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Wage Structure and Unionization in the U.S. Construction Sector

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Wage Structure and Unionization in the U.S. Construction Sector

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Abstract

This paper estimates the union effects on the wage gap and dispersion in two pooled samples of construction craftworkers (CPS 1983-88 and 2000-05) using decomposition analysis and kernel density estimation. It shows that despite the decline in the adjusted union wage gap declined over time, the unadjusted union wage premium remained high due to the divergence of returns to workforce characteristics in favor of union workers. This pattern was more marked in the basic trades in comparison with the mechanical trades. Unions also contributed to a wider wage dispersion because they created a union wage gap and this wage gap increased across the “competitive” wage distribution. Unions raised the wages of workers who were located in the middle of the wage density but did not have an effect on the lower wage workers.

Keywords: Wage structure, unionization, construction

JEL Classification: J3, J5, L7

1. Introduction

Technology, global trade patterns, minimum wage, and macroeconomic environment are often mentioned as contributors to changes in wage structure. But more than any other factor, the research has emphasized the declining rate of unionization as the cause of the declining wage gap between the union and nonunion workers and rising wage dispersion. This body of literature has frequently distinguished between the experiences of male and female workers, public and private sectors, or countries. Industries, however, usually entered the analysis as categorical controls in regressions. There have been relatively few studies dedicated to specific industries in spite of the fact that the latter exhibit distinct structural characteristics and types of union organization (Belman and Voos, 2004). Across industry variations in the evolution of industrial relations call for industry-specific studies of the wage structure.

The subject of this study is the construction industry. Historically, construction was one of the most heavily unionized sectors in the U.S., and had one of the highest union wage premia. Over the last three decades the U.S construction industry did not experience major dislocation (unlike manufacturing) or expansion (unlike services). Output and employment shares of construction remained relatively unchanged. Construction is relatively insulated from the impact of globalization and changing trade patterns (at least in the partial equilibrium sense) given the essentially non-tradable nature of its product. It is still primarily a local activity, generally unaffected by national or international competition. However, within the sector there were fundamental transformations. The most important change was in industrial relations (Allen, 1994; Philips, 2003; Weil, 2005). The open shop sector increasingly gained control of certain types of construction (residential, small commercial and modeling work), and also made inroads into industrial construction, which was historically dominated by union contractors. Government policies and deregulation increasingly favored the open shop sector. The union density declined steadily and sharply.

One distinguishing feature of construction, relative to manufacturing and much of services, is the predominance of craft unions. Craft unions differ from industrial unions because they represent groups of workers that are homogenous in terms of skill level and occupation. The common bond among the members of the industrial unions, on the other hand, is that they work in the same establishment or industry regardless of skill or occupation. This observation first implies that while industrial unions flatten the wage profile across skill levels and occupations,

this is not necessarily the case for the craft workforce. In fact, the crafts unions can sharpen wage differentials between occupations. Secondly, the bargaining power of craft unions is inseparable from the workers' skills because in an industry where employer-employee bond is loose and workers move continuously between contractors and worksites, the union's ability to deliver contractors high and uniformly skilled workers is the key in collective bargaining. Construction unions are therefore intimately engaged in training and ensuring the existence of a relatively homogenously skilled workforce. In contrast, nonunion contractors are more likely to use workers of various grades of skill across tasks and occupations. Thus, construction sector wage structure is more likely to be driven by occupational characteristics in comparison with other industries.

The objective of this paper is to examine the transformation of the blue-collar construction wage structure. It will estimate the effects of unions on mean wages and wage dispersion, and how these effects changed over time. This comparison is made over two six-year points, 1983-88 and 2000-05. The union impact on mean wages is estimated in terms of standard wage equations using individual level data from the Current Population Survey (CPS). I will use the Oaxaca-Blinder decomposition to assess the effects of union status, workforce characteristics and returns to workforce characteristics on the wage gap and their change over time. The union impact on wage dispersion, on the other hand, is measured alternatively by variance decomposition and semiparametric kernel density methods. There have been only few studies of union effect on average wages with a construction industry focus and by now they are quite dated (e.g. Allen, 1988, 1994; Perloff and Sickles, 1987). The question of wage dispersion in construction, to the best of my knowledge, is never addressed at the national scale.¹

The paper finds that the construction unions' ability to receive higher wages declined along with de-unionization. This was more marked in the basic trades, which have lower skill requirements, in comparison with the mechanical trades. In spite of this decline, high unadjusted wage premia persisted thanks to the divergence of average returns to workforce characteristics among the union and nonunion workers over time, especially among the basic trades workers. Unions are also found to increase wage dispersion. This outcome is due both to the union-nonunion wage gap and the positive covariation of the wage gap and skill levels. Unions affected

¹ Bourdon and Levitt's (1980) discussion of wage variation in construction based on local survey data.

especially the right hand side of the wage distribution, raising the wages of the workers who were in the middle of the wage density and leaving the lower wage workers mostly intact. These results hold for both the basic and the mechanical trades, although the magnitude of the union effect on wage dispersion in the mechanical trades was higher.

Owing to the methodology of reduced form wage equation estimations, the paper is fundamentally descriptive. It reports the changes in the wage distribution and decompositions of these changes, but in the absence of structural parameter estimates, falls short of explaining how these changes were related to the changes in market conditions. Thus, the limited number of comments that pertain to the latter question are inescapably tentative.

The next section provides information about how the wages and workforce characteristics changed in the industry from 1983-88 to 2000-05. It presents descriptive statistics on the dependent variables – union wage gap and wage dispersion – as well as a set of workforce characteristics. The third section is devoted to the wage gap. Following a brief review of the questions related to the theory and measurement of the union wage gap, it presents the union and nonunion wage equation estimates for the construction workforce, and the decomposition of the change in the union wage gap into the effects of union status, workforce characteristics, and returns to the latter. Symmetrically, section 4 presents a brief discussion of theory and measurement of the wage dispersion, and applies variance decomposition and kernel regression to measure the union effect on the wage distribution and its change over time. In sections 3 and 4, I adopt the usual practice of using occupation fixed effect as one of the skill variables. However, there is substantial variation across occupations in terms of the level and change in unionization. Mechanical trades that are commonly ranked higher than the basic trades in the skill hierarchy, are more unionized. In order to investigate variations in the union effect across different segments of the construction workforce, in section 5, I distinguish between the higher and lower skill occupations and examine whether the union effects on the wage premium and dispersion differ across these occupation groups. The final section summarizes the findings.

2. The Data

The empirical analysis uses on the 1983-2005 CEPR extracts of the CPS Outgoing Rotation Group (CPS-ORG) files that provide information on the union status (CEPR, 2007).

The focus of the analysis is on male production workers.² Thus, clerical, managerial, and engineering staff is removed from the sample. Also excluded are the apprentices, helpers, foremen, and supervisors. Finally, observations with imputed earnings are removed from the sample in view of the match-bias problem.³

For the purpose of over-time comparison, I identify two “points” in time. These are pooled samples of observations from 1983-88 and 2000-05. Relative sizes of the construction industry were similar during these periods, constituting 4.4 and 4.6 percent of the real GDP, respectively. After deleting duplicate observations (an outcome of the census design) and observations with missing information, the sample sizes were 28,190 in 1983-88 and 24,468 in 2000-05. Pooling has the obvious advantage of increasing the number of observations. Inconsistencies in CPS’s practice of flagging imputed observations also warrant dropping the observations of the in-between years; only a fraction of observations with imputed earnings were flagged during the intervening years (none in 1994).

2.1. Wage Gap and Wage Dispersion

The central problem in gauging the effects of unions on wages is the construction of the hypothetical wage distribution that would have prevailed in the absence of unions. The observable nonunion wage distribution is a flawed measure of this counterfactual because it overlooks the general equilibrium (e.g. spillover and threat) effects of the unions on the economy-wide wages. In spite of this flaw, in the absence of better measures, nonunion wage is used often as a proxy for the counterfactual. I adopt the same practice in this paper. It should also be borne in mind that general equilibrium effects in the U.S. are likely to be smaller than those in most other industrialized countries because the union sector is smaller, collective bargaining is not centralized and its provisions are not extended to the nonunion sector, and unions are not politically directly involved in wage determination (Card, Lemieux and Riddell, 2003). The usual caveat applies: in this paper, “union effect” refers to the deviation of the observed union wage from the nonunion wages, not from the hypothetical competitive wage.

² Female workers constituted less than 2 percent of the blue-collar construction workforce.

³ CPS imputes the earnings of workers who did not respond to earnings questions by matching a set of their attributes to workers who responded. Since the union status is not one of the matched attributes, exclusion of this criterion leads to the underestimation of the wage gap (Hirsch and Schumacher, 2004). This so-called match-bias problem is more serious in the blue-collar construction workforce which is more heavily unionized than the rest of the workforce. The average match-bias-uncorrected raw wage gap between 1983 and 1988 in construction workforce was 10 percentage points lower than the match-bias-corrected gap. During the 2000-05 period, the differential rose to 27 percentage points.

The wage variable is the hourly earnings in 2005 prices (CEPR-ORG's rw).⁴ One serious shortcoming of this measure is that it excludes pension and benefits, which are typically more generous in the union sector. Therefore, the union-nonunion wage differential does not capture fully the financial advantages of union membership.

Unadjusted union-nonunion wage gap can be calculated by the regression:

$$(1) \quad W = \beta_{0n} + U(\beta_{0u} - \beta_{0n}) + \varepsilon,$$

where W is the (natural log) real wage,⁵ U is a union-status dummy variable taking the value one for organized workers, subscripts u and n stand for union and nonunion, and ε is the error term.

The raw wage gap (unadjusted for workforce attributes) is $\hat{\beta}_{0u} - \hat{\beta}_{0n}$ (the “hat” denotes the parameter estimate) and raw wage premium in percentage terms is $100 \times (e^{\hat{\beta}_{0u} - \hat{\beta}_{0n}} - 1)$. Unadjusted annual premia estimates in the U.S. construction sector, alongside the union density, are illustrated in Figure 1. Unionization rate declined about by half over the period, with particularly sharp drops at the beginning and the end of the term. The raw union wage gap declined precipitously until the mid-1990s. It rose and stabilized afterwards, although it did not reach its early 1980s levels. Relatively lower values observed in the 1990s probably reflect the match-bias problem mentioned earlier. The parallel declines in the union density and the union wage premium were similar to the experience observed in the private sector overall, although the levels of both union premium and density are higher in construction. The horizontal lines in Figure 1 are the union wage premia for the focal sample periods of this study. Union wage premium in the 1983-88 sample was 87 percent. By the end of the period, it declined to 78 percent.

A common measure of wage dispersion, often interpreted as a measure of wage inequality, is the variance of wages. The unadjusted impact of unionization on the overall variance is defined as $V - V_n$, where V and V_n stand for variances of overall and nonunion wages. Unadjusted annual values of the union impact are illustrated in Figure 2, alongside the union density. The first notable observation is the direct relationship between unionization and wage dispersion. Secondly, although annual V and V_n both declined over time (not shown on the

⁴ This variable excludes overtime pay. The nominal hourly wage is topcoded (varying by year). Values of real wage less than \$1 and greater than \$100 are set missing.

⁵ Henceforth, “wage” will refer to the natural log of real wage unless otherwise noted.

figure), the latter did so at a slower pace and the union impact on wage variance declined over time.⁶ Horizontal lines indicate that the union impact on dispersion declined by 0.02 points from 0.07 to 0.05 from the 1983-88 to 2000-05 pooled samples. This result runs against the grain of the studies of wage dispersion in the overall workforce in the US (as well as the UK and Canada), which uniformly found an inverse relationship between union density and wage variance.

Kernel density estimates presented in Figure 3 (section 4) show that the shrinkage in dispersion was not symmetric. In the upper panel both tails thinned over time and the center gained mass, but the larger change was on the right-hand side tail. In fact, the 2000-05 sample density is more log-normal in comparison with the 1983-88 density.⁷

2.2. Workforce Characteristics

Union effects reported above did not make allowances for individual and work-related factors that may influence wages. The adjusted union wage gap is estimated using information on various workforce characteristics available from the CPS-ORG extracts. Age (in quadratic form), education, veteran status, and occupation are standard proxies for experience and skill level. Other observed individual attributes are race/ethnicity, part-time work, marital status, and metropolitan status. Age is a continuous variable. Education is a categorical variable: not high school graduate (the base), high school graduate, and more than high school education. So is race/ethnicity: white (base), Black, Hispanic and “Other” (Native and Asian Americans, and Pacific Islanders). Other individual attributes are binary dummy variables. In order to control for the occupation effect, I identified the largest 14 occupations and the residual “other occupation” category.⁸

Many studies on the union wage gap also include what Lewis called the “extent of unionism” variable measured as the union density in the “area” (e.g. industry, occupation,

⁶ Alternative single coefficient measures of dispersion, Gini coefficient and Theil index, arrive at the same results as the variance.

⁷ Kernel estimates of union and nonunion wages (not shown) indicate that nonunion wage dispersion is wider. This is attributable to the fact that the nonunion workforce exhibits a greater mix of skills. According to Bourdon and Levitt (1980, p. 42), wages in the nonunion sector also vary with the size of the firm and the segment of the market (e.g. residential vs. commercial). Separate union and nonunion estimates underscore that the loss of mass on the right hand tail in Figure 3 is exclusively attributable to the rightward shift of the middle of the union wage distribution. On the nonunion side, the number of lower-wage workers declined and the number of middle-wage workers increased, with no change in the upper tail of the density.

⁸ Occupations are: bricklayer, carpenter, concrete, electrician, HVAC, laborer, operating engineer, painter, plumber/pipefitter, roofer, sheet metal, structural steel, trucker, and welder.

geographic region) within which the worker is located (Freeman and Medoff, 1981; Lewis, 1986, p.146). It is anticipated that workers located in higher union density environments receive higher wages regardless of their union status, and empirical studies support this hypothesis (and the effect is higher for the union workers). I define the union density in terms of a combination of occupation, geographic area and year. The state is not the ideal geographical designation because construction work is intrinsically temporary and construction workers may cross state lines as they move from job to job. Therefore, I identified 17 state groups on the basis of distance.⁹ Pooling of states also alleviated to some extent the small sample problem encountered especially in smaller states and occupations. Then, I calculated the union density in each state group-occupation-year combination as the ratio of the unionized to total workers.

In order to account for the effects of the macroeconomic conditions, I calculated the annual state level construction industry unemployment rates from the CPS-ORG extracts and included them among the explanatory variables.

Table 1 provides summary statistics on all variables by union status in the initial and terminal periods. These figures show that there were significant differences between the average characteristics of the union and nonunion workforce. In the 1983-88 sample, union workers were more likely to be older, better educated, white, married, located in a metropolitan area, and less likely to be a part-timers than their nonunion peers. While these still held in 2000-05, there were several notable changes. Average age increased for all but more so for the nonunion workers, so that the average union-nonunion age differential declined. The share of workers who were not high school graduates dropped sharply and the shares of those with more than high school education increased, resulting overall in a better educated workforce. But these patterns were much stronger in the union workforce, further widening the union-nonunion education level differential. The most conspicuous change in the ethnic/racial composition of the workforce was the rise in the shares of Hispanic workers. The shares of Hispanic workers doubled among the union and more than tripled among the nonunion workers, at the expense of Whites and Blacks.

⁹ Group 1: ME, NH, VT; Group 2: CT, MA, RI; Group 3: NY, NJ, PA, DE, MD, DC; Group 4: OH, IN, MI; Group 5: IL, WI; Group 6: MN, ND, SD; Group 7: IA, MO, NE, KS; Group 8: VA, WV, NC, SC; Group 9: GA, FL; Group 10: KY, TN, AL, MS; Group 11: AR, LA, OK, TX; Group 12: MT, ID, WY; Group 13: CO, NM, AZ, UT; Group 14: WA, OR; Group 15: NV, CA; Group 16: AK; Group 17: HI.

It should also be noted that the reported nonunion Hispanic share is likely to be an underestimate in view of the widespread use of undocumented immigrants in construction.

Table 1 also underscores the substantial decline in the extent of unionization. As expected, nonunion workers were more likely to be located in areas with lower union densities and the union density is on the decline. There were no substantial occupational shifts. Occupation shares and rankings hardly changed for either group of workers.

3. Union Wage Gap

3.1. Theory and Measurement

Economic theory attributes the union effect on wages to the existence of economic rent in the product market and the ability of the union to exercise monopoly power to appropriate a portion of this rent. The size of the rent varies inversely with the degree of competition in the product market. The ability of the union to appropriate the rent, in turn, varies inversely with the elasticity of demand for labor and the intensity of labor in production, and directly with the degree of its monopoly in the labor market. Thus, both labor and product market parameters determine the size of the union wage differential.

In the craft union context another factor gains importance. Craft unions represent workers with similar skills/occupation. They also act as gatekeepers both to ensure that the new entrants to the unionized labor force have acquired the requisite skills, and to enforce jurisdictional boundaries. This requires active involvement of unions in training, testing, and certification of new recruits, as well as upgrading the skills of journey-level workers. Where nonunion workers cannot match the union skill level, a higher union wage would reflect the higher productivity of union workers. Thus, distinct from industrial unionization, the trade union effect on wages derives from both provisioning of training and the traditional economic rent.

Studies on union wage effects usually estimate a pair of reduced-form wage equations for the union and nonunion workers at the individual worker level:

$$(2) \quad W_u = \beta_{0u} + \beta_u X_u + \varepsilon_u,$$

$$(3) \quad W_n = \beta_{0n} + \beta_n X_n + \varepsilon_n$$

where X is a set of variables other than union status, and definitions of other terms follow from equation (1).¹⁰ Vector X typically includes individual and work-related factors that influence the wage rate, and the associated regression coefficients are returns to these characteristics. Now, the term $\hat{\beta}_{0u} - \hat{\beta}_{0n}$ measures the adjusted wage gap. Returns to variables included in vector X across union and nonunion sectors can be compared in terms of $\hat{\beta}_u$ and $\hat{\beta}_n$. If unions flatten wages, elements of $\hat{\beta}_n$ are expected to be larger (in absolute value) and statistically more significant than the corresponding elements of $\hat{\beta}_u$.

Once equations (2) and (3) are estimated, sources of the wage can be identified. Among several methods, the Blinder-Oaxaca decomposition (Blinder 1973; Oaxaca, 1973) is the most popular. This approach decomposes the observed mean wage gap between two samples into the direct effect of the union, the effect of the differences in mean characteristics, and the effect of differences in the returns to characteristics:

$$(4) \quad \bar{W}_u - \bar{W}_n = (\hat{\beta}_{0u} - \hat{\beta}_{0n}) + (\hat{\beta}_u - \hat{\beta}_n)\bar{X}_u + (\bar{X}_u - \bar{X}_n)\hat{\beta}_n.$$

The “bar” indicates the mean value of the variable. In this version of the decomposition, nonunion sector is taken as the reference point, on the presumption that nonunion mean wage approximates the mean wage that would prevail in the absence of unions. The first term on the right hand side of equation (4) is the union effect controlled for workforce characteristics. It is expected to be positive. The second term measures the wage gap that is attributable to the differences in the returns to characteristics of the labor force (excluding the union status). It is commonly referred to as the *price effect*. If unions indeed flatten the wage profile by making returns to worker characteristics more uniform, then $\hat{\beta}_u - \hat{\beta}_n$ (in absolute value) would be larger. The final term is the *quantity effect*, measuring the difference in the wage gap that is attributed to the differences in the workforce characteristics. Its magnitude indicates how similar the union and nonunion workforces are: the more comparable they are in terms of average characteristics, the smaller is the absolute magnitude of this term.

3.2. Econometric Issues

¹⁰ Alternatively, equation (1) can be extended to estimate the pooled regression:

$$(1') \quad W = \beta_{0n} + \beta_n X + U[(\beta_{0u} - \beta_{0n}) + (\beta_u - \beta_n)X] + \varepsilon.$$

Various technical problems afflict individual-level wage estimation, including measurement errors and misclassification of union status. However, the question of omitted variable has attracted by far the most attention. The list of explanatory variables in the wage equation hardly captures all workforce characteristics. Presence of unobserved individual attributes (e.g. ability, trait, motivation) subsumed under the error term in OLS estimates and their potential correlation with the union status variable bias the measured union effect on wages. For instance, if higher wages in the union sector induce lower quality workers to join unions, then the estimate of the union effect on the wages would be biased downward, unless the estimation methodology controls for this correlation. Conversely, if unions or union employers select the higher quality workers, then the estimate overstates the wage premium. Most studies agree that the second effect is likely to dominate and that the direction of the bias is upward.

Empirical research adopted two solutions to address this problem. The first line of attack emphasizes the bias created by the endogeneity of the union status and adds to the model a union status equation. Studies adopting this approach often estimate the probability of union status in a binary choice model, and then either augment the wage equations with the inverse Mills ratio or substitute the union status in the wage equation with the predicted probability of union membership. Alternatively, they estimate a simultaneous system of equations (of union status and wage determination) by maximum likelihood methods. In his review of the literature, Lewis (1986, p.47) concludes that these therapies are worse than the malady because they do not yield robust results and rely on arbitrary exclusion restrictions (on variables assumed to affect union status but not the wage). Furthermore, the union status equation itself typically suffers from misspecification and omitted variable bias (Booth, 1995, p. 174).

The second solution to the problem is to use panel data where changes in union status can be observed over time. The problem of unobserved heterogeneity is averted by estimating the union effect using a sample of individuals for whom both union and nonunion wages are observed. There is also skepticism about these longitudinal studies on grounds that they underestimate the wage gap due to measurement and classification errors, and selectivity of union status changers (Lewis, 1986 p. 61; Booth, 1995, p.176). Thus, it is quite common in the literature to follow Lewis's recommendation to estimate the union impact using OLS, with the appropriate caveat that the estimates are upper boundaries for the union wage gap. Some studies comparing the wage gap over time finesse the problem by pointing out that even if the wage gap

estimate itself is biased, the problem would not be as serious in measuring the *change* in the estimated wage gap over time (Blanchflower and Bryson, 2004).

At first glance, the omitted variable bias strikes with vengeance in construction. Since most of the crafts training takes place on the job, the standard “skill” variables –education and experience– fall short of capturing the quality of workers in construction; and micro data sets do not provide any information on crafts training. A closer inspection, however, suggests a more nuanced interpretation. As already pointed out, the skill level is not exogenous to selection into unions. Construction trade unions participate actively in the registered apprenticeship programs in cooperation with signatory contractors. These programs require, depending on the occupation, three to five years of formal training combining on-the-job training with in-class theoretical instruction. Statistical evidence indicates that since 1980s two-thirds of all new construction apprentices were in union-management sponsored programs organized under the auspices of the collective bargaining agreement, although unionized workforce accounted for less than one-fifth of the total. Apprenticeship retention rate is also higher in the union sector. In addition, many union journey workers undergo periodic retraining in order to upgrade their skills. Nonunion workers, on the other hand, are more likely to be trained informally, in a more limited set of skills (although often across trades), and for shorter periods of time. Thus, in the case of craft workforce the union effect is a composite of economic rent and union-provided training, and it is specious to construe the positive correlation between unobserved skill differential and union status as a source of bias in the measured union effect.¹¹

This is not to say that the omitted variable bias is totally absent. Unions may select workers with higher basic skills into the apprenticeship programs. It is also well-known that not all union workers go through the apprenticeship program. The skill level of workers who qualify for union membership without apprenticeship is probably higher than that of the nonunion workforce on average. Both of these second-order factors would cause the OLS estimates of union effect to be biased upwards.¹² In view of the concerns mentioned earlier, however, I refrain from using

¹¹ Data limitations do not permit an assessment of the relative contributions of economic rent and training on the union wage gap. While it is commonly accepted that union workers are on average more productive than the nonunion counterparts, there are very few studies (e.g. Allen 1988) that compare the productivity levels of union and nonunion workers.

¹² Partial data from Oregon do not suggest that union programs are more selective. Rejection rates of union and nonunion programs are similar. Among apprentices with duplicate applications to both union and nonunion programs, the chances of an applicant being rejected by a union program and admitted to a nonunion program were

simultaneous equation estimation. Data limitations also do not allow a longitudinal analysis of wages in the construction sector to address this residual unobserved heterogeneity problem. Thus, the OLS estimates reported below are still the upper bounds for the union effect.

3.3. Estimation

The OLS estimates of equation (2) are reported in Table 2. These do not indicate that unions flatten wages across workers with different attributes. Age-wage profiles of union and nonunion workers were very similar. In the case of education, union workers received even higher returns to education than nonunion workers. White-Black and White-other race wage differentials were also similar for union and nonunion workers. Although union wages were more uniform by marital status, Hispanic origin, and metropolitan location, these variables still have statistically significant and economically substantial effects on the union wage rate. Wages responded positively to the extent of unionization, but none of the coefficients are statistically significant. It is possible that the state dummies pick up the effect of this variable. Indeed, when four region dummies were used instead of the state dummies (results not shown), the extent of unionization had a statistically significant positive effect on wages and the impact was substantially stronger on the union wage in 2000-05 sample.

Blinder-Oaxaca decompositions of the wage gaps are reported in Table 3. The adjusted union wage gap was 0.534 (2.890–2.356) log points in the 1983-88 sample (a union premium of 70.6 percent), accounting for 85 percent of the raw wage gap. The magnitude of the union effect is substantially higher than the estimates reported in the literature, which is in part explained by the wage equation specifications.¹³ In the 2000-05 sample, the union effect fell sharply to 0.269 log points (a union premium of 30.9 percent). Notably, while the union effect declined by 0.265 points, the unadjusted raw wage declined only by 0.063 points. A sharp increase in the price effect, alongside a modest rise in the quantity effect, made up for most of the declining union effect in the raw wage gap.¹⁴

almost identical to being rejected by a nonunion program and admitted to a union program (Berik, Bilginsoy and Williams, 2008). These findings come from a very limited dataset and are therefore tentative.

¹³ Linneman et al.'s (1990) estimates of the annual union premium in construction over the 1983-1986 period average 41.5 percent. Hirsh's estimates over 1983-88 average 37.3 percent (reported in Allen (1994)). Blanchflower and Bryson's (2004) estimate of 52 percent over 1983-88 is closest to the figure reported in Table 3. For wage gaps estimates in earlier periods, see Freeman and Medoff (1981), Perloff and Sickles (1987), Allen (1988, 1994).

¹⁴ The Oaxaca decomposition is also carried out using the union wage as the reference, viz. in terms of equation (4), quantity and price effects are calculated using nonunion means and union prices, respectively. In this alternative calculation, change in the price effects was smaller and the quantity effect was larger. For the 1983-88 sample, the

At least three factors may account for the lower union effect: rising competition in the product market between the open and organized shop sectors, rising elasticity of labor demand in the union sector (relative to that in the open shop sector), and diminishing union-nonunion skill differentials. Since the present study did not estimate structural equations, is not equipped to assess the relative contributions of these factors. Conjectures based on available observations on the construction industry, however, may facilitate future research. The prime suspect in explaining the lower wage premium is the intensification of competition as emphasized in several accounts of the open shop contractors' success in dominating certain types of construction and penetrating market segments that were traditionally dominated by the unionized contractors (Allen, 1994; Philips, 2003; Weil, 2005). There is probably less of an agreement on other explanations. Relatively higher labor elasticity of demand in the union sector may be the outcome of skilled-labor-substituting technological improvements such as modular construction, off-site construction, and on-site assembly. In order to assess the relevance of the diminishing skill requirements hypothesis, it is necessary to determine the segments of construction market where new technologies are more prevalent. If new skill-saving technologies are applied primarily in the residential and commercial sectors where open-shop contractors dominate, but are of more limited use in industrial construction, then the union wage premium would be expected to increase, not decline. Third, inter-sector skill differentials might have declined due to better training in the open shop construction. However, union-nonunion differentials in enrollment and retention in apprenticeship programs did not change much since the mid-1980s. Anecdotal evidence also indicates that unions are more effective on journey worker training. More recently, open shop contractors attempted to coordinate their efforts through Wheels of Learning and subsequently Cantron, which devised standardized curriculum to produce highly skilled craft workers especially in the mechanical trades. The performance of these programs is yet to be assessed fully (at least in the case of apprenticeship they do not match the retention rates of union programs), but the available evidence suggests that it is unlikely for them yet to make a notable mark on the average skill level of the nonunion workforce.

Finally, it should not be forgotten that the lower premium may merely be an outcome of union bargaining strategy: if unions gradually substituted benefit raises for wage increases, then

union, price, and quantity effects were found to be 0.534, -0.112, and 0.209. The corresponding values in the 2000-05 sample were 0.269, 0.035, and 0.264.

the estimated union effects in the 2000-05 period would be biased downward. There is, indeed, anecdotal evidence supporting the hypothesis that unions increasingly focus on benefits rather than wages in collective bargaining.

As Table 3 shows, the price effect has increased dramatically by 0.183 points and offset almost 70 percent of the union effect. Contributions of selected variables to the price effect are reported in the middle panel of Table 3. Age and education are prominent factors among these. As also observed in Table 2, returns to age across sectors moved in opposite directions over time, diminishing slightly in the nonunion sector but rising substantially in the union sector. Consequently, the wage gap increased in favor of the union sector. Similarly, returns to college education increased for both college and high school graduates in the union sector, while they remain virtually unchanged or declined in the nonunion sector. Together, age and education explain 22 percent of the rise in the price effect which helped maintain higher average relative union wage. Another factor was the metropolitan status. While unions flattened the wage profile across metropolitan and nonmetropolitan workers in the 1983-88 sample, this was substantially less so in 2000-05, due both to the declining returns in the nonunion and rising returns in union returns (Table 2). This explains another 16 percent of the increase in the price effect. Two other factors are the wage response to the business cycle and Hispanic origin. First, while wages were procyclical in the 1983-88 sample, especially in the union sector, they became noncyclical in the 2000-05 period. Procyclical union wage is typically taken to suggest that the unions are able to take advantage of economic rents that rise during the expansion. By this token, rents were apparently no longer as important a source of premium during the more competitive 2000-05 period; the noncyclical union wage implies that the more traditional argument of union attempting to insulate wages from the vagaries of the business cycle applies. Secondly, the already high white-Hispanic wage differential in the nonunion sector rose even more in 2000-05, while in the union sector it started low and diminished over time. Both of these factors raised union wages relative to nonunion wages, further increasing the price effect.¹⁵

¹⁵ It is unwieldy to determine at the state level whether the wage gap varies with the geographical distribution of workforce, but when I used region level controls instead of states, geographic factors were observed to be important. First, over time, relatively more workers were located in the West and Midwest where the returns to union workers were higher. Secondly, returns to union workers especially in the East and the West increased substantially and contributed to the higher price effect. It is also noteworthy that there were a few states in which relative union wages increased sharply to raise the price effect: California (0.020), Illinois (0.012) and New York (0.011).

Divergence of returns in union and nonunion sectors signals hypotheses concerning the labor force transformation. Specifically, among the unionized workers, variables such as age, education, and metropolitan status may be increasingly positively correlated with desirable unobserved attributes including experience, skills, and construction workforce attachment. In the nonunion sector these correlations may be weaker due to a higher turnover of the workforce both by occupation and industry. Consequently, the workers with identical observed characteristics may be receiving widely different wages given the difference in their unobserved attributes.

The overall quantity effect reported in the lowest panel of Table 3 is much smaller but it masks substantial effects of redistribution of workforce characteristics on the wage gap, albeit operating in opposite directions. For instance, the negative effect of the declining age gap between the union and nonunion workforces is -165 percent of the overall quantity effect (more than offsetting the sum of the age and education price effects noted above). However, faster increase of representation of Hispanic workers in the nonunion sector lowers the wage gap by almost an equivalent amount.

4. Wage Dispersion

4.1. Theory and Measurement

Nonunion wage density is expected to be more dispersed than the union wage density but the impact of unions on the overall wage dispersion is theoretically indeterminate. The union wage gap between the otherwise comparable union and nonunion workers raises the wage dispersion. If unions are stronger among the better paid segments of the labor force, as Friedman (1962, p.124) argued to be the case, dispersion between the higher and lower paid groups of workers would be further accentuated. Conversely, self selection of lower skilled workers into unions would reduce the dispersion. Another channel through which unions affect wage dispersion is proposed by Freeman (1980): to the extent that unions implement “standard rates” across and within establishments, they reduce wage dispersion. Ability to implement standard rates across establishments would depend on the degree of product market competition: the more competitive is the product market, the more desirable are the uniform wages from both the firms’ and the workers’ perspectives because it “takes wages out of competition.” Within-establishment standardization of wages also attaches wages to jobs rather than individuals and helps in maintaining union solidarity. In the case of craft unionization, however, it is widely recognized that the standardization argument is not likely to apply. Currently, in the U.S. construction

industry, after many mergers, unions are organized in fourteen crafts. General or specialty contractors enter into collective bargaining agreements with each of the local craft unions individually. Contractors are required to hire union workers in crafts they are signatory to collective bargaining agreements, but in other parts of the project they can hire nonunion workers. Thus, on the jobsite members of different trade unions work side by side, and often alongside nonunion workers even when the contractors are “union” (Philips, 2003). Where workers are organized along craft lines and wages are not negotiated “across the board,” occupational wage differentials within an industry or establishment will persist so long as some individual unions can successfully negotiate relatively higher wages than the others (Rees, 1979, p.60; Freeman, 1980; Card, Lemieux and Riddell, 2003). Thus, wage gap effect is expected to dominate the standardization effect in construction. The magnitudes of these effects, however, remain thus far unmeasured.

The union effect on wage dispersion is defined as the difference between the observed overall wage dispersion and the hypothetical wage dispersion that would prevail in the absence of unions (Figure 2). As before, I will use the nonunion wage distribution as the counterfactual. Among others, two alternative methodologies are commonly used in gauging the impact of the union effect: variance decomposition and kernel density estimation.¹⁶

4.2. Dispersion Measurement: Variance decomposition

Freeman (1980) used conditional variance decomposition to express the union effect as the sum of within- and between-sector components. Within-sector component derives from the differences between the variances of the union and nonunion sectors. Standardization of wages implies that the wage dispersion is smaller in the union sector, reducing the overall variance. Between-sector effect, on the other hand, is attributable to the union-nonunion wage gap. A higher gap increases the overall wage dispersion.

Let V , V_u and V_n stand for the variances of overall, union and nonunion real wages, \bar{W}_u and \bar{W}_n be the mean wages in the union and nonunion sectors, and α be the fraction of workers unionized. The impact of unions on the overall wages, or difference between the overall and nonunion wage variances can be written as:

$$(5) \quad V - V_n = \alpha(V_u - V_n) + \alpha(1 - \alpha)(\bar{W}_u - \bar{W}_n)^2.$$

¹⁶ JMP (after Juhn, Murphy and Pierce (1993)) and quantile regression are two other methods used to decompose dispersion.

The first term on the right hand side of this equation is the within-sector effect, or the difference in the dispersion of wages in the union and nonunion sectors. Wage standardization argument predicts that this term is negative. The second term is the between-sector effect, which reflects the union wage premium. Its sign is expected to be positive: the higher is the difference between the mean wages in the union and nonunion sectors, the wider is the wage dispersion. Freeman's empirical work and many others following it found that in the overall workforce the within-sector effect is larger than the between-sector effect and concluded that unions reduce wage dispersion.

This so-called "two-sector model" assumes that the unionization rate, wage gap, and variance gap are constant for all workers, regardless of skills or characteristics. Later studies pointed out that ignoring these differences lead to spurious within- and between-sector effects. Suppose that union workers are on average more skilled than the nonunion workers. Then the within-sector effect as defined above may in part be the outcome of greater homogeneity of the union workforce. Between-sector effect would also be spurious to the extent that it reflects average skill differences between the union and nonunion workers. In order to incorporate the differences in the characteristics of the union and nonunion workforce into the decomposition, Card (2001) divided the entire sample into characteristic- (or what he called skill-) controlled groups. These groups were identified in terms of predicted wages from nonunion wage equation estimates. He then recalculated within- and between-sector effects as unionization rate-weighted averages across these groups. Card's extension of the Freeman decomposition is:

$$(6) \quad V - V_n = E[\alpha(c)(V_u(c) - V_n(c))] + E[\alpha(c)(1 - \alpha(c))(\bar{W}_u(c) - \bar{W}_n(c))^2] \\ + Var[\alpha(c)(\bar{W}_u(c) - \bar{W}_n(c))] + 2Cov[\bar{W}_n(c), \alpha(c)(\bar{W}_u(c) - \bar{W}_n(c))]$$

where c denotes the characteristic group and $\alpha(c)$ stands for the share of union workers in the group. The left-hand side of equation (6) is again the union effect on the variance of wages.¹⁷ First two terms on the right-hand side are, respectively, the within- and between-sector effects as in equation (5) but now they are union-share weighted averages over characteristic groups. Card's method adds two more components to the variance decomposition that capture the "across" group effects: variability of union wage gap across groups and changing union wage gap across the wage distribution. The third term on the right-hand side is the effect of unions on

¹⁷ This is not numerically the same as the union effect in equation (5) because V_n is now reweighted in view of the fact that the distributions of nonunion workers and the overall workforce across skill groups are identical.

mean wage gap variance across different skill groups. If the union wage effect is higher for the higher skilled workers then the unions raise wage gap variance and therefore the wage dispersion. The fourth term is the covariance of union wage gap with the nonunion (competitive) wage. Again, if higher skilled workers are more likely to be unionized, then union wage gap varies directly with the level of competitive wages. Card's empirical analysis concluded that the overall effect of unions on dispersion is still negative but smaller in absolute value relative to the findings of the two-sector model.

4.3. Measurement: Kernel Estimation

Dinardo, Fortin and Lemieux (1996) (henceforth DFL) pioneered semiparametric kernel estimation methods to inspect the impact of unionization, minimum wages and other factors on the entire wage distribution. In order to find the impact of changing unionization on wage density over two points in time, they compared the actual wage density of the terminal period with the counterfactual wage density that would have prevailed if unionization, but no other workforce characteristic, had maintained its initial period values, and workers were paid according to the terminal period wage schedule. Visual comparison of the actual and counterfactual kernel densities show which sections of the density function are most influenced by the union presence. The technical challenge is to reduce the multi dimensional conditional density to a unidimensional measure. DFL solved this problem by creating a reweighting function that maps the distribution of unionization in the initial period to the distribution of the latter period: terminal observations are reweighted such that those that were more likely to be unionized conditional on initial period characteristics are weighted up and those that are less likely are weighted down. Calculation of the weights requires first pooling the initial and the terminal period observations and using a binary choice model to estimate the union status as a function of workforce attributes. The weight for each union worker is determined as the ratio of the predicted probability of being a union member conditional on initial period characteristics to that conditional on the terminal period characteristics. The kernel estimate of this counterfactual weighted wage density is then compared with the actual to determine where in the wage density unions are the most influential.¹⁸

4.4. Dispersion of Construction Wages: Variance Decomposition

¹⁸ Problems of general equilibrium effects unionization and non-random selection of union status apply to the DFL as well. In addition, the method is assumes that the wage distribution is not affected by the unionization rate.

The union effect on construction industry wage dispersion has not been a subject of investigation and what follows is the first attempt at measurement of the nationwide wage dispersion and its change over time. The first panel of Table 4 presents the wage variances. Overall, nonunion, and union wage variances are reported in the top three rows. The union wage variances were smaller than the nonunion – a stylized fact since Freeman (1980) – although the gap narrowed. While this finding is usually explained in terms of wage standardization by unions, by industry or establishment, it is more likely to be attributable in the construction industry to the smaller range of skill differences among the union craft workers. Bourdon and Levitt (1980, p. 42) also argued that the nonunion wage variance is higher due to variations in wage scales by segments of the industry, e.g. lower in residential than in commercial construction.¹⁹ Both the overall and the nonunion wage variance declined over time, although the latter declined by a relatively smaller amount (both declines were statistically significant at less than 0.01 level). The union wage variance remained unchanged.

The second panel in Table 4 decomposes the union effect into within and between-sector components following equation (5). The first row of the panel is the union effect – the difference between the overall and nonunion wage variances. These figures show that unions raised the wage dispersion in construction trades, and over time, concomitant with the decline in unionization the union effect declined substantially, by almost one-third. The union effect accounts for 66 percent of the decline in variance ($-0.023/-0.035$). Within- and between-sector effects are reported in the next two rows. In contrast to the findings on the overall labor force, the between-sector effect is higher than the within-sector effect (in absolute values), and the decline in the union effect was due primarily to the declining mean wage gap between the union and nonunion sectors.

The third panel in Table 4 follows Card (2001) by identifying ten characteristics-based groups of equal size in terms of the predicted wages from nonunion wage equation estimates. The union effect now explains 69 percent of the decline in variance. Much of the previously reported within- and between-sector effects are indeed spurious and across group differences

¹⁹ Another potential factor that would influence the union wage variance is the inter-occupation dispersion of wages. If trade unions representing workers in higher skill occupations drive a wedge between their wages and other unionized workers, then the union wage variance would increase. Estimation of inter-occupation wage differentials is beyond the scope of this paper, but a comparison of wages in higher and lower-skill occupations is presented in section 5.

matter. In comparison with the two sector-model, between-sector effects in the characteristics-controlled model were lower by more than one-half. The total across-group effects were quite large, even matching the between-sector effect in the 2000-05 period. They were due primarily to the positive covariance between the weighted wage gap and the nonunion (competitive) wage. That is, unions increased the wage gap more for higher-skilled workers, resulting in a higher variance of wages. The declining covariance shows, however, their ability to raise relative wages in the higher skill groups weakened over time.

4.5. Dispersion of Construction Wages: Kernel Estimation

The solid lines in the upper panel of Figure 3 are the kernel estimates of actual wage densities in the two samples. In the lower panel of Figure 3, facilitates comparison by illustrating the differences in densities. The solid line is the difference between the actual 2000-05 and 1983-88 sample wage densities. Had there been no change in density the difference line would be horizontal. The counterfactual wage distribution where unionization is kept at its 1983-88 value and workers are paid under 2000-05 wage schedules is illustrated by the dashed line in the upper panel of Figure 3, and the difference between the actual and counterfactual 2000-05 densities is the dashed line in the lower panel. If unionization had played no role in the changing wage density, the actual and counterfactual 2000-05 densities would be identical and the dashed difference line would be horizontal at zero. Positive(negative) values of the dashed line in the lower panel of Figure 3 indicate that, there would have been fewer(more) workers at those wages in 2000-05 had unionization remained at its 1983-88 level.

Figure 3 confirms that the wage variance declined over time. The actual difference line illustrates that masses on either tail of the density shrank, adding to the mass in the center of the distribution, with the larger change taking place in the right hand side of the distribution. For purposes of comparison, kernel estimate variances are reported in the third panel of Table 5. According to the second line of the lowest panel, had unionization remained unchanged from 1983-88 to 2000-05, the variance would have declined by 0.021. Thus, declining unionization explains 38 percent $[1 - (-0.021 / -0.034)]$ of the decline in wage variance, which is quite a bit lower than the values obtained from the Freeman and Card variance decompositions.

The counterfactual density exercise indicates that had unionization remained at its 1983-88 level, the wage density in 2000-05 would have lied slightly to the right of the observed density. The union effect moves workers from roughly the lower two-thirds of the distribution to

the upper one-third, and the primary beneficiaries are those in the middle section of the distribution. Furthermore, while de-unionization was a major factor in explaining the loss of mass on the right tail, it was not relevant to the thinning of the left-hand tail. The lowest wage segment (log wage <2) is not at all affected by unionization.

5. Wage Dispersion in Mechanical and Basic Trades

So far the analysis adopted the standard practice of using occupation controls to capture the skill differences that are missed by the standard age and education variables. This approach is restrictive because it does not permit variability in the returns to workforce characteristics across occupations. For instance, one may question, following Table 3, whether the estimated union and price effects were occupation-specific; or, following Figure 3, whether the relatively smaller union effect on the lower segment of the wage density is attributable to the population of the left-hand tail of the distribution with less unionized occupations.²⁰

One way to address such questions is to carry out the analysis for each occupation separately. But this route is likely not to be robust given the relatively small numbers of observations in many occupations. As a compromise, I opted to perform estimations for two groups of occupations – mechanical and structural/finishing – to investigate whether union impact on wages varies by occupation group. Mechanical trades include electrical, plumbing/pipefitting, sheet metal, heating-air conditioning, and welding. Structural/finishing trades are bricklaying, concrete, laborer, operating engineer, painting, roofing, and structural steel trades. I also added truck driving to the latter category, and refer to this group as *basic trades*.²¹ There exists a pecking order among the trades in terms of the required skill levels. Mechanical trades are considered to require more skills. Term lengths of apprenticeship programs are four to five years in the mechanical and three years in the basic trades. In the 1983-88 sample, average real wages in basic trades were \$12.21 per hour in nonunion and \$22.56 in the union sector. In the 2000-05 sample, the corresponding figures were \$12.75 and \$22.48. Mechanical trades workers on average earned more: the average nonunion and union hourly wages were, respectively, \$14.97 and \$26.82 in 1983-88, and \$16.00 and \$25.61 in 2000-05.

²⁰ Blanchflower and Bryson (2003, p. 217), for instance, suggested that occupations are likely to simply be slices of the wage distribution and argued against inclusion of occupation controls in wage regressions.

²¹ Carpentry, the largest occupation, is excluded because it contains workers who are highly heterogeneous in terms of skill level.

There were clear differences in the union densities of the two occupation groups. Union densities in the mechanical occupation group were higher: 38.4 percent in 1983-88 and 31.2 percent in 2000-05. In basic trades, these figures were 25.4 percent and 16.1 percent, respectively. The mechanical trade workers were on average older, better educated, and more likely to be white.

5.1. The Wage Gap in Mechanical and Basic Trades

The upper panel of Table 6 reports the wage gap and its decomposition by occupation groups. Similar to the patterns observed for the whole construction workforce, union effects on wage gaps declined over time, and to some extent this was offset by the price effect changes. However, there are also several notable differences. Union wage gap was substantially higher in the basic trades in the 1983-88 sample, indicating that workers in lower skill occupations benefited relatively more from unionization. Over time, however, the union effect in basic trades collapsed from 0.589 to 0.027 log points. Only thanks to an almost commensurate increase in the price effect was there still a substantial raw wage gap in the basic trades. In mechanical trades, parallel changes in the union and, especially, price effects were modest in size. The adjusted union effect is smaller, but still substantial, explaining 40 percent of the raw wage gap. In fact, in the 2000-05 sample, the union effect in the mechanical trades was higher than that in the basic trades by an order of seven. The relatively small magnitude of the price effect in the mechanical trades indicates that the previously observed divergent returns to characteristics in Table 3, which are attributed to the correlation of observed and unobserved workforce attributes, explains the sustained union wage differential in the basic trades.

These results raise more questions on both the levels and the changes in the wage gap. The initial high union premium and its subsequent collapse in the basic trades suggest that although there were relatively fewer union workers they enjoyed substantial insulation from competition originally; but their situation changed drastically for the worse over time. The change may be attributable to more intensive use of these workers in market segments where de-unionization is more prevalent (e.g. residential, commercial). In contrast, relative concentration of mechanical workers in more specialized construction (e.g. industrial) that is relatively less competitive may have allowed the relatively higher union premium to continue. At the same time, the sharp rise in the price effect indicates that the structure of the union workforce in basic trades was transformed radically and started commanding higher returns.

5.2. Wage Dispersion in Mechanical and Basic Trades

The variances and variance decompositions are reported in Table 6. Except for the mechanical trades in the 2000-05, union wages had lower dispersion and the dispersion declined over time. In mechanical trades, union wage variance *increased*, and exceeded nonunion wage variance. Decompositions show that the union effects on dispersion were positive and they were attributable to the union wage gap and the positive covariance between the nonunion wage and the wage gap across groups. Unions raised the dispersion more in the mechanical trades. This is due to the greater “between sector” effects in the mechanized trades. In contrast, “across characteristics group effects” were similar in magnitude. Within each occupational group, wage gap-nonunion wage covariances indicate that unionization added to the dispersion because the more “skilled” (as measured by the predicted nonunion wage) workers were more heavily unionized. Over time the union effect waned due to the decreases in the between-sector effect and the across-group effects. The between-sector wage gap declined by 0.015 in the basic trades and by 0.021 in the mechanical trades. Covariances between the wage gap and the nonunion average wage declined by 0.012 and 0.011. In summary, wage dispersions in both occupational groups were similarly affected by declining unionization although the magnitude of change is higher for the mechanized trades. These results suggest that occupational groups were not simply slices of the wage distribution and the changes in mass observed in specific segments of density shown in Figure 3 is unlikely to be an occupation-specific phenomenon.

This point is underscored by Figures 6 and 7 that compare the actual wage densities in each trade group with the 2000-05 counterfactuals where the unionization is kept at its 1983-88 level. As the dashed difference line shows, in each case, had unionization remained at the 1983-88 level, the middle of the density would lose mass and the right-hand side would be fattened. Thus, similar to the overall labor force, unions benefited workers in the middle of the distribution, and did not affect those at the lowest end of the distribution. The union effect on dispersion was stronger for the mechanical trades, as the greater loss of mass of the right-hand tail of the difference line illustrates.

6. Conclusion

The change in unionization since the 1980s influenced both the union wage premium and wage dispersion in the construction industry. Comparing the mid-1980s and early 2000s, this paper finds that the union wage gap was sharply lower at the end of the period, although the raw

wage differential between the union and nonunion sectors declined only slightly. This is due to the fact that union-nonunion differentials in returns to various factors such as age, education, Hispanic ethnicity, metropolitan status increased sufficiently to mitigate the decline in union wage premium. Among the structural and finishing trades, the magnitudes of the change in adjusted union premium and the divergence of return to observed characteristics were both higher, relative to the mechanical grades. While union membership on average benefitted the basic trades more in the 1980s, the situation was reversed after 2000. While the methodology used in the paper does not pinpoint the underlying causes of the declining union premia, the most likely explanation lies in increasing competition in the construction industry. Results also suggest that, unobserved positive characteristics of union workers keep their wages higher relative to their peers in the nonunion sector and maintain a high unadjusted union wage gap in their favor.

In contrast to the entire economy, unions increased the overall wage dispersion in the construction industry. The joint effect of the union wage gap and the higher rate of unionization among the higher skilled workers dominated the union wage standardization effect. This is not an entirely surprising result. Most economists suspect that the craft unions, unlike industrial unions, are likely to increase wage dispersion in the trades, although no evidence had been presented thus far. The lower dispersion is the outcome of thinning of both ends of the wage distribution as the middle gained more mass. The impact of declining unionization, however, is felt primarily on the right-hand tail distribution; workers were pushed from the higher wage segment of the density to the middle. This was true for both mechanical and basic occupations, but the magnitude is larger in the former. The observed losses in the lower end of the distribution, on the other hand, had little relation to the change in unionization.

Table 1: Sample Means

	1983-88		2000-05	
	Nonunion	Union	Nonunion	Union
Wage (in 2000 prices)	13.30	24.49	13.78	23.94
ln(real wage)	2.51	3.14	2.55	3.11
Age (years)	30.4	37.4	33.4	37.8
>High school education (%)	19.8	21.9	21.2	35.5
High school graduate (%)	45.3	53.3	45.3	51.6
< High school graduate (%)	34.9	24.8	33.5	12.9
Veteran (%)	19.3	31.4	20.6	22.9
Married (%)	56.0	74.8	52.1	65.4
White (%)	79.0	83.2	56.7	74.9
Black (%)	9.0	7.7	5.3	6.4
Hispanic (%)	10.6	7.3	35.8	15.8
Other race (%)	1.4	1.9	2.2	2.9
Part-time (%)	10.8	7.2	7.9	4.9
Metro (%)	72.0	87.8	80.6	86.1
Union density (%)	22.8	37.9	15.9	31.6
Unemployment	15.7	17.4	9.2	10.5
Bricklayer (%)	2.9	4.1	3.2	3.8
Carpenter (%)	26.1	17.3	24.0	19.1
Concrete (%)	2.2	1.9	2.1	2.1
Electrician (%)	6.4	12.4	6.8	17.1
Heat/AC (%)	3.3	2.2	3.8	3.1
Laborer (%)	16.4	14.3	18.2	11.9
Operating eng. (%)	5.5	6.7	4.7	6.1
Painter (%)	7.6	3.4	8.2	3.2
Plumber (%)	5.7	9.4	5.6	10.4
Roofer (%)	4.3	2.1	4.1	1.9
Sheetmetal (%)	1.6	3.0	1.0	3.0
Structural met. (%)	0.6	3.2	0.4	3.0
Truck driver (%)	3.4	3.3	3.2	2.2
Welder (%)	1.5	3.3	1.6	1.9
Other (%)	12.6	13.3	13.1	11.1
N	17,867	6,630	17,191	4,953

Note: Observations are weighed by “earnings weight” as calculated by the CEPR (orgwgt).

Table 2: Wage Equation Estimates

	1983-88		2000-05	
	Nonunion	Union	Nonunion	Union
Constant	2.356 (0.041)**	2.890 (0.091)**	2.454 (0.035)**	2.723 (0.106)**
Age	0.049 (0.002)**	0.046 (0.002)**	0.040 (0.001)**	0.054 (0.003)**
Age-squared/100	-0.054 (0.002)**	-0.049 (0.003)**	-0.044 (0.002)**	-0.059 (0.004)**
>High school education	0.103 (0.007)**	0.105 (0.012)**	0.113 (0.007)**	0.151 (0.017)**
High school graduate	0.094 (0.006)**	0.087 (0.010)**	0.082 (0.006)**	0.096 (0.016)**
Veteran	-0.027 (0.007)**	-0.010 (0.009)	-0.006 (0.008)	-0.020 (0.013)
Married	0.089 (0.006)**	0.073 (0.009)**	0.076 (0.005)**	0.039 (0.010)**
Black	-0.167 (0.009)**	-0.163 (0.015)**	-0.135 (0.012)**	-0.135 (0.020)**
Hispanic	-0.172 (0.009)**	-0.123 (0.016)**	-0.185 (0.007)**	-0.100 (0.016)**
Other race	-0.086 (0.022)**	-0.081 (0.031)**	-0.096 (0.017)**	-0.081 (0.031)**
Part-time	-0.119 (0.008)**	0.019 (0.014)	-0.079 (0.009)**	-0.020 (0.021)
Metropolitan	0.090 (0.006)**	0.043 (0.011)**	0.075 (0.007)**	0.066 (0.015)**
% Organized	0.019 (0.050)	0.044 (0.063)	0.073 (0.051)	0.047 (0.074)
Unemployment	-0.126 (0.075)	-0.458 (0.125)**	-0.084 (0.106)	0.069 (0.185)
R ²	0.32	0.33	0.31	0.35
N	17,867	6,630	17,191	4,181

Notes: Each regression includes 14 occupation, 51 state, and 6 year controls.

Observations are weighted by “earnings weight” as calculated by the CEPR (orgwgt).

Standard errors in parentheses. ** and * indicate statistical significance at the 1% and 5% levels.

Table 3: Oaxaca-Blinder Decomposition of the Wage Gap

	1983-88	2000-05	Δ
<i>Decomposition</i>			
Raw wage gap	0.631	0.568	-0.063
Union effect	0.534	0.269	-0.265
Price effect	-0.075	0.108	0.183
Quantity effect	0.172	0.192	0.020
<i>Individual price effects¹</i>			
Age	0.003	0.019	0.016
>High school education	0.000	0.014	0.014
High school graduate	-0.004	0.007	0.011
Married	-0.012	-0.024	-0.012
Hispanic	0.004	0.013	0.009
Part time	0.010	0.003	-0.007
Metropolitan	-0.038	-0.008	0.030
Unemployment	-0.015	-0.003	0.012
<i>Individual quantity effects¹</i>			
Age	0.070	0.037	-0.033
College	0.002	0.016	0.014
Married	0.017	0.010	-0.007
Hispanic	0.006	0.037	0.031
Metropolitan	0.009	0.004	-0.005
Union Density	0.003	0.012	0.009

¹ Results on selected variables are reported.

Table 5: Variance Decompositions

	1983-88	2000-05	Δ
<i>Wage variances</i>			
Overall	0.231	0.196	-0.035
Nonunion wage variance	0.160	0.148	-0.012
Union wage variance	0.133	0.133	0.000
<i>Variance decompositions</i>			
Two-sector model:			
Total union effect	0.071	0.048	-0.023
within-sector	-0.007	-0.003	0.004
between-sector	0.079	0.051	-0.028
Characteristics-controlled model			
Total union effect	0.064	0.040	-0.024
within-sector	-0.005	-0.002	0.003
between-sector	0.036	0.021	-0.015
across-group	0.033	0.021	-0.012
wage gap variation	0.004	0.002	-0.002
wage gap-nonunion wage covariance	0.029	0.019	-0.010
<i>Kernel estimates</i>			
Actual variance	0.230	0.196	-0.034
Unionization adjusted variance	0.230	0.209	-0.021

Note: Observations are weighted by “earnings weight” as calculated by the CEPR (orgwgt).

Table 6: Wage Gap and Variance Decompositions in by Occupation Groups

	Basic Trades			Mechanical Trades		
	1983-88 ¹	2000-05 ¹	Δ	1983-88 ¹	2000-05 ¹	Δ
<i>Wage gap decomposition</i>						
Wage gap	0.634	0.583	-0.051	0.599	0.476	-0.123
Union effect	0.589	0.027	-0.562	0.331	0.187	-0.144
Price effect	-0.146	0.353	0.499	0.130	0.164	0.034
Quantity effect	0.191	0.203	0.012	0.139	0.125	-0.014
<i>Variances</i>						
Overall	0.219	0.176	-0.043	0.224	0.187	-0.032
Nonunion	0.156	0.140	-0.016	0.146	0.137	-0.007
Union	0.133	0.126	-0.007	0.129	0.141	0.012
<i>Variance Decomposition</i>						
Characteristics controlled model:						
Total union effect	0.060	0.035	-0.025	0.076	0.041	-0.035
within-sector	-0.005	-0.001	0.004	-0.003	-0.002	0.001
between-sector	0.032	0.017	-0.015	0.044	0.023	-0.021
across-group	0.033	0.019	-0.014	0.035	0.020	-0.015
wage gap variation	0.004	0.002	-0.002	0.006	0.002	-0.004
wage gap-nonunion wage covar.	0.029	0.017	-0.012	0.029	0.018	-0.011
<i>Kernel estimate</i>						
Actual variance	0.222	0.179	-0.043	0.227	0.191	-0.036
Unionization adjusted variance	0.222	0.195	-0.027	0.227	0.201	-0.026
<i>N</i>	10,147	8,905		5,271	4,781	

Notes: Observations weighted by “earnings weight” as calculated by the CEPR (orgwgt).

Figure 1: Union Density and Raw Union Wage Gap

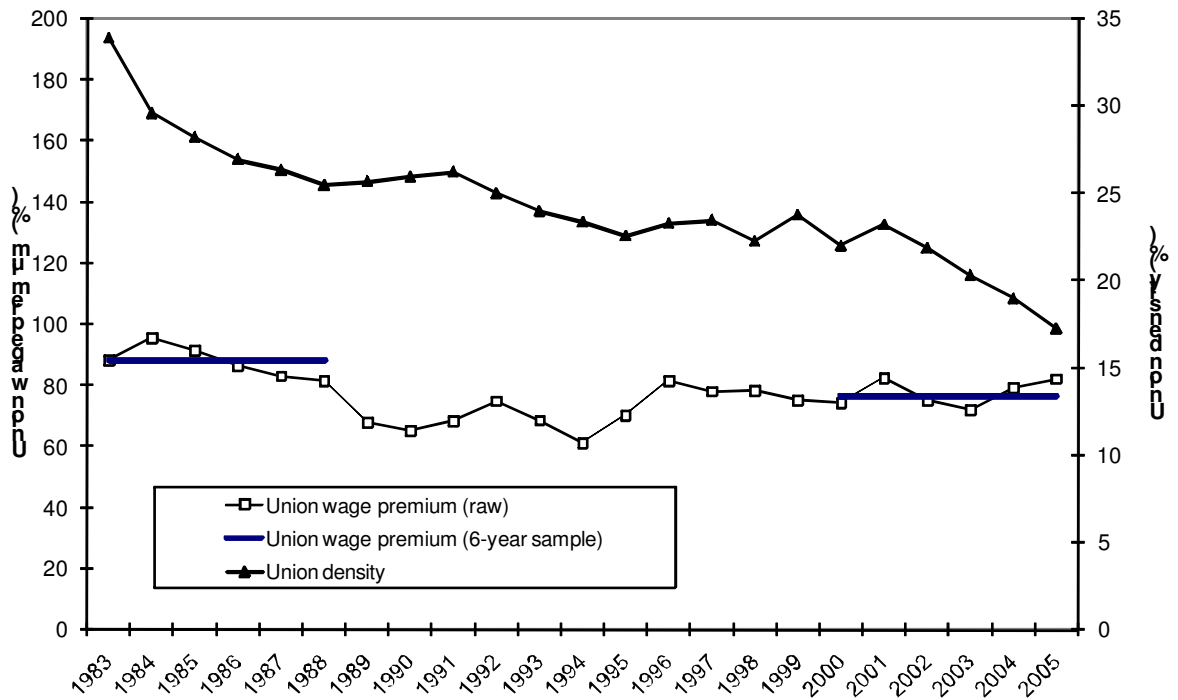


Figure 2: Union Density and Union Impact on Wage Dispersion

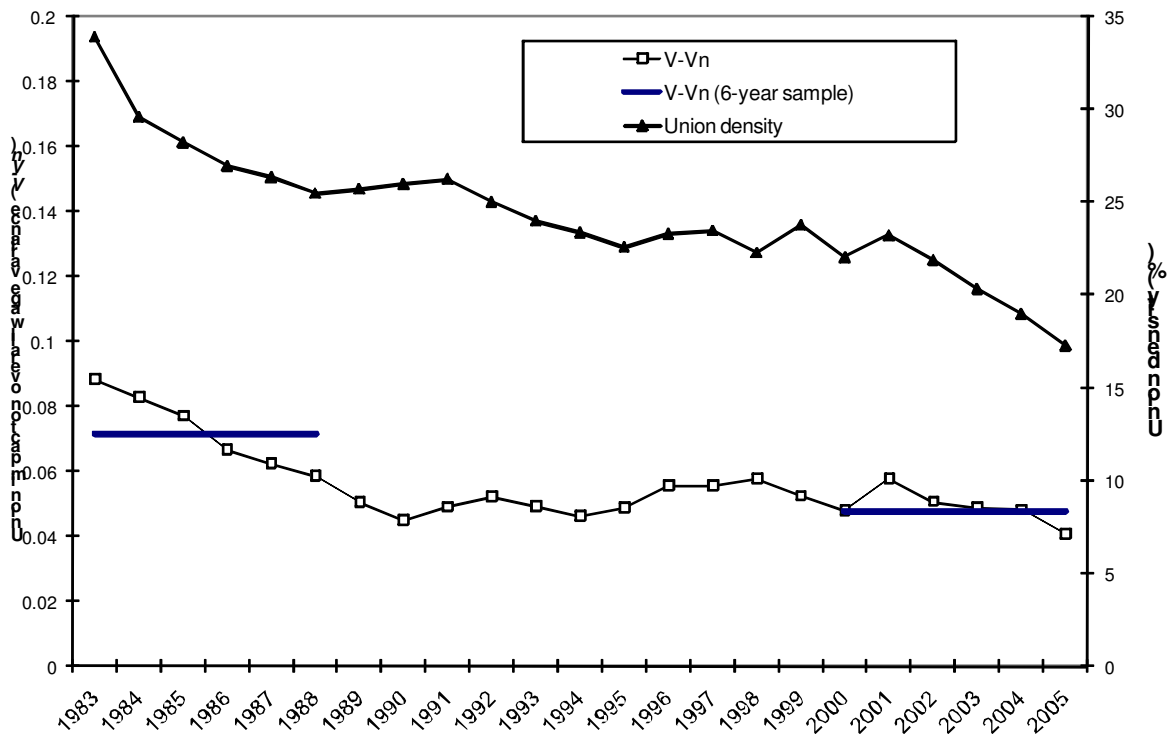


Figure 3: Union Impact on Wage Density

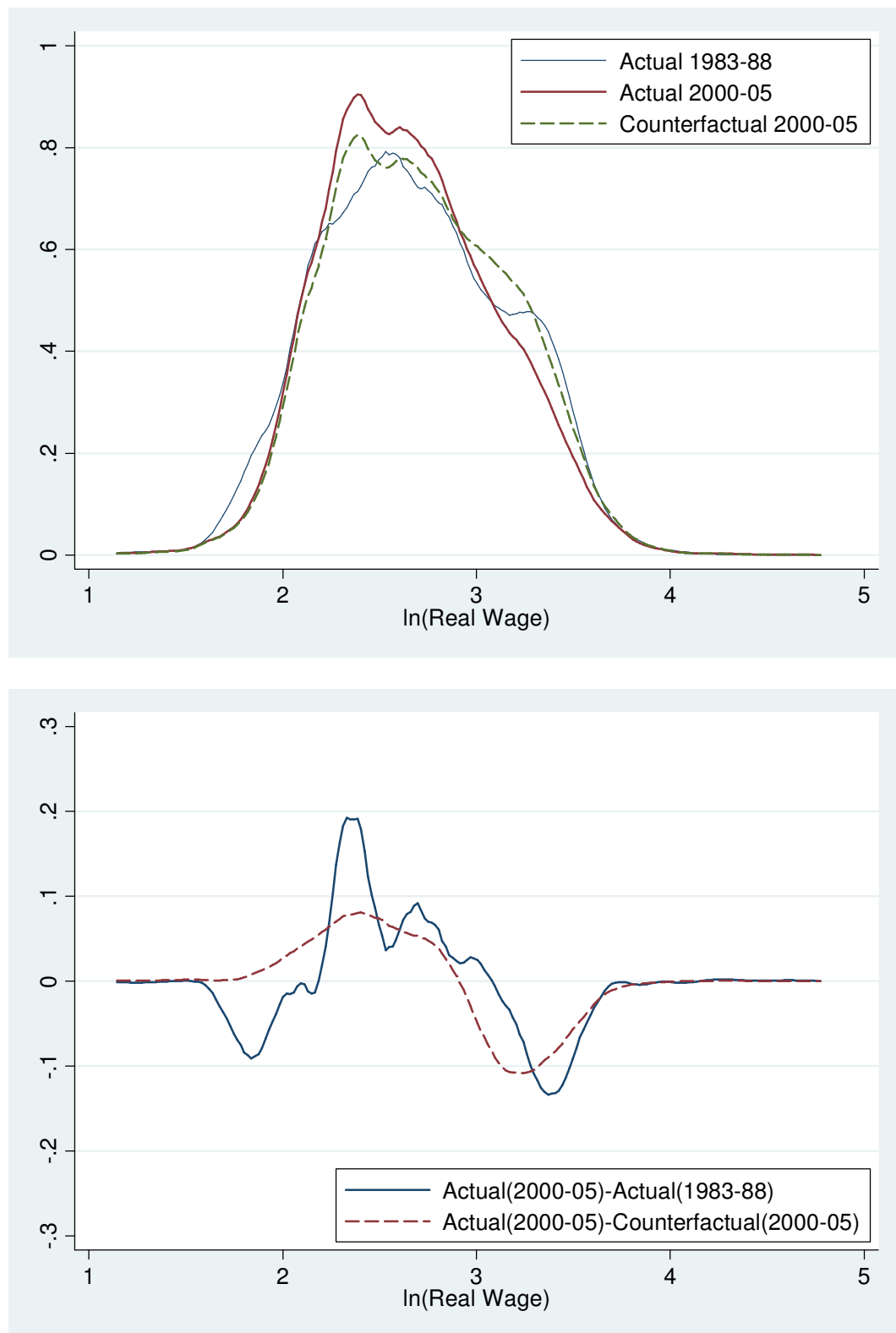


Figure 4: Union Impact on Wage Density -- Basic Trades

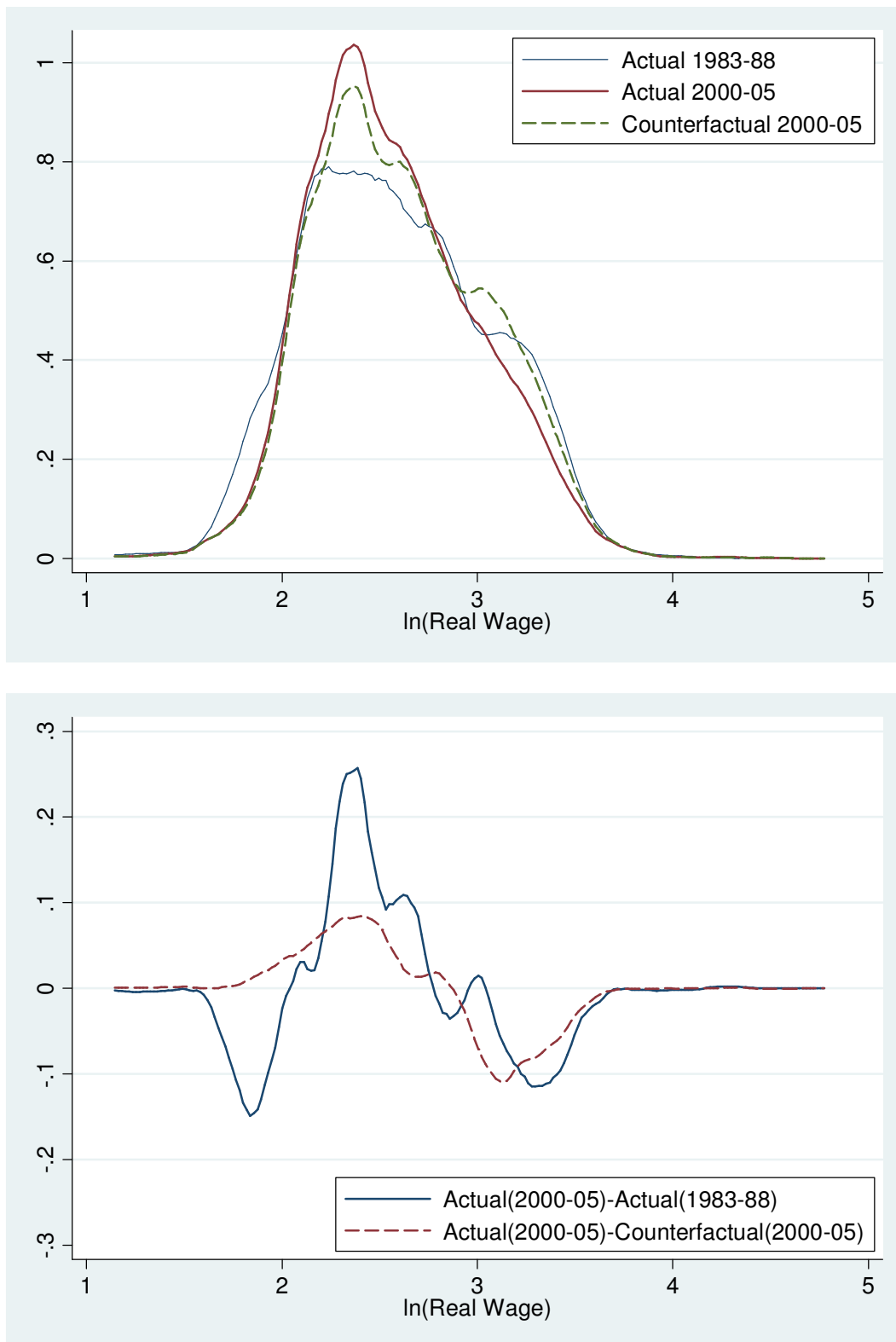
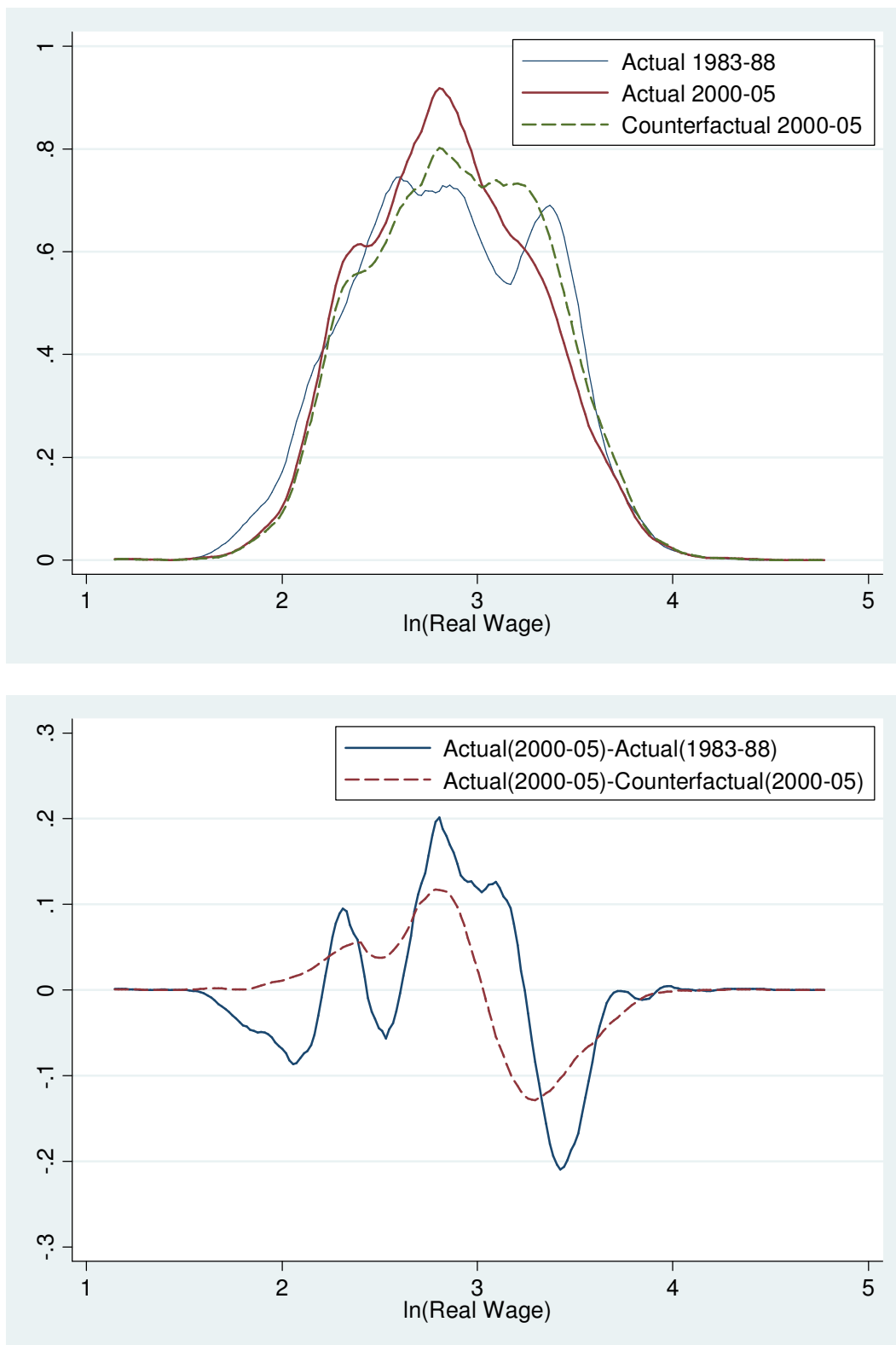


Figure 5: Union Impact on Wage Density -- Mechanical Trades



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