Unionism, Deunionization and Changing Wage Inequality

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1 Introduction

Several authors have addressed the question of what fraction of the the rise in US male wage inequality over the last three decades can be explained by the decline in unions over the same time period. Two ways of measuring the contribution of deunionization have been used in the literature. The first approach, employed by Freeman (1993) and Card (1998), begins by measuring the effect of unions on wage inequality in a given period t. Assumptions are made in order to identify the extent of wage inequality – as measured by some inequality measure $\mathfrak{I}$ – that would have prevailed in period t had unions been absent. Let $\mathfrak{I}(\bar{w}_t^n|\mathcal{I}_t)$ represent this magnitude. Here $\bar{w}_t^n$ stands for the period t wage structure in the absence of unions and $\mathcal{I}_t$ denotes the population of interest. Let $\bar{w}_t$ be the actually observed wage structure in period t. Then the effect of unions on wage inequality in period t is given by

$$\mathfrak{I}(\bar{w}_t|\mathcal{I}_t) - \mathfrak{I}(\bar{w}_t^n|\mathcal{I}_t). \tag{1}$$

The change of this magnitude over time

$$[\mathfrak{I}(\bar{w}_t|\mathcal{I}_t) - \mathfrak{I}(\bar{w}_t^n|\mathcal{I}_t)] - [\mathfrak{I}(\bar{w}_s|\mathcal{I}_s) - \mathfrak{I}(\bar{w}_s^n|\mathcal{I}_s)] \tag{2}$$

is then used to measure the contribution of deunionization to the change in wage inequality from period s to period t. By definition the term (2) is the change in the effect of unions on wage inequality as measured by $\mathfrak{I}$. Rewriting it as

$$[\mathfrak{I}(\bar{w}_t|\mathcal{I}_t) - \mathfrak{I}(\bar{w}_s|\mathcal{I}_s)] - [\mathfrak{I}(\bar{w}_t^n|\mathcal{I}_t) - \mathfrak{I}(\bar{w}_s^n|\mathcal{I}_s)] \tag{3}$$

suggests a second interpretation. Here the first term in square brackets is the actual change in wage inequality while the second term is the change in wage inequality if there had been no unions in both periods. From this perspective one might call it the effect of actual unionism on the change in wage inequality. This number tells a policymaker how he could have modified the change in wage inequality had he eliminated unions altogether in period s[1]. However, it is not clear how this measure is related to *deunionization*. If (2) is used to measure the contribution of deunionization to the change in wage inequality, then it should be equal to zero in the absence of deunionization. That is, no deunionization is implicitly defined as a situation in which

$$\mathfrak{I}(\bar{w}_t|\mathcal{I}_t) - \mathfrak{I}(\bar{w}_t^n|\mathcal{I}_t) = \mathfrak{I}(\bar{w}_s|\mathcal{I}_s) - \mathfrak{I}(\bar{w}_s^n|\mathcal{I}_s). \tag{4}$$

This is not a very attractive way of defining no deunionization as this definition depends on the inequality measure $\mathfrak{I}$. Is no deunionization a situation in which equation (4) holds for the variance of log wages, the 90-10 differential, the Gini coefficient or all members of a certain class of inequality measures? On a theoretical basis measure (2) does not appear to be a satisfactory measure of the contribution of deunionization to the change in wage inequality. For this reason and to distinguish measure (2) from the second approach discussed below, from now on I will refer to (2) as *the contribution of actual unionism to the change in wage inequality*.

The second approach, used in DiNardo, Fortin, and Lemieux (1996), starts by explicitly specifying a scenario of *no deunionization*. A constant aggregate unionization rate is an example for a no deunionization scenario, but many other scenarios could be considered. Assumptions are made so that the chosen no deunionization scenario is associated

[1]This of course assumes that all effects of eliminating unions are realized immediately. Although this is clearly unrealistic, I will not worry about dynamic effects in this paper. Assumptions that are much worse will follow.

with a unique hypothetical value of wage inequality $\mathcal{I}(\tilde{w}_t|\mathcal{I}_t)$. Then – in the absence of deunionization – the change in wage inequality would have been

$$\mathcal{I}(\tilde{w}_t|\mathcal{I}_t) - \mathcal{I}(\bar{w}_s|\mathcal{I}_t) \tag{5}$$

and the *contribution of deunionization to the change in wage inequality* is given by

$$\begin{aligned} &[\mathcal{I}(\bar{w}_t|\mathcal{I}_t) - \mathcal{I}(\bar{w}_s|\mathcal{I}_s)] - [\mathcal{I}(\tilde{w}_t|\mathcal{I}_t) - \mathcal{I}(\bar{w}_s|\mathcal{I}_t)] \\ =&[\mathcal{I}(\bar{w}_t|\mathcal{I}_t) - \mathcal{I}(\tilde{w}_t|\mathcal{I}_t)]. \end{aligned} \tag{6}$$

Measure (6) depends on the no deunionization scenario as well as the inequality measure $\mathcal{I}$.

One may expect that the two approaches will usually give the same answer in praxis, so that the distinction made above is unnecessarily fussy. The goal of this paper is to determine whether this is the case or if one has to be careful about which approach is adopted. Notice that

$$\begin{aligned} &[\mathcal{I}(\bar{w}_t|\mathcal{I}_t) - \mathcal{I}(\tilde{w}_t|\mathcal{I}_t)] \lesseqqgtr [\mathcal{I}(\bar{w}_t|\mathcal{I}_t) - \mathcal{I}(\bar{w}_s|\mathcal{I}_s)] - [\mathcal{I}(\bar{w}_t^n|\mathcal{I}_t) - \mathcal{I}(\bar{w}_s^n|\mathcal{I}_s)] \\ \iff &[\mathcal{I}(\bar{w}_s|\mathcal{I}_s) - \mathcal{I}(\bar{w}_s^n|\mathcal{I}_s)] \lesseqqgtr [\mathcal{I}(\tilde{w}_t|\mathcal{I}_t)] - \mathcal{I}(\bar{w}_t^n|\mathcal{I}_t)] \\ \iff &-\{[\mathcal{I}(\tilde{w}_t|\mathcal{I}_t)] - \mathcal{I}(\bar{w}_t^n|\mathcal{I}_t)] - [\mathcal{I}(\bar{w}_s|\mathcal{I}_s) - \mathcal{I}(\bar{w}_s^n|\mathcal{I}_s)]\} \lesseqqgtr 0. \end{aligned} \tag{7}$$

The expression on the left hand side of inequality (7) is positive if unions would reduce wage inequality more in period t than they did in period s if deunionization had not occured. I will refer to this magnitude as the change from period s to period t in *the effectiveness of unions to reduce wage inequality*. Similarly to the contribution of deunionization it depends on both the measure of inequality and the no deunionization scenario.

Now from the steps leading to inequality (7) it is seen that the contribution of deunionization to the change in wage inequality exceeds the contribution of actual unionism to the change in wage inequality if and only if the effectiveness of unions in reducing wage inequality has increased. Thus trying to determine whether the two approaches give different answers is the same as examining whether the effectiveness of unions to reduce wage inequality has changed.

The change in effectiveness may in general depend on the inequality measure, the no deunionization scenario, the identifying assumptions and of course the time period considered. There is a large number of possible combinations that could be analyzed in order to obtain a full picture. The number of combinations considered here is rather limited. First of all the analysis in this paper is restricted to the change from 1983 to 1993. The data will be discussed in section 2. Two sets of somewhat unsatisfying identifying assumptions will be used, called models 1 and 2 for short. In model 1 of section 3 it is assumed that union members are randomly selected and there are no general equilibrium effects of unionization. Model 2 of section 4 relaxes the first assumption by assuming random selection conditional on observables but maintains that unionization has no general equilibrium effects. For each model the change in effectiveness is analyzed for one no deunionization scenario and several inequality measures. The paper concludes with a discussion of the results.

2 Data

The data used in this paper are from the Current Population Survey (CPS) merged outgoing rotation group files for 1983 and 1993. Prior to 1983 information on the union status of workers was not collected in the outgoing rotation group supplements. Although the union status in 1979 could be obtained for a subset of workers, this was not done here as it would require some matching of observations across supplements. The 1993 CPS was the last survey before a substantially redesigned CPS was introduced in January 1994.

As a measure of the hourly wage I use hourly earnings on the main job for workers that are paid by the hour and usual weekly earnings on the main job divided by usual weekly hours on this job for all other respondents.

My preliminary data manipulations essentially follow DiNardo, Fortin, and Lemieux

(1996). All observations with allocated wages were eliminated. The GDP deflator for personal consumption expenditures was used to convert nominal wages into 1979 dollars. Only individuals of age 16 to 65 that reported an hourly wage from $1 to $100 (in 1979 dollars) were kept in the sample. It should be noted that I did not yet use a correction for the fact that usual weekly earnings were topcoded at $999 in 1983 while the top code was $1927 in 1993. The resulting sample sizes are 77529 observations for 1983 and 82186 observations for 1993, respectively.

As in DiNardo, Fortin, and Lemieux (1996) all inequality measures[2] are computed using as weights the product of the CPS sample weights[3] with usual hours of work.

Table 1 gives an overview over the development of wage inequality from 1983 to 1993. First note that the aggregate union density fell by about 6.5 percent. The mean log wage fell slightly, reflecting a falling mean log wage in the union sector and slight gains in the no union sector. The variance of log wages increased substantially. The absolut increase in the variance was somewhat smaller in the union sector than in the no union sector. The 90-10, 75-25 and 90-50 differentials all show a significant increase. There was no significant change in the 50-10 differential for all workers (although this differential expanded within both sectors).

3 Model 1

This section is devoted to the first set of assumptions – referred to as model 1 – that will allow me to evaluate the contributions of actual unionism and deunionization to the change in wage inequality. Subsection 3.1 begins with a reduced form model of the determination of wages and union status and introduces the identifying assumptions. Subsequently subsection 3.2 talks about estimation and the results are presented in

[2]While all inequality measures are computed using weights, I did not yet use weights in the probit estimation of section 4.

[3]My data set contained final weights as well as earnings weights. I decided to use the earnings weights.

subsection 3.3.

3.1 Identification

It is useful to begin the discussion of identification with a little reduced form model of the determination of wages and union status. This model will consists of the two functions u and w. Both functions have the same arguments, namely exogenous characteristics of the economy, which are split up into those related to unionism as an institution (such as laws regulating the process of union representation) and all other characteristics. The function u maps these characteristics into a union status structure and w yields a wage structure, both of which are random variables. For example it is

$$\bar{u}_t = u(\bar{U}_t, \bar{Z}_t) \qquad \text{and} \qquad \bar{w}_t = w(\bar{U}_t, \bar{Z}_t).$$

Here $\bar{u}_t$, $\bar{w}_t$, $\bar{U}_t$ and $\bar{Z}_t$ are the actual realizations of union status structure, wage structure, institutional characteristics of unionism and other characteristics in period t, respectively.

Next I will introduce sufficient assumptions on u and w that will allow me to evaluate expression (2). Instead of giving assumptions for each inequality measure separately I will aim at identifying the distribution $\text{distr}[\bar{w}_t^n|\mathcal{I}_t]$, so that any desired inequality measure of this distribution can be computed.

After that a no deunionization scenario is developed and the assumptions on u and w are strengthened so that a unique wage distribution $\text{distr}[\tilde{w}_t|\mathcal{I}_t]$ is associated with this no deunionization scenario and this distribution is also identified. This will allow me to evaluate the contribution of deunionization to the change in wage inequality as defined in (6) for any inequality measure $\mathfrak{I}$.

The first assumption aimed at identifying the distribution $\text{distr}[\bar{w}_t^n|\mathcal{I}_t]$ guarantees that the wage structure in the absence of unions is well defined. Let u_t^n be the union status structure in period t according to which no worker is unionized. Now consider a vector of other characteristics Z_t. Given Z_t there may be different vectors of institutional characteristics of unionism U_t that yield $u(U_t, Z_t) = u_t^n$. However, I assume that all such choices of U_t will lead to the same wage structure w_t^n, that is $w(U_t, Z_t) = w_t^n$. Of course w_t^n is allowed to vary with Z_t.

The second assumption is concerned with the selection of union members. Again take as given a vector of characteristics Z_t with associated no union wages w_t^n. By varying the institutional characteristics of unionism we can obtain a number of different union status structures. I assume that any union status structure u_t obtained in this way exhibits no selection with respect to the no union wage structure w_t^n, that is

$$\text{distr}[w_t^n | \mathcal{I}_t(u_t = u)] = \text{distr}[w_t^n | \mathcal{I}_t],$$

where the union status u is y for union members and n for non union members.

The third assumption will provide a link between no union wages and observed wages. Let Z_t be a vector of characteristics and u_t be a union-status structure that can be obtained given Z_t by choosing appropriate institutional characteristics of unionism U_t. I assume that for any choice of U_t the implied wage structure $w_t = w(U_t, Z_t)$ is such that individuals that are non union under u_t receive the same wage under both w_t and w_t^n, that is

$$w_t(i) = w_t^n(i)$$

for all $i \in \mathcal{I}_t(u_t = n)$. This assumption rules out spillover effects from the union to the non union sector and thereby limits the general equilibrium effects of unions. However, spillovers within the union sectors do not need to be disallowed at this point.

Now let $\bar{Z}_t$ and $\bar{U}_t$ be the actual characteristics of the economy in period t. Then $\bar{u}_t = u(\bar{U}_t, \bar{Z}_t)$ and $\bar{w}_t = w(\bar{U}_t, \bar{Z}_t)$ are the observed union-status structure and wage structure in period t, respectively. Moreover, by assumption there is a no union wage structure $\bar{w}_t^n$ associated with $\bar{Z}_t$. By the no selection assumption we have

$$\text{distr}[\bar{w}_t^n | \mathcal{I}_t(\bar{u}_t = n)] = \text{distr}[\bar{w}_t^n | \mathcal{I}_t].$$

The no spillovers assumption yields

$$\text{distr}[\bar{w}_t^n | \mathcal{I}_t(\bar{u}_t = n)] = \text{distr}[\bar{w}_t | \mathcal{I}_t(\bar{u}_t = n)].$$

Combining the last two results gives

$$\text{distr}[\bar{w}_t^n|\mathcal{I}_t] = \text{distr}[\bar{w}_t|\mathcal{I}_t(\bar{u}_t = n)], \tag{8}$$

and the right hand side of equation (8) is observable. Hence the three assumptions introduced up to now are sufficient to identify the contribution of actual unionism to the change in wage inequality as defined in (2) for all inequality measures $\mathfrak{I}$.

Next I turn to the contribution of deunionization to the change in wage inequality as defined in (6). To evaluate (6) a no deunionization scenario is needed. Formally a no deunionization scenario will be a collection of union-status structures. The simplest no deunionization scenario is the collection of union-status structures that yield the aggregate unionization rate prevailing in period s, that is

$$N_t^1 := \{\tilde{u}_t|\tilde{r}_t = \bar{r}_s\}. \tag{9}$$

Here $\tilde{u}_t$ denotes a hypothetical union-status structure in period t, $\tilde{r}_t$ is its associated unionization rate and $\bar{r}_s$ is the observed unionization rate in period s.

This scenario includes a wide variety of union-status structures. Moreover, given a union-status structure $\tilde{u}_t \in N_t^1$ there may be several configurations of institutional characteristics of unionism $\tilde{U}_t$ that implement $\tilde{u}_t$, holding fixed the other characteristics of the economy[4] at $\bar{Z}_t$. Collecting these profiles $\tilde{U}_t$ for all members of N_t^1 yields the set

$$P_t^1 := \{\tilde{U}_t|u(\tilde{U}_t, \bar{Z}_t) \in N_t^1\}.$$

The different members $\tilde{U}_t$ of P_t^1 could potentially induce very different wage distributions $\text{distr}[w(\tilde{U}_t, \bar{Z}_t)|\mathcal{I}_t]$. However, by strengthening the no selection and no spillovers assumptions made earlier in this section it can be achieved that they will all lead to the same wage distribution, so that the no deunionization scenario N_t^1 is associated with a unique wage distribution and expression (6) can be evaluated. First, I will now also rule

[4]The idea of fixing the other characteristics of the economy is that the problem is considered from the viewpoint of a policy maker who can only change the institutional characteristics of unionism but changing the other characteristics is not an option for him.

out spillovers within the union sector. Let Z_t be a vector of other characteristics of the economy. I assume that given Z_t there also exists a unique wage structure w_t^y such that union members always receive the wage they would receive under w_t^y. So if u_t is any union status structure and U_t is any vector of institutional characteristics of unionism such that $u(U_t, Z_t) = u_t$, then the wage structure $w_t = w(U_t, Z_t)$ satisfies

$$w_t(i) = \mathbf{1}(u_t(i) = y)w_t^y(i) + \mathbf{1}(u_t(i) = n)w_t^n(i)$$

for all $i \in \mathcal{I}_t$.

Now consider any $\tilde{U}_t \in P_t^1$ and let $\tilde{w}_t = w(\tilde{U}_t, \bar{Z}_t)$. Let $\bar{w}_t^y$ be the union wage structure associated with $\bar{Z}_t$. Then

$$\begin{aligned}\text{distr}[\tilde{w}_t|\mathcal{I}_t] &= \tilde{r}_t\text{distr}[\tilde{w}_t|\mathcal{I}_t(\tilde{u}_t = y)] + (1-\tilde{r}_t)\text{distr}[\tilde{w}_t|\mathcal{I}_t(\tilde{u}_t = n)] \\ &= \bar{r}_s\text{distr}[\bar{w}_t^y|\mathcal{I}_t(\tilde{u}_t = y)] + (1-\bar{r}_s)\text{distr}[\bar{w}_t^n|\mathcal{I}_t(\tilde{u}_t = n)].\end{aligned}$$

Now extending the no selection assumption with respect to no union wage structures w_t^n to union wage structures w_t^y allows to rewrite this as

$$\text{distr}[\tilde{w}_t|\mathcal{I}_t] = \bar{r}_s\text{distr}[\bar{w}_t^y|\mathcal{I}_t] + (1-\bar{r}_s)\text{distr}[\bar{w}_t^n|\mathcal{I}_t].$$

The right hand side of this equation no longer depends on which $\tilde{U}_t$ is selected from P_t^1.

The no spillovers and no selection assumptions also imply that

$$\begin{aligned}\text{distr}[\bar{w}_t|\mathcal{I}_t(\bar{u}_t = y)] &= \text{distr}[\bar{w}_t^y|\mathcal{I}_t(\bar{u}_t = y)] = \text{distr}[\bar{w}_t^y|\mathcal{I}_t] \quad \text{and} \\ \text{distr}[\bar{w}_t|\mathcal{I}_t(\bar{u}_t = n)] &= \text{distr}[\bar{w}_t^n|\mathcal{I}_t(\bar{u}_t = n)] = \text{distr}[\bar{w}_t^n|\mathcal{I}_t].\end{aligned}$$

Combining these results yields

$$\text{distr}[\tilde{w}_t|\mathcal{I}_t] = \bar{r}_s\text{distr}[\bar{w}_t|\mathcal{I}_t(\bar{u}_t = y)] + (1-\bar{r}_s)\text{distr}[\bar{w}_t|\mathcal{I}_t(\bar{u}_t = n)]. \tag{10}$$

The right hand side of equation (10) is observable, so that expression (6) can be evaluated for any inequality measure $\mathcal{J}$.

3.2 Estimation

All estimates of inequality measures reported in this paper are based on weighted empirical distribution functions[5]. Let $\mathcal{I}_t^*$ be a sample from the population $\mathcal{I}_t$. Then the weighted empirical distribution function corresponding to the actual period t wage distribution $\text{distr}[\bar{w}_t|\mathcal{I}_t]$ is given by

$$F(\omega|[\bar{w}_t|\mathcal{I}_t^*]) = \sum_{i \in \mathcal{I}_t^*} \theta_t(i)\mathbf{1}(\bar{w}_t(i) \leq \omega).$$

As mentioned in section 2 the $\theta_t(i)$'s are the CPS sampling weights times usual hours, normalized so that they sum to one. In analogy to equation (10) as an estimator of the counterfactual distribution $\text{distr}[\tilde{w}_t|\mathcal{I}_t]$ I use

$$\begin{aligned} F(\omega|[\tilde{w}_t|\mathcal{I}_t^*]) &= \bar{r}_s^* F(\omega|[\bar{w}_t|\mathcal{I}_t^*(\bar{u}_t = y)]) + (1-\bar{r}_s^*)F(\omega|[\bar{w}_t|\mathcal{I}_t^*(\bar{u}_t = n)]) \\ &= \frac{\bar{r}_s^*}{\sum_{i \in \mathcal{I}_t^*(\bar{u}_t = y)} \theta_t(i)} \sum_{i \in \mathcal{I}_t^*(\bar{u}_t = y)} \theta_t(i)\mathbf{1}(\bar{w}_t(i) \leq \omega) \\ &\quad + \frac{(1-\bar{r}_s^*)}{\sum_{i \in \mathcal{I}_t^*(\bar{u}_t = n)} \theta_t(i)} \sum_{i \in \mathcal{I}_t^*(\bar{u}_t = n)} \theta_t(i)\mathbf{1}(\bar{w}_t(i) \leq \omega). \end{aligned}$$

Noting that $\sum_{i \in \mathcal{I}_t^*(\bar{u}_t = y)} \theta_t(i) = \bar{r}_t^*$ (the estimator of for the actual unionization rate in period t) I obtain

$$\begin{aligned} F(\omega|[\tilde{w}_t|\mathcal{I}_t^*]) &= \sum_{i \in \mathcal{I}_t^*} \theta_t(i) \left\{ \mathbf{1}(\bar{u}_t(i) = y)\frac{\bar{r}_s^*}{\bar{r}_t^*} + \mathbf{1}(\bar{u}_t(i) = n)\frac{(1-\bar{r}_s^*)}{(1-\bar{r}_t^*)} \right\} \mathbf{1}(\bar{w}_t(i) \leq \omega) \\ &= \sum_{i \in \mathcal{I}_t^*} \tilde{\theta}_t(i)\mathbf{1}(\bar{w}_t(i) \leq \omega), \end{aligned} \tag{11}$$

where

$$\tilde{\theta}_t(i) = \theta_t(i) \left\{ \mathbf{1}(\bar{u}_t(i) = y)\frac{\bar{r}_s^*}{\bar{r}_t^*} + \mathbf{1}(\bar{u}_t(i) = n)\frac{(1-\bar{r}_s^*)}{(1-\bar{r}_t^*)} \right\}.$$

[5] DiNardo, Fortin, and Lemieux (1996) compute inequality measures from weighted kernel density estimates. I obtain relatively noisy estimates of differentials such as the 90-10 wage differential. Maybe this is partly due to the fact of using an empirical distribution function instead of kernel density estimates.

Hence the counterfactual weighted empirical distribution function is simply a reweighted version of the weighted empirical distribution function corresponding to the actual wage distribution. Similarly one obtains

$$F(\omega|[\bar{w}_t^n|\mathcal{I}_t^*]) = \sum_{i \in \mathcal{I}_t^*} \theta_t^n(i) \mathbf{1}(\bar{w}_t(i) \leq \omega) \tag{12}$$

where

$$\theta_t^n(i) = \theta_t(i) \mathbf{1}(\bar{u}_t(i) = n) \frac{1}{(1 - \bar{r}_t^*)}.$$

These formulas facilitate the computations as all inequality measures can be calculated in the usual way from a weighted empirical distribution function if the appropriate weights are used.

All standard errors and confidence intervalls in this paper are obtained by bootstrapping[6].

3.3 Results

Table 2 presents results for the variance of log wages. The first panel restates the changes in the mean log wage, the variance of log wages and the unionization rate for all workers already noted in table 1. The numbers from the second panel were also reported in table 1 as the changes within the nonunion sector. But the identifying assumptions of this section give them an additional interpretation. The change in the variance of log wages in the nonunion sector is now interpreted as the change that would have occured in the absence of unionism. In the absence of unions wage inequality would have been higher in 1983 but the subsequent increase would have been lower than the actual value.

[6] I did not try to verify whether the bootstrap is legitimate here. For some statistics it may have been easy to verify this or to come up with asymptotic distributions, for others it may be harder (e.g. for inequality measures computed from the counterfactual weighted empirical distribution functions with weights obtained by probit in the context of model 2). Typically no standard errors are provided in the papers of this literature, including those cited in the introduction. However, for the purpose of this paper it seemed crucial to have some idea of the variability of the estimates, and bootstrapping was the quickest way to get there.

The unionization rate in the absence of unions of course always equals zero. Finally the third panel gives the changes of the three variables in the absence of deunionization. By construction of the no deunionization scenario the aggregate unionization rate would have remained constant. The variance of log wages would have increased as well, but by a smaller amount than in the absence of unions.

In Table 3 the same numbers are used to compute the contributions of actual unionism and deunionization to the change in the variance of log wages. The point estimate for the contribution of actual unionism is 15 percent of the overall change in the variance of log wages, while with 23 percent the estimate of the contribution of deunionization is somewhat larger. From the discussion of equation (7) this suggests an increase in the effectiveness of unions to reduce wage inequality in the context of this specific combination of an inequality measure with a no deunionization scenario. Moreover, the bootstrap confidence interval indicates that this increase is significant at the 5 percent level.

Table 4 contains analogous computations for the 90-10 differential. Here the point estimate of the change in the differential in the absence of unionism actually exceeds the estimated actual change in the differential. This translates into a negative albeit insignificant contribution of actual unionism. However, as the contribution of deunionization is once again positive, the change in effectiveness is positive and significant.

As pointed out in section 2 there was little change in the 50-10 differential. From table 5 it is seen that unions had a substantial positive effect on this differential in 1983, an effect that had vanished by 1993. Consequently in the absence of unionism the change in the differential would have been larger than the observed change, implying a negative contribution of actual unionism to the change in wage inequality. The contribution of deunionization is also negative. But as it is of smaller magnitude than the contribution of unionism, again I find an increase in the effectiveness of unions to reduce wage inequality (although this is somewhat of a misnomer here as unionism tends to increase the 50-10 differential).

Table 6 for the 90-50 differential illustrates a case in which there appears to be no difference between the contributions of actual unionism and deunionization, but the estimates are quite noisy.

Quite dramatic differences between the two contributions mark the estimates for the 75-25 differential reported in table 7. While deunionization appears to account for the bulk of the increase in the differential, the contribution of actual unionism is rather small, but notice again the large standard errors.

All in all it seems that given the no deunionization scenario of this section there is reason to distinguish between the contributions of actual unionism and deunionization at least for some inequality measures. But the identifying assumptions that were used to obtain this result are unsatisfactory. While nothing will be done about possible general equilibrium effects of unions in this paper, at least the next section will take into account that union and nonunion workers differ with respect to their observable characteristics.

4 Model 2

The model of this section differs from the previous one only in its no selection assumptions. The unconditional no selection assumptions of the preceding section will now be replaced by analogous assumptions of no selection conditional on observable individual characteristics. Subsection 4.1 discusses identification and introduces a new no deunionization scenario. Subsection 4.2 is concerned with estimation and the results are presented in subsection 4.3.

4.1 Identification

Suppose that for each worker a profile of individual characteristics is observed. Let $\mathcal{X}$ be the collection of possible profiles. I assume that given a vector of exogenous characteristics Z_t with associated no union wage structure w_t^n and union wage structure w_t^y, any union status structure u_t obtained as $u_t = u(U_t, Z_t)$ for some institutional characteristics of unionism U_t satisfies

$$\text{distr}[w_t^n|\mathcal{I}_t(x, u_t = u)] = \text{distr}[w_t^n|\mathcal{I}_t(x)] \tag{13}$$

and

$$\text{distr}[w_t^y|\mathcal{I}_t(x, u_t = u)] = \text{distr}[w_t^y|\mathcal{I}_t(x)] \tag{14}$$

for $u \in \{y, n\}$ and all $x \in \mathcal{X}$.

Again only the no selection assumption with respect to no union wages is needed to identify the distribution $\text{distr}[\bar{w}_t^n|\mathcal{I}_t]$ and one obtains the relationship

$$\text{distr}[\bar{w}_t^n|\mathcal{I}_t] = \int \text{distr}[\bar{w}_t|\mathcal{I}_t(x, \bar{u}_t = n)]\text{distr}[x_t|\mathcal{I}_t].$$

Given these no selection assumptions the no deunionization scenario of the preceding section is no longer associated with a unique distribution of wages. In order to obtain a unique counterfactual wage distribution one would have to be more specific about what types of workers (in terms of their observable characteristics) are added to the union sector in order to keep the aggregate union density at its period s level. Certainly a number of different no deunionization scenarios could be developed along these lines. But a natural candidate for a no deunionization scenario is to keep the union density constant conditional on observable individual characteristics. This is the scenario I will focus on in this section. In analogy to the preceding section define

$$N_t^2 := \{\tilde{u}_t|\tilde{r}_t(x) = \bar{r}_s(x) \ \forall \ x \in \mathcal{X}\} \tag{15}$$

and

$$P_t^2 := \{\tilde{U}_t|u(\tilde{U}_t, \bar{Z}_t) \in N_t^2\}. \tag{16}$$

Then for all $\tilde{U}_t \in P_t^2$ the wage distribution induced by the wage structure $\tilde{w}_t = w(\tilde{U}_t, \bar{Z}_t)$ satisfies

$$\begin{aligned} &\text{distr}[\tilde{w}_t|\mathcal{I}_t] \\ &= \int [\bar{r}_s(x)\text{distr}[\bar{w}_t|\mathcal{I}_t(x, \bar{u}_t = y)] + (1 - \bar{r}_s(x))\text{distr}[\bar{w}_t|\mathcal{I}_t(x, \bar{u}_t = n)]] \, \text{distr}[x|\mathcal{I}_t]. \end{aligned} \tag{17}$$

The right hand side of equation (17) is again observable, so I can turn to the discussion of estimation in the next subsection.

4.2 Estimation

As in the preceding section all inequality measures are computed from weighted empirical distribution functions. The formulas (11) and (12) still apply when the weights are modified by substituting conditional for unconditional union densities:

$$\tilde{\theta}_t(i) = \theta_t(i) \left\{ \mathbf{1}(\bar{u}_t(i) = y) \frac{\bar{r}_s^*(x_t(i))}{\bar{r}_t^*(x_t(i))} + \mathbf{1}(\bar{u}_t(i) = n) \frac{(1 - \bar{r}_s^*(x_t(i)))}{(1 - \bar{r}_t^*(x_t(i)))} \right\},$$

$$\theta_t^n(i) = \theta_t(i) \mathbf{1}(\bar{u}_t(i) = n) \frac{1}{(1 - \bar{r}_t^*(x_t(i)))}.$$

I follow DiNardo, Fortin, and Lemieux (1996) in estimating the conditional union densities using probit[7]. The independent variables in the probit regression are a quartic in experience, years of education, experience and education interacted, 8 industry dummies, 3 occupation dummies and dummies for race, marital status and part-time status.

4.3 Results

Table 8 gives an overview over the changes of the mean log wage, the variance of log wages and the aggregate unionization rate in the absence of unionism and in the absence of deunionization as compared to the actual changes. Notice that in the absence of deunionization the aggregate unionization rate would have remained very close to its level in 1983. Thus the no deunionization scenario of this section is very similar to a specific no deunionization scenario that holds the aggregate unionization rate constant.

Comparing the tables 3 and 9 shows that controlling for differences in observable individual characteristics substantially reduces the magnitude of the effect of unionism

[7] At this point I did not use weighted probit, which would have been more consistent with the other estimation procedures of the paper. Another issue is that when the conditional union densities are estimated by probit, the weights no longer necessarily add up to one. Using weighted logit would insure that $\sum_{i \in \mathcal{I}_t} \theta_t(i)[\mathbf{1}(\bar{u}_t(i) = y) - \bar{r}_t(x_t(i))] = 0$, but still that does not guarantee that the weights add up to one. I decided to deal with this issue by always renormalizing weights.

on the variance of log wages in both periods. As the reduction is larger in 1983, the increase of the variance of log wages in the absence of unions turns out to be larger than in model 1, reducing the contribution of actual unionism to the change in the variance. But the contribution of deunionization drops by an even larger amount. While the point estimate of the increase in effectiveness is still positive, it is no longer significantly different from zero.

There is not much that can be said about the change in the effectiveness of unions to reduce wage differentials. Table 10 illustrates this for the case of the 90-10 differential. As with the variance of log wages, controlling for differences in observable individual characteristics reduces the magnitude of the effect of unions on inequality in both periods. But this time the reduction is larger in 1993, leading to an increase in the contribution of actual unionism as compared to model 1. At the same time the contribution of deunionization is unaffected by the change from model 1 to model 2. As a consequence the point estimate of the change in effectiveness drops from 33 percent of the overall change in the differential to only 10 percent, which is no longer significantly different from zero.

Controlling for observable characteristics has a rather dramatic impact on the results for the 75-25 differential, which are shown in table 11. Within model 1 the contribution of actual unionism was small and insignificant while the point estimate of the contribution of deunionization was estimated to be of the order of magnitude of the overall change in the differential. Within model 2 the point estimate for the contribution of deunionization is still very large. But the contribution of actual unionism is now estimated to be even larger. The source of this change is that moving to model 2 actually results in a larger estimate (in absolut value) of the effect of unionism on the 75-25 differential in 1983 while the effect of unionism in 1993 is estimated to be lower. All this results in a drop in the estimate of the change in effectiveness from 97 percent of the actual change in the differential to -19 percent, the latter point estimate being insignificant because of the high variability of the estimates.

The results change less dramatically for the 90-50 differential (not reported) and the point estimate of the change in effectiveness increases slightly but not by enough to become significant.

5 Discussion

The results of section 3 suggested that it may be important to distinguish the contribution of actual unionism and the contribution of deunionization to the change in wage inequality, at least for the no deunionization scenario and some of the inequality measures considered. However, taking into account the results of section 4, the main question of the paper must be left unanswered.

While I focussed on the estimation of the difference between the contribution of actual unionism and the contribution of deunionization and often obtained noisy results, the contributions themselves are frequently not estimated very precisely. So far these contributions are reported in the literature without providing standard errors. As the bootstrapping results of this paper suggest that these standard errors could be large, it may be desirable to obtain such standard errors in a more rigorous way.

The analysis of this paper could be extended by considering other time periods and no deunionization scenarios.

One could also try to better control of heterogeneity. Card (1998) provides estimates of the contribution of actual unionism to the change in the variance of log wages with and without taking into account differences in observable individual characteristics, but he also makes an attempt to control for unobserved heterogeneity by using longitudinal estimates of union wage effects. This strategy could be adapted to obtain estimates of the contribution of deunionization, again allowing to assess whether the two measures are different.

References

Card, David. 1998. “Falling Union Membership and Rising Wage Inequality: What’s the Connection?” Working paper 6520, NBER.

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Freeman, Richard B. 1993. “How much has De-Unionization Contributed to the Rise in Male Earnings Inequality?” In *Uneven Tides: Rising Inequality in America*, edited by Sheldon Danziger and Peter Gottschalk, 133–163. New York: Russell Sage Foundation.

Table 1: Changes in wage inequality from 1983 to 1993

		1983	1993	change, 0.95-conf.-int.
unionization rate	all	0.2536	0.1882	-0.06538 , [-0.0699,-0.0609]
mean log wage	all	1.818	1.808	-0.009236 , [-0.015,-0.003]
	union	2.007	1.988	-0.01944 , [-0.030,-0.009]
	nounion	1.753	1.767	0.01359 , [0.006,0.021]
variance log wage	all	0.2741	0.316	0.04194 , [0.038,0.046]
	union	0.1326	0.1666	0.03395 , [0.027,0.041]
	nounion	0.3058	0.3415	0.03569 , [0.031,0.040]
90-10 differential	all	1.386	1.49	0.1039 , [0.08,0.12]
	union	0.908	1.012	0.1036 ,[0.08,0.13]
	nounion	1.488	1.609	0.1214 , [0.10,0.14]
50-10 differential	all	0.7538	0.7545	0.000751 , [-0.022,0.023]
	union	0.5025	0.5895	0.08699 , [0.07,0.11]
	nounion	0.6931	0.7538	0.06062 , [0.037,0.085]
90-50 differential	all	0.6325	0.7357	0.1032 , [0.09,0.12]
	union	0.4055	0.4221	0.01661 , [0.001,0.032]
	nounion	0.7949	0.8557	0.06081 , [0.035,0.086]
75-25 differential	all	0.7802	0.8225	0.04234 , [0.029,0.056]
	union	0.4463	0.5169	0.07058 , [0.046,0.095]
	nounion	0.8329	0.8725	0.03958 , [0.025,0.054]

Confidence intervalls obtainded by naive bootstrapping, the number of repetitions is 200.

Table 2: The variance of log wages in model 1: part one

actual realizations

period	mean	variance	unionization rate
initial	1.837	0.2754	0.2455
	(0.001)	(0.001)	(0.001)
final	1.818	0.315	0.1895
	(0.002)	(0.001)	(0.001)
change	-0.01955	0.03958	-0.05601
	(0.002)	(0.002)	(0.002)

in the absence of unionism

period	mean	variance	unionization rate
initial	1.774	0.3063	0
	(0.002)	(0.001)	
final	1.776	0.3416	0
	(0.002)	(0.001)	
change	0.002098	0.03535	0
	(0.003)	(0.002)	

in the absence of deunionization

period	mean	variance	unionization rate
initial	1.837	0.2754	0.2455
	(0.001)	(0.001)	(0.001)
final	1.83	0.3065	0.2455
	(0.002)	(0.001)	(0.001)
change	-0.007372	0.03107	0
	(0.002)	(0.001)	

Number of bootstrap repetitions: 200

Table 3: The variance of log wages in model 1: part two

	estimate	standard error	0.95 confidence int.
variance in initial period			
actual	0.2754	(0.001)	
in absence of unionism	0.3063	(0.001)	
effect of unionism in initial period	-0.03084	(0.0005)	
variance in final period			
actual	0.315	(0.001)	
in absence of unionism	0.3416	(0.001)	
in absence of denunionization	0.3065	(0.001)	
effect of unionism in final period	-0.0266	(0.0005)	
change in variance			
actual	0.03958	(0.002)	
in absence of unionism	0.03535	(0.002)	
in absence of deunionization	0.03107	(0.001)	
contribution of unionism	0.004235	(0.0007)	
as percentage of actual change	0.107	(0.02)	[0.07,0.15]
contribution of deunionization	0.008512	(0.0003)	
as percentage of actual change	0.215	(0.009)	[0.20,0.23]
change in effectiveness	0.004277	(0.0008)	[0.0029,0.0057]
as percentage of actual change	0.1081	(0.02)	[0.07,0.14]

Number of bootstrap repetitions: 200

Table 4: The 90-10 differential in model 1

	estimate	standard error	0.95 confidence int.
differential in initial period			
actual	1.403	(0.0036)	
in absence of unionism	1.492	(0.0022)	
effect of unionism in initial period	-0.08832	(0.0026)	
differential in final period			
actual	1.497	(0.005)	
in absence of unionism	1.587	(0.0038)	
in absence of denunionization	1.468	(0.0051)	
effect of unionism in final period	-0.08961	(0.0059)	
change in differential			
actual	0.09372	(0.0062)	
in absence of unionism	0.09502	(0.0041)	
in absence of deunionization	0.06457	(0.006)	
contribution of unionism	-0.001293	(0.0066)	
as percentage of actual change	-0.0138	(0.065)	[-0.18,0.15]
contribution of deunionization	0.02916	(0.005)	
as percentage of actual change	0.3111	(0.047)	[0.21,0.42]
change in effectiveness	0.03045	(0.0054)	[0.020,0.041]
as percentage of actual change	0.3249	(0.069)	[0.19,0.46]

Number of bootstrap repetitions: 200

Table 5: The 50-10 differential in model 1

	estimate	standard error	0.95 confidence int.
differential in initial period			
actual	0.7713	(0.0045)	
in absence of unionism	0.7195	(0.0053)	
effect of unionism in initial period	0.0518	(0.0039)	
differential in final period			
actual	0.7621	(0.0041)	
in absence of unionism	0.731	(0.0047)	
in absence of denunionization	0.7716	(0.0046)	
effect of unionism in final period	0.03114	(0.0063)	
change in differential			
actual	-0.009125	(0.0064)	
in absence of unionism	0.01154	(0.0071)	
in absence of deunionization	0.0003426	(0.0068)	
contribution of unionism	-0.02066	(0.0071)	
as percentage of actual change	2.265	(NaN)	[-25.02,29.55]
contribution of deunionization	-0.009468	(0.0052)	
as percentage of actual change	1.038	(NaN)	[-8.85,10.93]
change in effectiveness	0.0112	(0.005)	[0.001,0.021]
as percentage of actual change	-1.227	(NaN)	[-22.26,19.81]

Number of bootstrap repetitions: 200

Table 6: The 90-50 differential in model 1

	estimate	standard error	0.95 confidence int.
differential in initial period			
actual	0.6321	(0.0051)	
in absence of unionism	0.7722	(0.0055)	
effect of unionism in initial period	-0.1401	(0.0048)	
differential in final period			
actual	0.7349	(0.0029)	
in absence of unionism	0.8557	(0.0055)	
in absence of denunionization	0.6963	(0.0055)	
effect of unionism in final period	-0.1207	(0.0058)	
change in differential			
actual	0.1028	(0.0061)	
in absence of unionism	0.08348	(0.0076)	
in absence of deunionization	0.06422	(0.0075)	
contribution of unionism	0.01937	(0.007)	
as percentage of actual change	0.1884	(0.066)	[0.05,0.33]
contribution of deunionization	0.03862	(0.0048)	
as percentage of actual change	0.3755	(0.051)	[0.28,0.48]
change in effectiveness	0.01925	(0.0065)	[0.007,0.032]
as percentage of actual change	0.1872	(0.068)	[0.06,0.32]

Number of bootstrap repetitions: 200

Table 7: The 75-25 differential in model 1

	estimate	standard error	0.95 confidence int.
differential in initial period			
actual	0.7625	(0.0065)	
in absence of unionism	0.8713	(0.005)	
effect of unionism in initial period	-0.1088	(0.0075)	
differential in final period			
actual	0.8054	(0.0072)	
in absence of unionism	0.8527	(0.0059)	
in absence of denunionization	0.7949	(0.0053)	
effect of unionism in final period	-0.0473	(0.0046)	
change in differential			
actual	0.04288	(0.0093)	
in absence of unionism	-0.01859	(0.0074)	
in absence of deunionization	0.03236	(0.0079)	
contribution of unionism	0.06147	(0.0086)	
as percentage of actual change	1.434	(0.22)	[1.01,1.86]
contribution of deunionization	0.01051	(0.0043)	
as percentage of actual change	0.2452	(0.09)	[0.08,0.41]
change in effectiveness	-0.05096	(0.0084)	[-0.066,-0.036]
as percentage of actual change	-1.188	(0.23)	[-1.63,-0.74]

Number of bootstrap repetitions: 200

Table 8: The variance of log wages in model 2 (probit weighting function): part one

actual realizations			
period	mean	variance	unionization rate
initial	1.837	0.2754	0.2455
	(0.001)	(0.001)	(0.001)
final	1.818	0.315	0.1895
	(0.002)	(0.001)	(0.001)
change	-0.01955	0.03958	-0.05601
	(0.002)	(0.002)	(0.002)

in the absence of unionism			
period	mean	variance	unionization rate
initial	1.787	0.2947	0
	(0.002)	(0.001)	
final	1.783	0.331	0
	(0.002)	(0.001)	
change	-0.004052	0.03634	0
	(0.003)	(0.002)	

in the absence of deunionization			
period	mean	variance	unionization rate
initial	1.837	0.2754	0.2455
	(0.001)	(0.001)	(0.001)
final	1.829	0.3092	0.2467
	(0.002)	(0.001)	(0.001)
change	-0.007886	0.03376	0.001133
	(0.002)	(0.001)	(0.001)

Number of bootstrap repetitions: 200

Table 9: The variance of log wages in model 2 (probit weighting function)

	estimate	standard error	0.95 confidence int.
variance in initial period			
actual	0.2754	(0.001)	
in absence of unionism	0.2947	(0.001)	
effect of unionism in initial period	-0.01925	(0.0005)	
variance in final period			
actual	0.315	(0.001)	
in absence of unionism	0.331	(0.001)	
in absence of denunionization	0.3092	(0.001)	
effect of unionism in final period	-0.016	(0.0005)	
change in variance			
actual	0.03958	(0.002)	
in absence of unionism	0.03634	(0.002)	
in absence of deunionization	0.03376	(0.001)	
contribution of unionism	0.003244	(0.0007)	
as percentage of actual change	0.08196	(0.02)	[0.05,0.12]
contribution of deunionization	0.005826	(0.0002)	
as percentage of actual change	0.1472	(0.007)	[0.132,0.162]
change in effectiveness	0.002582	(0.0008)	[0.0010,0.0042]
as percentage of actual change	0.06522	(0.02)	[0.03,0.10]

Number of bootstrap repetitions: 200

Table 10: The 90-10 differential in model 2 (probit weighting function)

	estimate	standard error	0.95 confidence int.
differential in initial year			
actual	1.403	(0.0035)	
in absence of unionism	1.451	(0.0092)	
effect of unionism in initial year	-0.0481	(0.01)	
differential in final year			
actual	1.497	(0.0044)	
in absence of unionism	1.571	(0.0055)	
in absence of denunionization	1.481	(0.0069)	
effect of unionism in final year	-0.07356	(0.006)	
change in differential			
actual	0.09372	(0.0057)	
in absence of unionism	0.1192	(0.011)	
in absence of deunionization	0.07742	(0.0079)	
contribution of unionism	-0.02546	(0.011)	
as percentage of actual change	-0.2717	(0.12)	[-0.53,-0.01]
contribution of deunionization	0.0163	(0.0052)	
as percentage of actual change	0.1739	(0.057)	[0.04,0.31]
change in effectiveness	0.04176	(0.013)	[0.011,0.073]
as percentage of actual change	0.4456	(0.14)	[0.08,0.81]

Number of bootstrap repetitions: 200

Table 11: The 75-25 differential in model 2 (probit weighting function)

	estimate	standard error	0.95 confidence int.
differential in initial year			
actual	0.7625	(0.0061)	
in absence of unionism	0.8431	(0.0073)	
effect of unionism in initial year	-0.0806	(0.0055)	
differential in final year			
actual	0.8054	(0.0072)	
in absence of unionism	0.8422	(0.0022)	
in absence of denunionization	0.8065	(0.0042)	
effect of unionism in final year	-0.03678	(0.0077)	
change in differential			
actual	0.04288	(0.0099)	
in absence of unionism	-0.0009397	(0.0075)	
in absence of deunionization	0.04395	(0.0075)	
contribution of unionism	0.04382	(0.0094)	
as percentage of actual change	1.022	(0.19)	[0.67,1.38]
contribution of deunionization	-0.001074	(0.0058)	
as percentage of actual change	-0.02504	(0.12)	[-0.33,0.28]
change in effectiveness	-0.04489	(0.0068)	[-0.061,-0.029]
as percentage of actual change	-1.047	(0.22)	[-1.47,-0.63]

Number of bootstrap repetitions: 200

Table 12: The 50-10 differential in model 2 (probit weighting function)

	estimate	standard error	0.95 confidence int.
differential in initial year			
actual	0.7713	(0.0043)	
in absence of unionism	0.7074	(0.0096)	
effect of unionism in initial year	0.06384	(0.0097)	
differential in final year			
actual	0.7621	(0.0038)	
in absence of unionism	0.7419	(0.0026)	
in absence of denunionization	0.7785	(0.0064)	
effect of unionism in final year	0.0202	(0.0047)	
change in differential			
actual	-0.009125	(0.0055)	
in absence of unionism	0.03452	(0.0099)	
in absence of deunionization	0.007213	(0.0077)	
contribution of unionism	-0.04364	(0.0099)	
as percentage of actual change	4.783	(NaN)	[-35.15,44.72]
contribution of deunionization	-0.01634	(0.0051)	
as percentage of actual change	1.79	(NaN)	[-9.71,13.29]
change in effectiveness	0.0273	(0.011)	[0.005,0.050]
as percentage of actual change	-2.992	(NaN)	[-30.56,24.58]

Number of bootstrap repetitions: 200

Table 13: The 90-50 differential in model 2 (probit weighting function)

	estimate	standard error	0.95 confidence int.
differential in initial year			
actual	0.6321	(0.0051)	
in absence of unionism	0.744	(0.0064)	
effect of unionism in initial year	-0.1119	(0.0057)	
differential in final year			
actual	0.7349	(0.0027)	
in absence of unionism	0.8287	(0.0057)	
in absence of denunionization	0.7023	(0.0047)	
effect of unionism in final year	-0.09376	(0.0044)	
change in differential			
actual	0.1028	(0.0057)	
in absence of unionism	0.08467	(0.0086)	
in absence of deunionization	0.07021	(0.0071)	
contribution of unionism	0.01818	(0.0072)	
as percentage of actual change	0.1767	(0.07)	[0.03,0.32]
contribution of deunionization	0.03264	(0.0045)	
as percentage of actual change	0.3173	(0.047)	[0.22,0.41]
change in effectiveness	0.01446	(0.0087)	[-0.006,0.035]
as percentage of actual change	0.1406	(0.087)	[-0.06,0.34]

Number of bootstrap repetitions: 200