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Unions and Wage Inequality in Mexico

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Abstract: This paper offers empirical evidence on the impact of trade unions on wage inequality in Mexico. The results indicate that unions were a strongly equalizing force in the dispersion of wages in 1984, but only half as effective at reducing wage inequality in 1996. Not only did the unionized percentage of the labor force fall rather significantly over the period, but unions also lost some of their ability to reduce wage dispersion among the workers they continued to represent. Had unions maintained in 1996 the same structural power that they possessed in 1984, the rise in wage inequality between the two periods would have been reduced by roughly one-third.

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Rising wage inequality has become an international phenomenon in recent decades, affecting both developed and developing economies – Mexico among them. The explanations offered for increased wage dispersion in Mexico have focussed on the rather dramatic liberalization of trade the country undertook beginning in the mid-1980s.[1] Among the various explanations are the following. Trade liberalization has placed Mexico into competition with other developing economies that possess even lower wages for less-skilled labor, thereby depressing the wages of unskilled workers relative to skilled workers (Hanson and Harrison 1999; Cragg and Epelbaum 1996). Trade liberalization has resulted in a disproportionate reduction in the appropriable rents of blue-collar workers, whose jobs suffered the largest reductions in protection levels (Revenga 1997). The reduction of restrictions on the inflow of foreign capital has resulted in an increased demand for skilled labor by foreign firms, and therefore rising relative wages of skilled workers (Feentra and Hanson 1997).

This paper explores a different explanation for growing wage dispersion in Mexico. It builds on the research of economists who have studied the role of domestic institutional changes – such as declining unionization and the declining real value of the minimum wage – in rising U.S. wage inequality (Freeman 1993; DiNardo, Fortin, and Lemieux 1996). In this paper, I focus on the effect of declining union power on growing wage dispersion in Mexico.[2] The results suggest that unions reduce wage dispersion in Mexico, but that this dispersion-reducing effect was much more pronounced in the mid-1980s – prior to certain institutional changes that contributed to the declining power of unions – than in the mid-1990s. Unions reduced formal sector wage inequality, for instance, by roughly 6.5% in 1984, but by only 3.2% in 1996. Had the power of unions to affect wage dispersion remained the same over the period, the rise in wage inequality would have been reduced by roughly one-third.

The first section of the paper reviews the theoretical and empirical literature on the relationship between unions and wage inequality, and presents suggestive empirical evidence of a link between declining union power and rising wage inequality in Mexico over the period of the 1980s and 1990s. The second section spells out an estimation procedure for examining this relationship in greater depth, and discusses the data to be used in the empirical work. The third section discusses the empirical results. The final section offers concluding thoughts and some recommendations for future research.

Unions and the Dispersion of Wages

Unions may affect the dispersion of wages in three distinct ways (Lewis 1963; Freeman 1980): First, unions raise the wages of union workers relative to a group of nonunion workers who, either because of legal prevention or because of the extreme difficulty of organizing, are “beyond the pale” of unionization. In developed economies such as the U.S., this latter group is typically higher-paid white-collar workers, whereas in developing countries such as Mexico this group is typically lower-paid informal sector workers. Whether this effect of unions leads to an increase or decrease in the overall dispersion of wages depends on the average wage level of the workers who are untouched by unionization. In the U.S., this effect of unions lowers overall wage dispersion by bringing the average wage of blue-collar workers closer to the average wage of higher-paid white-collar workers. In Mexico, the effect is most likely the opposite –unions raise the wages of higher-paid formal sector workers relative to those of lower-paid informal sector workers.

The second effect of unions on wage dispersion comes from the higher wages union workers obtain relative to comparable nonunion workers in the same industries and occupations. This effect increases wage dispersion by creating wage differences between similarly situated workers that would not exist except for the actions of unions.

Finally, unions may reduce wage dispersion within the union sector as compared to the comparable nonunion sector, thereby contributing to a reduction in overall wage dispersion. Unions may do this in a variety of ways and for a variety of reasons. Unions typically strive to “take wages out of competition” both within and across plants with so-called standard rate policies (Reynolds and Taft 1956; Slichter, Healy, and Livernash 1960). The effect of such policies may be to reduce or eliminate arbitrary pay differences across workers (based on gender or geographical region, for example) and to level the wage structure with regard to the productive characteristics of workers (such as education and occupation).

Within plants, standard rate policies take the form of collective bargaining language establishing a system of formal job classifications that tie wages to jobs as opposed to people. Because these union policies create “equal pay for equal work” they may eliminate arbitrary pay differences, such as those based on gender. In addition, unions may negotiate, for ethical reasons or to foster solidarity, a leveling of the pay structure across these job classifications, thereby raising the pay of less-skilled workers by proportionately more than that of high-skilled workers.

Unions also typically strive to standardize pay across plants both within and even across industries. In the best of contexts, multi-plant and multi-employer bargaining can result in similar pay for similar job classifications within an industry, and may even reduce wage differentials between similarly-skilled workers across industries. This can act to reduce arbitrary pay differences across workers based, for example, on geographical region.

These activities will reduce the dispersion of wages in the union sector as compared to the comparable nonunion sector to the extent the latter possesses arbitrary pay differences across workers within plants, or across plants, employers, or industries.[3] These activities will also reduce overall wage dispersion if – because of ethical motives, the desire to foster worker solidarity, or collective decision making based on a median voter model – unions negotiate a leveling of the pay structure across skills as compared with the pay structure in the nonunion sector.

There is significant empirical evidence to suggest that the dispersion-decreasing effects of unions dominate their dispersion-increasing effects in developed countries. Freeman (1980) was the first to clarify, and then to offer empirical evidence on, these three distinct effects of unions on wage dispersion in the U.S. The lack of micro-survey data has, until very recently, prevented studies of this kind in the developing world. In fact, I am aware of no empirical study, patterned after that of Freeman's, on the impact of unions on wage inequality in a developing economy. A recent paper on unions in South Africa employing a different empirical technique finds evidence for an equalizing effect of unions on wage inequality in that country (Schultz and Mwabu 1998).

If unions are an equalizing force in the overall distribution of wages, and if the power of unions to equalize wages has waned at the same time that the distribution of wages has become more unequal, it is legitimate to ask to what extent the rising wage inequality may be attributable to declining union power.[4] The empirical evidence for the U.S., where the above conditions are met, suggests that the declining power of unions can account for a significant share of the increased wage inequality over the period of the 1980s (Freeman 1993; DiNardo et al. 1996). Freeman's results suggest that roughly 20% of the rise in wage inequality can be accounted for by the declining percent of unions in the U.S. labor force.

Are these conditions met in the case of Mexico? Income inequality in Mexico, as in many other Latin American countries, is and historically has been high compared to income inequality in other developing economies. Whereas the Gini coefficient for total income in Mexico is about 53.7, it is 48.4 in Malaysia and only 36.5 in Indonesia (Griffin and Ickowitz forthcoming). During the course of the 1980s and 1990s, inequality in the distribution of income went from bad to worse in Mexico. The share of household income possessed by the bottom 60% of households fell from 26.8% to 25.5% between 1984 and 1998, while the share going to the top 10% of households rose from 34.1% to 38.1% (La Reforma 2000).

Labor income accounts for roughly 50% of total household income in Mexico, and almost two-thirds of household income if non-monetary income – such as the imputed rental value of owner-occupied housing – is excluded. Moreover, the increased dispersion of labor income has been a major factor in rising income inequality in Mexico over the 1980s and 1990s, the largest component of which is wages (Alarcon and McKinley 1997).

Figure 1 shows the changing distribution of wages in the overall economy (i.e., the formal and informal sectors combined), and the changing distributions in the union and nonunion segments of the formal sector, between 1984 and 1996.[5] Looking first at the changes over time, the graph reveals that the variance of the natural log of wages – a common measure of dispersion – in the overall economy rose from 0.555 to 0.663, or roughly 19%, during this period. To get some sense of what this means, let us compare the percentage difference between various quintiles and the median ln wage across the two years. In 1984, the first quintile was 10% below the median ln wage, and the fifth quintile was 11% above the median ln wage. In 1996, the first quintile was 39% below the median ln wage, and the fifth quintile was 47% above the median ln wage.

The results also reveal that while dispersion rose in both the union and nonunion segments of the formal sector, the rise in the variance of ln wages in the union sector was greater – 81% compared to 37% among nonunion workers. The ratio of nonunion/union dispersion fell from 3.92 to 1.39, or roughly 65%, over the period. Thus, the evidence indicates that wage dispersion grew substantially in Mexico over the 1980s and 1990s, but more so in the union sector than in the nonunion sector.

A second important feature of the Figure 1 results is that the variance of ln wages is lower in the union sector than in the nonunion sector for both years. Figures 2 and 3 give comparisons of wage dispersion in the union and nonunion sectors of more detailed industry and occupation categories. The results are similar – the variance of ln wages is always lower in the union sector.[6] This evidence suggests that unions are able to reduce within-group dispersion, and therefore may well have an equalizing effect on the overall dispersion of wages.

What happened to union power over this same period? To begin with, unions appear to have suffered a significant decline in coverage over the period. The data I use in this study reveal a decline from roughly 31% to 21% in the unionized percentage of the formal sector labor force between 1984 and 1996. The factors accounting for this decline are not yet well understood. However, most labor experts would cite changing political forces as one of the most important contributing factors.

The success of the labor movement in Mexico has been linked historically to its alignment with the ruling Institutional Revolutionary Party (PRI), which held power for most of the twentieth century (Middlebrook 1995). State and federal labor authorities can exert substantial influence over both the union registration process and contractual relations between unions and employers. Beginning with the shift in Mexico's development strategy in the 1980s, unions fell into disfavor among influential members of the PRI. As a consequence, union organizing and maintenance of membership became more difficult in the 1980s and 1990s. The government also acted to contain union bargaining demands, forcing wage concessions or minimal wage increases in particularly crucial industries. In reducing the benefits of unionization, these government actions may have quelled support for unions among existing members and made union organizing less attractive for those who were considering membership.

Not only did unions lose coverage among the labor force, they also appear to have lost the power to level wages and to reduce or eliminate arbitrary pay differences through collective bargaining. While the influence of centralized bargaining structures has never been great in Mexico, beginning in the 1980s, various developments appear to have affected the ability of unions to take wages out of competition across plants. These developments include: the break up of bargaining structures in the newly-privatized industries (De la Garza 1990); the concerted efforts of the political authorities to maintain lower wages in the foreign assembly plants of the north (Shaiken and Herzenberg 1987); and modifications to the *contratos-leyes*, which are legal mandates initiated by unions that create wage uniformity across plants in particular industries and geographical regions.

Perhaps even more importantly, the newly emerging assembly plants in the north brought demands for greater labor flexibility in production. This led to wider job classifications, or to their absence altogether, in many unionized plants, and to greater discretion on the part of management to attach wages to people as opposed to jobs. Pay-for-skills payment schemes may increase wage dispersion across individuals in a plant, while gain-sharing arrangements may do the same for work groups. Moreover, once wages are no longer attached to jobs, there is increased scope for arbitrary pay differences to arise across workers. These institutional reforms soon spread from the northern plants to their industry counterparts in the south, and then to other industries as well. This made union efforts to level wages and to reduce arbitrary pay differences within plants more difficult.

The evidence thus far presented is at least consistent with the view that unions reduce wage dispersion in Mexico, and that lost union power accounts for some of the rising wage inequality over the period of the 1980s and 1990s. However, the evidence is suggestive at best. The lower dispersion of wages in the union sector could be the result of union policies that foster greater wage equality, but it could also be reflective of less dispersion in the wage-determining characteristics (such as the education level or the occupational mix) of union workers compared to nonunion workers. Moreover, lowering within-group wage dispersion is only one of three ways in which unions may affect overall wage dispersion. A method must be found for assessing the importance of the other two effects of unions on wage inequality, and then a procedure must be established for combining the three effects to arrive at an assessment of the union impact on overall wage inequality.

Even if union policies are found to be genuinely responsible for the lower within-group wage dispersion in the union sector, and even if this dispersion-reducing effect of unions is found to dominate the two dispersion-increasing effects, the evidence presented thus far does not firmly implicate lost union power in the rising wage inequality of the 1980s and 1990s. The decline in unionization may result from changes in the structural power of unions to organize new members and to maintain old ones, but it may also be due to the changing industry and occupation mix of the labor force. Similarly, the increased relative dispersion of wages in the unionized segment of the formal sector might be due to changing contract language and the breakdown of centralized bargaining structures, but it might also be merely the result of increased relative dispersion in the wage-determining characteristics of union workers. And, finally, a method must be found for combining the changes in the three effects of unions on wage dispersion in order to discern how important changes in union power have been to the growing dispersion of wages. It is to these various tasks that I now turn.

Empirical Specifications and Data

This section of the paper spells out procedures for combining the three effects of unions on wage inequality and for separating wage dispersion owing to wage-determining characteristics from dispersion owing to the structural or institutional determinants of wages. These procedures allow for the calculation of a set of counterfactuals that sheds light on the effects of unions on wage inequality. In particular, I shall be able to answer the following kinds of counterfactual questions: What would the dispersion of wages be in the economy if union workers were transferred to the nonunion sector and subject to the pay policies that exist there? What would the dispersion of wages be in the economy in 1996 if unions possessed the same structural power to influence wages as they did in 1984?

The metric I use to capture the dispersion of wages is the variance of ln wages, whose square root – the standard deviation of ln wages – is a common measure of wage inequality. The variance of ln wages is an appropriate metric to use when wages are set by the standard ln wage equation of empirical labor economics, when the union impact on wages is measured in relative terms, and when wages follow a lognormal distribution.

My initial focus will be on the union and nonunion segments of the formal sector labor market. I shall make use of the following conditional variance formula:

$$\text{(1)} \qquad \sigma_F^2 = \overline{U}\sigma_U^2 + (1-\overline{U})\sigma_N^2 + \overline{U}(1-\overline{U})(\overline{W}_U - \overline{W}_N)^2,$$

where $\overline{U}$ is the percent unionized in the formal sector, σ_U^2 is the variance of ln wages among unionized workers, σ_N^2 is the variance of ln wages among nonunion workers, $\overline{W}_U$ is the mean of ln wages in the union segment, and $\overline{W}_N$ is the mean of ln wages in the nonunion segment.

The formula allows for the derivation of the formal sector variance of ln wages based on information about the union and nonunion segments that together compose this sector. Note that the formula neatly displays the two distinct ways in which unions can influence formal sector wage dispersion. Unions increase formal sector wage dispersion by increasing the mean ln wage for union workers compared to nonunion workers. Unions decrease wage dispersion to the extent they reduce the variance of ln wages in the union segment compared to the nonunion segment.

The conditional variance formula for the formal and informal sectors combined is:

$$\textbf{(2)} \qquad \boldsymbol{\sigma_T^2 = \overline{F}\sigma_F^2 + (1-\overline{F})\sigma_I^2 + \overline{F}(1-\overline{F})(\overline{W}_F - \overline{W}_I)^2},$$

where

$$\textbf{(3)} \qquad \boldsymbol{\overline{W}_F = \overline{U}\,\overline{W}_U + (1-\overline{U})\overline{W}_N},$$

$\overline{\mathbf{F}}$ is the percent of the labor force in the formal sector, $\boldsymbol{\sigma_F^2}$ is the variance of ln wages in the formal sector (equation 1 above), $\boldsymbol{\sigma_I^2}$ is the variance of ln wages in the informal sector, $\overline{\mathbf{W}}_{\mathbf{F}}$ is the mean of ln wages in the formal sector, and $\overline{\mathbf{W}}_{\mathbf{I}}$ is the mean of ln wages in the informal sector.

The two effects of unions on formal sector wage dispersion discussed above are captured in equation (2) through the $\boldsymbol{\sigma_F^2}$ term. The third effect of unions on total wage dispersion stems from their influence on formal sector wages, and thus on the difference between the mean of ln wages in the formal and informal sectors. If informal sector wages are higher on average than formal sector wages, then unions reduce this spread, and this third union effect is dispersion reducing. If, on the contrary, formal sector wages are higher on average than informal sector wages, then this third union effect is dispersion increasing.

The counterfactuals I generate make use of the conditional variance formulas and the estimated coefficients and residual variances from ln wage equations of the following form:

(4) $$\mathbf{W_U = \alpha_U + \beta_U X_U + \varepsilon_U},$$

(5) $$\mathbf{W_N = \alpha_N + \beta_N X_N + \varepsilon_N},$$

where **W** is ln wages, $\boldsymbol{\beta}$ is a vector of estimated coefficients (capturing the structural and institutional determinants of wages), **X** is a vector of wage-determining characteristics, $\boldsymbol{\varepsilon}$ is the error term, and **U** and **N** are subscripts that refer to the union and nonunion sectors respectively.

A common concern in the estimation of equations (4) and (5) is bias due to sample selection. The difference in estimated coefficients across these two equations may reflect the impact of union pay policies, or simply the differential selection of workers into the union and nonunion sectors.[7] Differential selection may result in bias if selection is based on wage-determining characteristics that are unobserved to the researcher and yet correlated with observed wage-determining characteristics on the right-hand-side of estimated wage equations. There are different theories about the nature of the selection process. One theory posits that unionized employers may seek to reduce the union wage rent by hiring from the labor pool those workers with high unobserved skills (Lewis 1986). In this case, the estimated union relative wage effect is likely to be biased upward, but estimates of the dispersion of wages across skill categories in the two sectors may be unaffected.

However, a second theory regarding the selection process suggests that the pattern of selection may bias downward the estimated returns to skills in the union sector relative to the nonunion sector, which could lead us to erroneously conclude that unions reduce wage dispersion across skill categories. The theory runs as follows: Assume that unions do indeed reduce the returns to skills, that workers with the highest union wage premiums are more likely to seek jobs in the union sector, and that employers seek workers with the lowest union wage premiums (Abowd and Farber 1982). In this case, high-skill workers that choose to locate in the union sector are more likely to possess low unobserved skills, and low-skill workers that are hired by unionized employers are more likely to possess high unobserved skills. The rate of return to observed skills is therefore likely to be biased downward in the union sector, with the higher pay of low-skill workers reflecting their higher unobserved skills, and the lower pay of high-skill workers reflecting their lower unobserved skills. The selection process is reversed, producing an upward bias in the return to observed skills, in the nonunion sector.

I am unable to correct for selection bias with panel data techniques, which recently have been shown to be particularly useful for addressing the second pattern of selection discussed above (Lemieux 1998). I rely instead on the standard two-step selection correction procedure of Heckman (1979).[8] That is, I model the process by which workers are selected into the union sector, and then estimate equations (4) and (5) jointly with the union selection equation:

(6) $$\mathbf{U = \alpha + \delta Y + \varepsilon},$$

where **U** is a dichotomous (0,1) variable indicating union status, $\boldsymbol{\delta}$ is a vector of estimated coefficients, **Y** is a vector of variables determining the union status of workers, and $\boldsymbol{\varepsilon}$ is the error term.

Armed with the conditional variance formulas and information from the estimation of equations (4) and (5) for the two years 1984 and 1996, we can proceed to the construction of appropriate counterfactuals that shed light on the two central questions posed by this paper: Do unions reduce the dispersion of wages? Has the changing power of unions contributed to rising wage inequality over the period?

Do unions reduce the dispersion of wages?

To explore the effect of unions on wage dispersion, I compare the actual variance of ln wages with various counterfactual measures of this variance. There are two counterfactuals to consider: First, union workers are transferred (as it were) to the nonunion sector, where they are subject to the pay policies that exist there. Second, nonunion workers are transferred (as it were) to the union sector, where they experience the pay policies that exist there. In the first scenario, I utilize the estimation results from equation (5) and the characteristics of union workers to generate the variance of ln wages and the mean of ln wages that union workers would possess were they to be located in the nonunion sector. In the second scenario, I utilize the estimation results from equation (4) and the characteristics of nonunion workers to generate the variance of ln wages and the mean of ln wages that nonunion workers would possess if they were to be located in the union sector.

In each case, the counterfactual measures of the within-group variance of ln wages and the mean of ln wages are placed into equations (1) – (3) to derive counterfactual measures for the variance of ln wages in the formal sector and the variance of ln wages in the formal and informal sectors combined. If unions are instrumental in reducing wage dispersion, the counterfactual variance ln wage measure under the first counterfactual scenario should be significantly larger than the actual measure. Under the second counterfactual scenario, it should be significantly smaller than the actual measure. The difference between the actual and counterfactual measures is purely reflective of structural differences in pay policies across the union and nonunion sectors, and not of differences in wage-determining characteristics across the sectors.

The precise derivation of the two counterfactual components – i.e., the within-group variance of ln wages and the mean of ln wages – under the first counterfactual scenario is as follows. (The derivation of these measures under the second counterfactual scenario follows an analogous procedure.) The counterfactual union variance of ln wages is:

(7) $\hat{\sigma}_U = \hat{\sigma}^2_{\text{exp lained}} + \hat{\sigma}^2_{\text{residual}}$,

where

(8) $\hat{\sigma}^2_{\text{exp lained}} = \Sigma_i \hat{\beta}^2_{N_i} \sigma^2(X_{U_i}) + \Sigma_i \Sigma_j \hat{\beta}_{N_i} \hat{\beta}_{N_j} \sigma(X_{U_i} X_{U_j})$,

(9) $\hat{\sigma}^2_{\text{residual}} = \sigma^2(\hat{\varepsilon}_N)$,

and the $\hat{\beta}_{N_I}$ and $\hat{\beta}_{N_j}$ are estimated coefficients from the nonunion ln wage equation, $\sigma^2(X_{U_i})$ is the variance in characteristic **i** among union members, $\sigma(X_{U_i} X_{U_j})$ is the covariance between characteristics **i** and **j** among union members, and $\sigma^2(\hat{\varepsilon}_N)$ is the variance of the estimated residual in the nonunion ln wage equation. In practice, the counterfactual explained variance can be derived by placing the wage-determining characteristics of union workers into the nonunion estimated wage equation, and then solving for the variance of predicted ln wages. The counterfactual residual variance is simply the regression sum of squared errors divided by the number of observations from the nonunion estimated wage equation. [9]

The second component of the counterfactual calculation is the counterfactual union mean of ln wages:

$$\textbf{(10)}\quad \hat{\overline{\mathbf{W}}}_{\mathbf{U}} = \hat{\boldsymbol{\alpha}}_{\mathbf{N}} + \hat{\boldsymbol{\beta}}_{\mathbf{N}}\overline{\mathbf{X}}_{\mathbf{U}}$$

where $\hat{\boldsymbol{\beta}}_{\mathbf{N}}$ is the vector of estimated coefficients from the nonunion ln wage equation, and $\overline{\mathbf{X}}_{\mathbf{U}}$ is the vector of mean characteristics from the union sample.

When these two counterfactual components are placed into the conditional variance formulas, the other components of the formulas – namely, the variance of nonunion ln wages, the mean of nonunion ln wages, the variance of informal sector ln wages, the mean of informal sector ln wages, and the distribution of employment across the formal and informal sectors – are assumed to remain unchanged. Thus, I am ignoring the spillover and threat effects of unions on employment and wages in those sectors of the economy not covered by unions. This is a common approach in the literature on union relative wage effects (Lewis 1986).[10]

Has changing union power contributed to rising wage inequality over the period?

If unions are found to decrease wage dispersion, then it is reasonable to ask whether the decline in union power between 1984 and 1996 can account for any of the rise in wage inequality over the period. To shed light on this issue, I derive counterfactual measures that address the following question: What would the variance of ln wages have been in 1996 had the structural power of unions remained the same as it was in 1984?

Two counterfactual measures are considered. The first simply takes the 1984 values of the dispersion of wages in the union sector, the difference between the mean of ln wages in the union and nonunion sectors, and the extent of unionization as the appropriate counterfactual measures for 1996. This approach assumes that the changes in these values over the period are entirely attributable to changes in union power. However, some of the increase in the dispersion of wages in the union sector might reflect the increased dispersion of wage-determining characteristics, and not changes in the structural power of unions. Likewise, the decline in unionization might be the result of changing occupational and industry mix, and not the result of fundamental changes in the organizational capabilities of unions.

The second counterfactual measure derives these various counterfactual values using the appropriate structural equations from 1984. In doing so, it is possible to remove the influence of changes in wage-determining and union-status-determining characteristics, and to focus solely on changes in structural determinants. This approach assumes that the changes in structural determinants are entirely attributable to changes in union power. Specifically, the estimation results for the 1984 union ln wage equation are used, in combination with the 1996 wage-determining characteristics of union workers, to generate the 1996 counterfactual variance of ln wages and mean ln wage in the union sector. The procedure is analogous to that described above.

To obtain the counterfactual extent of unionization, the following derivation is performed:

$$\textbf{(11)} \quad \hat{\bar{\mathbf{U}}}_{96} = \hat{\boldsymbol{\alpha}}_{84} + \hat{\boldsymbol{\delta}}_{84}\bar{\mathbf{Y}}_{96},$$

where $\hat{\boldsymbol{\delta}}_{84}$ is the vector of estimated coefficients from the 1984 union status equation (equation 6 above) and $\overline{\mathbf{Y}}_{96}$ is a vector of the 1996 means of characteristics that determine union status.

Under both counterfactual scenarios, the various components are substituted into equations (1) – (3) and the resulting counterfactual measures of formal sector wage dispersion and overall wage dispersion are compared to their actual measures. If the declining power of unions accounts for some of the rise in wage inequality over the period, then restoring union power to its level at the outset of the period will yield a counterfactual measure of dispersion that is less than the actual measure. The difference between the counterfactual and actual measures of dispersion reflects that portion of the overall change in dispersion that is attributable to declining union power.

Data

The data for the analysis come from the 1984 and 1996 *Encuesta Nacional de Ingresos y Gastos de los Hogares* (INEGI 1986, 1998). These are national household surveys that began in 1984 and continued in 1989, 1992, and every two years thereafter (the 1998 survey has not yet been made available to researchers). Each survey is a stratified sample based on city size, with a similar sampling distribution across the survey years, and weights that render the sample representative of the national experience. I utilize information on working individuals from the surveyed households. The data contain good information on certain labor market characteristics of workers. Most importantly for my analysis, however, these are the only micro-surveys in Mexico that contain information on the union status of workers.

The samples utilized in this analysis are composed of wage earners who are sixteen years of age or older and who work at least twenty hours per week.[11] The earnings variable is the hourly wage, and is computed from reported earnings during the month prior to the survey and reported hours of work. To insure an accurate measure of the wage, I delete from the sample those who are self-employed or working without pay. Reported earnings for the self-employed are likely to include returns on owned capital, which would bias upward the measured wage. Because information is available on union status only for the primary job of a respondent, I also exclude from the analysis those who are engaged in secondary employment (i.e., who hold more than one job). The 1984 sample contains 3531 respondents and the 1996 sample contains 11,610 respondents. Table 1 gives the definitions for the full set of variables used in the analysis.

The nonunion sample is divided into two segments for purposes of analysis – those workers who work in jobs that are typically “beyond the pale of unionization” and those who work in jobs that are comparable to union jobs elsewhere in the economy. I refer to the first segment as the “informal” sector and to the second, in combination with the union sample, as the “formal” sector. These two sectors are distinguished based on industry and occupation categories. The informal sector is composed of workers in agriculture, forestry, and fishing, and those who engage in domestic service or who are sellers of goods or services without a fixed or stable establishment.

Although I use these familiar labels to describe the divisions in the labor market of my sample, it is not really my purpose to accurately capture the formal and informal sectors of the economy. I have divided the sample in this way merely to insure that the union-nonunion comparisons are legitimate. That is, when asking whether unions decrease wage dispersion for a comparable group of workers, by eliminating what I refer to here as the informal sector from the analysis, I hope to have achieved a truly meaningful nonunion comparison group. Holding the informal sector out for separate analysis biases against the finding that unions decrease wage inequality. If the informal sector workers were included as part of the comparison nonunion segment, the structural differences between the sectors – which drive the finding that unions decrease wage inequality – would be far more pronounced than those reported below.

Results

Table 2 gives the estimated coefficients from the union and nonunion ln wage equations. Table 3 gives the variance components that are derived from these equations. These two tables offer evidence on the structural differences in pay policies between the union and nonunion sectors, and thus on the likelihood that the lower union wage dispersion reported in Figures 1-3 is due to pay polices as opposed to a lower dispersion of wage-determining characteristics. The results suggest that union pay policies level the pay structure for productive characteristics and reduce or eliminate arbitrary pay differences across workers. [12]

Looking first at the 1984 results, the finding that is perhaps most immediately striking is that there is no significant gender discrimination in pay in the union sector. This compares with a significant and quite sizeable pay difference based on sex in the nonunion sector.[13] To the extent such pay differences are indeed arbitrary, and not a reflection of uncaptured worker productive characteristics, union pay policies that stress "equal pay for equal work" appear to completely eliminate arbitrary pay differences based on the gender of the worker.

The north-south pay differential is also lower – by roughly one-half – in the union sector than in the nonunion sector. The north-south pay differential may reflect differences in productive characteristics, but it may also be the result of differences in the relative power of workers in the north. In this case, the results suggest that collective bargaining is able to partially counteract the power imbalance between north and south, and to reduce the arbitrary pay differentials that result from this imbalance.

There is evidence in these findings of a leveling of the pay structure with regard to worker productive characteristics as well. The return to education, for instance, is flatter in the union sector than in the nonunion sector. For example, the return to having completed *superior* (college) relative to having completed *secundaria* (junior high school) is 2 in the nonunion sector, but only 1.4 in the union sector. While the numerical estimates of the returns to education are lower in the union sector, an interactive specification reveals that few of the estimated coefficients on the schooling variables are statistically significantly different across the two equations.

The evidence for wage structure leveling is statistically much stronger with regard to occupational differences in pay. (The inter-occupation and inter-industry differentials in pay are not reported in Table 2 in order to economize on space.) Occupation differentials reflect returns to skills that are acquired informally, through on-the-job training for example. The standard deviation of the set of estimated occupation coefficients is 0.25 in the union sector and 0.31 in the nonunion sector, and many of the estimated coefficients on the occupational categorical variables are statistically significantly different across the union and nonunion estimated wage equations.

There is also evidence of leveling with regard to the returns to technical training in these results. In fact, the results reveal that, in contrast to the nonunion sector, the union sector possesses virtually no significant return to the acquisition of formal technical schooling. In addition to union pay policies that level the wage structure for technical training, this finding could be the result of differences in on-the-job training in the two sectors. Suppose, for example, that union workers receive on-the-job training that is either comparable or superior to formal technical schooling and that unionized employers absorb the full cost of on-the-job training, but that this is not true of the nonunion sector. In this case, the return to a technical degree could be negligible in the union sector but positive and significant in the nonunion sector.

Among the various determinants of wages, the only two that fail to display a leveling effect are age and the inter-industry wage differentials. The age and age-squared variables proxy years of labor market experience and tenure with the firm. The return to age displays the characteristic increasing-at-a-decreasing-rate pattern commonly found for experience and tenure variables in estimated wage equations. The higher return to age in the union sector is plausibly explained by greater worker-firm attachment, more on-the-job training, and the greater use of seniority as a criterion for promotion in the union sector.

In the absence of highly centralized bargaining structures, the differential power of unions across industries is likely to result in different union relative wage effects, and thus higher wage dispersion across industries in the union sector. The results reveal that the standard deviation of the inter-industry wage differentials is 0.14 in the union sector and 0.11 in the nonunion sector, although the estimated coefficients are not statistically significantly different across the union and nonunion wage equations.

I turn now to the 1996 results presented in the last four columns of Table 2. The leveling of the wage structure and the reduction of arbitrary pay differences due to union pay policies are apparent in these results as well. For example, the lower return to technical schooling in the union sector persists in the 1996 results. The difference in the north-south pay differential between the union and nonunion sectors is virtually unchanged from the 1984 findings. And the extent of wage structure leveling with regard to inter-occupation wage differentials even increases a bit over the period. The standard deviation of the set of estimated inter-occupation wage differentials is 0.342 in the nonunion sector and .207 in the union sector in 1996, compared with 0.312 and 0.25 respectively in 1984.

However, there are also signs that the ability of unions to level wages and to reduce arbitrary pay differences may have diminished over the period. The most striking example of the latter is that the gender pay differential is now positive and statistically significant in the union sector, and much closer to the gender differential in the nonunion sector. Unions appear to have a more difficult time establishing policies of “equal pay for equal work” in collective bargaining in 1996.

Union power across industries has become more disperse over the period as well. The standard deviation of the inter-industry wage differentials increases from 0.141 to 0.189 in the union sector, but remains roughly unchanged at 0.114 in the nonunion sector.

Moreover, in contrast to the 1984 results, many of the inter-industry effects are statistically significantly different across the union and nonunion sectors in 1996.

Even some of the observed tendencies towards greater wage structure leveling in 1996 are probably reflective of lost union power. For example, the return to age, which proxies for returns to labor market experience and tenure with the firm, is lower in the union sector than in the nonunion sector in 1996 – the exact opposite of the 1984 results. Recent demands for greater labor flexibility in production have posed a challenge to collective bargaining language that stipulates seniority as a criterion for promotion.

One of the more striking features of the 1996 results compared to the 1984 results is the well-documented rise in the return to education. The Table 2 results reveal that the increased return to formal education has occurred in both the union and nonunion sectors. In both sectors, the returns to low levels of education (the completion of *secundaria* and below) are dramatically lower in 1996 than they were in 1984, while the returns to high levels of education (the completion of *superior* and above) are dramatically higher. In 1996, the return to completion of *superior* relative to the completion of *secundaria* was 1.8 in the union sector and 3.2 in the nonunion sector. This compares to relative returns of 1.4 and 2, respectively, in 1984. The return to education in the union sector is flatter than that in the nonunion sector, but once again the differences are not collectively statistically significant.

Table 3 repeats the union and nonunion variance of ln wage measures reported in Figure 1, but also gives their breakdown into explained and residual variance components derived from the estimated wage equations in Table 2.[14] The differences in explained variances across the union and nonunion sectors reflect both pay policies and wage-determining characteristics. Below, I parse out the separate effects of pay policies and use these in the counterfactual derivations. The differences in residual variances, however, reflect differences in wage dispersion among workers with the same wage-determining characteristics. Thus, these results reflect the differential effect of union pay policies only.

The significantly smaller residual variances in the union sector in both 1984 and 1996 are yet another piece of evidence suggesting that union policies act to reduce wage dispersion through either leveling or reductions in arbitrary pay differences. However, the changes in the residual variances over time also suggest that union policies may have lost some of their dispersion-reducing effect over the intervening period. The union/nonunion residual variance ratio rose over the period from roughly 0.5 to 0.65. The union residual variance rose by roughly 63%, while the nonunion residual variance rose by only 25%.

A question that must be addressed before proceeding to the derivation of counterfactuals is whether the differences in union-nonunion pay structures in Table 2 reflect actual differences in pay policies, or rather the differential selection of workers into the two sectors based on unobserved wage-determining characteristics. Table 4 reports the selection-corrected results of the union and nonunion estimated wage equations for 1984 and 1996. The union wage equations appear not to be affected by selection bias, as indicated by the lack of statistical significance of the Inverse Mill's Ratios (denoted in the table as "lambda") in these estimated equations.

The nonunion estimated coefficients, however, do appear to suffer from selection bias. Focussing just on the returns to education, and comparing the ordinary least squares results from Table 2 with the selection-corrected results of Table 4, the general nature of the bias appears to be upward, suggesting that there is disproportionate selection of workers with high unobserved skills into the nonunion sector. Moreover, contrary to expectations, the selection-corrected results indicate an even higher return to formal education in the nonunion sector than do the ordinary least squares results.[15] Whether this dispersion-increasing pattern of selection holds true once the larger set of estimated coefficients is considered is best discerned from the counterfactual exercises to which I now turn.

Do unions reduce formal sector wage dispersion?

Table 5 gives the actual and counterfactual variance ln wage measures for the formal sector for 1984 and 1996, as well as the various components that are used to derive these measures as indicated by equation (1). For any given counterfactual variance of ln wage measure, the components that are themselves counterfactuals depend on the particular exercise being conducted. For example, in row 2 the counterfactual exercise involves subjecting union workers to the pay policies of the nonunion sector (as indicated by U→N). Thus, the variance of ln wages and the mean of ln wages in the union sector are the two counterfactual components making up the counterfactual variance of ln wage measure listed in the last column of row 2. The counterfactual in row 3 subjects nonunion workers to the pay policies of the union sector, and thus has the variance of ln wages and the mean of ln wages in the nonunion sector as the two counterfactual components.

The row 2 results reveal that if union workers had been subject to the pay policies of the nonunion sector in 1984, the dispersion of wages in the formal sector would have risen from 0.443 to 0.474. This suggests that union pay policies decreased formal sector wage dispersion by roughly 6.5%. Recall that there are two distributive effects to consider: Unions increase dispersion through union relative wage effects. Thus, transferring union workers to the nonunion sector will reduce wage dispersion by reducing the difference in mean ln wages between sectors, as seen by comparing the difference between the fifth and sixth columns of rows 1 and 2. However, union pay policies may also reduce within-group dispersion by leveling the wage structure for worker productive characteristics and reducing or eliminating arbitrary pay differentials. And this is precisely what the column two results indicate; subjecting union workers to nonunion pay policies raises within-group dispersion from 0.255 to 0.437. The results suggest that the dispersion-decreasing effect of unions dominates the dispersion-increasing effect.

The two counterfactual components in row 2 contain interesting information in their own right. The counterfactual mean ln wage in the union sector is a measure of what union workers would earn if they were transferred to the nonunion sector. Thus, the difference between the counterfactual mean ln wage (5.21) and the actual mean ln wage (5.38) is a measure of the union relative effect, which in this case is roughly 18%.[16] The difference between the counterfactual mean ln wage (5.21) and the actual mean ln wage for nonunion workers (4.95) is a measure of the extent to which the wage-determining characteristics of union workers are different from those of nonunion workers.

The counterfactual variance of ln wages in the union sector is a measure of the dispersion of wages that union workers would experience were they to be employed in the nonunion sector. Thus, the difference between the counterfactual (0.437) and actual (0.255) measures reflects differences in pay policies between the union and nonunion sectors. The difference is the sum of the differences in explained and residual variances (see equations (7)-(9)). The difference in residual variances can be read directly from Table 3 (i.e., 0.249-0.124), and the actual explained variance of ln wages appears there as well (i.e., 0.131). Thus, the counterfactual explained variance of ln wages (0.188) may be derived by simple manipulation.[17] (Of course, this number was originally derived from the nonunion estimated regression coefficients and the wage-determining characteristics of union workers – see equation (8).)

The difference between the counterfactual (0.188) and actual (0.131) measures of the explained variance of ln wages derives entirely from the structural differences in estimated coefficients across the union and nonunion wage equations. The difference between the counterfactual measure of the variance of ln wages in the union sector (0.437) and the actual variance of ln wages in the nonunion sector (0.47) in row 2 derives, on the other hand, entirely from differences in wage-determining characteristics across the two sectors. Thus, these comparisons suggest that while the dispersion of wage-determining characteristics is indeed lower in the union sector, this cannot entirely account for the lower dispersion of wages in that sector. Union pay polices contribute to the lower within-group dispersion as well.

Having explained in some detail the results of row 2, we can now move more rapidly through the remaining rows of results. In row 3, I consider the counterfactual in which nonunion workers are subject to the pay policies of the union sector. The results suggest that the formal sector variance of ln wages would fall substantially under this scenario, offering yet further support for the claim that unions reduce wage dispersion.[18] The results presented in row 4 reveal that correcting the nonunion estimated wage equation for selection bias does not alter the fundamental conclusion that unions reduce formal sector wage dispersion.[19] Formal sector wage dispersion still rises significantly if union workers are transferred to the nonunion sector. Comparing these results to those of row 2, it is apparent that selection bias has only a minimal impact on the estimated impact of unions on formal sector wage dispersion.

A closer comparison of the row 2 and row 4 results sheds further light on the nature of the bias due to differential selection. The selection corrected counterfactual mean ln wage in the union sector is slightly larger than that of row 2, suggesting that selection bias leads to an exaggerated union relative wage effect and thus to an overestimate of the dispersion-increasing effect of unions in the formal sector. This result is consistent with the theoretical prediction that unionized employers select from the labor pool those workers with higher unobserved skills.

Solving for the counterfactual explained variance of ln wages in the union sector, I find that it is 0.178 in the selection-corrected results of row 4, as compared with 0.188 in the ordinary least squares results of row 2. This suggests that the wage dispersion associated with nonunion pay policies is biased upward as a result of differential selection, just as theory predicts. Thus, the dispersion-reducing effect of unions is biased upward. However, the overall results suggest that this bias is insufficient, especially when coupled with upward bias in the dispersion-increasing effect of unions, to alter our central finding: unions reduce formal sector wage dispersion.

Turning to the results of similar counterfactual exercises using the 1996 data, as shown in rows 6 through 8, the conclusion is the same – unions reduce formal sector wage dispersion. Subjecting union workers to the pay policies of the nonunion sector increases wage dispersion (row 6). Subjecting nonunion workers to the pay policies of the union sector decreases wage dispersion (row 7). Correcting for selection bias (row 8) lowers the estimated dispersion-reducing impact of unions on formal sector wages found in row 6, but not sufficiently to alter the fundamental conclusion.

While the analyses from both periods suggest that unions reduce wage dispersion, there are also clear signals in the data that unions have lost some of their power to do so in the later years. Compare, for example, the 6.5% rise in formal sector wage dispersion in 1984 under the counterfactual in which union members are transferred to the nonunion sector (rows 1 and 2) with the 3.2% rise under a similar scenario in 1996 (rows 5 and 6).

The smaller increase in 1996 is accounted for by two factors. First, the percent of the formal sector labor force that is unionized has fallen from 31% to 21%. This means that, for any given amount by which the union sector is able to reduce dispersion relative to the nonunion sector, the impact on formal sector wage dispersion is less because a smaller percentage of the workforce is "transferred" to nonunion jobs. Second, unions appear to have lost some of their ability to reduce within group dispersion. When union workers were subject to nonunion pay policies in 1984, within group dispersion rose from 0.255 to 0.437, or roughly 70%. A similar exercise in 1996 raises within group wage dispersion by only 40%, from 0.463 to 0.648.[20]

Has changing union power contributed to rising wage inequality?

In Table 6, I present a set of counterfactual variance calculations that allow for greater insight into the effect of these features of changing union power on the actual rise in formal sector wage inequality over the period 1984 to 1996. The counterfactual exercise involves granting the union movement in 1996 the power it possessed in 1984. I utilize two different counterfactuals in this exercise. The first posits that unions in 1996 possess the exact same extent of unionization in the labor force, the exact same within-group variance of ln wages, and the exact same gap in mean ln wages between the union and nonunion sectors as existed in 1984.

Of course, these three features of the unionized labor market may have changed over the period not as a result of any structural change in the power of unions per se, but rather as a result of changes in the union-status-determining and wage-determining characteristics of workers. Thus, in the second counterfactual exercise I generate the extent of unionization, the within group variance of ln wages, and the mean ln wage gap holding these characteristics constant. I do this by utilizing the structural determinants of the union status and union wage equations in 1984 in combination with the 1996 worker characteristics.

The components of the first counterfactual derivation are shown in row 3 of Table 6. If the unionized labor market in 1996 had possessed the exact same features as in 1984, the variance of ln wages in the formal sector would have been 0.55 instead of 0.629. The components of the second counterfactual derivation are shown in row 4. If the structural features of union power – such as government policies regarding union organizing and union bargaining policies regarding wage leveling and the reduction of arbitrary pay differences – had been the same in 1996 as they were in 1984, the formal sector variance of ln wages would have been 0.578. There is very little difference in the two counterfactual measures. In both counterfactuals, the most significant factor accounting for the decreased formal sector wage dispersion is the decreased wage dispersion in the union sector rather than the increased extent of unionization.

A comparison of the actual components for 1984 and the counterfactual components for Counterfactual II is illuminating because the structural power of unions is held constant across the two rows of components. The difference in the extent of unionization is minimal (0.308 versus 0.306), suggesting that if, for example, governmental policy regarding union organizing or the willingness of workers to join unions had not changed over the period, the extent of unionization also would have changed by very little. Thus, the bulk of the decline in unionization from 0.308 in 1984 to 0.205 in 1996 can be accounted for by the changing structural ability of unions to organize and maintain members, and not by changing demographic, occupational, or industrial determinants of unionization.

The difference between the actual and counterfactual within-group variance of ln wages (0.255 versus 0.275) is also rather small. This suggests that if the power of unions to level the wage structure and to reduce arbitrary pay differences had not changed over the period, the within-group variance of ln wages would have risen by very little. Thus, the bulk of the rise in the variance of union ln wages from 0.255 to 0.463 can be accounted for by changing union policies and not by the changing demographic, occupational, or industrial determinants of union wages.

Granting union workers in 1996 the structural power possessed by their union counterparts in 1984 also gives them the ability to extract greater wage rents from employers. Thus, the union-nonunion mean ln wage gap grows from 2.37-1.89 to 2.42-1.89. Since this is a dispersion-increasing feature of unions, it puts upward pressure on the counterfactual variance of formal sector ln wage measure in row 4 compared to the actual measure in row 2.

The overall results suggest that had the union movement in 1996 possessed the structural power it possessed in 1984 the variance of ln wages in the formal sector would have been 0.578 instead of 0.629 – roughly 8% lower. The difference between these two measures accounts for 27% of the total change in the variance of ln wages (0.629-0.443) over the period. Thus, almost one-third of the rise in formal sector wage inequality over the period 1984 to 1996 can be accounted for by the decreased power of unions. The remaining two-thirds is accounted for by other factors such as the changing occupational or industrial mix resulting from trade liberalization or the changing skill endowments of the work force resulting from the changing demand for worker skills.

Do unions reduce overall wage inequality?

The results presented in Table 7 bring wage earners from the informal sector into the analysis.[21] Using the counterfactual calculations from Tables 5 and 6, and equations (2) and (3), I generate actual and counterfactual measures of the total (formal and informal sectors combined) variance of ln wages similar to those of previous analyses. Here I am primarily interested in discovering whether accounting for the third effect of unions on wage dispersion – namely, the extent to which unions increase wage dispersion by increasing the gap between formal and informal sector mean ln wages – alters my fundamental findings. The results suggest that it does not.

The various counterfactual measures of the variance of ln wages follow the exact same pattern as those of earlier analyses. The first eight rows reveal that unions reduce total wage dispersion. Total wage inequality rises if union workers are transferred to the nonunion sector. Total wage inequality falls if nonunion workers are transferred to the union sector. In each case, the dispersion-reducing effect of unions found in the formal sector analyses is slightly dampened by the introduction of the third, dispersion-increasing effect of unions on overall wage dispersion. But the dampening effect is slight.

The last two rows reveal that total wage inequality would have been lower in 1996 if unions had possessed the same structural power to maintain and organize members and to influence wages as they possessed in 1984. Interestingly, because the informal sector witnessed a rather substantial decrease in wage dispersion over the period – from 0.567 to 0.45 – the loss of union power in the formal sector accounts for a greater percentage of the rise in the variance of overall ln wages (30%) than it does for the rise in the variance of ln wages in the formal sector alone (27%).

Conclusion

The results of this study suggest that unions are an equalizing force in the dispersion of wages in Mexico. Unions reduce formal sector wage dispersion, and, as a consequence, the overall dispersion of wages among wage earners in the formal and informal sectors combined. They do this by negotiating collective bargaining agreements that level wages with regard to the productive skills of workers and reduce arbitrary pay differences across workers.

Losses in the structural power of unions to maintain and organize workers and to influence wage dispersion through collective bargaining have diminished the dispersion-decreasing effect of unions over the period of the mid-1980s to mid-1990s. Wage dispersion in the economy has risen as a result. The empirical results presented in this paper suggest that roughly one-third of the increase in wage inequality over the period can be accounted for by the changing power of unions.

This study, like most others that have tackled the causal influences of rising wage inequality over the 1980s and 1990s in Mexico, has taken a “mono-causal” approach – that is, it looks at one and only one determinant of the increased inequality. Future research should be devoted to specifying what portion of the increased wage dispersion can be legitimately accounted for by each of the various causal factors that thus far have been shown to matter. Until this is done, we can only speculate as to the final results.

Proponents of the view that the increased wage dispersion in Mexico is attributable to the decline in capturable rents resulting from the elimination of trade protections (e.g., Ravenga 1997) might well claim that my analysis overstates the role of declining union power in the increased wage inequality. They might argue, for example, that if my analysis had included profit rates or concentration ratios on the right-hand-side of estimated wage equations, a larger role would have been granted to changes in wage-determining characteristics, and a correspondingly smaller role for structural changes in union power.

Of course, every “mono-causal” analysis is open to similar criticisms. Feenstra and Hanson (1997) claim that increased wage dispersion in Mexico is related to the increased demand for skilled workers by the large number of foreign firms that have entered the country since Mexico’s relaxation of restrictions on foreign investments in the early 1980s. Their evidence is based on cross-sectional findings regarding the higher returns to skills along the border region, where foreign assembly plants are congregated. I might argue, however, that the higher return to skills has less to do with the skill demands of employers than it does with the quality of industrial relations in this region. There is a well-noted absence of unions along the border, and those that do exist are more likely to be "ghost unions" or to possess collective bargaining agreements that prevent unions from leveling wages for productive skills.

Future research should undertake a comparative exploration of the various causal factors that have been linked to rising wage inequality in Mexico over the past two decades. The results of the present paper suggest that declining union power warrants consideration as one of those causal factors.

Endnotes

[1] Between 1985 and 1989, import licensing was eliminated in all but a few strategic sectors of the economy; the average tariff declined from roughly 25% to 12% and the maximum tariff declined from 100% to 20%; and reference prices were eliminated entirely (Ten Kate 1992). In addition, Mexico joined GATT in 1986, and signed a free trade agreement with the U.S. and Canada which took effect in January of 1994.

[2] The notion that domestic institutional changes might be a factor in the growing wage inequality in the U.S. was first raised after scholars noted that certain European countries – West Germany, for example – did not suffer increases in wage inequality even though they were subject to the same growing international competition and technological changes in production that the U.S. experienced. Similar suggestive evidence of the importance of domestic institutional factors exists for the case of trade liberalization and growing wage inequality in developing countries. Robbins (1995) found, for example, that trade liberalization led to increased relative wages of skilled workers in Chile, Columbia, Costa Rica, and the Philippines, but not in Argentina or Malaysia.

[3] The typically low explanatory power of estimated wage equations employing a simple human capital specification, and the existence of significant inter-industry wage differentials and discriminatory pay based on gender and race provide suggestive evidence that nonunion pay structures display such arbitrary differences.

[4] By "the power of unions to equalize wages" I am referring to the prevalence of unions across the labor force as well as to the ability of unions to reduce within-group dispersion through wage leveling and the reduction or elimination of arbitrary pay differences. While I shall continue to employ the term "power" in this context, it really refers to both the "structural ability" and the "desire" to engage in these activities. For example, while

unions may continue to possess the ability to engage in wage leveling over a given period, they may, for a host of reasons, have lost the desire to do so, thereby contributing to an increase in wage inequality.

[5] The data upon which this figure is based will be discussed in greater detail below.

[6] This same pattern exists in virtually all of the industry and occupation categories that I use in the empirical work to follow.

[7] The fact that union wage dispersion is lower in more detailed industry and occupation categories (as shown in Figures 2 and 3) is suggestive evidence against an explanation for differences in estimated coefficients based on differential selection.

[8] A drawback to this procedure for the correction of selection bias is that it utilizes a single threshold model for the determination of union status, which implies that selection biases should go in the same direction for all workers. This, of course, is inconsistent with the second pattern of selection discussed above. See Lemieux (1998) for a discussion of the comparative pros and cons of the various techniques for addressing selection bias.

[9] Note that if the union and nonunion estimated wage equations contained all the important productive characteristics of workers, and only those productive characteristics, then the difference in estimated coefficients would reflect the ways in which union pay policies level the pay structure for worker skills. The difference in residual variances would then reflect the ways in which union pay policies reduce or eliminate arbitrary pay differences. Because the difference between the actual and the counterfactual explained variance of ln wages captures only the structural differences in estimated coefficients, it would reflect the extent to which unions level the pay structure for worker skills. The difference between the actual and the counterfactual residual variance of ln wages would then reflect the extent to which union pay policies reduce or eliminate arbitrary pay differences. In practice, of course, capturing the entire set of productive worker characteristics is virtually impossible, and thus distinguishing between these two effects of union pay policies is extremely difficult.

[10] Lewis argued that because we do not know how the emergence of unions and higher wages in the union sector affect the nonunion sectors of the economy, our estimates of the union relative wage effect (and, by analogy, my estimates of the union relative variance effect) are a special kind of counterfactual measure. They are a measure of the wage "gap" between sectors – the difference between what union workers currently earn and what their wages would be if they were transferred to the nonunion sector *as it currently exists* – and not a measure of the wage "gain" between sectors – the difference between what union workers currently earn and what their wages would be if unions were completely absent from the industrial landscape.

[11] Few workers under the age of sixteen or who work less than twenty hours per week are unionized. Including this group of individuals in the nonunion sample raises the variance

of ln wages for this segment significantly, and leads to even stronger results regarding the dispersion-reducing activities of unions than those reported below.

[12] A Chow test reveals that the estimated coefficients from the union and nonunion wage equations are statistically significantly different (1 percent level) in both 1984 and 1996.

[13] Panagides and Patrinos (1994) reported similar findings in their empirical exploration of the union relative wage effect in Mexico.

[14] The differences are all statistically significant (5 percent level).

[15] The 1984 nonunion return to some *primaria* falls by roughly 60% when corrected for selection bias, whereas the return to *postgrado* falls by only 10%. The 1996 nonunion return to some *primaria* falls by roughly 75% when corrected for selection bias, but the return to *postgrado* falls by only 1%.

[16] Calculated as $e^{5.38-5.21} - 1$.

[17] The precise calculation is 0.437-0.255= (0.249-0.124) + (X-0.131). Solving for X yields 0.188.

[18] The larger difference between actual (0.443) and counterfactual (0.282) measures in this scenario as opposed to that of row 2 is due in large part to the higher proportion of nonunion workers in the economy. Under the first counterfactual scenario only 31% of the labor force is “transferred” to a new sector, whereas under the second counterfactual scenario 69% of the labor force is “transferred.” If the difference between the actual and counterfactual measures of wage dispersion is calculated on a *per unit change in unionization* basis, the difference under the first scenario is roughly one-half the difference under the second.

[19] Note that since the union estimated wage equation does not suffer from selection bias, there is no need to redo the counterfactual exercise of row 3.

[20] Lost union power also reduces the union relative wage effect (from 18% in 1984 to 13% in 1996, to be exact), and this reduces the dispersion increasing effect of unions. Thus, the fact that unions are less successful in decreasing wage dispersion in later years is reflective of the depth of their losses in the areas of union coverage and union pay policies.

[21] Most reports suggest that the informal sector of the Mexican economy has grown over the period of my analysis (Pastor and Wise 1997, p. 436). Note that this is not true of the particular segment of the informal sector that I have isolated – namely, full-time workers who are over the age of 16 and not self-employed.

TABLE 1
VARIABLE DEFINITIONS

Wage	= natural log of respondent's hourly wage
Age	= Age of respondent
Age^2	= Age squared of respondent
Sex	= 1 if respondent's sex is male; 0 if female
$Primaria_1$	= 1 if respondent attended primaria, but did not complete; 0 otherwise
$Primaria_2$	= 1 if respondent completed primaria and nothing more; 0 otherwise
$Secundaria_1$	= 1 if respondent attended secundaria, but did not complete; 0 otherwise
$Secundaria_2$	= 1 if respondent completed secundaria and nothing more; 0 otherwise
$Preparatoria_1$	= 1 if respondent attended preparatoria, but did not complete; 0 otherwise
$Preparatoria_2$	= 1 if respondent completed preparatoria and nothing more; 0 otherwise
$Superior_1$	= 1 if respondent attended superior, but did not complete; 0 otherwise
$Superior_2$	= 1 if respondent completed superior and nothing more; 0 otherwise
Postgrado	= 1 if respondent attended postrado; 0 otherwise
$Technical_1$	= 1 if respondent completed a technical or commercial training program which required no formal education; 0 otherwise
$Technical_2$	= 1 if respondent completed a technical or commercial training program which required completion of primaria; 0 otherwise
$Technical_3$	= 1 if respondent completed a technical or commercial training program which required completion of secundaria; 0 otherwise
$Technical_4$	= 1 if respondent completed a technical or commercial training program which required completion of preparatoria; 0 otherwise
South	= 1 if respondent lives in the south of Mexico; 0 otherwise
Union	= 1 if respondent is affiliated with a union; 0 otherwise
Occupation	= 13 categorical variables (0,1) indicating respondent's occupation
Industry	= 19 categorical variables (0,1) indicating respondent's industry

Figure 1. Variance of Ln Wage Comparisons, 1984 and 1996

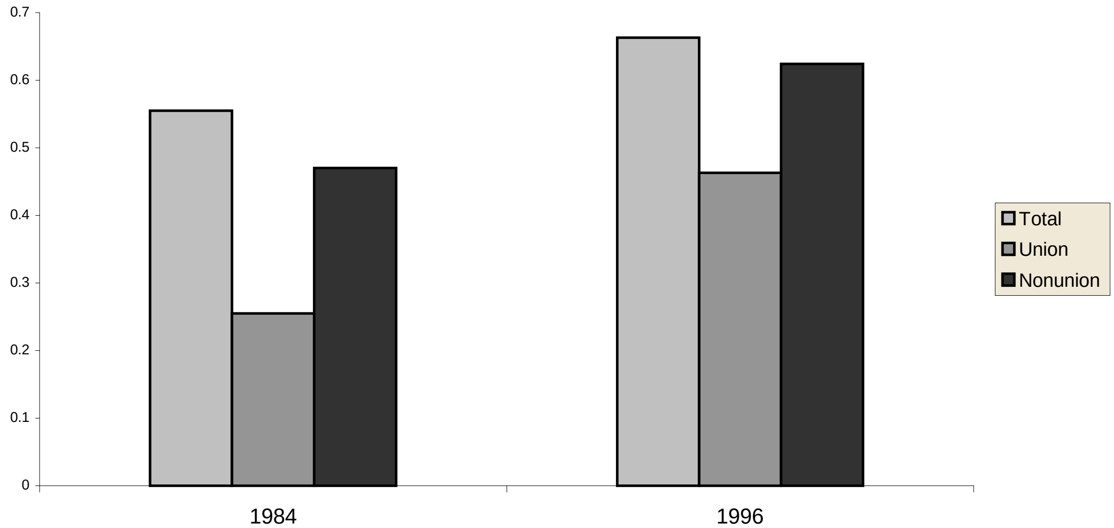

Figure 2. Variance of Ln Wages Across Industries

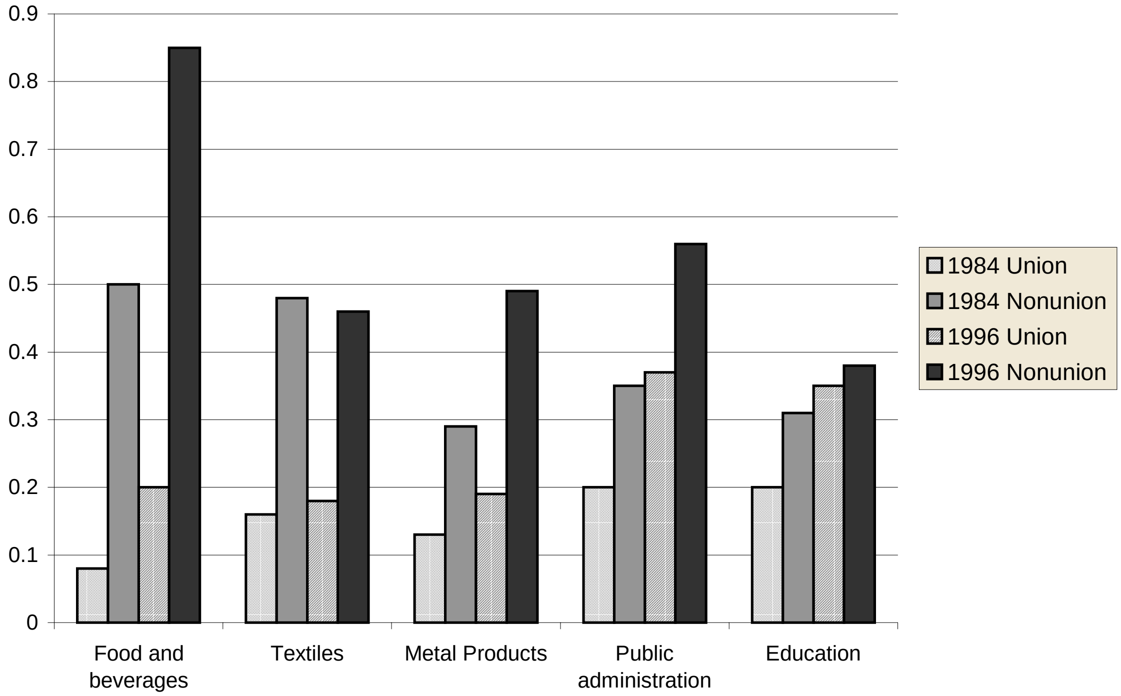

Figure 3. Variance of Ln Wages Across Occupations

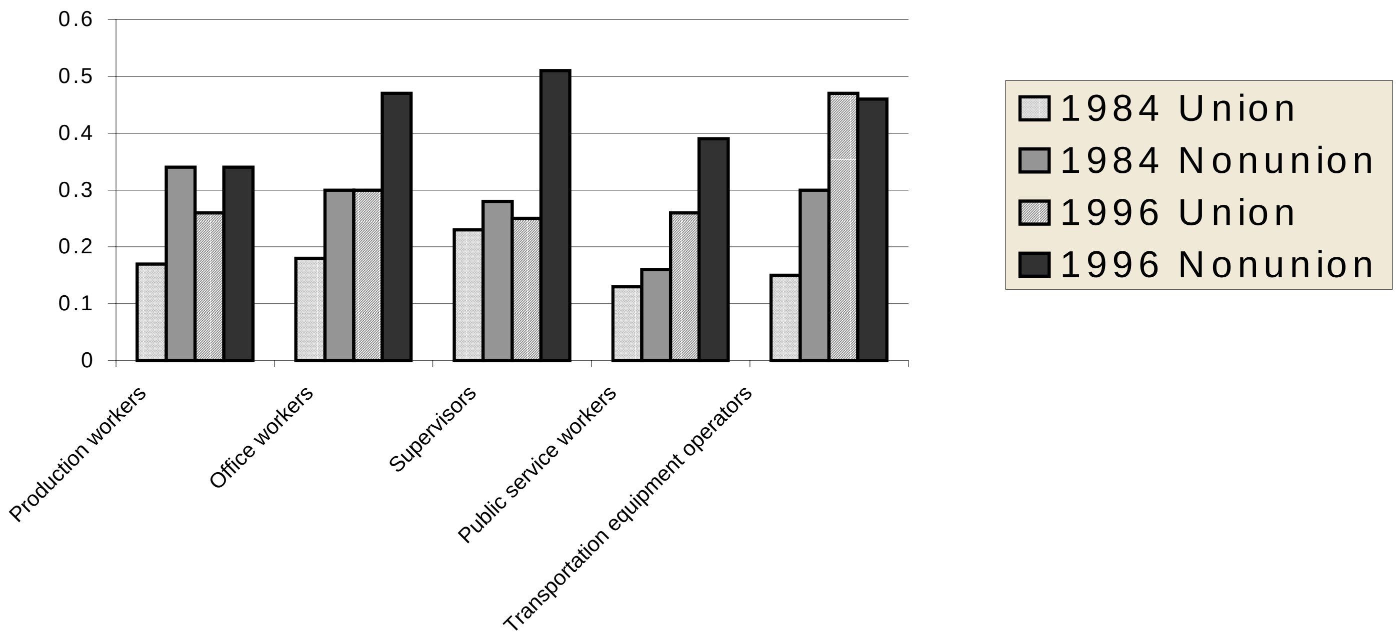

TABLE 2

DESCRIPTIVE STATISTICS AND ESTIMATED LOG WAGE EQUATIONS

Explanatory Variables	**1984** *Means and Std. Deviations*		*Estimated Coefficients*		**1996** *Means and Std. Deviations*		*Estimated Coefficients*	
	Union	*Nonunion*	*Union*	*Nonunion*	*Union*	*Nonunion*	*Union*	*Nonunion*
Age	34.69 (11.21)	31.21 (11.86)	0.068* (0.007)	0.051* (0.005)	35.66 (10.52)	31.273 (11.05)	0.036* (0.005)	0.056* (0.003)
Age^2	1329.2 (872.93)	1114.7 (920.97)	-0.0008* (0.00008)	-0.0005* (0.0001)	1382.2 (819.25)	1100.1 (839.44)	-0.0003* (0.0001)	-0.0006* (0.00004)
Sex	0.63 (0.483)	0.734 (0.442)	0.036 (0.032)	0.192* (0.03)	0.574 (0.495)	0.695 (0.46)	0.075* (0.023)	0.152* (0.016)
$Primaria_1$	0.148 (0.355)	0.208 (0.406)	0.096 (0.098)	0.215* (0.069)	0.069 (0.253)	0.117 (0.321)	0.045 (0.094)	0.049 (0.041)
$Primaria_2$	0.261 (0.439)	0.263 (0.44)	0.354* (0.097)	0.398* (0.071)	0.164 (0.37)	0.194 (0.395)	0.265* (0.091)	0.21* (0.04)
$Secundaria_1$	0.066 (0.248)	0.067 (0.251)	0.508* (0.105)	0.471* (0.076)	0.032 (0.177)	0.055 (0.227)	0.395* (0.097)	0.272* (0.047)
$Secundaria_2$	0.216 (0.412)	0.21 (0.407)	0.482* (0.102)	0.44* (0.076)	0.25 (0.433)	0.295 (0.456)	0.425* (0.092)	0.317* (0.04*
$Preparatoria_1$	0.036 (0.187)	0.035 (0.183)	0.515* (0.109)	0.485* (0.088)	0.052 (0.222)	0.056 (0.23)	0.532* (0.097)	0.459* (0.051)
$Preparatoria_2$	0.116 (0.321)	0.051 (0.22)	0.594* (0.107)	0.62* (0.083)	0.160 (0.367)	0.102 (0.302)	0.521* (0.094)	0.564* (0.045)
$Superior_1$	0.046 (0.209)	0.047 (0.212)	0.457* (0.116)	0.628* (0.088)	0.097 (0.296)	0.063 (0.243)	0.787* (0.098)	0.775* (0.05)

Table 2 Continued

	1984				1996			
	Means and Std. Deviations		*Estimated Coefficients*		*Means and Std. Deviations*		*Estimated Coefficients*	
	Union	*Nonunion*	*Union*	*Nonunion*	*Union*	*Nonunion*	*Union*	*Nonunion*
Superior$_2$	0.097 (0.296)	0.062 (0.24)	0.674* (0.118)	0.878* (0.091)	0.153 (0.36)	0.081 (0.273)	0.78* (0.097)	1.00* (0.054)
Postgrado	0.001 (0.03)	0.005 (0.067)	0.72 (0.699)	1.19* (0.151)	0.016 (0.124)	0.007 (0.084)	0.995* (0.132)	1.39* (0.098)
Technical$_1$	0.01 (0.1)	0.012 (0.11)	-0.175 (0.158)	-0.0002 (0.072)	0.007 (0.083)	0.01 (0.101)	0.146 (0.111)	0.04 (0.055)
Technical$_2$	0.069 (0.253)	0.055 (0.228)	-0.005 (0.047)	0.17* (0.048)	0.037 (0.19)	0.024 (0.153)	-0.054 (0.051)	0.026 (0.043)
Technical$_3$	0.128 (0.334)	0.075 (0.264)	0.056 (0.048)	0.183* (0.044)	0.171 (0.376)	0.128 (0.334)	0.042 (0.031)	0.109* (0.022)
Technical$_4$	0.004 (0.065)	0.003 (0.053)	-0.037 (0.294)	0.154 (0.177)	0.026 (0.16)	0.013 (0.114)	-0.014 (0.088)	0.138+ (0.0733)
South	0.175 (0.38)	0.186 (0.389)	-0.141* (0.033)	-0.293* (0.029)	0.088 (0.283)	0.071 (0.256)	-0.14* (0.036)	-0.331* (0.024)
Occupation	-	-	YES	YES	-	-	YES	YES
Industry	-	-	YES	YES	-	-	YES	YES
Constant	-	-	4.31* (0.238)	3.87* (0.192)	-	-	1.38* (0.16)	0.648* (0.122)
Adjusted R^2	-	-	0.486	0.458	-	-	0.554	0.496
Observations	-	-	934	2099	-	-	2041	7396

Note: Standard errors are in parenthesis. + indicates significant at 10% level; * indicates significant at 5% level. Results are corrected for heteroscedasticity using "White's heteroscedasticity-consistent estimator."

TABLE 3

IS WAGE DISPERSION LOWER IN THE UNION SECTOR?

ACTUAL VARIANCE COMPONENTS FROM TABLE 2 RESULTS

			$\sigma^2_{explained}$	$\sigma^2_{residual}$	σ^2
1984	(1)	Union Sector	.131	.124	.255
	(2)	Nonunion Sector	.221	.249	.470
1996	(3)	Union Sector	.261	.202	.463
	(4)	Nonunion Sector	.312	.312	.624

Note: $\sigma^2_{explained}$ = SSR/N; $\sigma^2_{residual}$ = SSE/N; and σ^2 = SST/N. All of the union-nonunion variance comparisons are statistically significantly different at the 5% level.

TABLE 4
ESTIMATED LOG WAGE EQUATIONS CORRECTED FOR SAMPLE SELECTION BIAS

	1984			1996		
Explanatory	*Union Selection*	*Selection-Corrected Coefficients*		*Union Selection*	*Selection-Corrected Coefficients*	
Variables	*Equation*	*Union*	*Nonunion*	*Equation*	*Union*	*Nonunion*
Age	0.084*	0.062*	0.036*	0.088*	0.06*	0.047*
	(0.013)	(0.019)	(0.009)	(0.008)	(0.021)	(0.005)
Age^2	-0.0008*	-0.0007*	-0.0004*	-0.0009*	-0.0005*	-0.0005*
	(0.0002)	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.00005)
Sex	-0.057	0.039	0.201*	0.006	0.078*	0.152*
	(0.072)	(0.033)	(0.03)	(0.04)	(0.023)	(0.016)
$Primaria_1$	0.543*	0.052	0.133+	0.261+	0.117	0.028*
	(0.167)	(0.159)	(0.082)	(0.139)	(0.113)	(0.041)
$Primaria_2$	0.693*	0.301+	0.289*	0.604*	0.426*	0.156*
	(0.168)	(0.18)	(0.094)	(0.135)	(0.171)	(0.044)
$Secundaria_1$	0.919*	0.438+	0.326*	0.682*	0.569*	0.21*
	(0.192)	(0.23)	(0109)	(0.152)	(0.188)	(0.052)
$Secundaria_2$	0.791*	0.421*	0.315*	0.704*	0.608*	0.253*
	(0.178)	(0.203)	(0.103)	(0.136)	(0.188)	(0.046)
$Preparatoria_1$	0.798*	0.454*	0.358*	0.774*	0.732*	0.39*
	(0.215)	(0.21)	(0.111)	(0.151)	(0.203)	(0.055)
$Preparatoria_2$	1.21*	0.506+	0.414*	0.831*	0.729*	0.484*
	(0.201)	(0.279)	(0.132)	(0.143)	(0.208)	(0.053)
$Superior_1$	0.74*	0.399*	0.521*	0.632*	0.958*	0.723*
	(0.214)	(0.202)	(0.104)	(0.151)	(0.182)	(0.053)
$Superior_2$	0.793*	0.614*	0.746*	0.608*	0.944*	0.947*

	(0.21)	(0.22)	(0.115)	(0.149)	(0.179)	(0.057)

Table 4 Continued

	1984			1996		
	Union Selection Equation	*Selection-Corrected Coefficients* Union	*Selection-Corrected Coefficients* Nonunion	*Union Selection Equation*	*Selection-Corrected Coefficients* Union	*Selection-Corrected Coefficients* Nonunion
Postgrado	0.797 (0.56)	0.652 (0.73)	1.08* (0.159)	0.334 (0.221)	1.1* (0.159)	1.368* (0.097)
Technical$_1$	-0.001 (0.279)	-0.176 (0.158)	-0.002 (0.071)	-0.264 (0.176)	0.082 (0.125)	0.066 (0.056)
Technical$_2$	-0.081 (0.123)	-0.0008 (0.047)	0.179* (0.049)	-0.013 (0.101)	-0.057 (0.051)	0.03 (0.044)
Technical$_3$	0.131 (0.108)	0.047 (0.055)	0.156* (0.046)	0.001 (0.053)	0.045 (0.031)	0.11* (0.022)
Technical$_4$	-0.581+ (0.334)	0.0004 (0.318)	0.268 (0.183)	-0.056 (0.135)	-0.021 (0.089)	0.145* (0.073)
South	0.019 (0.068)	-0.141* (0.033)	-0.295* (0.029)	-0.057 (0.071)	-0.155* (0.038)	-0.327* (0.024)
Lambda	-	-0.115 (0.337)	-0.455* (0.201)	-	0.369 (0.333)	-0.325* (0.123)
Occupation	YES	YES	YES	YES	YES	YES
Industry	YES	YES	YES	YES	YES	YES
Constant	-2.45* (0.444)	4.6* (0.906)	3.85* (0.191)	-2.745* (0.251)	0.353 (0.938)	0.665* (0.121)
Adjusted R^2	-	0.486	0.46	-	0.554	0.496
Observations	3034	934	2099	9438	2041	7396

Note: Standard errors are in parenthesis. + indicates significant at 10% level; * indicates significant at 5% level. Least-squares results are corrected for heteroscedasticity using "White's heteroscedasticity-consistent estimator."

TABLE 5

DO UNIONS REDUCE FORMAL SECTOR WAGE DISPERSION?
ACTUAL AND COUNTERFACTUAL VARIANCE COMPONENTS

			$\overline{U}$	σ^2_U	$(1-\overline{U})$	σ^2_N	$\overline{W}_U$	$\overline{W}_N$	σ^2_F
1984	(1)	Actual	.308	.255	.692	.47	5.38	4.95	.443
	(2)	Counterfactual $U \rightarrow N$	.308	.437	.692	.47	5.21	4.95	.474
	(3)	Counterfactual $N \rightarrow U$	.308	.255	.692	.277	5.38	5.14	.282
	(4)	Counterfactual $U \rightarrow N_{SC}$	.308	.426	.692	.47	5.23	4.95	.473
1996	(5)	Actual	.205	.463	.795	.624	2.37	1.89	.629
	(6)	Counterfactual $U \rightarrow N$	.205	.648	.795	.624	2.25	1.89	.650
	(7)	Counterfactual $N \rightarrow U$	.205	.463	.795	.40	2.37	2.03	.431
	(8)	Counterfactual $U \rightarrow N_{SC}$	.205	.527	.795	.624	2.36	1.89	.640

Note: $U \rightarrow N$ indicates that union workers experience the pay structure in the nonunion sector; $N \rightarrow U$ indicates that nonunion workers experience the pay structure in the union sector; $U \rightarrow N_{sc}$ indicates that union workers experience the selection-corrected pay structure in the nonunion sector. The counterfactual variance calculations are derived using equations (7)-(9). The counterfactual means are derived using equation (10).

TABLE 6

HAS THE CHANGING POWER OF UNIONS CONTRIBUTED TO RISING WAGE INEQUALITY?
ACTUAL AND COUNTERFACTUAL VARIANCE COMPONENTS

			$\overline{U}$	σ^2_U	$(1-\overline{U})$	σ^2_N	$\overline{W}_U$	$\overline{W}_N$	σ^2_F
1984	(1)	Actual	.308	.255	.692	.47	5.38	4.95	.443
1996	(2)	Actual	.205	.463	.795	.624	2.37	1.89	.629
	(3)	Counterfactual I	.308	.255	.692	.624	2.32	1.89	.55
	(4)	Counterfactual II	.306	.275	.694	.624	2.42	1.89	.578

Note: Counterfactual I assumes $\hat{\overline{U}}_{96} = \overline{U}_{84}$; $\hat{\sigma}^2_{96} = \hat{\sigma}^2_{84}$; and $(\hat{\overline{W}}_U - \overline{W}_U)_{96} = (\overline{W}_U - \overline{W}_U)_{84}$. Counterfactual II generates predicted values for each of these from 1984 estimated regression equations (4) and (11).

TABLE 7

DO UNIONS REDUCE OVERALL WAGE INEQUALITY?
ACTUAL AND COUNTERFACTUAL VARIANCE COMPONENTS

			$\overline{F}$	σ_F^2	$(1-\overline{F})$	σ_I^2	$\overline{W}_F$	$\overline{W}_I$	σ_T^2
1984	(1)	Actual	.842	.443	.158	.567	5.08	4.25	.555
	(2)	Counterfactual $U \rightarrow N$	.842	.474	.158	.567	5.03	4.25	.570
	(3)	Counterfactual $N \rightarrow U$	.842	.282	.158	.567	5.21	4.25	.450
	(4)	Counterfactual $U \rightarrow N_{SC}$	.842	.473	.158	.567	5.04	4.25	.571
1996	(5)	Actual	.847	.629	.153	.45	1.99	1.30	.663
	(6)	Counterfactual $U \rightarrow N$	.847	.650	.153	.45	1.96	1.30	.676
	(7)	Counterfactual $N \rightarrow U$	.847	.431	.153	.45	2.10	1.30	.517
	(8)	Counterfactual $U \rightarrow N_{SC}$	.847	.640	.153	.45	1.98	1.30	.671
	(9)	Counterfactual I	.847	.55	.153	.45	2.02	1.30	.602
	(10)	Counterfactual II	.847	.578	.153	.45	2.05	1.30	.631

Note: See notes from previous tables.

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