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LABOR MARKET INSTITUTIONS AND THE DISTRIBUTION OF WAGES, 1973-1992 : A SEMIPARAMETRIC APPROACH

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RÉSUMÉ

Les études antérieures des changements dans l'inégalité des salaires dans le temps ont généralement utilisé des mesures agrégées de la distribution des salaires comme la variance du logarithme des salaires, les différences entre le 10^{ième} et le 90^{ième} centile ou le coefficient de Gini. Dans cet article, nous proposons l'utilisation d'une méthode semi-paramétrique qui nous permet d'estimer les changements dans la densité des salaires et de décomposer ces changements en divers facteurs. Nous utilisons les données du *Current Population Survey* des États-Unis de 1973 à 1992. Lorsque nous tentons d'expliquer la croissance de l'inégalité de 1979 à 1988, nous constatons, comme d'autres, que le déclin de la syndicalisation et les chocs d'offre et demande sont des facteurs explicatifs importants. Cependant, nous constatons aussi que le déclin de la valeur réelle du salaire minimum de 1979 à 1989 explique une partie importante de la croissance dans l'inégalité des salaires, surtout pour les femmes et les individus qui se trouvent dans le bas de la distribution des salaires. Nous concluons que les institutions du marché du travail sont au moins aussi importantes que les changements d'offre et de demande pour expliquer les changements dans la distribution des salaires aux États-Unis.

Mots clés : inégalité des salaires, estimation de densités par méthode de noyau, salaire minimum, syndicalisation.

ABSTRACT

Previous studies of changes in wage inequality over time have typically focused on summary measures of the distribution of wages such as the variance of log wages, the 90-10 wage differential, and the Gini coefficient. In this paper, we propose a semiparametric procedure that allows us to estimate changes in the entire density of wages and decompose these changes into various factors. Using data from the Current Population Survey, we find, like others, that de-unionization and supply and demand shocks were important factors in explaining the rise in wage inequality from 1979 to 1988. We also find, however, that the decline in the real value of the minimum wage from 1979 to 1989 explains a substantial proportion of the increase in wage inequality, particularly for women and for persons in the lower tail of the wage distribution. We conclude that labor market institutions are at least as important as supply and demand considerations in explaining changes in the U.S. distribution of wages from 1973 to 1992.

Key words : wage inequality, kernel density estimation, minimum wage, union.

1. INTRODUCTION

Several recent studies of changes in wage inequality in the United States have focused on summary measures of the distribution of wages such as the variance of log wages, the Gini coefficient, various percentiles of the wage distribution, or class shares ("lower class", "middle class" or "the rich").[2] A related set of studies by Bound and Johnson (1992), Blackburn, Bloom and Freeman (1992), Katz and Murphy (1992), and Murphy and Welch (1992) have tried to explain changes in the structure of wages by changes in the supply and demand for workers. Remarkably, none of these studies have attempted to analyze changes over time in the entire *distribution* of wages.[3]

To fully appreciate the importance of considering the entire distribution of wages, we display hours-weighted kernel estimates of the density of hourly wages for men and women over the period 1973–1992 in Figure 1. These densities are estimated using data from the May dual job holding supplement of the Current Population Survey (CPS) from 1973 to 1978, and from the outgoing rotation group files of the CPS from 1979 to 1992. A vertical line indicating the value of the federal minimum wage is also shown in each of the figures. The value of the federal minimum wage in real 1979 dollars is also reported in Table 1. It is useful to note that the federal minimum rose and fell in erratic steps from 1973 to 1978 and coverage expanded in 1976; minimum wages fell in real terms every

[2]See for example Blackburn and Bloom (1986), Blackburn and Bloom (1987), Danziger, Gottschalk and Smolensky (1989), Dooley and Gottschalk (1984), Dooley and Gottschalk (1985), Harrison and Bluestone (1988), Juhn, Murphy and Pierce (1993), Karoly (1992), Karoly (1993), Levy (1987), Levy and Murnane (1992), and Slottje (1989). Karoly (1993) has a particularly valuable discussion of the variety of approaches that have been taken.

[3]Pudney (1993) has used nonparametric techniques to study the age/income and age/wealth distributions using 1987 Chinese data.

year from 1979 to 1989.[4]

Far from the "constancy" described by Blinder (1980) in a survey of the U.S. economy in 1980, the distribution of wages has undergone some fairly dramatic changes over the last two decades. In particular, there is clear evidence that the minimum wage compresses the lower tail of the densities of wages. This large "visual" impact of the minimum wage on the distribution of wages, in fact, prompts us to more formally investigate its effect.

In this study we propose and implement a semi-parametric procedure to analyze changes over time in the density of wages. The aforementioned studies have considered the effect of changes in supply and demand conditions but paid little attention to the role of labor market institutions. Several other studies have looked at the effect of minimum wages on the distribution of income in isolation[5], while a few studies have focused on the role of unionization.[6] In contrast, our method allows us to decompose changes in the entire density of hourly wages both into factors related to labor market institutions and into factors related to supply, demand, and changes in the composition of the workforce. Moreover, the methodology we propose in this paper is potentially useful for a wide variety of applications.

In contrast with previous research that has concluded that "the minimum wage ... [has] had little impact on the overall wage structure" (Freeman and Katz 1994), we find that changes in minimum wages in the United States explain from nineteen

[4]As we explain in detail below, there was (in some years) substantial variation across states in the value of the minimum wage. We exploit this variation specifically in Section 5.3.

[5]See for example Krumm (1981), Gramlich (1976), Parsons (1980), Welch (1978), and Wessels (1980).

[6]See Card (1992b), Freeman (1993), or Lemieux (1993b) for some recent evidence.

to twenty-five percent of the change in the standard deviation of men's wages over the period 1979-1988 and from thirty to thirty-six percent of the change in the standard deviation of women's wages over the same period. Changing minimum wages, union density, and composition of the workforce explain most of the change in the standard deviation of wages for men and women over this period. After also accounting for changes in the supply and demand for labor, we explain up to 88 percent of the change in the standard deviation of wages for men and up to 80 percent of the change for women. For men and women respectively, 12 and 20 percent of the change in the standard deviation of log wages remains unexplained by the factors we consider. This represents an upper bound to the possible importance of factors such as skill–biased technological change which have been the focus of some recent research.

The rest of the paper is organized as follows. In Section 2 we describe the data, the weighted kernel density estimator, and the changes in the density of wages we seek to explain. In Section 3, we propose a semiparametric approach to decompose changes in the density of wages into various components of interest. In Section 4 we present our primary results focusing on changes in the density of wages in the period 1979–1988. In Section 5 we consider additional evidence from different time periods and from regional differences of the impact of unions and minimum wages. The final section concludes.

2. Data and Issues in Density Estimation

2.1. The Data

This paper uses data from the Current Population Survey (CPS) to analyze changes in the distribution of wages in the United States from 1973 to 1992. Starting in May 1973, the CPS surveys have regularly collected hourly or weekly earnings for each respondent's main job. From 1973 to 1978, these wage questions were asked only in May, as part of the dual job holding supplement. Beginning in 1979, the wage questions were asked in each month of those people in the outgoing rotation groups (which represent one-quarter of all individuals). Relatively large samples of workers are thus available to estimate changes in the distribution of hourly wages over the last two decades. The sample sizes are approximately 40,000 workers per year from 1973 to 1978, and 150,000 workers per year from 1979 to 1992. The exact sample sizes are reported in Table 1.

In addition to their large size, one advantage of these samples is that the wage measure collected is a good measure of a "point-of-time" price of labor. By contrast, most previous studies on changes in the structure of wages in the United States have relied on the average weekly earnings measure on all jobs available in the March CPS.[7] One problem with this alternative earnings concept is that it also depends on labor supply decisions and on the choice of holding more than one job. It combines the earnings from several jobs in the case of workers who held more than one job during the previous year. The hourly wage measure used in this paper is thus more closely connected to theories of wage determination based on supply and demand that focus on the hourly price of labor. Similarly, the connection between the minimum wage and the hourly wage on the main job is more direct than the connection between the minimum wage and average weekly earnings on all jobs.

Another key advantage of the 1973–78 dual job supplements and of the 1983–92 outgoing rotation group supplements is that they contain information on the

[7]For examples of studies using the March CPS, see Katz and Murphy (1992) and Murphy and Welch (1992). These studies restrict most of their analysis to full-time/full-year workers to minimize the confounding effect of labor supply decisions. One problem with the distribution of wages of these workers is that it is not necessarily representative of the distribution for the whole workforce.

union status of workers. This variable is essential to any attempt to evaluate the effect of labor market institutions on the distribution of wages. One shortcoming of these data, however, is that information on the union status of workers was not collected in the outgoing rotation group supplements of the CPS from 1979 to 1982. It is still possible, however, to obtain the union status of workers for the subsample of workers who were in an outgoing rotation group in May 1979, 1980, or 1981 since they were asked about their union status in the dual job holding supplement. A larger sample can be obtained in 1979 by matching answers about union status from the 1979 Pension Supplement of the May 1979 CPS to the corresponding wage data collected in the outgoing rotation group supplements of May, June, July or August 1979. About 35,000 observations are available in this matched sample. We thus use this matched 1979 sample whenever the estimation requires using the union status for 1979.

Several other preliminary data manipulations were performed to insure enough year to year continuity in our CPS samples. One issue is that usual weekly earnings are topcoded at $999 per week from 1973 to 1985, and at $1923 per week thereafter. A relatively low value of the topcode in real terms is likely to understate wage dispersion in the upper tail of the distribution. To avoid this type of bias, we use the upper tail of the 1986 distribution of wages to impute a wage distribution to the observations censored at the topcode in other years. This imputation procedure is only used for years in which more than 0.5 percent of the workforce is topcoded (1981 to 1985 and 1990 to 1992). Note that we impute a whole distribution as opposed to a simple average wage conditional on being topcoded. Imputing a whole distribution is necessary when estimating the entire distribution of wages.[8] Our imputation procedure is similar to the procedure

[8]Imputing a fixed wage as opposed to a distribution is appropriate when the only parameters of interest of the distribution are conditional means. This is the case in Bound and Johnson (1992)

we use to estimate the effect of the minimum wage on the distribution of wages (section 3.3).

As is well known, a significant fraction of interviewees in the CPS fail to answer questions about wages (Lillard, Smith and Welch 1986). While the Census Bureau used a "hot deck" procedure to replace missing wages by an allocated value in the 1979–92 outgoing rotation group files, it simply coded the wage as missing in the 1973–78 dual job holding supplements. All observations with allocated wages were thus eliminated from the 1979–1988 outgoing rotation group files to keep these samples comparable to the 1973 to 1978 samples.[9] In addition, only individuals between the age of 16 and 65 and reporting an hourly wage between \$1 and \$100 (in 1979 dollars) were kept in the sample. The GDP deflator for personal consumption expenditures was used to convert wages into 1979 dollars.

Note finally that the CPS sample weights are used to compute all estimates reported in this paper. For example, we consistently estimate the distribution of hourly wages of workers using a weighted kernel estimator of the density of wages. We use hours–weighted estimates of the distribution of hourly wages obtained by using the product of sample weights with usual hours of work as weights. These "hours-weighted" estimates put more weight on the wages of workers who supply many hours to the labor market. This gives a better representation of the dispersion of wages for each and every hour worked in the labor market, irrespective of who is supplying this hour.

Summary statistics of the final CPS samples are reported in Table 1. While

and Murphy and Welch (1992).

[9]Note that, because of a coding error in the CPS, it is impossible to identify most workers with allocated wages from 1989 to 1992 (see Devine (1993)). This error is of little consequence for most of our analysis that focuses on the 1979-88 period.

real wages remained relatively constant over the 1973–1992 period, the workforce became increasingly more female, educated, and nonwhite. Potential labor market experience (age–education–five) follows a U–shaped curve as the baby boom generation first enters the labor market during the 1970s and then ages during the 1980s. Table 1 also indicates that while the minimum wage and the unionization rate were relatively stable or even increasing during the 1970s, they both fell precipitously during the 1980s. Table 2 shows, however, that the relative constancy of the unionization rate during the 1970s masks a decline in private sector unionization that was more than offset by a steep increase in public sector unionization. Public sector unionization then remained constant while private sector unionization declined sharply during the 1980s, thus explaining the pronounced decline in the aggregate unionization rate over this period.

2.2. Weighted Kernel Density Estimation

The density estimates reported in this paper are obtained by adapting the kernel density estimator introduced by Rosenblatt (1956) and Parzen (1962) to the case in which sample weights are attached to each observation. The kernel density estimate $\hat{f}_h$ of a univariate density f based on a random sample $W_1, \ldots, W_n$ of size n, with weights $\theta_1, \ldots, \theta_n$ ($\Sigma_i \theta_i = 1$), is

$$\hat{f}_h(w) = \sum_{i=1}^{n} \frac{\theta_i}{h} K\left(\frac{w - W_i}{h}\right), \tag{1}$$

where h is the bandwidth and $K(\cdot)$ is the kernel function. The critical issue in kernel density estimation is the choice of bandwidth. The development and comparison of optimal bandwidth selectors is a topic of continuing research (Turlach 1993). Park and Turlach (1992) conducted simulation experiments to evaluate the performance of various bandwidth selectors in terms of minimizing the Expected Integrated Square Error ($EISE = E\int(\hat{f}_h - f)^2$) and the Expected Integrated

Absolute Error ($EIAE = E\int|\hat{f}_h - f|$). The plug-in method of Sheater and Jones (1991) and the bandwidth factorized smoothed cross-validation of Jones, Marron and Park (1991) were shown to be the best selectors for densities with complex structures exhibiting more than one mode. In this paper, we use Sheater and Jones' plug-in method since it does not exhibit the discretization problems associated with cross-validation methods (Silverman 1986). Unless otherwise indicated, all kernel density estimates presented here use the optimal bandwidth calculated using Sheater and Jones' selector. The densities of real log wages in Figures 1a and 1b were estimated using the kernel density estimator shown in equation (1). The kernel function used is Gaussian while the weights θ_i are the CPS sample weights multiplied by usual hours of work and normalized to add up to one. The estimates reported in Figure 1 are thus "hours-weighted" density estimates of the distribution of log wages. The optimal bandwidths for these densities range from 0.05 to 0.08, depending on the range of the support and the sparsity of the observations. While the estimated densities become less smooth when the bandwidth becomes small relative to the optimal value, the general shape of the densities remains the same for a large range of bandwidths.

2.3. Sources of Changes in the Density of Wages: 1979-88

We take as our point of departure the study by Bound and Johnson (1992) which considers the following competing explanations for the dramatic changes in the structure of wages over the period 1979–1988.

1. Increases in the relative demand for better–educated workers (Murphy and Welch 1992).
2. A slowdown in the rate of growth of relatively educated workers (Murphy and Welch 1992).

3. Changing technology brought upon by the computer revolution (Mincer 1992).

4. Decline in manufacturing employment and the power of unions (Harrison and Bluestone 1988)

To evaluate these explanations Bound and Johnson (1992) analyze changes in wage differentials among thirty–two experience/education/gender groups. They conclude that although each of the other three explanations contributed slightly to observed relative wage movements, the primary cause of changes in the structure of wages was technical change.

Decomposing changes in wage differentials into components attributable to a variety of factors is standard in applied economics.[10] Similar techniques have also been used to decompose summary measures of the distribution of wages into separate components.[11] In this paper, we propose a new semiparametric approach to perform similar decompositions of the overall distribution of wages. After presenting the basic facts to be explained, we introduce an alternative taxonomy for decomposing changes in the distribution of wages and explain how we do these decompositions.

In the analysis that follows we focus, as in Bound and Johnson (1992), on the changes in the distribution of wages over the period 1979–1988.[12] Figure 2 illustrates our hours-weighted kernel density estimates of the distribution of real log wages for 1979 and 1988 (in 1979 dollars) for both men and women. Several aspects of these distributions are particularly noteworthy. For men (Figure 2a)

[10] An early reference is Oaxaca (1973) who decomposes the gender wage gap into a component due to differences in factor endowments and a residual component due to discrimination.

[11] For example, see Freeman (1980b) who decomposes the variance of log wages into the sum of a union effect and of the variance that would prevail in the absence of unions.

[12] We also consider 1973-1979 and 1988-1992 changes for the sake of completeness in Section 5.

the lower tail grew significantly fatter over the period. Most of this fattening occurred in the area between the old (higher) 1979 federal minimum wage and the lower 1988 minimum. In addition, the upper tail of the distribution grew somewhat fatter.

Figure 2b, which plots the distribution of wages for women, clearly illustrates the primary role that minimum wages must play in any complete discussion of changes in the distribution of wages. Note the fattening of the lower tail in the area between the 1979 and the 1988 minimum. Even more striking is the fact that the mode of the 1979 distribution is near the minimum wage. A closer examination of the data indicates that this spike contains a roughly equal proportion of workers paid the minimum wage ($2.90) and of workers paid an integer wage just above the minimum wage ($3.00). Note also that the upper tail of the women's wage distribution grew more than did the upper tail of the men's distribution. This is not surprising since women's real wages grew by 4.2 percent over this period while men's real wages fell by 4.7 percent (the mean log wage for the entire sample declined by 1.7 percent).

In light of these considerations, we introduce the following taxonomy to understand the changes in the distribution of wages:

1. The effect of changes in the minimum wage.
2. The effect of changes in unionization.
3. The effect of changes in the distribution of workers' attributes, including industry affiliation.
4. The effect of changes in the supply and demand for workers of various skill categories.
5. Residual or "unexplained" changes.

The first factor (the effect of changes in the minimum wage) is suggested by the dramatic changes in the shape of the distribution around the minimum wage which is illustrated clearly in Figures 1 and 2. The second factor (the effect of changes in unionization) is suggested by the work of Freeman (1993) and Card (1992b) who found that the decline in unionization explains twenty percent of changes in the variance of wages from the 1970s to 1987. Our third factor (the effect of changes in the distribution of workers' attributes including industry affiliation) essentially captures changes in the distribution of the characteristics of the workforce that could affect the distribution of wages for a given structure of wages. For example, a polarization of schooling into dropouts and university graduates would widen the distribution of wages even if the returns to schooling remained constant. Bound and Johnson's explanations 1 and 2 are collapsed into a single factor in our taxonomy — supply and demand (our category 4) — while changes in the industrial structure have been combined with changes in other attributes of workers in our third category. In this taxonomy, factors such as "skill-biased technological shocks" (explanation 3 in Bound and Johnson) fall into the last category — residual changes.

3. Estimation Methods

In order to assess the role of different factors on changes in the distribution of wages, we must estimate the effect of each of these factors on the distribution of wages. In general, it is useful to view the estimation procedure as two separate steps. In step one, the effect of a given factor is estimated for workers with a given set of attributes. In step two, this estimated effect is integrated over the distribution of individual attributes. This yields the overall effect of the factor on the distribution of wages.

A few assumptions are required to make the estimation in a non-parametric setting tractable. The assumptions made in step one tend to be economic assumptions. For example, some assumptions must be made on the determination of the union status of workers, on the employment effects of the minimum wage, etc. By contrast, the assumptions required in step two tend to be statistical assumptions to facilitate the integration of the estimated effects over the distribution of workers attributes. We detail both types of assumptions below. Since discussion of our estimation procedure is greatly simplified by establishing some notation, we turn to that task first.

3.1. Notation

In order to discuss the estimation procedure, it is useful to view each individual observation as a vector (w, z, t) made up of a wage w, (which is virtually a continuous variable), a vector z of individual attributes (some of which are discrete variables), and a date t, which will take on only two values in the following comparisons. Each individual observation belongs to a joint distribution $F(w, z, t)$ of wages, individual attributes, and dates. The distribution of wages and attributes at one point in time is the conditional distribution $F(w, z|t)$. The density of wages at one point in time, $f_t(w)$ — which is the density of the marginal conditional distribution of w — can be written as the integral of the density of wage, $f(w|z, t_w)$, conditional on a set of individual attributes and on a date t_w over the distribution of individual attributes $F(z|t_z)$ at date t_z

$$\begin{aligned} f_t(w) &= \int_{z\in\Omega_z} dF(w, z|t_{w,z}=t; m_t) = \int_{z\in\Omega_z} f(w|z, t_w=t; m_t)\, dF(z|t_z=t) \qquad (2) \\ &\equiv f(w; t_w=t, t_z=t, m_t), \end{aligned}$$

where Ω_z is the domain of definition of the individual attributes, and m_t denote distributional characteristics of interest which we discuss below. Since the estimation and the decomposition exercise involves the combination of different "datings", the last line introduces the notation that accounts for these. For example, while $f(w; t_w=88, t_z=88, m_{88})$ represents the actual density of wages in 1988,

$f(w; t_w = 88, t_z = 79, m_{88})$ represents the density of wages that would have prevailed in 1988 had the distribution of individual attributes remained as in 1979. The difference between the actual density and this hypothetical density represents the effect of changes in the distribution of workers' attributes.

The conditional density of w given z at date t_w, $f(w|z, t_w = t; m_t)$, also depends on some distributional characteristics, m_t. In a parametric context, the distributional characteristics would summarize the particular functional form and its distributional parameters. The level of the minimum wage, which appears to compress the lower tail of the densities in Figure 1, can be viewed as a distributional characteristic of the densities. The conditional density of wages $f(w|z, t_w; m_t)$ in equation (2) is best viewed as identifying the "structure of wages" at date t_w when the minimum wage is equal to m_t. For example, a wage differential is simply a difference between two first order moments of this conditional distribution (i.e. conditional means). An actual (or hypothetical) distribution will depend on this structure of wages and the conditional distribution $F(z|t_z)$ which represents the distribution of workers' attributes z at date t_z. This notation enables us to formally write a large variety of hypothetical wage distributions that can then be used to decompose actual changes in the density of wages into a series of components.

Equation (2) also helps illustrate the two step estimation procedure sketched at the beginning of this section. The first step consists of estimating the effect of various factors on the density of wages for workers with similar attributes, that is the conditional density of wages $f(w|z, t_w; m_t)$. The second step consists of integrating the conditional density over the distribution of z, $F(z|t_z)$, to obtain the overall effect.

A direct application of this two step procedure, however, would be difficult. It would first require estimating a conditional density by kernel methods, and then numerically integrating this density over the (estimated) distribution of worker attributes. Considerable simplification can be achieved by a judicious choice of a reweighing function. We next turn to an explanation of these reweighing functions for each factor we consider. To aid the reader, we consider these reweighing functions in ascending order of complexity.

3.2. Effect of Changes in Unionization

Starting with Freeman (1980b), several studies based on micro data have established that unions tend to reduce wage inequality, at least among men. A potential explanation for the increase of wage inequality between 1979 and 1988 could thus be the 32% decline in unionization rates.

To investigate the effect of unions on the distribution of wages using our framework, let $z = (u, x)$, where u is a dummy variable that takes on the value 1 if the worker belongs to a union and the value 0 otherwise, and x is a vector of other attributes. The conditional distribution of unionization $F(u|x, t_u)$ indicates how the probability that workers are unionized in year t_u depends on their other attributes x. By analogy with the "structure of wages", we can call this conditional distribution the "structure of unionization". The structure of wages conditional of the vector of individual attributes other than union status can then be written as

$$f(w|x, t_w{=}\,88, t_u{=}\,88; m_{88}) = \int f(w|u, x, t_w{=}\,88; m_{88})\, dF(u|x, t_u{=}\,88). \qquad (3)$$

Since u can only take on two discrete values (0 or 1), equation (3) can be rewritten as the weighted sum of the conditional density in the non–union sector and in the union sector:

$$\begin{aligned} f(w|x, t_w{=}\,88, t_u{=}\,88; m_{88}) = {} & \text{Prob}(u{=}\,0|x, t_u{=}\,88)\, f(w|u{=}\,0, x, t_w{=}\,88; m_{88}) \\ & + \text{Prob}(u{=}\,1|x, t_u{=}\,88)\, f(w|u{=}\,1, x, t_w{=}\,88; m_{88}). \end{aligned} \tag{4}$$

Figure 3a illustrates these two weighted conditional densities while the density represented by a dotted line in Figure 3a and 3b represents their sum as defined in equation (4). [13] One representation of an increase in the unionization rate in this framework is an increase in the weight attached to the conditional density in the union sector. For example, increasing the unionization rate back to its 1979 level would yield a new weighted sum of densities

$$\begin{aligned} f(w|x, t_w{=}\,88, t_u{=}\,79; m_{88}) = {} & \text{Prob}(u{=}\,0|x, t_u{=}\,79)\, f(w|u{=}\,0, x, t_w{=}\,88; m_{88}) \\ & + \text{Prob}(u{=}\,1|x, t_u{=}\,79)\, f(w|u{=}\,1, x, t_w{=}\,88; m_{88}). \end{aligned} \tag{5}$$

This new density is represented by the solid line in Figure 3b. The estimated effect of an increase in the unionization rate is thus simply the difference between the two densities in Figure 3b. Note that for this estimated effect to be valid, the union status of workers with similar attributes z must be determined at random. Although this economic assumption seems strong, recent research suggests that the "selection" bias it introduces may be small.[14]

An explicit link between equations (3),(4), and (5) is obtained by writing the equation

$$\begin{aligned} f(w|x, t_w{=}\,88, t_u{=}\,79; m_{88}) & = \int f(w|u, x, t_w{=}\,88; m_{88})\, dF(u|x, t_u{=}\,79) \\ & = \int f(w|u, x, t_w{=}\,88; m_{88})\, \psi_u(u, x)\, dF(u|x, t_u{=}\,88). \end{aligned} \tag{6}$$

[13]The densities of real log wages in 1988 of Figure 3 were estimated for men with 12 years of education, and between 10 and 30 years of experience.

[14]This observation is based on the panel data estimates reported by Card (1992b) and Lemieux (1993a) and semi-parametric estimates of selection model reported by Lanot and Walker (1993).

where $\psi_u(u,x)$ is a reweighing factor defined as follows:

$$\begin{aligned}\psi_u(u,x) &\equiv dF(u|x,t_u{=}79)/dF(u|x,t_u{=}88)) \\ &= I(u{=}1)\,\frac{\text{Prob}(u{=}1|x,t_u{=}79)}{\text{Prob}(u{=}1|x,t_u{=}88)} + [1-I(u{=}1)]\,\frac{\text{Prob}(u{=}0|x,t_u{=}79)}{\text{Prob}(u{=}0|x,t_u{=}88)}.\end{aligned} \tag{7}$$

This equation restates the fact that we can estimate the effect of changes in the unionization rate by simply reweighing the conditional densities of wages in the union and in the non–union sector. Using this reweighing factor also simplifies the estimation of the overall effect of unions, that is, the effect for each set of attributes integrated over the whole distribution of attributes z.

The overall 1988 density of wages is obtained by integrating the product of the "structure of wages" by the "structure of unionization" over the distribution of individual attributes

$$\begin{aligned}&f(w;t_w{=}88,t_u{=}88,t_x{=}88,m_{88}) = \\ &\qquad \int\!\!\int f(w|u,x,t_w{=}88;m_{88})\,dF(u|x,t_u{=}88)\,dF(x|t_x{=}88).\end{aligned} \tag{8}$$

The hypothetical density of wages that would have prevailed in 1988 had the structure of unionization remained as in 1979 is then

$$\begin{aligned}&f(w;t_w{=}88,t_u{=}79,t_x{=}88,m_{88}) = \\ &\qquad \int\!\!\int f(w|u,x,t_w{=}88;m_{88})\,dF(u|x,t_u{=}79)\,dF(x|t_x{=}88).\end{aligned} \tag{9}$$

Using the reweighing factor $\psi_u(u,x)$ defined in equation (7), equation (9) can be rewritten as

$$\begin{aligned}&f(w;t_w{=}88,t_u{=}79,t_x{=}88,m_{88}) = \\ &\qquad \int\!\!\int f(w|u,x,t_w{=}88;m_{88})\,\psi_u(u,x)\,dF(u|x,t_u{=}88)\,dF(x|t_x{=}88).\end{aligned} \tag{10}$$

Except for $\psi_u(u,x)$, the density $f(w;t_w{=}88,\ t_u{=}79,\ t_x{=}88,m_{88})$ is thus identical to the density $f(w;t_w{=}88,t_u{=}88,t_x{=}88,m_{88})$. This fact greatly simplifies the estimation procedure. The estimation of the density in equation (9) could theoretically be performed by nonparametrically estimating the conditional densities of w, u, and x and numerically integrating their product over the values of u and x although it would be very difficult. It is much simpler to estimate $\psi_u(u,x)$ and use its estimated value to "reweigh" the 1988 sample.

The only remaining substantive estimation problem is to estimate $\psi_u(u,x)$. From equation (7), it is clear that we only need to estimate the conditional probability $\text{Prob}(u{=}1|x,t_u)$ for $t_u = 79$ and 88. One standard model for this conditional probability is the probit model

$$\text{Prob}(u{=}1|x,t_u{=}t){=}\text{Prob}(\epsilon > -\beta_t' H(x)) = 1 - \Phi(-\beta_t' H(x)), \tag{11}$$

where $\Phi(.)$ is the cumulative normal distribution and $H(x)$ is a vector of covariates that is a function of x, the vector of individual attributes including age, education, industry, occupation, etc. The vector $H(x)$ typically used is a low order polynomial in x. If x only took on a limited number of values, the best $H(x)$ to use would be a full set of dummy variables indicating each possible value of x. In this special case, the probit model would be a equivalent to a "cell-by-cell" non-parametric model.

The estimated value $\hat{\psi}_u(u,x)$ can then be used as a reweighing factor in the kernel density estimation. Expanding on the weighted kernel estimation described above, one finds that

$$\hat{f}(w;t_w{=}88,t_u{=}79,t_x{=}88,m_{88}) = \sum_{i\in S_{88}} \frac{\theta_i}{h^*}\hat{\psi}_u(u,x)\,K\left(\frac{w-W_i}{h^*}\right), \tag{12}$$

where S_{88} is the set of indices of the 1988 sample.

3.3. *Effect of Changes in the Minimum Wage*

Another potential explanation for the increase in wage inequality between 1979 and 1988 is the 27 percent decline in the real value of the minimum wage. In order to understand how we estimate the effect of the minimum wage, consider the examples of densities of wages in 1979 and 1988 that are illustrated in Figure 4a and 4b for workers with similar attributes. [15] The vertical lines represent the real value of the minimum wage in 1979. How could we estimate the effect on the 1988 density of increasing the minimum wage back to its 1979 value?

One simple estimator consists in replacing the section of the 1988 density at or below the 1979 minimum by the corresponding section of the 1979 density. Note that this imputed section of the 1979 density has to be scaled appropriately to make certain that the overall density still integrates to one. The resulting 1988 density with the minimum wage at its 1979 level is represented in Figure 4c. The effect of the minimum wage on the density of wages is the difference between Figure 4a and 4c.

One advantage of this proposed estimation procedure is that it is implementable in a nonparametric setting. A simple procedure that applies to a sample of workers with similar attributes could be implemented in the four following steps on a 1979 and a 1988 sample:

1. Throw out all observations from the 1988 sample with $w \leq m_{79}$, where m_{79} is the 1979 minimum wage.
2. Replace these observations with observations in the 1979 sample that satisfy $w \leq m_{79}$.
3. Reweigh the observations obtained from the 1979 sample so that the (weighted) number of observations in the new 1988 sample is the same as before.
4. Estimate the density of wages in this modified 1988 sample by weighted kernel methods.

This procedure captures the essence of the estimation procedure we propose below, which in addition accounts for workers with different attributes. In what

[15] The densities of real log wages in 1988 and in 1979 of Figure 4 were estimated for women with 12 years of education or less, and with less than 20 years of experience. The apparent spill over effect of the minimum wage in Figure 4c is in fact a result of the smoothing.

follows we more formally present this estimation procedure and the conditions that are required to ensure its validity.

Note first that the idea of replacing a section of the 1988 density with a section of the 1979 density can be justified on economic grounds as long as the three following assumptions, similar to those of Blackburn et al. (1992), are satisfied.

ASSUMPTION 1: The minimum wage has no spill over effects on the distribution of wages above the minimum wage. This implies that for any two values m_0 and m_1 ($m_0 \leq m_1$) of the minimum wage m_t, the structure of wages, $f(w|z,t_w;m_t)$, for wages above the higher value of the minimum wage, that is for w such that $[1 - I(w \leq m_1)] = 1$, is the same in both cases:

$$[1 - I(w \leq m_1)]f(w|z,t_w;m_0) = [1 - I(w \leq m_1)]f(w|z,t_w;m_1), \qquad (13)$$

where $I(.)$ is an indicator function that takes on the value 1 if the condition in parentheses is satisfied, and 0 otherwise, and where, to simplify the exposition, the vector z regroups all attributes.

ASSUMPTION 2: The shape of the conditional density of real wages is time stationary for values of the real wage smaller or equal to the real value of any minimum wage (m). This implies that, for wages below the value of the minimum wage, that is for w such that $I(w \leq m) = 1$, the structure of wages in 1988 is proportional to the structures of wages in 1979

$$I(w \leq m)f(w|z,t_w{=}\,79;m) = \kappa I(w \leq m)f(w|z,t_w{=}\,88;m), \qquad (14)$$

where κ is an integration factor to be specified below.

ASSUMPTION 3: The minimum wage has no effects on employment probabilities.

In terms of Figure 4a, ASSUMPTION 1 guarantees that the distribution to the right of the vertical line remains the same when the minimum wage increases.

ASSUMPTION 2 states that the new part of the density to the left of the vertical line has the same shape as in 1979, while ASSUMPTION 3 guarantees that the fraction of workers to the left of the vertical line is not affected by changes in the minimum wage. In other words, the surface underneath the density to the left of the vertical line remains constant, notwithstanding changes due to smoothing.

Note that ASSUMPTIONS 1 to 3 are conservative in the sense of minimizing the effect (in absolute value) of an hypothetical increase in the minimum wage on measures of wage dispersion such as the variance. The wage compression effect would be larger if the minimum wage had some positive spill over effects on wages just above the minimum wage (Grossman 1983). In addition, ASSUMPTION 3 implies that an increase in the minimum wage causes no attrition in the lower tail that would contribute to make the variance even smaller.[16]

Using ASSUMPTIONS 1 to 3, the conditional density of wages that would prevail in 1988 if the minimum wage was at its 1979 level may be written as

$$\begin{aligned} f(w|z,t_w=88;m_{79}) = {} & I(w \leq m_{79})\,\psi_w(z,m_{79})\,f(w|z,t_w=79;m_{79}) \\ & + [1 - I(w \leq m_{79})]\,f(w|z,t_w=88;m_{88}) \end{aligned} \tag{15}$$

where

$$\psi_w(z,m_{79}) = \frac{\text{Prob}(w \leq m_{79}|z,t_w=88)}{\text{Prob}(w \leq m_{79}|z,t_w=79)} \tag{16}$$

is a weighting function that ensures that the density integrates to one.

Equation (16) is a formal statement of the procedure discussed above. To construct a 1988 conditional density with the minimum wage at its 1979 value

[16]The assumption of no employment effects is consistent with the results of studies using micro data (see Card (1992a), Card (1992c), Card and Krueger (1993), Katz and Krueger (1992) and an earlier set of studies using an event study approach (see Lester (1964) and Lester (1960)). On the other hand, evidence from U.S. time-series studies (see Brown (1988), Brown, Gilroy and Kohen (1982)) and one panel data study (see Neumark and Wascher (1992)) suggests small disemployment effects.

$[f(w|z, t_w = 88; m_{79})]$, we select the part of the 1988 density above m_{79} and the part of the 1979 density at or below m_{79} using the indicator function. We also premultiply the 1979 density by a "reweighing" or "scaling" factor $\psi_w(z, m_{79})$ to make sure the overall density integrates to one.

To obtain the effect of the minimum wage on the overall distribution of wages in 1988, it is necessary to integrate the conditional density in equation (16) over the distribution of attributes:

$$\begin{aligned} f(w; t_w = 88, t_z = 88, m_{79}) &= \int f(w|z, t_w = 88; m_{79})\, dF(z|t_z = 88) \\ &= \int I(w \leq m_{79})\, \psi_w(z, m_{79})\, f(w|z, t_w = 79; m_{79})\, dF(z|t_z = 88) \\ &\quad + [1 - I(w \leq m_{79})]\, f(w|z, t_w = 88; m_{88})\, dF(z|t_z = 88) \\ &= \int I(w \leq m_{79})\, \psi_w(z, m_{79})\, f(w|z, t_w = 79; m_{79})\, \psi_z(z) dF(z|t_z = 79) \\ &\quad + [1 - I(w \leq m_{79})]\, f(w|z, t_w = 88; m_{88})\, dF(z|t_z = 88), \end{aligned} \tag{17}$$

where $\psi_w(z, m_{79})$ is as defined in equation (16), and where

$$\psi_z(z) = \frac{\text{Prob}(t_z = 88|z)\, \text{Prob}(t_z = 79)}{\text{Prob}(t_z = 79|z)\, \text{Prob}(t_z = 88)} \tag{18}$$

is another weighting function that allows us to write the density $f(w; t_w = 88, t_z = 88, m_{79})$ as the sum of a component that is a reweighing of the 1979 density and a component from the 1988 density. This strategy will facilitate the estimation. After applying Bayes' rule, the product of the two weighing functions simplifies to

$$\psi(z, m_{79}) \equiv \psi_w(z, m_{79}) \cdot \psi_z(z) = \frac{\text{Prob}(t_w = 88|z, w \leq m_{79})}{\text{Prob}(t_w = 79|z, w \leq m_{79})} \cdot \frac{\text{Prob}(t_z = 79)}{\text{Prob}(t_z = 88)}. \tag{19}$$

The probability of being in period t, given certain individual attributes and a wage below the 1979 minimum wage, can be estimated using a probit model

$$\text{Prob}(t_w = t|z, w \leq m_{79}) = \text{Prob}(\epsilon > -\beta' H(z)) = 1 - \Phi(-\beta' H(z)), \tag{20}$$

where $\Phi(.)$ is the cumulative normal distribution and $H(z)$ is a vector of covariates that is a function of z, the vector of individual attributes. In practice, this probit model is estimated by pooling observations from the 1979 and 1988 samples that have real wages smaller or equal to the 1979 real minimum wage.[17] The actual vector z chosen for the estimation consists of relatively unrestricted combinations of the individual attributes including age, education, union status, industry, occupation, etc. Given that we view the two dates as the two possible events in the date space, the unconditional probability $\text{Prob}(t_z = 79)$ is equal to the weighted number of observations in 1979 divided by the weighted number of observations in both 1979 and 1988. The unconditional probability $\text{Prob}(t_z = 88)$ is similarly defined.

Once again, the use of a reweighing factor simplifies the estimation. Equation (17) can be translated directly into a weighted kernel format as the sum of weighted kernel functions over two sets. The first set (L_{79}) consists of the 1979 observations with a wage smaller or equal to the 1979 minimum wage, appropriately reweighed by the factor $\hat{\psi}(Z_i, m_{79})$. The second set (U_{79}) consists of the 1988 observations that have a wage larger than the 1979 minimum wage. The density estimate of $f(w; t_w = 88, t_z = 88, m_{79})$ in the weighted kernel format is thus written as

$$\begin{aligned} \hat{f}(w; t_w = 88, t_z = 88, m_{79}) = \sum_{i \in L_{79}} \frac{\theta_i}{h^*} \hat{\psi}(Z_i, m_{79}) K\left(\frac{w - W_i}{h^*}\right) \\ + \sum_{i \in U_{79}} \frac{\theta_i}{h^*} K\left(\frac{w - W_i}{h^*}\right), \end{aligned} \tag{21}$$

where $L_{79} = \{i : W_i \in S_{79} \text{ and } W_i \leq m_{79}\}$ and $U_{79} = \{i : W_i \in S_{88} \text{ and } W_i > m_{79}\}$, and S_t is the set of indices of the sample of date t. The weighting factor $\hat{\psi}(Z_i, m_{79})$ is simply the value of $\psi(Z_i, m_{79})$ estimated via the probit model. The estimation of the density $f(w; t_w = 88, t_z = 88, m_{79})$ is thus very similar to the four step procedure

[17] In the estimation, we use the log of \$3.00 instead of the minimum wage of \$2.90 as the value of m_{79} (in 1979 dollars). As discussed in section 4.3, this choice is driven by the abnormal concentration of workers at \$3.00 in 1979 which suggests some small spill over effects of the minimum wage. The sensitivity of our results to this choice is reported in Appendix Table A2.

described early in this subsection. We replace the 1988 observations at or below the 1979 minimum wage with the corresponding 1979 reweighed observations. The only difference here is that our reweighing factor is more than a simple rescaling factor. It puts a different weight on 1979 observations with different attributes z to insure that the distribution of attributes remains as in 1988.

3.4. *Effect of Changes in the Distribution of Individual Attributes*

As a next step in estimating the sources of change in the distribution of wages, we estimate the effect of changes in the distribution of individual attributes other than union status. We focus on the following attributes: experience, schooling, race, full-time or part-time status, SMSA dummy, 3 occupational categories, and 19 industry categories. We view this part of the estimation as mostly mechanical. We simply ask whether changes in the weights attached to different workers in the workforce can explain some of the changes in the overall distribution of wages. The hypothetical density of wages in 1988 with the 1979 distribution of individual attributes, other than union status, is simply

$$\begin{aligned} &f(w; t_w = 88, t_u = 88, t_x = 79, m_{88}) \\ &\qquad = \iint f(w|u, x, t_w = 88; m_{88})\, dF(u|x, t_u = 88)\, dF(x|t_x = 79) \\ &\qquad = \iint f(w|u, x, t_w = 88; m_{88})\, dF(u|x, t_u = 88)\, \psi_x(x)\, dF(x|t_x = 88), \end{aligned} \tag{22}$$

where $\psi_x(x) \equiv dF(x|t_u = 79)/dF(x|t_u = 88))$. Applying Bayes' rule, this ratio can be written as

$$\psi_x(x) = \frac{\text{Prob}(t_x = 88)}{\text{Prob}(t_x = 79)} \cdot \frac{\text{Prob}(t_x = 79|x)}{\text{Prob}(t_x = 88|x)}. \tag{23}$$

The probability of being in period t, given individual attributes x can once again be estimated using a probit model

$$\text{Prob}(t_x = t|x) = \text{Prob}(\epsilon > -\beta' H(x)) = 1 - \Phi(-\beta' H(x)), \tag{24}$$

where $\Phi(.)$ is the cumulative normal distribution. The probability of an observation belonging to period t is once again the appropriate pooled sample proportion. It is thus straightforward to compute an estimate $\hat{\psi}_z(x)$ and use it as a reweighing factor in the weighted kernel density estimation.

The weighted kernel estimate of the hypothetical density is

$$\hat{f}(w; t_w = 88, t_u = 88, t_z = 79, m_{88}) = \sum_{i \in S_{88}} \frac{\theta_i}{h^*} \hat{\psi}_z(x) K\left(\frac{w - W_i}{h^*}\right), \tag{25}$$

where S_{88} is the set of indices of the 1988 sample.

3.5. *Effect of Changes in Supply and Demand*

The effects that we have described to this point have been primarily "institutional". What remains is at the heart of many discussions on the causes of changes in the wage structure, that is, the changes in the supply and demand for workers of different skill categories. In most analyses of changes in the structure of wages, the supply and demand for various categories of labor is simply assumed to shift the mean of the distribution of wages for each of these categories. The situation is slightly more complex, however, when we consider the possibility of workers been paid the minimum wage. We use a standard approach to estimate how changes in supply and demand shift the distribution of wages other than the minimum wage. We label this shift Δw. It is also necessary, however, to estimate the effect of demand and supply shocks on the *probability* of being paid the minimum wage. We call this effect ΔP.

In the appendix, we show how the approach suggested by Bound and Johnson (1992) to estimate the effect of supply and demand conditions on the structure of wages can be used in the presence of minimum wage laws. Like Bound and Johnson, we divide the workforce into 32 experience–education–gender cells and construct measures of supply (N_{jt}) and demand (D_{jt}) for each of cell j. Linear regression methods are used to estimate the shift, $\Delta\widehat{w}_j$, in the mean of the distribution of wages other than the minimum wage due to changes in supply and demand. Grouped data logit methods are then used to estimate the effect of changes in supply and demand on the probability of being paid the minimum wage, $\Delta\widehat{P}_j$. Note that the supply variable N_{jt} simply reflects the share of the total workforce in cell j. The demand variable D_{jt} is a "fixed–coefficient manpower requirements index" that reflects between–sector shifts in relative labor demands.

Let $d_t = (\ln D_{1,t}, \ldots, \ln D_{32,t})$ and $n_t = (\ln N_{1,t}, \ldots, \ln N_{32,t})$, be two vectors of distributional characteristics summarizing demand and supply conditions. Expanding the list of distributional characteristics to include d_t and n_t, the conditional density of wages in 1988 can be written as

$$\begin{aligned} f(w|z, t_w{=}\,88; m_{88}, d_{88}, n_{88}) &= I(m_{88}{-}\epsilon \leq w \leq m_{88}{+}\epsilon)\, f(w|z, t_w{=}\,88; m_{88}, d_{88}, n_{88}) \\ &\quad + [1 - I(m_{88}{-}\epsilon \leq w \leq m_{88}{+}\epsilon)]\, f(w|z, t_w{=}\,88; m_{88}, d_{88}, n_{88}) \end{aligned} \tag{26}$$

where ϵ is set to a small value such as \$0.05. The indicator function is used to divide the conditional density into two components: a first component for values of the wage at or very close (within \$0.05) to the minimum wage, and a second component for other values of the wage. If supply and demand conditions had remained at their 1979 level, the mean of the conditional density of wages at values other than the minimum wage would have been $\Delta\widehat{w}_j$ lower. The probability of being at the minimum wage would have been $\Delta\widehat{P}_j$ lower. The conditional density of wages if supply and demand conditions had remained at their 1979 level can thus be written as

$$\begin{aligned} & f(w|z, t_w{=}\,88; m_{88}, d_{79}, n_{79}) \\ &\quad = I(m_{88}{-}\epsilon \leq w \leq m_{88}{+}\epsilon)\, \varphi(z, m_{88})\, f(w|z, t_w{=}\,88; m_{88}, d_{88}, n_{88}) \\ &\quad + [1 - I(m_{88}{-}\epsilon \leq w \leq m_{88}{+}\epsilon)]\, \varphi^{-1}(z, m_{88})\, f(w - \Delta\widehat{w}_j|z, t_w{=}\,88; m_{88}, d_{88}, n_{88}) \end{aligned} \tag{27}$$

where

$$\varphi(z, m_{88}) \equiv \frac{\text{Prob}(m_{88}-\epsilon \leq w \leq m_{88}+\epsilon | z, d_{79}, n_{79})}{\text{Prob}(m_{88}-\epsilon \leq w \leq m_{88}+\epsilon | z, d_{88}, n_{88})}, \tag{28}$$

and where

$$\text{Prob}(m_{88}-\epsilon \leq w \leq m_{88}+\epsilon | z, d_{79}, n_{79}) = \text{Prob}(m_{88}-\epsilon \leq w \leq m_{88}+\epsilon | z, d_{88}, n_{88}) - \Delta\widehat{P}_j. \tag{29}$$

The overall effect of supply and demand shocks is the difference between the 1988 density of wages and the hypothetical density obtained by integrating equation (27) over the distribution of individual attributes z:

$$\begin{aligned} & f(w; t_w = 88, t_z = 88, m_{88}, d_{79}, n_{79}) \\ = & \int I(m_{88}-\epsilon \leq w \leq m_{88}+\epsilon)\, \varphi(z, m_{88})\, f(w|z, t_w = 88; m_{88}, d_{88}, n_{88}) \\ & +[1 - I(m_{88}-\epsilon \leq w \leq m_{88}+\epsilon)]\, \widehat{\varphi}^{-1}(z, m_{88})\, f(w - \Delta\widehat{w}_j | z, t_w = 88; m_{88}, d_{88}, n_{88}) \\ & \quad dF(z|t_z = 88). \end{aligned} \tag{30}$$

The density estimate of $f(w; t_w = 88, t_z = 88, m_{88}, d_{79}, n_{79})$ is translated into a weighted kernel format as the sum of weighted kernel functions over two sets. The first set (M_{88}) consists of the 1988 observations with a wage close to the 1988 minimum wage, reweighed by a factor that accounts for the shift in probability $\widehat{\varphi}(z, m_{88})$. The second set ($O_{88}$) consists of the 1988 observations with a wage different from the 1988 minimum wage, reweighed by the inverse of the previous factor.

$$\begin{aligned} \widehat{f}(w; t_w = 88, t_z = 88, m_{88}, d_{79}, n_{79}) = & \sum_{i \in M_{88}} \frac{\theta_i}{h^*} \widehat{\varphi}(z, m_{88})\, K\left(\frac{w - W_i}{h^*}\right) \\ & + \sum_{i \in O_{88}} \frac{\theta_i}{h^*} \widehat{\varphi}^{-1}(z, m_{88})\, K\left(\frac{w - W_i + \Delta\widehat{w}_j}{h^*}\right), \end{aligned} \tag{31}$$

where $M_{88} = \{i : W_i \in S_{88} \text{ and } m_{88}-\epsilon \leq W_i \leq m_{88}+\epsilon\}$ and $O_{88} = \{i : W_i \in S_{88}$ and $m_{88}-\epsilon \leq W_i \leq m_{88}+\epsilon\}$.

4. Accounting for Changes in the Density of Wages 1979-88

4.1. Decomposition Method

In this section, we use the estimation methods described in section 3 to decompose changes in the density of wages between 1979 and 1988. Using the taxonomy we proposed in section 2.3, and considering the explanations in the same order as previously enumerated, the change in density from 1979 to 1988 can be decomposed into the five following components:

$$\begin{aligned} f_{88}(w) - f_{79}(w) &= [f(w; t_w{=}88, t_z{=}88, m_{88}) - f(w; t_w{=}88, t_z{=}88, m_{79})] \\ &+ [f(w; t_w{=}88, t_u{=}88, t_z{=}88, m_{79}) - f(w; t_w{=}88, t_u{=}79, t_z{=}88, m_{79})] \\ &+ [f(w; t_w{=}88, t_u{=}79, t_z{=}88, m_{79}) - f(w; t_w{=}88, t_u{=}79, t_z{=}79, m_{79})] \\ &+ [f(w; t_w{=}88, t_z{=}79, m_{79}, d_{88}, n_{88}) - f(w; t_w{=}88, t_z{=}79, m_{79}, d_{79}, n_{79})] \\ &+ [f(w; t_w{=}88, t_z{=}79, m_{79}, d_{79}, n_{79}) - f(w; t_w{=}79, t_z{=}79, m_{79}, d_{79}, n_{79})]. \end{aligned} \tag{32}$$

Note that we have omitted the distributional factors d_{88} and n_{88} in the first three components of the decomposition to simplify the notation.

The decomposition in equation (32) is sequential. The first component indicates the effect of changes in the minimum wage holding all other factors at their 1988 level. The second component indicates the effect of changes in unionization, holding the minimum wage at its 1979 level but leaving all other factors at their 1988 level. By contrast, our discussion in section 3.2 considered estimation of the effect of unionization when all factors, including the minimum wage, were at their 1988 level. The estimation approach proposed in that section can be used, however, provided that it is applied to an appropriately modified sample.

More specifically, equation (12) has to be estimated over a 1988 sample in which observations at or below the 1979 minimum have been replaced by corresponding reweighed 1979 observations instead of over the raw 1988 sample (the set of indices in S_{88}). In other words, the estimator in equation (12) is valid for estimating the density $f(w; t_w = 88, t_u = 79, t_z = 88, m_{79})$ provided that the 1988 sample is appropriately modified. A similar approach can be used to estimate the other components of equation (32) sequentially.

The decomposition in equation (32) is sequential to insure that the change in density between 1979 and 1988 is exactly equal to the sum of the five components corresponding to the factors in our taxonomy. While the total change in density is in fact equal to the sum of the five corresponding partial effects (the effect of one factor holding all the others at their 1988 level) plus ten interactions between the effect of the various factors, it would be cumbersome to include all of these terms. Rather, the sequential decomposition implicitly accounts for the interaction terms. One drawback of the sequential decomposition, however, is that the estimated effect of each factor may depend on the order of the decomposition. We thus later report results of performing the decomposition in a different order to insure that our findings are robust to such modifications of procedure.

4.2. *Estimated Effect of Changes in Factors on the Density of Wages*

We report kernel estimates of the densities used in the decomposition equation (32) in Figure 5 for men, and in Figure 6 for women. We use the following convention: for each of the four factors considered (minimum wage, unions, attributes, and demand and supply) we report the density obtained by holding the factor to be explained at its 1988 value with a dotted line, and the density obtained by holding the factor at its 1979 value with a solid line. The difference between the two lines represents the effect of changes in the factor on the density of wages.

For example, the dotted line in Figure 5a is the raw 1988 density of wages $f(w; t_w = 88,\ t_z = 88, m_{88})$ while the solid line represents $f(w; t_w = 88, t_z = 88,\ m_{79})$.

The influence of the minimum wage on the distribution of wages is clearly seen. There is considerably more mass at the bottom of the wage distribution in 1988 (at and slightly above $2 (1979$)) than if the minimum wage had remained at its 1979 level. This is true both for men and women (Figure 6a). It is clear from these figures that the decline in the real value of the minimum wage between 1979 and 1988 is responsible for a significant fraction of the rise in wage inequality over the 1980s.

The estimated densities with the unionization rate at its 1979 and its 1988 levels for men and women are presented in Figures 5b and 6b respectively. In these figures, the dotted line represents estimates of the 1988 density of wages with the 1988 structure of unionization [$f(w; t_w = 88, t_u = 88, t_x = 88, m_{79})$]. The solid line represents a hypothetical 1988 density that would prevail if the structure of unionization had remained as in 1979 [$f(w; t_w = 88, t_u = 79, t_x = 88, m_{79})$]. Figure 5b shows that changes in unionization had a substantial effect on the distribution of men's wages. The decline in unionization between 1979 and 1988 contributed to the decline of the "middle" of the distribution and the fattening of the lower tail of the distribution. This is easily explained by the fact that unions provide the most help to relatively unskilled workers by moving them toward the middle of the wage distribution. As unionization declines, these workers slide back in the lower tail of the distribution.

By contrast, Figure 6b indicates that changes in the structure of unionization had a negligible effect on the distribution of women's wages. This result is consistent with the fact that the unionization rate did not decline very much for women (Table 2) and that unions have little impact on women's wage inequality (see Lemieux (1993b)).

The estimated densities corresponding to individual attributes are shown in

Figure 5c (men) and Figure 6c (women). The estimated densities with the 1988 distribution of individual attributes [$f(w; t_w = 88, t_u = 79, t_x = 88, m_{79})$] are shown in dotted lines. The densities with the 1979 distribution of attributes [$f(w; t_w = 88, t_u = 79, t_x = 79, m_{79})$] qualitatively look like a translation to the left of the densities with the 1988 distribution of attributes. This suggests there was an upgrade in the attributes or "skills" of workers between 1979 and 1988. This is consistent with secular increases in the number of years of schooling of workers.

The density estimates corresponding to the effect of demand and supply are reported in Figure 5d for men and in Figure 6d for women. The dotted line represents the estimated density with the actual supply and demand conditions of 1988 [$f(w; t_w = 88, t_z = 79, m_{79}, d_{88}, n_{88})$] while the solid line represents the density that would have prevailed had the supply and demand conditions remained at their 1979 level ($f(w; t_w = 88, t_z = 79, m_{79}, d_{79}, n_{79})$). The figures clearly show that, overall, workers in the lower tail of the wage distribution were adversely affected by supply and demand shocks. This is particularly clear for men. Most of the concentration of workers at the 1979 real minimum wage (spike on the left of the dotted density) would not have occurred had demand and supply conditions remained at their 1979 level.

4.3. Residual Differences in Densities

A clearer illustration of the contribution of the different factors in explaining changes in the structure of wages is obtained by plotting the residual difference in densities that cannot be explained after accounting for the effect of a specific factor. For example, after having accounted for changes in the minimum wage — the first component in equation (32) — there are still four other components to be accounted for. By the time we look at the effect of supply and demand, there is only one component unaccounted for. When viewed this way, the goal of the decomposition exercise is simply to get a "flat line" difference in densities once all the factors have been accounted for.

Looking explicitly at the difference between two densities also provides a complete description of changes in the distribution of wages. For example, a mean-preserving spread of the distribution would result in a positive density difference in the tails of the distribution and in a negative density difference in the middle of the distribution. By contrast, measures like the variance, the 10–90 differential, or the Gini coefficient only summarize differences between two distributions with a single number. When looked at in isolation, they fail to indicate the region of the wage distribution in which most of the changes are happening. Note also that the difference between the 1979 and the 1988 densities are not very smooth, which reduces their "visual impact". The differences in densities presented in Figures 7 and 8 were thus all smoothed further using a Gaussian kernel and a bandwidth of 0.07.

Our estimates of residual differences for men are displayed in Figure 7; for women the results are displayed in Figure 8. Considering men first, Figure 7a displays our estimates of the difference between the distribution of wages in 1979 and 1988. The vertical line marks the location of the 1979 minimum wage. One of the most important differences between the two periods is the considerable additional density in the 1988 distribution at wages below the 1979 minimum. In Figure 7b, we remove changes associated with the fall in the minimum wage. The difference between the two densities at values below the 1979 minimum falls considerably, the tall "hump" of panel a) shifts to the right in panel b). This indicates that even after accounting for changes in the minimum wage, the 1988 density still would have had much greater weight in the low wage part of the distribution, albeit at levels around the 1979 minimum.

Next we remove changes attributable to changes in union density in Figure 7c. Note that the unionization rate fell by 32 percent during this period (see Table 2). This apparently played a significant role in explaining the decline in the density of wages in the middle of the distribution. After removing the effect due to changes in unionization, therefore, the difference between the two densities near the center of the distribution diminishes.

Over the period, the sample of men grew slightly older, and slightly more educated. The effects of this change can be seen in the estimates in Figure 7d. The net effect of changes in demographic characteristics and industrial structure ran somewhat counter to the effect of changes in union density. In particular, while changes in the distribution of individual attributes almost completely explains the difference between the two densities in their respective upper tails, it actually exacerbates the difference between the two distributions everywhere else.

As shown in Figure 7e, changes in the supply and demand for various classes of workers explain almost all of the remaining difference between the two densities. In particular, it appears that changing patterns of supply and demand factors most adversely affected workers in the lower part of the distribution.

In Figure 8, we perform a similar decomposition for the changes in the distribution of women's wages over the 1979–1988 period. Clearly, the single most important cause of changes in the distribution of women's wages was the decline in the federal minimum wage. As it turns out, there is a considerable amount of difference between the two densities in the area immediately to the right of the 1979 minimum. This indicates the possibility of additional spill over effects of minimum wage legislation. It does appear that we may still be attributing some minimum wage effects to the residual or unexplained variance.[18]

[18] The importance of correctly allowing for spill over effects is addressed in Appendix Table A2

The remaining decompositions are presented in panels b) to e) of Figure 8. In comparison to men, the effect of unions on the distribution of women's wages presented in Figure 8c is quite slight. As documented by Lemieux (1993b), this is largely a consequence of the pattern of unionization among women. First, women are less likely to be unionized (see Table 2). Second, whereas low skill men are more likely to be unionized than their high skill counterparts, for women just the reverse is true.[19]

Like men, women workers also became more skilled between 1979 and 1988. The effect of changes in the distribution of individual attributes displayed in Figure 8d is somewhat more important than the effect of unions. Changes in supply and demand also had some impact on the difference in densities, and what is left (the residual change) is plotted in Figure 8e.

4.4. Quantitative Measures

By way of comparison with the rest of the literature and to provide some numerical values for the changes documented in the previous subsection, it is useful to compute a few summary measures of the distribution of wages. Computing these measures is straightforward once the density of wages has been estimated.

For example, the 10th percentile of the estimated density of wages for 1988,

where we compare our estimates assuming that minimum wages affect those at or below a $3.00 minimum, to estimates computed assuming that minimum wages affect those at or below the $2.90 minimum. It is evident that using a $2.90 cutoff would have led to some underestimation. Using the $3.00 cutoff increases our estimate of the percentage change in the standard deviation of women's wages explained by minimum wages from 15.1 percent to 30.2 percent. For men, the percentage explained jumps from 9.7 percent to 24.8 percent.

[19]The situation is actually more pronounced for women in Canada, where the effect of unions on women's wages is to increase the level of inequality.

$w_{.1}$, is such that

$$\int_{-\infty}^{w_{.1}} \hat{f}_{88}(w)\, dw = 0.10, \tag{33}$$

while the 90th percentile , $w_{.9}$ is such that

$$\int_{-\infty}^{w_{.9}} \hat{f}_{88}(w)\, dw = 0.90. \tag{34}$$

The 10–90 wage differential is simply $w_{.9} - w_{.1}$. The 1979–88 change in the 10-90 wage differential is obtained by computing similar statistics from the estimated density for 1979.

It is also easy to derive other inequality measures. Among the more widely used are Theil's entropy coefficient and the Gini coefficient. Let $\omega = \exp(w)$ be the real wage, and denote the density of its distribution by $\hat{f}_{88}(\omega) = \hat{f}_{88}(w)/\omega$. Theil's entropy coefficient is the negative of the expectation of the logarithm of the 1988 density

$$T = -\int_{0}^{\infty} \ln[\hat{f}_{88}(\omega)]\, \hat{f}_{88}(\omega)\, d\omega. \tag{35}$$

If the distribution is normal, with mean μ and variance σ^2, for example, it is easily shown that this measure of inequality depends only on the variance $T = \frac{1}{2}\ln 2\pi e\sigma^2$.

The Lorenz curve is another device commonly used to measure inequality. It is the locus of points with the cumulative population share on the abscissa, $F(\omega) = \int_0^\omega f(\omega)d\omega$, and the cumulative wage share, $F_1(\omega)$, which is the first moment distribution

$$F_1(\omega) = \int_0^{\omega} \omega f_{88}(\omega) d\omega \Big/ \int_0^{\infty} \omega f_{88}(\omega)\, d\omega, \tag{36}$$

on the ordinate. The Gini coefficient is the ratio of the area enclosed by the Lorenz curve and the diagonal line to the total area below the diagonal. It thus takes on a minimum value of zero (perfect equality) when the Lorenz curve coincides with the diagonal and a maximum value of one (perfect inequality). For the estimated density of wages for 1988, the Gini coefficient is given by

$$G = 1 - 2\int_0^\infty \widehat{F}_1(\omega)\hat{f}_{88}(\omega)\,d\omega \tag{37}$$

where $\widehat{F}_1(\omega)$ is the estimated first moment distribution.

Table 3 presents our decomposition results for various measures. In parentheses underneath each estimate, we present the percentage of the total change explained by each specific factor. From the top half of the table, which presents our estimates for the sample of men, note that the effect of the minimum wage is greatest on those measures that consider the lower part of the distribution. Minimum wages explain 25 of the change in the 10–90 differential, 66 percent of the 10–50 differential, and 49 percent of the change in the 5–95 differential. Minimum wages on the other hand, explain very little of the change in the 50–90 or the 25–75 differential.

Union impacts tend to be important everywhere except in the tails and explain about 14 percent of the change in the standard deviation of log wages of men. The same holds for changes in the distribution of individual attributes, which also explain about 14 percent of the change in the standard deviation of log wages. Supply and demand effects are also quite important for men. These effects explain almost half of the change in the 25–75 differential, and 30 percent of the change in the standard deviation of log wages.

In the final column, we list the unexplained or residual part of the change in differentials. In the differentials we examine, the proportion unexplained ranges from a high of 32 percent for the 50–90 differential to a low of 5.4 percent for the 5–95 differential. The unexplained portion of the change in the standard deviation of log wages is about 10 percent.

It is interesting to note that the importance of the various factors is quite different in explaining changes in Theil's coefficient versus changes in the Gini coefficient for men. For Theil's coefficient, for example, changes in individual attributes explain the greatest proportion of the change over the period—62 percent. On the other hand, the most important factor in explaining changes in the Gini coefficient is changes in supply and demand which explains about 37 percent of the total. This highlights the inadequacy of simple summary measures of the distribution of income to fully capture changes in the multi-modal distributions actually encountered in U.S. data.

The second half of Table 3 displays a similar decomposition for women. Most of the patterns are very similar although not surprisingly changes in the density of unionization generally explain less of the change in wage dispersion than they do for men. For example, changes in union density explain only 3 percent of the change in the standard deviation of log wages of women, compared to 14 percent for men.

In Table 4, we present a more detailed summary of changes in the standard deviation of log wages for two demographic groups: 16–24 years of age and 25–65 years of age. Again the patterns are not too surprising although the effect of changes in the minimum wage on young men is rather large. Also worthy of note is the fact that changes in minimum wages have quite important effects on the wages of more mature women.

5. Supporting Evidence

Having considered the period 1979-1988 in rather extensive detail we now turn to some additional evidence.

In section 4.1 we noted that sequential decomposition results might depend on the order of the decomposition. For example, by proceeding first to remove minimum wage effects and then union effects, we were essentially labeling the interaction between the two as a minimum wage effect. It might then be argued that we run the risk of attributing too much of the change in the distribution of wages to a minimum wage effect. We therefore first consider the consequences of performing our sequential decompositions in "reverse order", hoping to minimize any "over attribution" of wage distribution changes to the minimum wage.

We next consider the 1973–79 and 1988–92 periods. The 1973–79 period is particularly interesting because the minimum wage and, to some extent, the unionization rate went in opposite directions in 1973–79 compared to 1979-88. If our results for 1979–88 are not spurious consequences of the particular period analyzed, the effect of labor market institutions on changes in wage inequality from 1973 to 1979 should be the reverse of what we observe for 1979–88. That is, changes in minimum wages and in unionization rates should have contributed to reduce wage inequality.

Next, we also consider a related piece of evidence for minimum wage effects by comparing its effect in "low" and "high" wage states. Finally, we look at a similarly related piece of evidence for the effect of unionization by comparing the United States to Canada. This is likely to be instructive about the effects of unionization, since by contrast with the U.S., there was little drop in the union density over the 1980s in Canada.

5.1. Reversing the Order of the Decomposition

By beginning our sequential decomposition with minimum wage effects and then union status, we attribute any interaction effect to minimum wages. The extent or existence of this "over attribution", depends on the covariances between being the risk of being affected by the minimum wage or union membership, or any of the other factors we consider. As a consequence, in Table 5 we repeat the analysis of Table 3, this time performing our sequential analysis with union status and minimum wage effects last, and beginning with individual attributes and supply and demand effects.

As before, we consider a variety of different measures. Not surprisingly, perhaps, the results are similar. Considering the standard deviation of log wages for examples, the share of changes explained by minimum wages falls slightly for men, but rises by a similar amount for women. The share of wage changes explained by changes in union density remains virtually unchanged for both men and women. Although the precise percentages explained by the various are factors are not identical when we reverse the order of the decompositions, it is clear that doing so does not qualitatively affect the results.

5.2. *Changes in the Distribution of Wages: 1973–1979 and 1988–1992*

In the 1973–79 period, the real value of the minimum wage rose substantially as did its coverage. Union density showed a slight increase for men and for women. In panel a) of Table 6, we decompose the changes in the standard deviation of log wages for the period 1973–1979. For men, minimum wage changes explain 35 percent of the decline in the standard deviation of log wages, while unions explain 26 percent. The contribution of minimum wages is still important for women, and explains 29 percent of the decline. The change in union density had virtually no impact on the distribution of wages for women. These results thus support our prediction that labor market institutions should have had opposite effects over the 1973–79 and the 1979–88 periods.

The period 1988-1992 is less interesting for this type of analysis since there was virtually no change in union density over the period for men or for women, and minimum wages were 11 percent higher in 1992 than they were in 1988.[20] We present the results for completeness. In panel b) of Table 6 we decompose changes in the standard deviation of log wages from 1988 to 1992 for both men and women. For men, there was virtually no change in the standard deviation of log wages. The results for women in the period actually show a decrease in wage inequality as measured by the standard deviation of log wages with neither unions nor minimum wages explaining much of the effect.

5.3. "High wage" vs. "Low wage" states

As has been noted in recent research, states vary greatly in the average level of wages that prevail.[21] In general, a larger proportion of workers in states with low average wages is likely to be affected by the minimum wage than in states with high average wages. As a consequence we would expect the minimum wage to have a larger effect in "low wage" states than in "high wage" states. We thus divide our sample into "high wage" and "low wage" states.[22]

In Figure 9, we illustrate the distribution of wages in 1979 and 1988 for women in low wage states and women in high wage states.[23] The estimates conform to

[20] By 1988, many states had legislated minimum wage increases above the federal minimum. When appropriate, we thus use these state minimum wages to compute the effect of changes in the minimum wage on changes in the distribution of wages between 1988 and 1992.

[21] See Card (1992c) and Neumark and Wascher (1992) for example.

[22] Low wage states are: Alabama, Arkansas, Florida, Georgia, Mississippi, North Carolina, North Dakota, New Mexico, South Carolina, South Dakota, and Tennessee. High wage states are: Alaska, California, Connecticut, Delaware, District of Columbia, Illinois, Maryland, Michigan, New Jersey, New York, Nevada, Oregon, and Washington.

[23] To obtain similar degrees of smoothness for these different densities, we used the bandwidth that was optimal for the whole sample.

expectations. The compression effect of the minimum wage is more pronounced for women in low wage states and somewhat less pronounced for women in high wage states.

In Table 7, we present a similar comparison using the standard deviation of the log wage. Not surprisingly, the equalizing effect of minimum wages is apparent in the low wage states, where the effect of the minimum wage explains 7 percent of the change for men and 21 percent of the change for women. By contrast, the equalizing effect of the minimum wage is negligible in high wage states.

5.4. The United States vs. Canada

Finally, Figure 10 is a comparison between the United States, where unionization rates declined in the 1980s, and Canada, where the unionization rate remained relatively constant. Because of the limitations of the Canadian data, we were only able to look at changes in the distribution of wages between 1981 and 1988.[24] Table 8 reports the estimated changes in the standard deviation of log wages for men and women in the United States and in Canada over the 1981–88 period. While wage inequality remained relatively constant in Canada over this period, it steeply increased in the United States. The table also indicates that the decline in unionization among U.S. men contributed to the increase in wage inequality. By contrast, the effect of changes in unionization is three times smaller among Canadian men. It is very small for women in both Canada and the United States.

[24] For additional details on the Canadian data set used and on changes in the distribution of wages in Canada, see DiNardo and Lemieux (1993).

6. Conclusions

In this paper, we have proposed and implemented a semi–parametric procedure that allows us to estimate changes in the entire density of wages and decompose these changes into various factors.

As has been documented in previous research, we find that de-unionization and supply and demand shocks were important factors in explaining the rise in wage inequality from 1979 to 1988. We also find, however, that the decline in the real value of the minimum wage from 1979 to 1989 explains a substantial proportion of the increase in wage inequality, particularly for women and for others in the lower tail of the wage distribution. Furthermore, we conclude that labor market institutions are at least as important as supply and demand considerations in explaining changes in the U.S. distribution of wages from 1973 to 1992.

It is interesting to note that our findings about the importance of labor market institutions are in concordance with analyses that were common before the advent of widespread use of "marginal productivity analysis" as it was called, and the focus on "economic" factors such as supply and demand in empirical labor economics. In Lester's analysis of the wage structure for example (Lester 1964), supply and demand played an important but secondary role to such factors as unionization and changes in the minimum wage. Our findings suggest that this earlier emphasis was not misplaced.

Appendix

This appendix describes in detail the approach we use to construct the measures of demand and supply for various categories of workers. It also explains how we use these measures to estimate the impact of changes in supply and demand on the mean of the distribution of wages other than the minimum wage, $\Delta\widehat{w}_j$, and on the probability of being paid the minimum wage, $\Delta\widehat{P}_j t$.

Bound and Johnson (1992) provide the point of departure for our investigation. Like them, we divide our data into 32 experience–education–gender cells based on four values of completed years of schooling S, (dropouts: $S < 12$, high school: $S = 12$, some college: $12 < S < 16$; and college: $S \geq 16$), four levels of potential labor market experience, X, (0–9, 10–19, 20–29, and 30+ years), and two sexes.

We first estimate the effect of supply and demand changes on the mean of the distribution of wage, for wages other than the minimum wage, for each of the 32 cells . We will later explain how to estimate the effect of supply and demand on the probability of being paid the minimum wage. Let $z = (z_1, z_2)$, where z_1 denotes the industry and union status, and z_2 is a vector of other attributes. For each experience–education–gender class j, we first calculate a measure of the change in the cell mean log wage which is purged of changes in individual attributes and changes in industry or union premiums. These "corrected" mean cell wages are computed by running a micro regression of log wages on z_1, z_2, and dummy variables for the 32 experience–education–gender classes. We also fully interact z_1 and z_2 with a gender dummy. These wage regressions are estimated on the sample of workers paid a wage other than the minimum wage.

For each experience–education–gender class, we then compute a predicted wage for a representative person with a vector of attributes z_2 equal to the mean cell attributes in a base period $\bar{z}_{2j}$, and with the average union status and industry affiliation of the entire sample $\bar{z}_1$ in a base period. The corrected changes in mean cell wages are then

$$\Delta w_j = \widehat{w}_{88}(\bar{z}_1, \bar{z}_{2j}) - \widehat{w}_{79}(\bar{z}_1, \bar{z}_{2j}). \tag{A1}$$

where $\widehat{w}_t(\bar{z}_1, \bar{z}_{2j})$ is the cell j predicted wage from a period t wage regression.

Following the model of competitive wage determination of Bound and Johnson, these changes are modeled as a linear function of changes in supply and demand conditions in each cell

$$\Delta w_j = (1/\sigma)\Delta \ln D_j - (1/\sigma)\Delta \ln N_j + (1 - 1/\sigma)\Delta \ln b_j, \tag{A2}$$

where $\Delta \ln N_j$ is the change in the supply of a given type of labor, $\Delta \ln D_j$ is the change in demand, $\Delta \ln b_j$ is a technological shock (the residual error in the equation), and σ is the elasticity of intrafactor substitution, which is assumed to be constant across industries. The supply shock $\Delta \ln N_j$ is merely the change in the log of the number of observations in cell j

$$\Delta \ln N_j = \ln N_{j88} - \ln N_{j79}, \tag{A3}$$

where N_{jt} denotes the number of persons in experience–education–gender cell j at time t.

Earlier research has used a "fixed–coefficient manpower requirements index" to measure between–sector shifts in relative labor demands (see Freeman (1980a) and Katz and Murphy (1992) for a formal development). Letting N_{kjt} be the number of persons of experience–education–gender class j in industry k at time t, one measure of demand is:

$$\Delta \ln D_j = \sum_k \phi_{kj}\, \Delta \ln N_k, \tag{A4}$$

where $\phi_{kj} = N_{kj79}/N_{j79}$ and $\Delta \ln N_k = \ln N_{k88} - \ln N_{k79}$ is a measure of changes in industry i's workforce. Bound and Johnson (1992) note that using $\Delta \ln N_k$ as a proxy for product demand shifts may confound these shifts with relative-supply changes. As a consequence we follow the modification in the appendix to their paper (their equations (A10) to (A12)) and use a measure that nets–out supply shifts.

The contribution of supply and demand shocks to changes in the structure of wages is estimated by fitting equation (A2) with ordinary least squares. The estimated coefficients are then used to construct a measure of predicted change in estimated mean cell wages, $\Delta\widehat{w}_j$, attributable to changes in supply and demand conditions between 1979 and 1988.

We next run grouped data logits to estimate the effect of supply and demand factors on the probability of being paid the minimum wage $(\Delta\widehat{P}_j)$. For each cell j, we model the log odds of being paid the minimum wage as:

$$\ln\left[\frac{\text{Prob}(w = m_t|z, d_{jt}, n_{jt})}{(1 - \text{Prob}(w = m_t|z, d_{jt}, n_{jt}))}\right] = \gamma_t + \alpha_j + \beta_1 d_{jt} + \beta_2 n_{jt} + \epsilon_{jt}, \qquad \text{(A5)}$$

where $d_{jt} = \ln D_{j,t}$ and $n_{jt} = \ln N_{j,t}$. In this model, we let the probability of being paid the minimum wage arbitrarily vary across groups of workers by introducing a group fixed effect α_j. Other than the time effect γ_t, other systematic changes over time in the log odds of being paid the minimum wage depend on changes in supply and demand conditions.

The group fixed effect is eliminated by taking first differences of equation (A5):

$$\Delta\ln\left[\frac{\text{Prob}(w = m_t|z, d_{jt}, n_{jt})}{(1 - \text{Prob}(w = m_t|z, d_{jt}, n_{jt}))}\right] = \Delta\gamma_t + \beta_1\Delta d_{jt} + \beta_2\Delta n_{jt} + \Delta\epsilon_{jt}, \qquad \text{(A6)}$$

The weighted least squares estimates of the parameters in equation (A6) can then be used to predict changes in the log odds of being paid the minimum wage attributable to demand and supply shocks. It is easily shown that the predicted effect of demand and supply on the probability of being paid the minimum wage is given by:

$$\Delta\widehat{P}_j = \text{Prob}(w = m_t|z, d_{jt}, n_{jt}) - e^K/(1 + e^K), \qquad \text{(A7)}$$

where

$$K = \ln[\text{Prob}(w = m_t|z, d_{jt}, n_{jt})/(1 - \text{Prob}(w = m_t|z, d_{jt}, n_{jt}))] - \beta_1\Delta d_{jt} - \beta_2\Delta n_{jt}. \qquad \text{(A8)}$$

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Table 1 – Sample Means from the *Current Population Survey* 1973–1992

Year	Minimum Real Log Wage[a]	Real Log Wage	Union[b]	Female	Nonwhite	Education	Experience	Number of Observations
1973	0.926	1.680 (0.534)	0.240	0.409	0.109	11.93 (2.92)	19.43 (14.13)	37698
1974	1.051	1.659 (0.528)	0.238	0.411	0.110	12.05 (2.89)	18.98 (13.99)	36087
1975	1.022	1.664 (0.530)	0.225	0.416	0.111	12.18 (2.89)	19.01 (13.95)	36182
1976	1.055	(1.666 (0.525)	0.225	0.422	0.110	12.23 (2.85)	18.43 (13.76)	36587
1977	0.989	1.660 (0.525)	0.241	0.428	0.110	12.32 (2.86)	18.32 (13.71)	44989
1978	1.060	1.657 (0.517)	0.235	0.434	0.114	12.35 (2.82)	17.96 (13.56)	43007
1979	1.065	1.662 (0.507)	0.251	0.444	0.124	12.44 (2.79)	17.64 (13.57)	140284
1980	1.033	1.651 (0.505)		0.452	0.123	12.52 (2.77)	17.65 (13.41)	160534
1981	1.027	1.648 (0.513)	0.222	0.455	0.124	12.57 (2.75)	17.65 (13.21)	155129
1982	0.974	1.654 (0.527)		0.461	0.123	12.69 (2.74)	17.73 (13.06)	148441
1983	0.933	1.648 (0.541)	0.203	0.464	0.125	12.77 (2.72)	17.62 (12.95)	146272
1984	0.897	1.651 (0.551)	0.191	0.463	0.128	12.81 (2.69)	17.49 (12.73)	147625
1985	0.860	1.654 (0.559)	0.182	0.466	0.130	12.84 (2.69)	17.60 (12.63)	150561
1986	0.830	1.658 (0.562)	0.179	0.467	0.134	12.88 (2.67)	17.71 (12.53)	156651
1987	0.790	1.650 (0.565)	0.174	0.471	0.135	12.90 (2.68)	17.66 (12.36)	152424
1988	0.748	1.645 (0.567)	0.171	0.473	0.136	12.93 (2.71)	17.70 (12.27)	144750
1989	0.699	1.645 (0.569)	0.167	0.470	0.142	12.99 (2.70)	18.19 (12.28)	167863
1990	0.774	1.645 (0.570)	0.163	0.473	0.142	13.03 (2.70)	18.40 (12.16)	175337
1991	0.843	1.642 (0.567)	0.163	0.475	0.143	13.08 (2.69)	18.64 (12.05)	169327
1992	0.813	1.641 (0.565)	0.161	0.477	0.145	13.09 (2.52)	18.78 (11.92)	167244

Note: Standard deviations in parentheses.

[a] 1979 Constant Dollars

[b] Union coverage is not available in 1980 or 1982.

Table 2 – Proportion Unionized by Gender and by Sector[a]

Year	Men	Women	Private	Public
1973	0.308	0.142	0.243	0.227
1974	0.306	0.140	0.236	0.246
1975	0.287	0.139	0.219	0.251
1976	0.288	0.138	0.216	0.259
1977	0.302	0.160	0.218	0.336
1978	0.294	0.158	0.210	0.343
1979	0.317	0.170	0.222	0.383
1981	0.283	0.149	0.194	0.352
1983	0.250	0.149	0.166	0.373
1984	0.235	0.141	0.155	0.363
1985	0.224	0.135	0.145	0.363
1986	0.220	0.132	0.141	0.366
1987	0.214	0.129	0.135	0.364
1988	0.209	0.129	0.129	0.373
1989	0.201	0.128	0.125	0.370
1990	0.196	0.127	0.121	0.367
1991	0.196	0.128	0.119	0.370
1992	0.190	0.129	0.116	0.368

[a] Union coverage is not available in 1980 or 1982.

Table 3 – Decomposing Changes in Measures of Wage Dispersion 1979–1988

Statistic	Total Change	Effect of: Minimum Wage	Unions	Individual[a] Attributes	Supply & Demand	Unexplained[b] Change
Men:						
Standard[c] Deviation	0.072	0.018 (24.8)	0.010 (14.3)	0.010 (14.3)	0.023 (32.6)	0.010 (14.1)
10-90[d]	0.195	0.049 (25.3)	0.021 (10.7)	0.040 (20.7)	0.065 (33.5)	0.019 (9.9)
10-50	0.076	0.050 (65.7)	-0.019 (-25.6)	0.038 (49.7)	0.026 (34.8)	-0.019 (-24.6)
50-90	0.119	-0.000 (-0.4)	0.040 (33.7)	0.003 (2.3)	0.039 (32.6)	0.038 (31.8)
25-75	0.109	-0.001 (-0.6)	0.031 (28.7)	-0.000 (-0.0)	0.050 (46.3)	0.028 (25.7)
5-95	0.290	0.141 (48.6)	0.025 (8.7)	0.051 (17.5)	0.057 (19.8)	0.016 (5.4)
Theil's Coefficient	0.113	0.038 (34.0)	0.003 (3.0)	0.070 (61.6)	0.021 (18.6)	-0.020 (-17.3)
Gini Coefficient	0.041	0.005 (11.3)	0.009 (21.1)	0.003 (7.9)	0.015 (36.9)	0.009 (22.1)
Women:						
Standard[c] Deviation	0.090	0.027 (30.2)	0.003 (3.2)	0.023 (25.9)	0.014 (15.8)	0.023 (25.0)
10-90[d]	0.328	0.148 (45.1)	0.004 (1.3)	0.084 (25.6)	0.060 (18.3)	0.032 (9.7)
10-50	0.243	0.150 (61.7)	-0.010 (-4.1)	0.078 (32.1)	0.026 (10.6)	-0.001 (-0.0)
50-90	0.085	-0.002 (-2.5)	0.014 (16.9)	0.006 (7.0)	0.034 (40.4)	0.032 (38.2)
25-75	0.146	0.011 (7.4)	0.001 (0.7)	0.049 (33.7)	0.049 (33.8)	0.035 (24.4)
5-95	0.380	0.169 (44.3)	0.008 (2.2)	0.083 (21.9)	0.005 (1.3)	0.115 (30.3)
Theil's Coefficient	0.302	0.078 (25.9)	-0.008 (-2.8)	0.148 (48.9)	0.069 (22.7)	0.016 (5.1)
Gini Coefficient	0.049	0.011 (23.3)	0.003 (5.1)	0.012 (23.7)	0.011 (22.9)	0.012 (24.9)

Note. Percent of total variation explained in parenthesis.

[a] The individual attributes are experience, experience squared, education, SMSA, marital status, full–time or part–time, 3 occupational and 19 industry classes.

[b] "Unexplained" is the residual not accounted for by all other factors. See text for further explanation.

[c] Standard deviation of log wage distribution.

[d] Difference between the 90th and the 10th percentiles of the log wage distribution. The 10-50, 50-90, 25-75 and 5-95 statistics are similarly defined.

Table 4 – Decomposition of Changes in Standard Deviations of Log Wages by Demographic Groups 1979–1988

Age Group	Total Change	Effect of : Minimum Wage	Unions	Individual[a] Attributes	Supply & Demand	Unexplained[b] Change
Men:						
16-24	0.0146	0.0460	-0.0107	-0.0014	-0.0099	-0.0095
		(315.1)	(-73.3)	(-9.6)	(-67.8)	(-65.1)
25-65	0.0659	0.0069	0.0141	0.0132	0.0177	0.0142
		(10.5)	(21.4)	(20.0)	(26.9)	(21.5)
Women:						
16-24	0.0691	0.0353	-0.0026	0.0125	-0.0094	0.0332
		(51.1)	(-3.8)	(18.1)	(-13.6)	(48.0)
25-65	0.0774	0.0225	0.0042	0.0208	0.0169	0.0131
		(29.1)	(5.4)	(26.9)	(21.8)	(16.9)

Note. Percent of total variation explained in parenthesis.

[a] The individual attributes are experience, experience squared, education, SMSA, marital status, full–time or part–time, 10 occupational and 10 industry classes.

[b] "Unexplained" is the residual not accounted for by all other factors. See text for further explanation.

Table 6 – Decomposition of Changes in Standard Deviations of Log Wages 1973-1979 and 1988–1992

Gender	Total Change	Effect of : Minimum Wage	Unions
a) 1973–1979:			
Men	-0.0130	-0.0045	-0.0034
		(34.6)	(26.2)
Women	-0.0322	-0.0092	0.0006
		(28.6)	(-1.9)
b)1988–1992:			
Men	0.0038	0.0002	0.0027
		(4.0)	(70.2)
Women	0.0106	0.0022	0.0009
		(20.7)	(8.4)

Note: Percent of total variation explained in parenthesis.

Table 5 – Reversing the Order of the Decomposition 1979–1988

Statistic	Total Change	Effect of: Individual[a] Attributes	Supply & Demand	Unions	Minimum Wage	Unexplained[b] Change
Men:						
Standard[c] Deviation	0.072	0.014 (20.2)	0.025 (34.5)	0.010 (14.2)	0.013 (18.8)	0.009 (12.3)
10-90[d]	0.195	0.041 (21.0)	0.074 (38.0)	0.030 (15.5)	0.024 (12.2)	0.026 (13.5)
10-50	0.076	0.029 (37.9)	0.038 (50.4)	-0.009 (-11.8)	0.024 (32.1)	-0.006 (-8.6)
50-90	0.119	0.012 (10.1)	0.036 (30.0)	0.039 (32.9)	-0.001 (-0.5)	0.033 (32.7)
25-75	0.109	0.016 (15.1)	0.035 (32.0)	0.026 (24.0)	-0.001 (-0.6)	0.032 (29.6)
5-95	0.290	0.053 (18.4)	0.097 (33.6)	0.026 (9.1)	0.103 (35.5)	0.009 (3.4)
Theil's Coefficient	0.113	0.073 (64.9)	0.031 (27.6)	-0.003 (-3.0)	0.034 (30.5)	-0.022 (-20.0)
Gini Coefficient	0.041	0.006 (14.6)	0.014 (34.8)	0.008 (20.3)	0.004 (8.9)	0.008 (7.7)
Women:						
Standard[c] Deviation	0.090	0.021 (23.2)	0.015 (16.2)	0.003 (2.8)	0.033 (36.1)	0.020 (21.7)
10-90[d]	0.328	0.061 (18.7)	0.044 (13.5)	0.007 (2.0)	0.176 (53.5)	0.040 (12.2)
10-50	0.243	0.050 (20.7)	0.022 (8.8)	-0.003 (-1.1)	0.178 (73.3)	-0.004 (-1.7)
50-90	0.085	0.011 (12.8)	0.023 (26.9)	0.009 (10.8)	-0.002 (-2.8)	0.044 (52.2)
25-75	0.146	0.035 (23.7)	0.024 (16.6)	0.002 (1.2)	0.036 (24.8)	0.048 (33.5)
5-95	0.380	0.079 (20.6)	0.057 (14.9)	0.008 (2.3)	0.166 (43.6)	0.070 (18.5)
Theil's Coefficient	0.302	0.125 (41.4)	0.068 (22.4)	-0.004 (-1.3)	0.107 (35.5)	0.006 (2.0)
Gini Coefficient	0.049	0.009 (18.6)	0.009 (17.4)	0.002 (4.2)	0.016 (32.5)	0.013 (27.2)

Note. Percent of total variation explained in parenthesis.

[a] The individual attributes are experience, experience squared, education, SMSA, marital status, full–time or part–time, 3 occupational and 19 industry classes.

[b] "Unexplained" is the residual not accounted for by all other factors. See text for further explanation.

[c] Standard deviation of log wage distribution.

[d] Difference between the 90th and the 10th percentiles of the log wage distribution. The 10-50, 50-90, 25-75 and 5-95 statistics are similarly defined.

Table 7 – Effect of Changes in the Minimum Wage on Changes in the Standard Deviation of Log Wages in High Wage and Low Wage States[a] 1979-1988

	Total Change	Effect of Minimum Wage
Low Wage States:		
Men	0.0539	0.0039 (7.2)
Women	0.1041	0.0222 (21.3)
High Wage States:		
Men	0.0737	-0.0013 (-1.8)
Women	0.0824	0.0021 (2.6)

Note. Percent of total variation explained in parenthesis.

[a]The low wage states are AL, AR, FL, GA, MS, NC, ND, SC, SD. The high wage states are AK, CA, CN, DC, DL, IL, MI, NJ, NY, NV, OR, WA. See text for details.

Table 8 – Effect of Changes in Unionization on Changes in the Standard Deviation of Log Wages in the United States and Canada 1981-1988

	Total Change	Effect of Unions
United States:		
Men	0.0631	0.0090 (14.3)
Women	0.0912	0.0013 (1.4)
Canada:		
Men	-0.0002	0.0029 (—)
Women	0.0053	0.0013 (24.5)

Note. Percent of total variation explained in parenthesis.

Table A1 – Measures of Wage Inequality

Year	Standard Deviation of Log Wage	Percentiles of Log Wage Distribution: 10-90	10-50	50-90	25-75	5-95	Theil's Entropy Coefficient	Gini Coefficient
				a) Men				
1973	0.516	1.269	0.674	0.595	0.625	1.630	2.575	0.274
1974	0.514	1.253	0.667	0.586	0.627	1.604	2.471	0.274
1975	0.520	1.293	0.697	0.596	0.654	1.643	2.575	0.279
1976	0.517	1.292	0.685	0.606	0.662	1.616	2.566	0.277
1977	0.519	1.289	0.704	0.585	0.676	1.626	2.572	0.275
1978	0.509	1.290	0.700	0.590	0.685	1.594	2.561	0.275
1979	0.501	1.277	0.692	0.585	0.682	1.572	2.546	0.270
1980	0.508	1.284	0.702	0.583	0.695	1.598	2.555	0.276
1981	0.519	1.313	0.715	0.598	0.717	1.621	2.566	0.282
1982	0.537	1.365	0.742	0.623	0.746	1.693	2.604	0.290
1983	0.554	1.414	0.762	0.652	0.770	1.755	2.628	0.301
1984	0.561	1.435	0.770	0.665	0.769	1.792	2.646	0.305
1985	0.568	1.451	0.780	0.671	0.777	1.817	2.661	0.308
1986	0.570	1.463	0.790	0.673	0.784	1.829	2.665	0.306
1987	0.572	1.465	0.777	0.687	0.785	1.843	2.658	0.307
1988	0.576	1.471	0.767	0.704	0.791	1.862	2.659	0.311
1989	0.575	1.462	0.762	0.700	0.791	1.848	2.650	0.309
1990	0.579	1.472	0.756	0.717	0.791	1.849	2.653	0.315
1991	0.576	1.478	0.755	0.723	0.795	1.836	2.637	0.316
1992	0.577	1.491	0.762	0.729	0.809	1.841	2.654	0.316
				b) Women				
1973	0.460	1.062	0.492	0.569	0.592	1.369	2.049	0.250
1974	0.449	1.009	0.446	0.564	0.578	1.340	1.992	0.243
1975	0.452	1.044	0.467	0.577	0.594	1.328	2.025	0.249
1976	0.450	1.021	0.434	0.587	0.586	1.310	2.017	0.246
1977	0.446	1.039	0.446	0.593	0.591	1.305	1.949	0.248
1978	0.437	0.991	0.402	0.589	0.576	1.264	1.980	0.241
1979	0.429	0.984	0.394	0.590	0.573	1.241	1.962	0.238
1980	0.422	1.007	0.417	0.590	0.582	1.245	1.969	0.239
1981	0.432	1.045	0.431	0.614	0.601	1.281	2.004	0.245
1982	0.450	1.104	0.482	0.623	0.638	1.345	2.080	0.255
1983	0.467	1.154	0.521	0.633	0.659	1.396	2.123	0.264
1984	0.482	1.202	0.557	0.645	0.678	1.456	2.163	0.272
1985	0.494	1.242	0.588	0.654	0.696	1.499	2.201	0.278
1986	0.502	1.268	0.613	0.655	0.706	1.538	2.225	0.279
1987	0.511	1.295	0.629	0.665	0.713	1.583	2.249	0.284
1988	0.516	1.312	0.637	0.675	0.718	1.621	2.264	0.287
1989	0.523	1.324	0.638	0.685	0.728	1.649	2.281	0.289
1990	0.524	1.327	0.630	0.697	0.727	1.640	2.287	0.291
1991	0.523	1.318	0.613	0.704	0.726	1.616	2.287	0.292
1992	0.526	1.331	0.618	0.712	0.731	1.627	2.324	0.294

Table A2 – Sensitivity of Changes in Measures of Log Wage Dispersion to Changes in the Level of the Minimum Wage

Gender	Statistic	Total Change	Effect of : Minimum Wage	Unions
a) 1979-1988 with Minimum Wage= $3.00				
Men	Standard Deviation	0.072	0.018 (24.8)	0.010 (14.3)
	10-50	0.076	0.050 (65.7)	-0.019 (-25.6)
	50-90	0.119	-0.000 (-0.4)	0.040 (33.7)
Women	Standard Deviation	0.090	0.027 (30.2)	0.003 (3.2)
	10-50	0.243	0.150 (61.7)	-0.010 (-4.1)
	50-90	0.085	-0.002 (-2.5)	0.014 (16.9)
b) 1979-1988 with Minimum Wage= $2.90				
Men	Standard Deviation	0.072	0.007 (9.7)	0.012 (16.2)
	10-50	0.076	0.033 (43.7)	-0.017 (-22.4)
	50-90	0.114	-0.001 (-1.3)	0.042 (35.3)
Women	Standard Deviation	0.090	0.0136 (15.1)	0.003 (3.8)
	10-50	0.243	0.118 (48.7)	-0.009 (-3.7)
	50-90	0.085	-0.003 (-3.3)	0.014 (16.9)
c) 1978-1988 with Minimum Wage= $2.65				
Men	Standard Deviation	0.064	0.008 (13.0)	0.006 (9.3)
	10-50	0.067	0.032 (47.6)	-0.011 (-16.8)
	50-90	0.114	0.001 (0.9)	0.022 (19.2)
Women	Standard Deviation	0.085	0.020 (23.4)	0.002 (2.6)
	10-50	0.235	0.127 (54.0)	-0.005 (-2.3)
	50-90	0.086	0.002 (2.3)	0.007 (8.0)

Note. Percent of total variation explained in parenthesis.

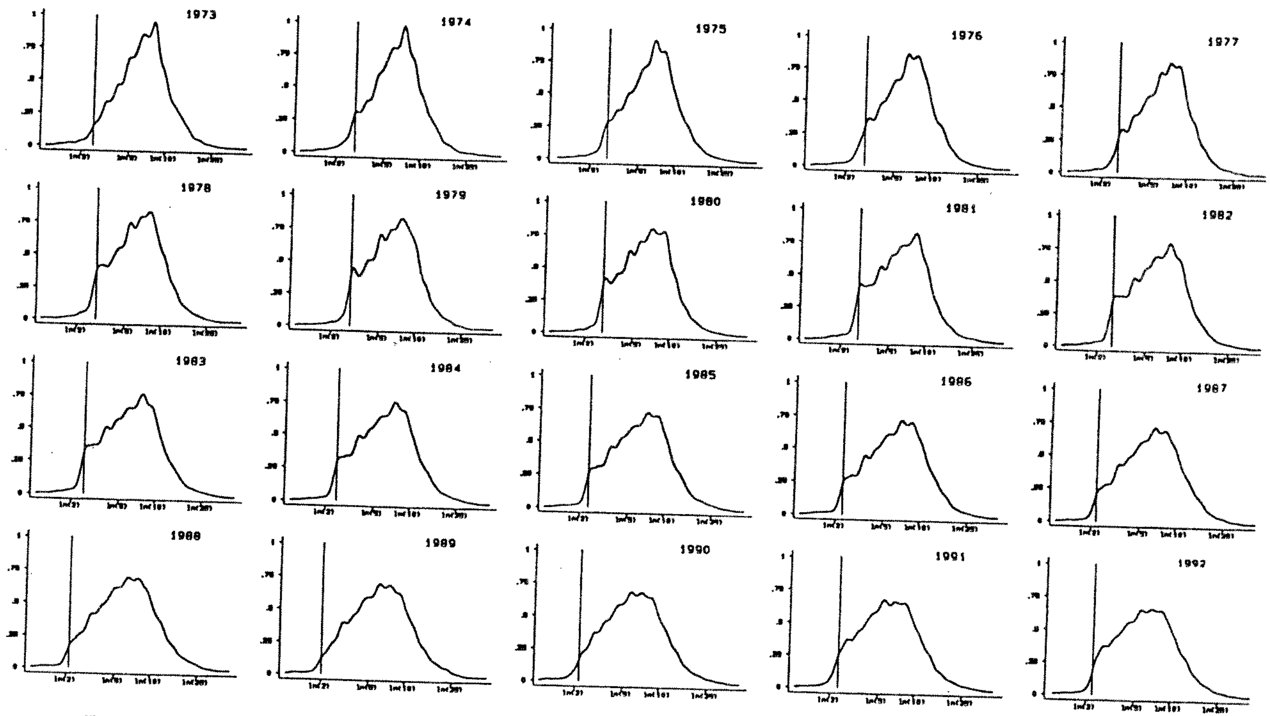


Figure 1a. Kernel Density Estimates of Men's Real Log Wages 1973-1992 ($1979)

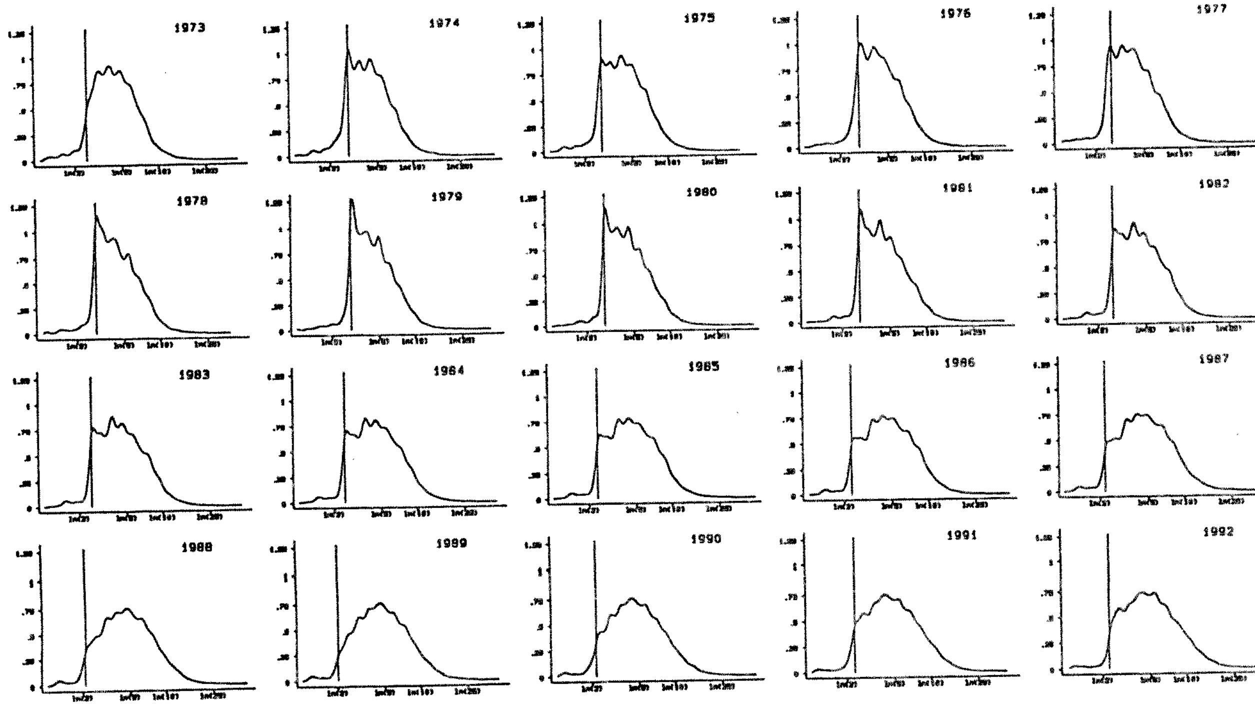
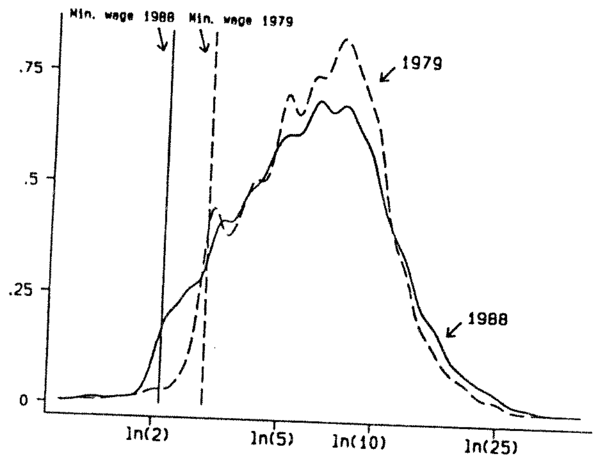
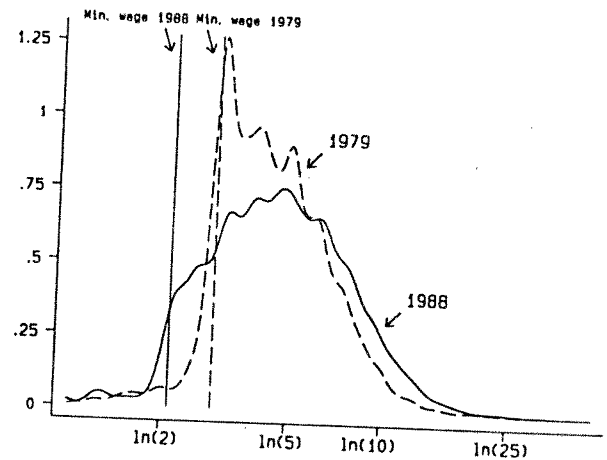


Figure 1b. Kernel Density Estimates of Women's Real Log Wages 1973-1992 ($1979)

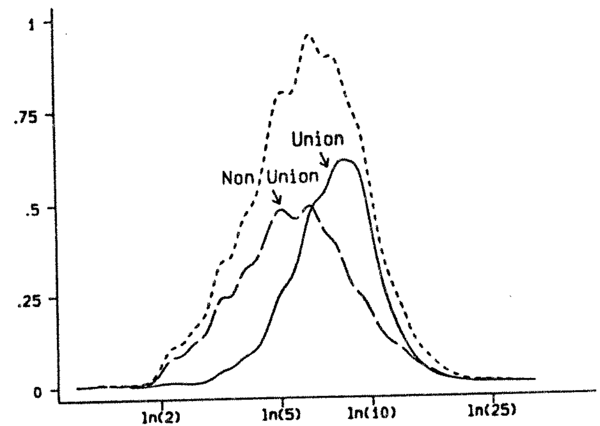


a) Men

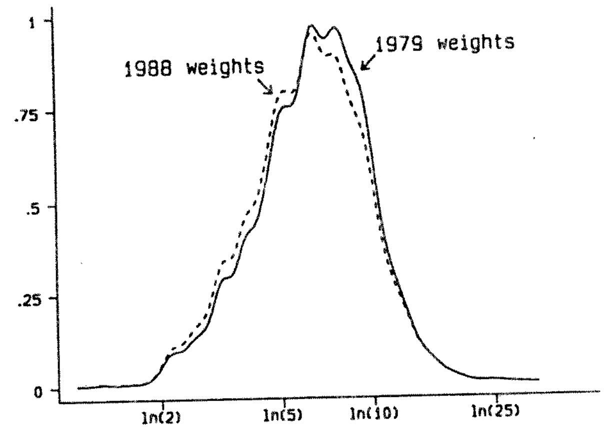


b) Women

Figure 2. Kernel Density Estimates of Real Log Wages ($1979) in 1979 and 1988

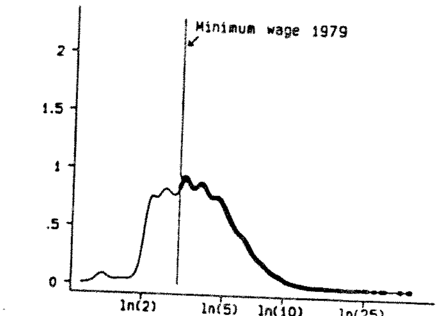


a) Densities from the union and non-union sector

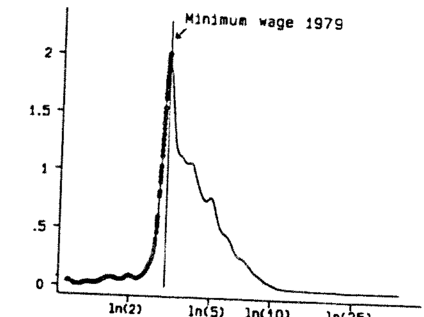


b) Weigthed sum of the union and non-union densities

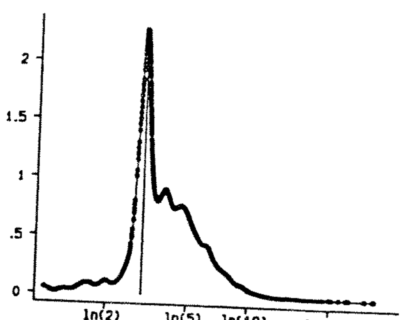
Figure 3. An Illustration of the Estimation of the Effect of Unions



a) Density Estimate of Real Log Wages in 1988



b) Density Estimate of Real Log Wages in 1979



c) Density Estimate of Real Log Wages in 1988
Adjusted for 1979's Minimum Wage

Figure 4. An Illustration of the Estimation of the Effect of the Minimum Wage

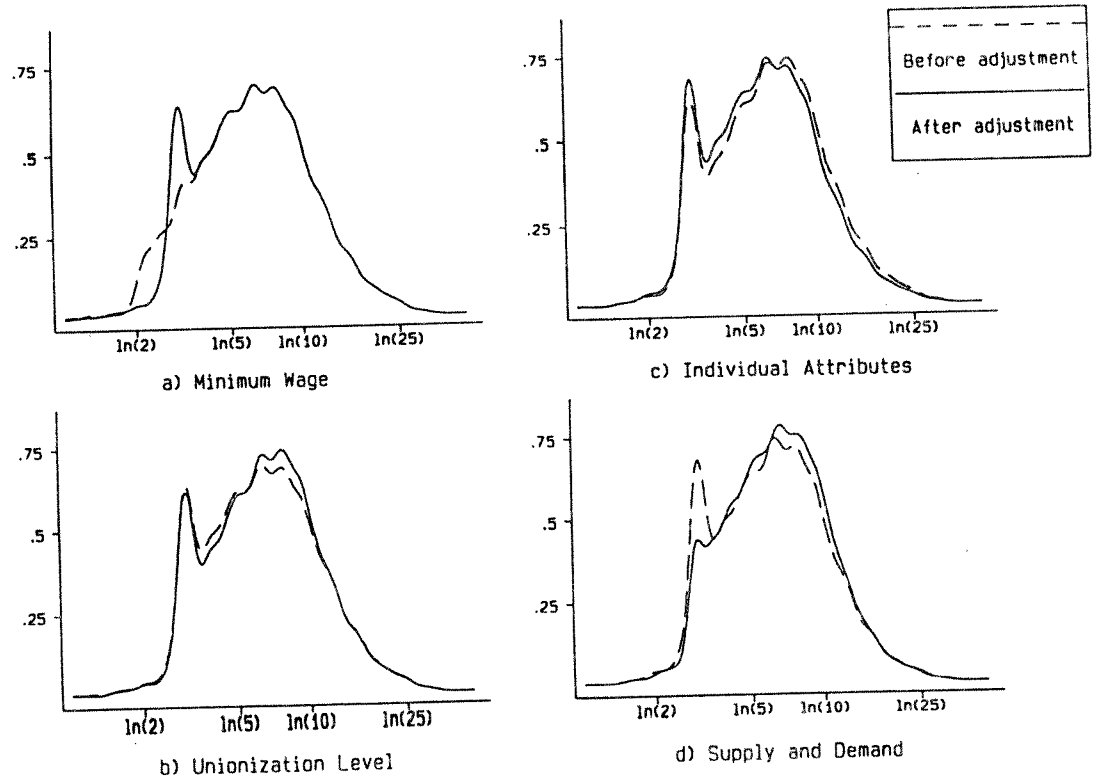


Figure 5. Effects of Indicated Changes on the 1988 Density of Men's Real Log Wages ($1979)

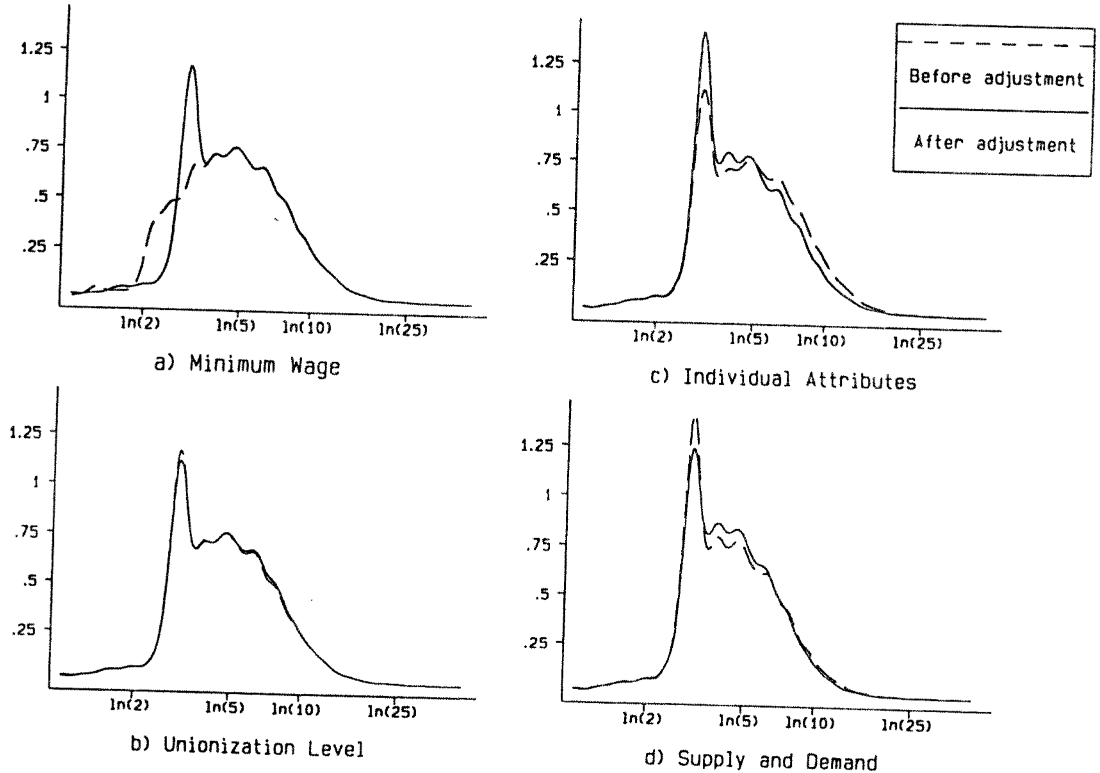
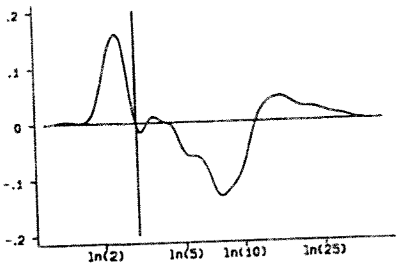
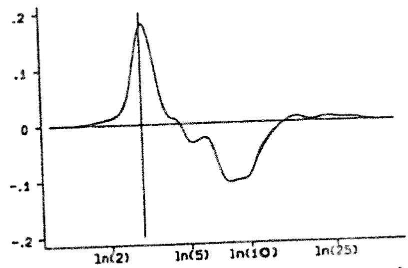


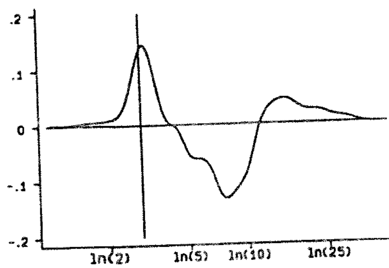
Figure 6. Effects of Indicated Changes on the 1988 Density of Women's Real Log Wages ($1979)



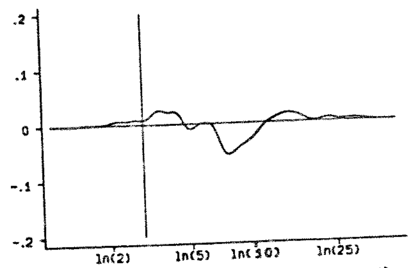
a) Actual Difference



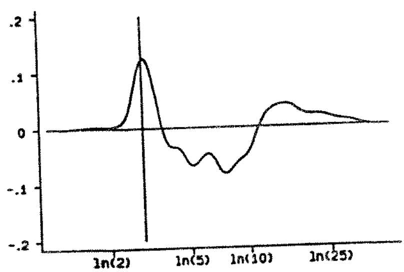
d) Individual Attributes and c)



b) Minimum Wage

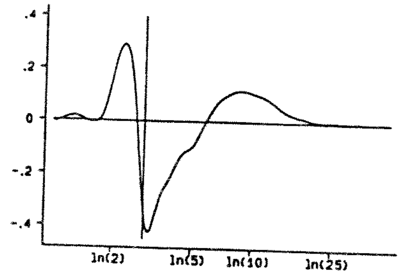


e) Supply and Demand and d)

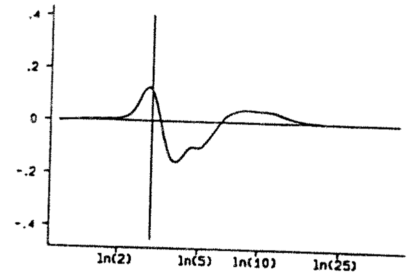


c) Unionization Level and b)

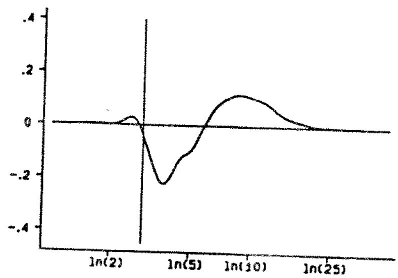
Figure 7. Smoothed Differences between Men's 1979 and 1988 densities, adjusting for



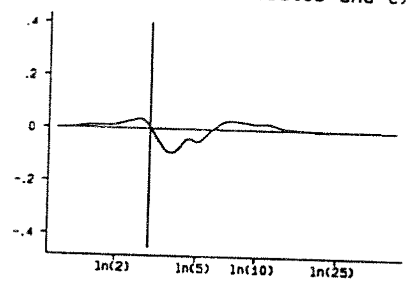
a) Actual Difference



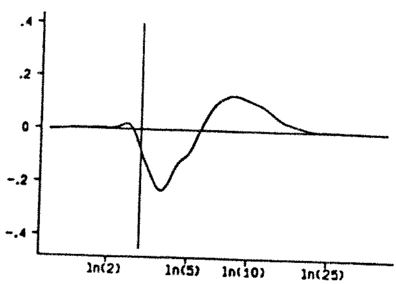
d) Individual Attributes and c)



b) Minimum Wage



e) Supply and Demand and d)



c) Unionization Level and b)

Figure 8. Smoothed Differences between Women's 1979 and 1988 densities, adjusting for

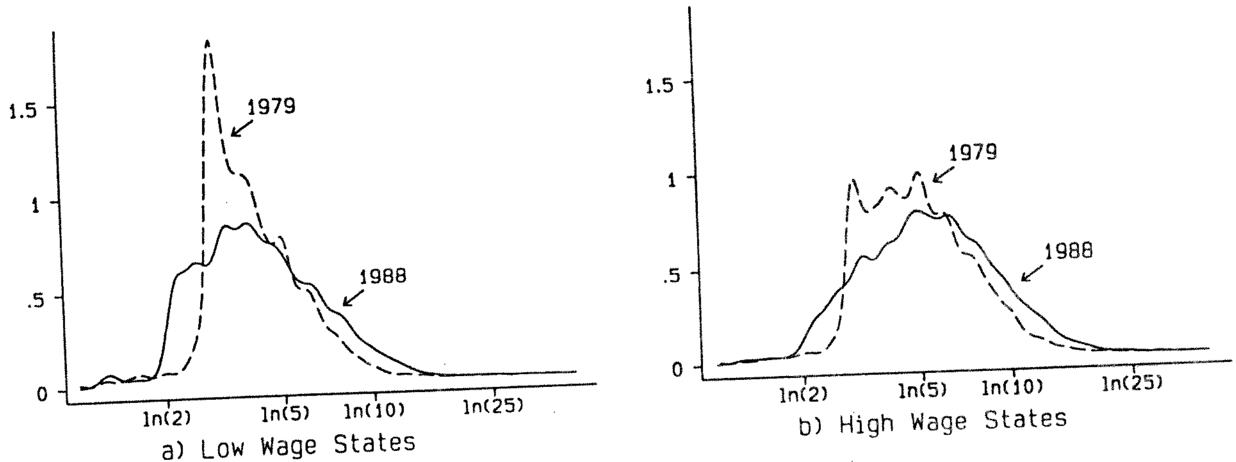


Figure 9. Density Estimates of Women's Real Log Wages ($1979) in Low Wage and High Wage States

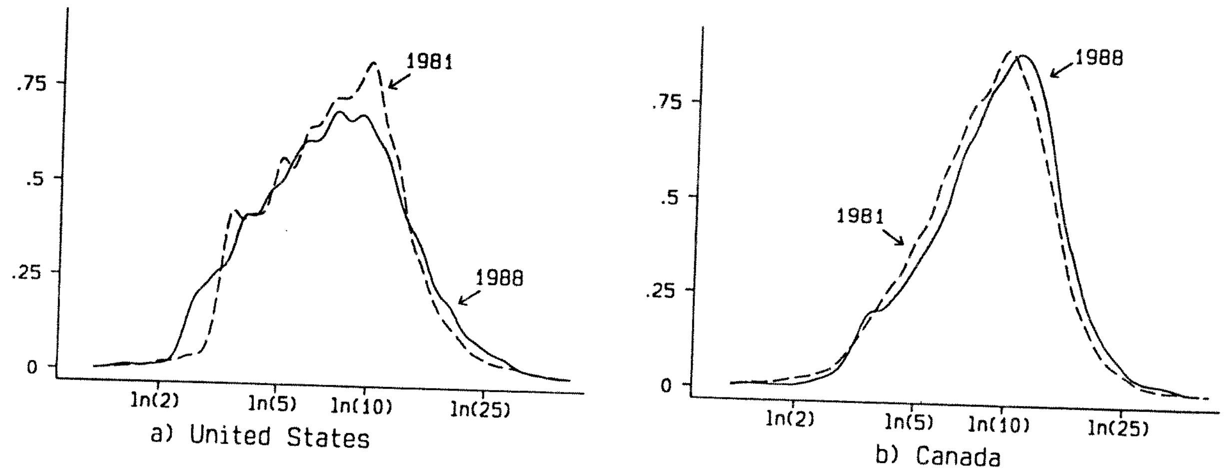


Figure 10. Density Estimates of Men's Real Log Wages ($1981)
in the United States and Canada

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