other changes in the economy.

We account for some rival explanations in a regression that controls for education, demographic characteristics, and region. The increasing education gradient accompanying skill-biased technological change, immigration, and trade is directly incorporated into the analysis. The regression also allows a comparison of the effects of deunionization to the effects of education. Still, omitted variable bias likely remains and this affects our interpretation of the results, as we discuss below.

Data and Methods

We link union decline to rising inequality by decomposing the variance of log wages. The decomposition uses a variance function regression in which the mean and the variance of an outcome depend on independent variables, providing a model for between-group and within-group inequality (Western and Bloome 2009).

The decomposition is based on a regression on the log hourly wage, y_i, for respondent i ($i = 1, \ldots, N$) for a given year of the CPS. Our key predictors are an indicator for union membership, u_i, and a continuous variable, $\bar{u}_i$, that records for each respondent the unionization rate for the industry and region in which they work. Covariates, including schooling, age, race, ethnicity, and region, are collected in the vector, $\boldsymbol{x}_i$. The model includes equations for the conditional mean of log wages,

$$\hat{y}_i = \boldsymbol{x}_i'\boldsymbol{\alpha}_1 + u_i\alpha_2 + \bar{u}_i\alpha_3,$$

and the conditional variance,

$$\log \sigma_i^2 = \boldsymbol{x}_i'\boldsymbol{\beta}_1 + u_i\beta_2 + \bar{u}_i\beta_3.$$

We expect union membership and industry-region unionization to have positive effects on average wages, but negative effects on the variance of log wages. Economy-wide changes in average wages and within-group inequality, perhaps due to general shifts in norms or the macroeconomy, are captured through the regression intercepts.

Between-group wage inequality is measured by the variance of the conditional means,

$$B = \sum_{i=1}^{N} w_i (\hat{y}_i - \bar{y})^2, \tag{1}$$

where w_i is the sample weight for respondent i, normed to sum to 1, and $\bar{y}$ is the grand mean of log wages. Within-group inequality is measured by the residual variance,

$$W = \sum_{i=1}^{N} w_i \sigma_i^2. \tag{2}$$

That each respondent has a variance, σ_i^2, may be counterintuitive, but it is simply estimated by the squared residual from the regression on y_i. Because unions are mostly associated with within-group inequality, we expect deunionization to be closely associated with an increase in the within-group variance, W. Summing within-group and between-group components yields the total variance, a measure of overall inequality:

$$V = B + W.$$

We decompose the rise in inequality with three adjusted variances that fix in a baseline year either the coefficients of some predictors or the distribution of some predictors. First, focusing on individual union membership, we calculate the level of inequality assuming unionization had remained at its 1973 level. This compositional effect is estimated by reweighting the data to preserve the 1973 unionization rate across all years, from 1973 to 2007. The adjusted weights are then used in equations (1) and (2) to calculate adjusted variances for all years. Second, we add the effects of union threat and equitable wage norms on nonunion workers by calculating wage inequality assuming industry-region unionization and its coefficients remain constant at the 1973 level. Fixing industry-region unionization, coefficients, and reweighting to hold union density constant, yields adjusted values of $\hat{y}_i$ and σ_i^2 that are plugged into equations (1) and (2) to obtain adjusted measures of between-group and within-group inequality. Finally, we assess the contribution of education to rising inequality by fixing education coefficients for the mean and variance equations at their 1973 values. Fixed education coefficients adjust the values of $\hat{y}_i$ and σ_i^2 and the resulting measures of between-group and within-group inequality. (See Appendix 1 for more details.)

The adjusted variances quantify the growth in inequality statistically attributable to changes in unionization rates or changes in coefficients. The adjustments describe the association between rising inequality, on the one side, and union decline and educational inequality, on the other. The adjusted variances might be interpreted as counterfactual measures of inequality, prevailing if unionization rates or regression coefficients remained fixed at 1973 levels. The counterfactual assumes, however, that the regression coefficients accurately estimate causal effects, that all other predictors and their coefficients change as observed, and that there are no broader effects on employment of fixing, say, unionization rates or other quantities.

Counterfactual interpretation is likely implausible because of omitted variable bias. Because we analyze the change in inequality rather than its level at a point in time, biases of constant magnitude will not affect our analysis of the trend. However if union members are increasingly positively selected, say, upward bias in the wage effects will grow because of the rising productivity of union workers. In the context of rising inequality, shifts in technology, industry regulation, and the use of performance pay, for example, might be correlated with industry-region unionization and wages. Industry and region fixed effects might reduce omitted variable bias, though fixed effects account for nearly all the variation in industry-region unionization, leaving little variance for the effects of interest. We try to reduce bias with region fixed effects. Given a clear trade-off between population description and causal analysis, we prefer this simple model because our interest centers on the population trend in rising wage inequality. We thus aim to shows how the wage distribution moves with shifts in unionization and other labor market conditions, rather than providing a causal analysis of wage determination.

Measurement error in union membership also adds bias. Using data from a 1977 employer survey, Card (1996) found that 2.5 percent of CPS respondents mis-report their union status.[3] With this estimate of measurement error, at 50 percent unionization, the number of nonunion workers misclassified as union equals the number of union workers misclassified as nonunion. With observed unionization at 10 percent, a large number of nonunion workers (2.5 percent of 90 percent) report they are union members. Given a misclassification rate, λ, and observed unionization, $\bar{u}$, the true unionization rate is $(\bar{u} - \lambda)/(1 - 2\lambda)$. With observed unionization at .10 (10 percent) and a misclassification rate of $\lambda = .025$ (2.5 percent), the true unionization rate is .079, a measurement error of about 20 percent. Measurement error biases estimates of union effects and increases

[3]Card (1996) studied a sample of 1,718 men with valid wage data and nonmissing employer and employee union membership responses.

with declining unionization.

We account for measurement error in union status by augmenting the likelihood for the wage model with an extra term for the probability of misclassification (for other approaches, see Card 1996 and Hirsch 2004). Because the misclassification rate may vary across surveys, we specify an average rate of 2.5 percent but allow it to vary from 2 to 3 percent with a Bayesian prior distribution. Correcting for misclassification significantly reduces bias in the estimated union effects. (See Appendix 2.)

Data are compiled from the annual May files of the CPS from 1973 to 1981, and the annual Merged Outgoing Rotation Group files of the CPS from 1983 to 2007. We exclude 1982, when union questions were omitted from the survey, and 1994 and three-quarters of 1995, when allocation flags for wages were missing. The analysis includes men and women working full-time (30 hours a week or more) in the private sector. The dependent variable is log hourly wages adjusted for inflation to 2001 dollars.

Several adjustments improve the quality and continuity of the wage data. Nonresponse to wage and income questions increases over time, and by 2007, about a third of CPS workers did not report wages. The CPS imputes wages to non-respondents, but regression coefficients for non-matched criteria are attenuated and the residual variance for wages is sensitive to the imputed data (Hirsch 2004; Mouw and Kalleberg 2010). We omit imputed wages from the analysis.[4] Earnings imputation flags change

[4]Analysis of earnings nonresponse reveals modest positive selection, but bias in regression coefficients is small if imputed data are excluded (Bollinger and Hirsch 2009). If unobserved high earners are nonunion, union wage effects and the union decline effect on between-group inequality will be overestimated. But if high-pay unobserved nonunion workers are drawn from low-union industries and regions, inequality and the union effect on within-group inequality will be underestimated. In short, the effects of non-response are likely modest and small biases in union effects are offsetting.

across surveys and we follow Hirsch and Schumacher (2004, esp. Table 2) to create a consistent series of non-imputed earners. A few respondents with wages less than one dollar are excluded. Top-codes for wages and the proportion top-coded varies over surveys, biasing estimates of inequality. For each year, we impute the top two percent of wages from a Pareto distribution. Other common methods for top-codes yielded similar results to those reported here. Incomes greatly increased in the top percentile since the 1990s, a trend measured by administrative rather than survey data (Piketty and Saez 2003). Similar to other research on wage inequality (Autor, Katz, and Kearney 2008; Mouw and Kalleberg 2010), we cannot analyze the top fractions of a percentile of the income distribution with these data. (Appendix 3 details our coding.)

Table 1 reports descriptive statistics for the CPS data. Wage stagnation is indicated by trends in average pay for union and nonunion men. Average pay increased for women reflecting the entry of women into professional and other skilled occupations. Inequality trends were similar for men and women, increasing everywhere and most rapidly among union workers. Rising wage inequality among union workers tends to reduce the effect of union decline on inequality as the union wage distribution comes to resemble the nonunion. Educational attainment also increased across the labor market as the proportion of high school dropouts in the workforce dropped from around 1 in 4 workers in the 1970s to about 1 in 10 by the 2000s. Fewer than 20 percent of private sector workers were college graduates in the 1970s and more than a quarter had college degrees three decades later. The college graduation gap between union and nonunion workers also narrowed from the 1970s to the 2000s.

Other covariates might also be analyzed although coding inconsistency

Table 1. Descriptive statistics for the analysis of unionization and hourly wage inequality among private sector, full-time, men and women, 1973–2007.

	1973–1984		1985–1995		1996–2007	
	N	U	N	U	N	U
Male workers						
Log Hourly Wages						
Mean	2.84	3.05	2.81	3.04	2.87	3.07
Variance	.29	.12	.32	.14	.35	.19
Level of Schooling (proportion)						
Less than high school	.23	.31	.14	.17	.13	.10
High school graduate	.33	.47	.35	.51	.31	.46
Some college	.23	.17	.24	.24	.26	.33
BA or more	.21	.04	.27	.07	.29	.11
Demographic characteristics (proportion)						
White	.86	.83	.83	.81	.76	.77
Black	.07	.09	.06	.09	.06	.08
Other	.08	.08	.11	.10	.18	.15
Age (years)	36.11	39.04	36.39	40.56	38.79	42.43
Female workers						
Log Hourly Wages						
Mean	2.46	2.66	2.56	2.74	2.67	2.82
Variance	.19	.12	.24	.18	.29	.23
Level of Schooling (proportion)						
Less than high school	.20	.33	.10	.17	.08	.10
High school graduate	.47	.48	.40	.48	.32	.36
Some college	.22	.14	.29	.22	.33	.30
BA or more	.11	.05	.20	.13	.27	.23
Demographic characteristics (proportion)						
White	.84	.74	.81	.68	.76	.64
Black	.09	.16	.09	.18	.09	.16
Other	.07	.10	.10	.14	.15	.20
Age (years)	35.59	38.83	36.41	40.05	39.20	42.50

Note: N=nonunion workers, U=union workers.

and missing data prevent us from greatly augmenting the model. In additional analysis, we explored the effects of marital status, Hispanics, urban residence, and occupations. Results were largely unchanged for men. There is greater confounding for women but their union effects are relatively small in any case.

To estimate the effect of unions on nonunion workers, earlier research measured unionization in industries and localities (e.g., Freeman and Medoff 1981; Neumark and Wachter 1995). Indeed, regions and industries are two key dimensions along which wage comparisons are made by workers and employers (Foulkes 1980; Hyman and Brough 1976). To capture the influence of unions on the nonunion sector, we measured unionization in 18 industry categories combining broad (one-digit) sectors with several detailed (two-digit) industries. We recode the several revisions of CPS industry classifications to produce a relatively consistent series whose distribution changes smoothly across survey redesigns. We disaggregate further by measuring industry unionization in four census regions (the northeast, the south, the midwest, and the west). Regional disaggregation helps account for time-varying changes in regional pay scales and allows union strength to be spread unevenly across the country. Smaller spatial units, like states, might correspond better to local labor markets, though a full set of state identifiers are only available since 1977 and region codes are available for the whole series. In large capital-intensive industries where unions are concentrated, pay norms, like unionization, are also likely to stretch across state lines.

Dividing the national labor market into 18 industries by 4 regions yields 72 annual industry-region unionization rates. On average, unionization rates were highest in utilities, transportation, and communications, particularly in the northeast and the midwest (Figure 3). Unionization was lowest in agriculture, finance, and retail trade, and throughout the south. Women's unionization rates in the private sector were everywhere lower than men's.

Evidence for the relationship between industry-region unionization and wages is given by plotting unionization for the 72 industry-regions nonunion wages for each year from 1973 to 2007. Average nonunion wages are higher in regions and industries where unionization is higher (Figure 4). Higher wages may reflect the effect of union threat, the skills of workers in different industries, or industry rents. Preliminary support for the waning normative influence of unions on nonunion wages is given by Figure 5. For men and women, nonunion wages are more compressed in local labor markets that are highly unionized, and highly-unionized local labor markets become less common over time.

Results

To study the effects of declining union membership on wage inequality, we fix the unionization rate at 1973 levels. Similar to earlier research, we find the largest effect of declining union membership on men's within-group inequality (Figure 6, panel a). The observed within-group variance increases by .046 points, but the adjusted variance with the 1973 unionization rate increases by only .028 points. Thus, over a third of the increase in within-group inequality is associated with declining union membership ($[.046-.028]/.046 = .40$). Between-group inequality increases by .055 variance points and the adjusted series increases by almost as much, by .053, indicating that union decline explains little of the rise in men's between-group inequality. Summing the between-group and within-group effects,

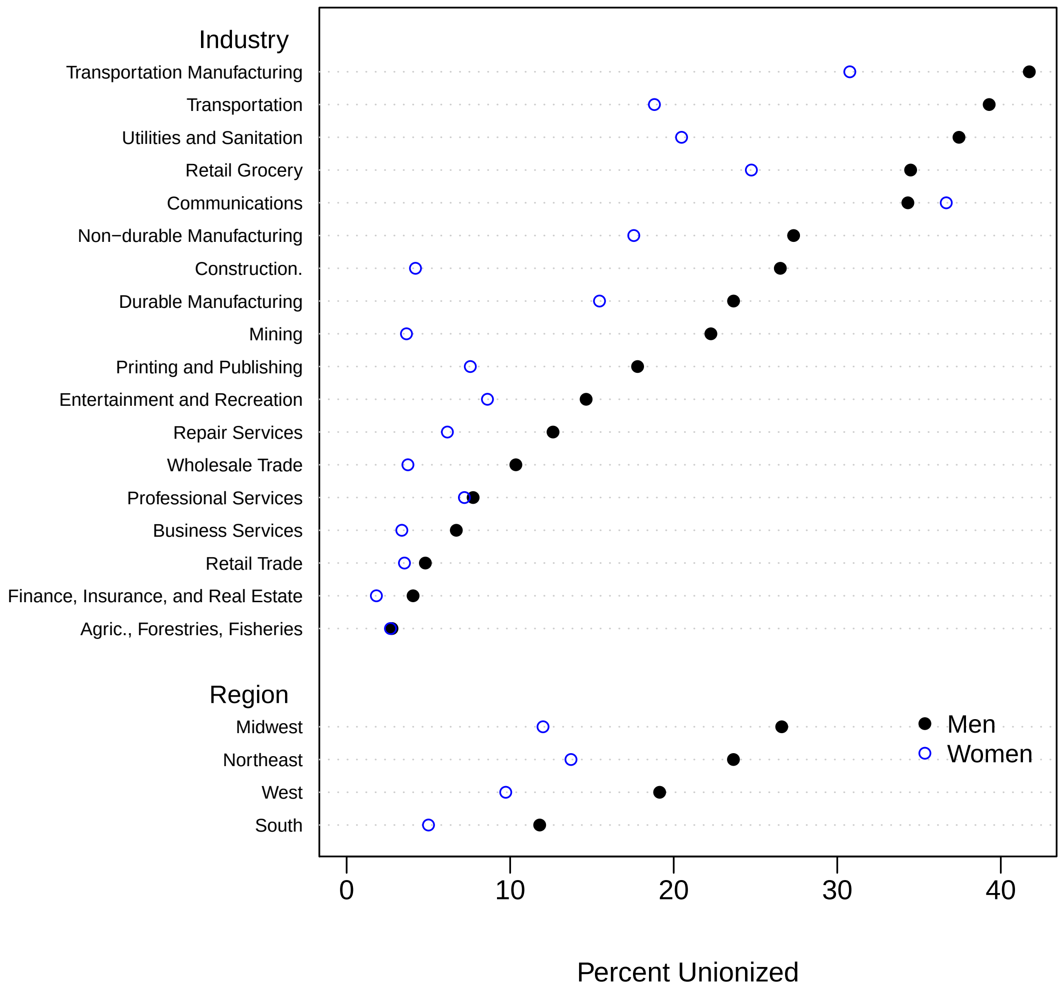


Figure 3. Average unionization rates for full-time, male and female private sector workers, in 18 industries and 4 regions, 1973–2007.

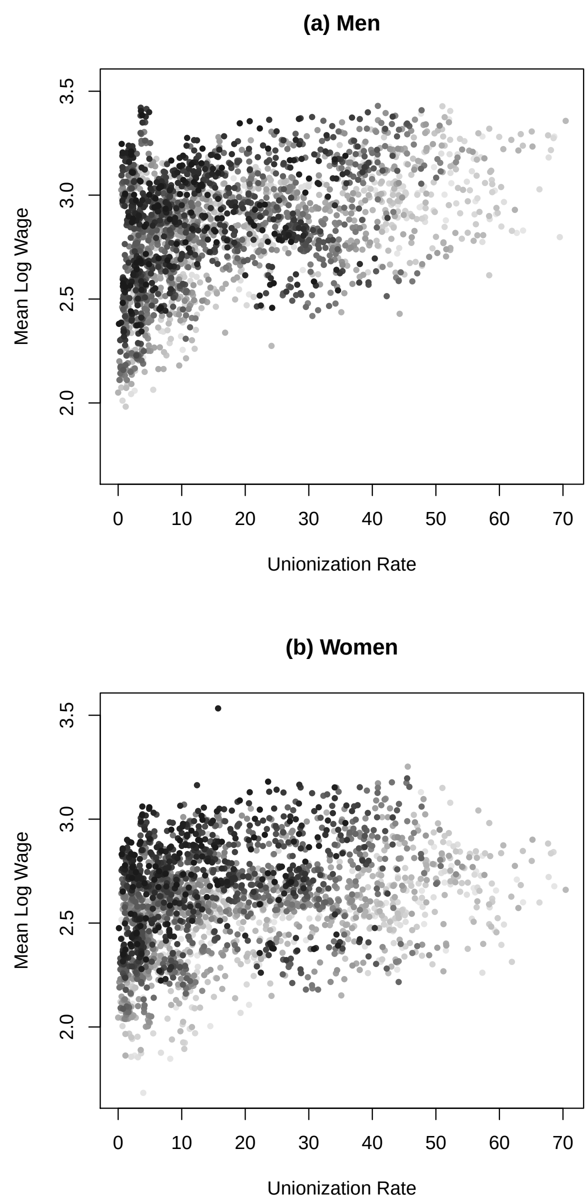


Figure 4. Unionization rates and mean log hourly wages for full-time, private sector, male and female nonunion workers, by 72 industry-regions, 1973–2007. Points go from light gray to black, from 1973 to 2007.

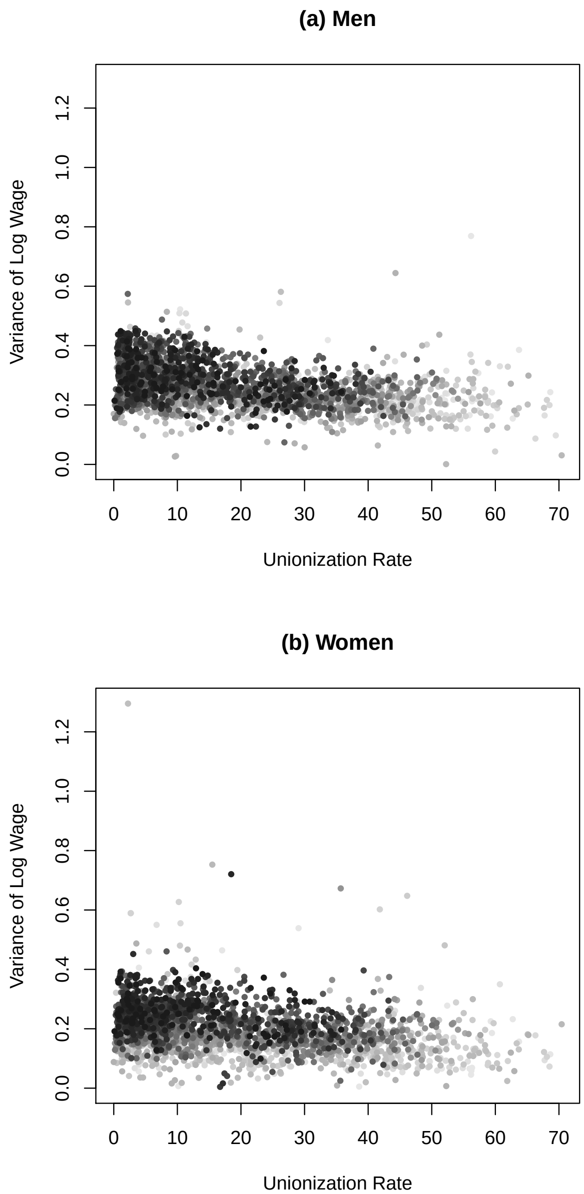


Figure 5. Unionization rates and the variance of log hourly wages for full-time, private sector, male and female nonunion workers, by 72 industry-regions, 1973–2007. Points go from light gray to black, from 1973 to 2007.

the decline in unionization from 34 to 8 percent explains about a fifth of the rise in inequality in hourly wages among full-time, private-sector men.

The effect of declining union membership on wage inequality among women is smaller (Figure 6, panel b). Inequality grew slightly more for women than men, but increasing women's wage inequality is unrelated to union membership. Within-group inequality increased by about .047 points among women but the adjusted series, with 1973 unionization held constant, increases nearly as much, by .043 points. Similarly, between-group inequality increased by .051 points and the adjusted between-group variance, with union membership fixed, increased by .054 points. Consistent with other research, holding the unionization rate constant explains almost none of the rise in women's wage inequality.

Union decline explains more of the rise in wage inequality once we account for the link between unions on nonunion wages. The link between organized labor and nonunion workers is captured by the effects of industry-region unionization on the mean and variance of wages. The total effect of deunionization can be measured by fixing the 1973 unionization rate as before, and also fixing industry-region unionization rates and their coefficients at 1973 values (Figure 7, panel a). This adjustment indicates the strong relationship between union decline and rising within-group inequality. Among men, the adjusted within-group inequality does not increase between 1973 and 2007, suggesting the effects of union threat and eroding norms of equity on the wage distribution. Deunionization's effect on union and nonunion wages is associated with about a third of the rise in wage inequality. This accounting, that includes the effects of unions on nonunion wages, is 50 to 100 percent larger than the union effects on inequality reported in other decompositions (cf. DiNardo, Fortin,

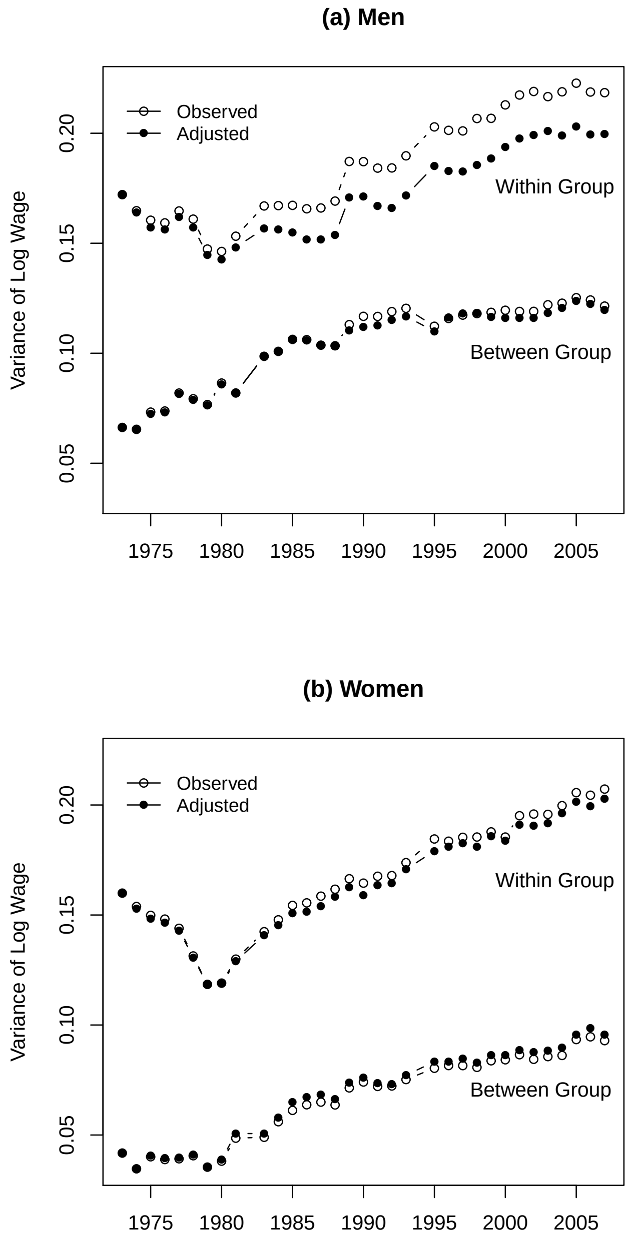


Figure 6. Observed and adjusted within- and between-group variances of log hourly wages, full-time, private sector men and women, 1973–2007. Adjusted variances fix union membership at the 1973 level.

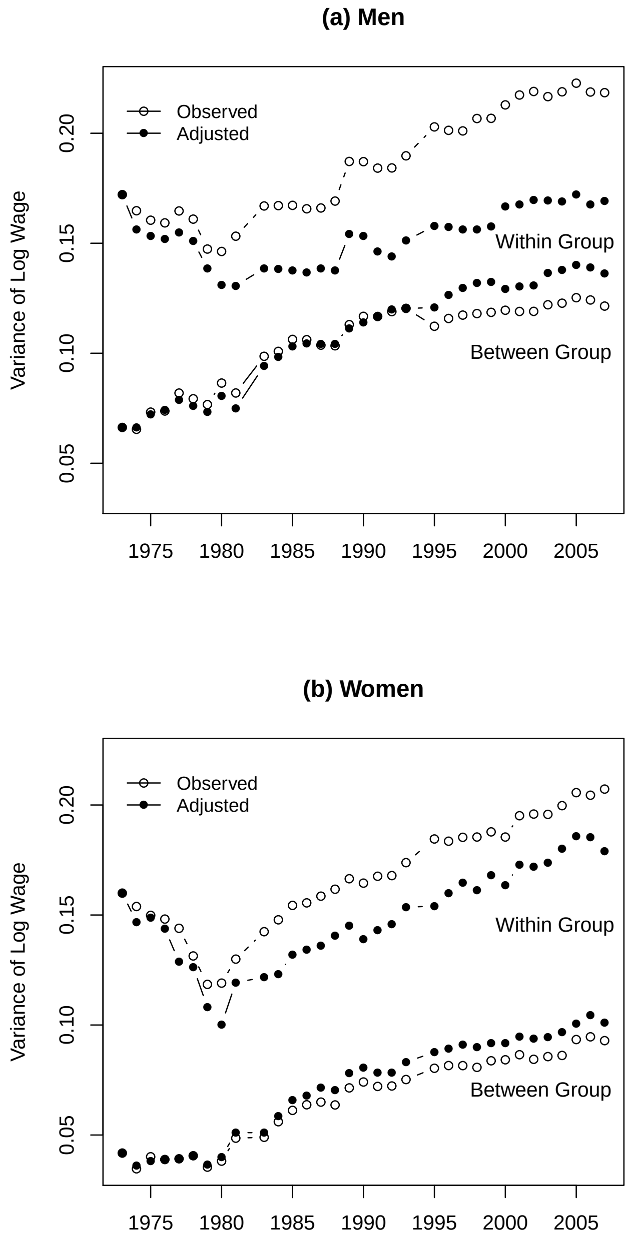


Figure 7. Observed and adjusted within- and between-group variances of log hourly wages, full-time, private sector men and women, 1973–2007. Adjusted variances fix at the 1973 level: union membership, industry-region unionization rates, and industry-region unionization effects.

and Lemieux 1996; Card 2001).

The association between union decline and wage inequality among women is smaller, but greater than zero, once nonunion wages are taken into account (Figure 7, panel b). Declining industry-region unionization accounts for over half of the increase in women's within-group inequality. Adding the between-group and within-group effects together, union decline is associated with about a fifth of the rise in wage inequality among women.

How do the effects of union decline compare to the growing inequality of wages by education? Counting the union and nonunion wage effects, deunionization explains about a third of the rise in men's earnings inequality (Figure 8, panel a). Increasing returns to education and increasing wage inequality among highly educated workers explains a similar share of the rise in wage inequality. Among women, union decline explains about a fifth of the rise in wage inequality (Figure 8, panel b). Rising educational inequality in pay explains nearly twice as much. In short, the effects of deunionization on inequality are only half as large as the effects of education for women, but union and education effects are equally large for men.

Finally, the combined effects of union decline and education on wage inequality can be calculated by holding constant the unionization rate, industry-region unionization and its coefficients, and the education coefficients (Figure 9). From 1973 to 2007, the adjusted variance for men's wage inequality increased only 10 percent compared to a 40 percent increase in the observed variance. Thus three-quarters of the rise in men's hourly wage inequality is associated with union decline and increasing inequality by education. Women's wage inequality increased by nearly 50

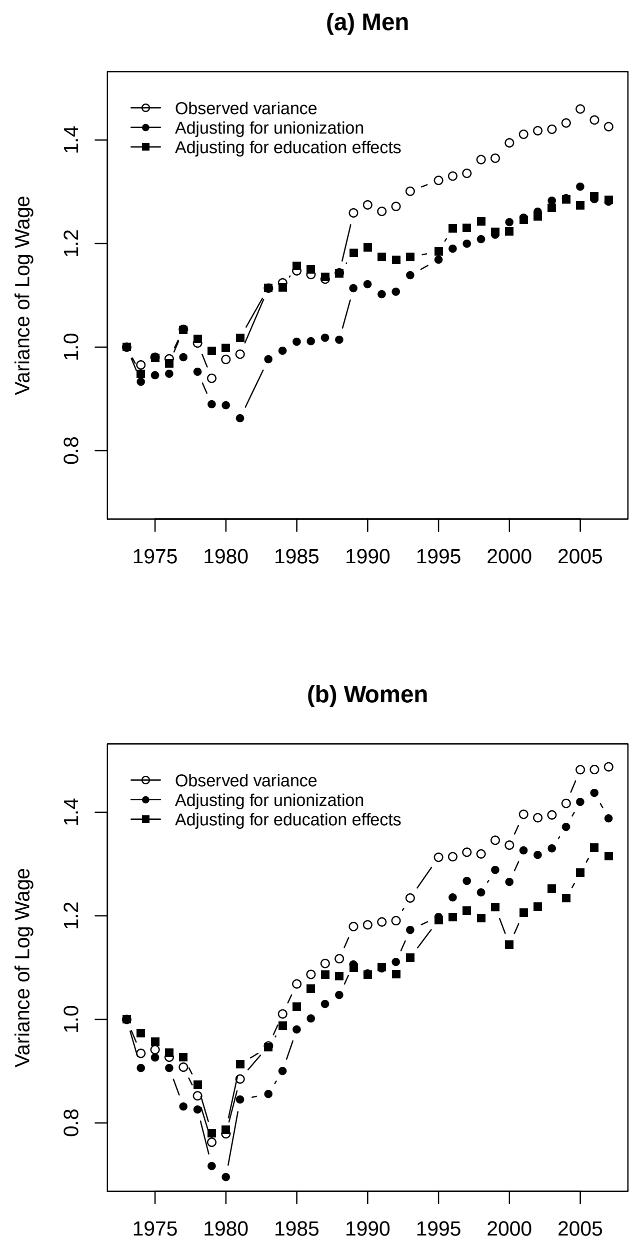
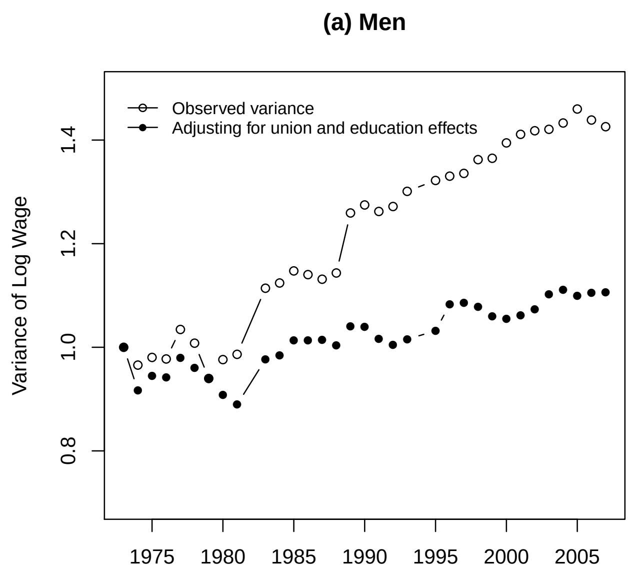


Figure 8. Total variance of log hourly wages, full-time, private sector men and women, 1973–2007. The variances adjust for unionization and education. Variances are set to equal one in 1973.



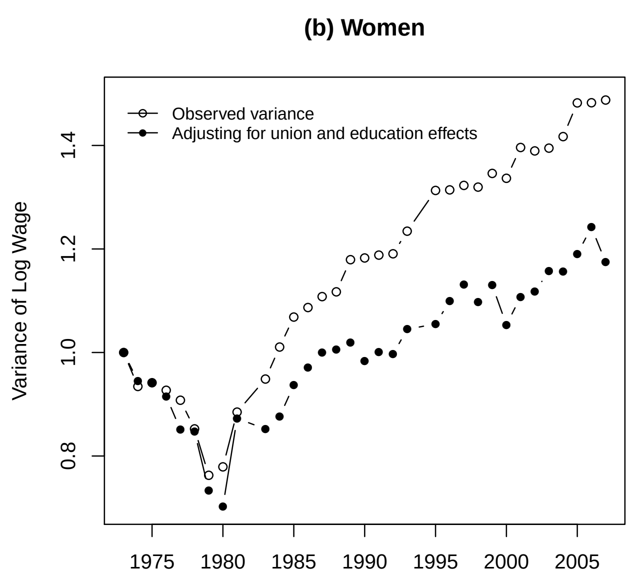


Figure 9. Total variance of log hourly wages, full-time, private sector men and women, 1973–2007. The adjusted variance fixes at the 1973 level: the unionization rate, regional unionization and its effects, and education effects. Variances are set to equal one in 1973.

Table 2. Summary of the decomposition of the change in the variance of log hourly wages, private sector, full-time men and women, 1973–2007.

	Between	Within	Total
Male workers			
Observed change in the variance of log wages	.055	.046	.102
Change in variance fixing:			
Union membership rate	.053	.028	.081
+ Regional unionization	.070	-.003	.067
+ Education Effects	.030	-.005	.025
Percentage of change explained:			
Union membership rate	3.2	40.3	20.2
+ Regional unionization	-26.9	106.3	33.9
+ Education effects	45.8	110.1	75.2
Female workers			
Observed change in the variance of log wages	.051	.047	.098
Change in variance fixing:			
Union membership rate	.054	.043	.097
+ Regional unionization	.059	.019	.078
+ Education effects	.018	.017	.035
Percentage of change explained:			
Union membership rate	-5.2	9.2	1.7
+ Regional unionization	-16.1	59.8	20.4
+ Education effects	64.9	63.6	64.3

Note: Adjusted variances are obtained by successively adding the listed effects. For example, adjusted variances in line 2 are obtained by fixing union membership, adjusted variances in line 3 are obtained fixing union membership and industry-regional unionization and related coefficients, and so on.

percent, but adjusted inequality—accounting for inequality by education and unionization—increased 17 percent. The decomposition of women's wage inequality indicates that unions and returns to schooling together account for nearly two-thirds of the rise in wage inequality.

The results of the variance decomposition are summarized in Table 2. Two findings stand out. First, the effect of declining unionization on private sector wage inequality is much larger for men than women. This result is consistent with the large fall in private sector unionization among men. Second, most of the rise in men's wage inequality can be explained by deunionization and the increasing returns to schooling. Educational inequality in wages explains a large fraction of the increase in between-group inequality. Increasing inequality in highly-unionized industries and regions explains a large fraction of rising within-group inequality.

Discussion

We revisited the effect of declining union membership on wage inequality, arguing that unions not only equalized the wages of union members; they also equalized the nonunion wage distribution by threatening union organization and buttressing norms for fair pay. We found strong evidence that unionization rates in detailed industries for geographic regions were positively associated with wage equality among nonunion workers. As unionization rates fell in the national labor market, industry-region unionization rates also declined, and and within-group inequality increased among union and nonunion workers.

A variance decomposition analysis estimated the effect of union membership decline and the effect of declining industry-region unionization rates. When individual union membership is considered, union decline accounts for a fifth of the growth in men's earnings inequality. Adding the normative and threat effects of unions on nonunion pay increases the effect of union decline on wage inequality from a fifth to a third. By this measure, the decline of the American labor movement has added as much to men's wage inequality as the relative increase in pay for college graduates. Among women, union decline and inequality are only related through through the link between industry-region unionization and nonunion wage dispersion. Union decline contributes just half as much as education to the overall rise in women's wage inequality. These results suggest unions were a normative presence that helped sustain the labor market as a social institution, in which norms of equity shaped the allocation of wages outside the union sector.

Of course, not all unions of the 1970s were in the vanguard of egalitarianism. In skilled trades and construction, unions often reinforced racial and ethnic inequalities. Industrial unions originating in the Congress of Industrial Organizations more than the craft unions of the American Federation of Labor supported redistributive public policy that extended to civil rights and the welfare state (Quadagno 1994; Draper 1989). Neither were all nonunion workplaces of the 1970s dominated by norms of fairness and limited markets. Nonunion employers were often virulently anti-union (Jacoby 1997), and wage inequality was clearly higher in the nonunion sector.

Could the union-inequality connection we observed spuriously result from economic changes that helped eradicate organized labor, increase inequality, and perhaps transform labor market norms? Computerization, firm and industry deregulation, and economic globalization may have fostered a more dynamic, and unequal, American capitalism to which unions were poorly adapted (Hirsch 2008). In this case, deunionization is not directly implicated in the process of growing inequality; it is just the visible by-product of new kinds of employment relations in new, nonunion, industries. At least some of the association between industry-region unionization and nonunion wages is likely related to technological, organizational, and institutional changes that eliminated rents and fueled deunionization. Still, changes in the economy are themselves unlikely to be purely exogenous, as the process of economic change spread unevenly, with unions likely slowing or cushioning the impact on wages.

Union decline, too, captures much of the growth in inequality, with just two explanatory variables—union membership and the industry-region unionization rate. Given the parsimony and empirical power of the union effects compared to a multiplex and hard-to-measure process of economic change, we view the empirical analysis as supporting the broad influence of unions on American labor market inequality. Having established the population-level association, more research is now needed to better understand the mechanisms connecting deunionization to nonunion wages in key industries and labor markets.

More generally, the analysis contributes to a political account of rising economic inequality in the United States. The analysis suggests that unions helped shape the allocation of wages not just for their members, but across the labor market. The decline of American labor and the associated increase in wage inequality signaled the deterioration of the labor market as a political institution. Workers became less connected to each other in their organizational lives, and less connected in their economic fortunes. The de-politicization of the American labor market appears self-reinforcing: as the political power of organized labor dissipates, economic interests in the labor market are dispersed and policymakers have fewer incentives to strengthen unions or otherwise equalize economic rewards.

Though industry and regional variation plays an important role, the key comparison implied by our analysis is fundamentally historical, from the early 1970s to the 2000s. In the earlier period, unions offered an alternative to an unbridled market logic, and this institutional alternative employed over a third of all male private sector workers. The social experience of organized labor bled into the nonunion sectors, contributing to greater equality overall. As unions declined, not only did the logic of the market encroach on what had been the union sector, the logic of the market deepened in the nonunion sector too, contributing to the rise in wage inequality.

Appendix 1. Calculating the Adjusted Variances

Between-group and within-group variances in year t can be written in matrix form:

$$B_t = \boldsymbol{w}_t'(\hat{\boldsymbol{y}}_t - \bar{y}_t)^2 \qquad A.1$$

and,

$$W_t = \boldsymbol{w}_t'\boldsymbol{\sigma}_t^2 \qquad A.2$$

where $\boldsymbol{w}_t$ is a vector of survey weights, $\hat{\boldsymbol{y}}_t$ is a vector of conditional means of log wages, $\bar{y}_t$ is the grand mean, and $\boldsymbol{\sigma}_t^2$ is the vector of residual variances. The conditional means and variances are given by the variance function regressions, in matrix form,

$$\hat{\boldsymbol{y}}_t = \boldsymbol{X}_t\boldsymbol{\alpha}_{1t} + \boldsymbol{u}_t\alpha_{2t} + \bar{\boldsymbol{u}}_t\alpha_{3t}$$

and,

$$\log\boldsymbol{\sigma}_t^2 = \boldsymbol{X}_t\boldsymbol{\beta}_{1t} + \boldsymbol{u}_t\beta_{2t} + \bar{\boldsymbol{u}}_t\beta_{3t}.$$

The first adjusted variance is based on the reweighted data:

$$w_{ti}^* = \begin{cases} w_{ti}\bar{u}_b/\bar{u}_t & \text{if } u_{ti} = 1 \\ w_{ti}(1-\bar{u}_b)/(1-\bar{u}_t) & \text{otherwise,} \end{cases}$$

where b is a baseline year set here to $b = 1973$. The adjusted weights, $\boldsymbol{w}_t^*$, are then plugged into the between-group and within-group equations, A.1 and A.2.

The second adjusted variance adds to the union effect by fixing industry-region unionization terms at their 1973 values:

$$\hat{\boldsymbol{y}}_t^* = \boldsymbol{X}_t\boldsymbol{\alpha}_{1t} + \boldsymbol{u}_t\alpha_{2t} + \bar{\boldsymbol{u}}_b\alpha_{3b}$$

and,

$$\log\boldsymbol{\sigma}_t^{*2} = \boldsymbol{X}_t\boldsymbol{\beta}_{1t} + \boldsymbol{u}_t\beta_{2t} + \bar{\boldsymbol{u}}_b\beta_{3b}.$$

The adjusted means and variances, with weights adjusted to 1973 unionization rates, are then plugged into the equations for the between and within-group variance, $\boldsymbol{w}_t^{*\prime}(\hat{\boldsymbol{y}}_t^* - \bar{y}_t)^2$ and $\boldsymbol{w}_t^{*\prime}\boldsymbol{\sigma}_t^{*2}$.

The third adjusted variance fixes the education effects in $\boldsymbol{\alpha}_{1t}$ and $\boldsymbol{\beta}_{1t}$ at their 1973 values. The modified coefficient vectors yields an adjusted set of conditional means, which are plugged into the between-group and within-group equations.

Appendix 2. Correcting for Union Misclassification

Card (1996) analyzed validation data from an employer survey that indicated that the misclassification rate of true union status in the 1977 May CPS is about 2.5 percent ($\lambda = .025$). If 10 percent of respondents are union members, 2.5 percent of the 10 percent mis-report that they are nonunion, and 2.5 percent of the 90 percent who are nonmembers mis-report that they are union members. Formally, if an indicator for true union status, $u^* = 1$, then $p(u_i = 0|u_i^* = 1) = p(u_i = 1|u_i^* = 0) = \lambda$.

Given the observed unionization rate, $\bar{u}$, and the true unionization rate, $\bar{u}^*$,

$$\bar{u} = (1 - \lambda)\bar{u}^* + \lambda(1 - \bar{u}^*).$$

Rearranging terms yields an adjustment to the observed rate, providing the true rate of unionization:

$$\bar{u}^* = \frac{\bar{u} - \lambda}{1 - 2\lambda}.$$

Evidence for misclassification of union status is based on just one year and measurement error in other years may be larger or smaller. To accommodate this uncertainty, we place a prior distribution on λ. True union status, u_i^*, is also unobserved, and uncertainty about union status should be incorporated in the final results. To describe this uncertainty, we can write a probability distribution for u_i^*, given the observed union status.

Rearranging terms we can find the probability distribution for true union status, u_i^*,

$$p(u^*|u = 0, \lambda) = \pi_0{}^{u^*}(1 - \pi_0)^{u^*-1},$$

where,

$$\pi_0 = \frac{\lambda\bar{u}^*}{1 - \bar{u}}$$

and

$$p(u^*|u = 1, \lambda) = \pi_1{}^{u^*}(1 - \pi_1)^{u^*-1},$$

where,

$$\pi_1 = \frac{(1 - \lambda)\bar{u}^*}{\bar{u}}.$$

To incorporate uncertainty about the misclassification parameter, we write

$$p(u^*|u) = \int p(u^*|u, \lambda)p(\lambda)d\lambda,$$

where $p(\lambda)$ is a prior distribution. In our analysis, we specify this distribution to be uniform on the interval, [.02,.03].

If log wages, y_i, are normal, the variance function model, conditional on true union status, u_i^*, can be written,

$$y_i \sim N(\hat{y}_i, \sigma_i^2),$$

where

$$\hat{y}_i = \boldsymbol{x}_i'\boldsymbol{\alpha}_1 + u_i^*\alpha_2 + \bar{u}_i^*\alpha_3$$

and

$$\log \sigma_i^2 = \boldsymbol{x}_i'\boldsymbol{\beta}_1 + u_i^*\beta_2 + \bar{u}_i^*\beta_3$$

Because y_i and u_i^* are independent, conditional on u_i, we can write the likelihood,

$$p(y_i|\theta, u_i) = \sum_{u^*} p(y_i|u_i^*, \theta) \times p(u_i^*|u_i),$$

where all the model parameters are collected in the vector θ. A full Bayesian model is completed by supplying proper priors to θ. Posterior inference for θ is obtained by Markov Chain Monte Carlo simulation.

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