DIVERGING MALE WAGE INEQUALITY IN THE UNITED STATES AND CANADA, 1981–1988: DO INSTITUTIONS EXPLAIN THE DIFFERENCE?

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The U.S. and Canadian economies have much in common, including similar collective bargaining structures. During the period 1981–88, however, although both countries witnessed a decline in the percentage of workers belonging to unions and an increase in hourly wage inequality, those changes were much more pronounced in the United States than in Canada. Using data on men in Canada and the United States in 1981 and 1988 (from the Labour Force Survey and supplements to the Current Population Survey), the authors study the effect of labor market institutions on changes in wage inequality by computing simple counterfactuals such as the distribution of wages that would prevail if all workers were paid according to the observed nonunion wage schedule. Their results suggest that much more severe declines in the unionization rate in the United States than in Canada account for two-thirds of the differential growth in wage inequality between the two countries.

In a burgeoning literature, researchers have attempted to provide explanations for changes in the structure of wages, particularly in the United States.[1] A large increase in wage inequality in the United States in the 1980s has been so well documented by recent research that it has attained the rare status of a "fact" (Freeman and Katz 1994). Education and age differentials increased, as well as wage dispersion within narrow demographic and skill groups.

The U.S. experience during the 1980s, though not unique, is not shared by all

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[1]See, for example, Blackburn, Bloom, and Freeman (1992); Bound and Johnson (1992); Goldin and Margo (1992); Katz and Murphy (1992); Levy and Murnane (1992); and Murphy and Welch (1992).

The data used in this paper are from the public use files of the May 1981 Current Population Survey (CPS), the 1981 and 1988 Outgoing Rotation Group Supplements of the CPS, the 1981 Canadian Survey of Work History, and the 1988 Canadian Labour Market Activity Survey. These data are publicly available. The computer programs used for data extraction and analysis are available upon request to the authors.

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OECD countries for which detailed micro data are available. While the United Kingdom also saw dramatic increases in wage inequality during the 1980s, Japanese wage inequality grew only modestly, and French inequality grew even less (Katz, Loveman, and Blanchflower 1995).

Most of this research has attempted to explain changes in the wage structure by changes in the supply and demand for different skill categories of workers. On the other hand, the role of labor market institutions-in particular, the structure of collective bargaining agreements and the level of unionization-has gone relatively unexamined. Recent research suggests, however, that between 10% and 20% of the increase in wage inequality among men in the United States and the United Kingdom can be explained by the decline in unionization in these two countries.[2] There is also indirect evidence that cross-country differences in unionization play an important role in cross-country differences in the level of wage inequality.[3] In addition, recent research by DiNardo, Fortin, and Lemieux (1996) (henceforth DFL) suggests that another institution, the minimum wage, played a major role in the recent changes in U.S. wage inequality.

Using comparable micro data on men in the United States and Canada in 1981 and 1988, in this paper we obtain direct evidence on how relative changes in unionization are linked to relative changes in the distribution of wages. We also look at the impact of the minimum wage on changes in wage inequality. Extending techniques described in DFL and applying them in a "comparative" setting, we document the effect of these institutional forces on the entire distribution of wages instead of focusing on a few summary measures of wage inequality like the variance of log wages or the rate of return to education.

By comparing two countries at two points in time, we implicitly control for common underlying changes in wage inequality and for intrinsic differences in the level of wage inequality in the two countries. Our estimates, therefore, are less likely to reflect spurious correlations than estimates based either on cross-country comparisons at a point in time or on comparisons over time in a single country.

Micro Data on Wages and Union Status

Our analysis is based on data from supplements to the 1981 and 1988 Current Population Survey (CPS) in the United States, and from supplements to the 1981 and 1988 Labour Force Survey (LFS) in Canada. The CPS and the LFS are very similar surveys.

Since 1979, earnings information has been collected for one-quarter of individuals in the CPS. Large samples of about 90,000 men can thus be constructed from the "outgoing rotation groups files" of the CPS in both 1981 and 1988. One drawback of these samples, however, is that information on the union status of workers is only available after 1983. Fortunately, a question on union status was asked in the *Multiple Job Holding and Premium Pay Supplement* of the May 1981 CPS. We have therefore combined the May 1981 CPS data with the outgoing rotation group data to obtain comparable samples in both years.[4]

[2]For the United States, see Card (1992); DiNardo, Fortin, and Lemieux (1996); and Freeman (1993). For the United Kingdom, see Gosling and Machin (1995).

[3]See Blau and Kahn (1996) and Lemieux (1993). The latter found that differences between Canada and the United States in unionization rates explain 40% of the Canada/U.S. difference in wage inequality in 1986.

[4]Note that the *Multiple Job Holding and Premium Pay Supplement* was only administered to one-fourth of individuals in the May 1981 CPS. These individuals constitute roughly one-twelfth of the sample from the outgoing rotation group files. All the statistics that are computable even when union status is not available are thus calculated using the larger CPS sample. For example, the overall variance of log wages is always computed using the larger 1981 CPS sample, while the effect of unions on wages is computed using the 1981 May CPS data only.

The choice of the period 1981–88 is solely based on the more limited availability of Canadian data. The first large-scale survey to contain information on the union status of workers was conducted by Statistics Canada in early 1982. This survey, called the 1981 Survey of Work History (SWH), was a supplement to the January and February 1982 LFS. It contains retrospective information on the year 1981. A more recent survey comparable to the 1981 SWH is the 1988 Labour Market Activity Survey (LMAS).[5] Individuals in the 1988 LMAS were surveyed in early 1989 but the questions are again retrospective and pertain to labor market activities during the year 1988. Information on both earnings and union status is available for samples of 20,000 to 25,000 men in both 1981 and 1988.

In addition to the availability of union status information, these U.S. and Canadian samples have another advantage for studying changes in the structure of wages. Unlike the March CPS, the Canadian Survey of Consumer Finances, or the census data for both countries, our samples contain direct information on hourly or weekly wages and on usual hours of work on the main job. An hourly wage on the main job can thus be computed for each worker in our sample. This is a better measure of a point-of-time price of labor than the wage measures available in those other surveys.[6]

Our U.S. and Canadian samples are not strictly comparable, however, since the earnings information is collected differently in the two samples. Earnings in the CPS are defined as usual (weekly or hourly) earnings on the main job in the week previous to the survey date. By contrast, the 1981 SWH and the 1988 LMAS are work history surveys that ask for usual earnings on up to five jobs held during the previous year. We have therefore edited the Canadian samples to obtain a sample of wages comparable to the samples in the CPS (see Appendix 1 for details).

The Canadian samples have other limitations compared to the CPS samples. First, only broad age and education categories are available in the public use versions of the 1981 SWH and 1988 LMAS. We have only four comparable education categories in the two countries, which is nevertheless satisfactory, since years of education in Canada and in the United States are not strictly comparable anyway.[7] The four education categories we use are primary education or less, some or completed high school, some college, and a university degree or more. In terms of years of schooling in the United States, these education categories correspond to 0–8 years, 9–12 years, 13–15 years, and 16 and more years, respectively. We also have six age categories (17–19, 20–24, 25–34, 35–44, 45–54, 55–64) for our samples of men aged 17 to 64. Excluding university graduates aged 17 to 19, we can thus divide the data into 23 comparable age-education groups in the two countries.

Another limitation of the Canadian data is that there is no earnings allocation flag in the public use samples. We thus keep allocated wage information in both U.S. and Canadian samples. This should have limited consequences for the analysis, since the U.S. Bureau of the Census and Statistics Canada use very similar (so-called "hot deck") imputation procedures for missing earnings data. Furthermore, we only keep observations with a wage greater than or equal to 1.75 constant 1981 dollars. This corresponds to half the Canadian federal minimum wage in 1981.

[5]Individuals in the 1988 LMAS were followed up in 1989 and 1990. The 1989 and 1990 samples are not strictly comparable to the 1988 sample, however, because of sample attrition.

[6]In fact, the hourly wage rate cannot be computed from the Survey of Consumer Finances or the Canadian census because of a timing problem. These data sources contain information on earnings and weeks worked during the previous year, but hours worked per week during the survey week.

[7]For example, one needs from 11 to 13 years of schooling to obtain a high school degree in Canada, depending on the province. In the United States, states uniformly require 12 years of schooling for high school graduation.

Note also that there is essentially no top-coding in the Canadian data.[8] By contrast, weekly earnings are top-coded at $999 in the 1981 CPS and at $1,923 in the 1988 CPS. The percentage of workers with top-coded weekly earnings in the CPS is about 1% in 1981, but less than .5% in 1988.[9]

Summary statistics on the U.S. and Canadian samples are reported in Table 1. Unless otherwise indicated, all the statistics presented in this paper are weighted using the CPS (or LFS) sample weights.[10]

Canada and the United States: Similarities and Differences in Labor Market Institutions

As summarized by Card and Freeman (1993) in a volume dedicated to examining the effect of "small differences" in the labor market institutions of the United States and Canada, the two countries have much in common: similar cultures, a similar standard of living, and similar economic institutions. U.S. citizens own a large share of Canadian business assets and vice-versa. Immigration between the two countries is substantial and the countries are each other's largest trading partners.

The two countries' highly decentralized collective bargaining systems are also similar. Since the 1930s Canadian institutions have followed changes in U.S. practices fairly closely, although U.S. labor history has typically been more violent. Union densities in the two countries tracked each other until the late 1960s and early 1970s, when unionization began to decline in the United States while rising in Canada. Attitudes about unionization are fairly similar in the two countries (Riddell 1993). Likewise, despite the higher levels of unionization in Canada, the sizes of U.S. and Canadian union wage differentials are similar, at least for men (Lemieux 1993).

These important similarities between the two countries allow us to focus more sharply on the small differences between them. In particular, Freeman (1990) has argued that it is minor differences in labor laws that have led to the sharp divergence in the percentage of the work force that is unionized in the 1980s: U.S. unionization has fallen precipitously, whereas unionization rates in Canada have remained roughly constant. Perhaps not surprisingly, erosion of the inflation-adjusted minimum wage was much greater in the United States than in Canada. Over the period 1981–88, the U.S. minimum fell 23%, while Canadian minimum wages fell only 12%.

The extent and pattern of unionization in the two countries are documented in detail in columns 3 and 4 (United States) and 7 and 8 (Canada) of Table 1. In the United States, union density declined by 7% (from 28% in 1981). This decline is almost three times larger than the 2.4% decline in unionization in Canada. The level of union membership is also much higher in Canada. The most striking difference is for those men with 9–12 years of school. In the United States in 1981, the unionization rate for this group was 33%; by 1988 it had fallen to 26%. By contrast, the unionization rate for the same group of workers in Canada was 43% in 1981, and in 1988 it was still about 42%.

The unionization rate also decreased faster for younger workers than for older workers in the United States. Interestingly, the unionization rate declined almost as fast among young Canadians as it did among young Americans. So while the decline in U.S. unionism was more pronounced for young and less educated workers, the decline in Canada was solely concentrated among younger workers.

[8] The 1988 LMAS user guide mentioned that "two records with computed total earnings from all jobs in excess of $150,000 have had their hourly wage rates reduced to values which yield totals close to $150,000."

[9] The edited (and allocated) earnings variable is still top-coded at $999 in 1988, but an alternative unedited earnings variable top-coded at $1,923 is also available. We use the unedited earnings variable for workers who report earnings from $999 to $1,923. For workers with missing earnings data, we impute the average wage of workers with non-missing earnings from $999 to $1,923. We also impute this wage to workers top-coded at $999 in 1981.

[10] The weights for Canada are actually these sample weights multiplied by our computed "job" weights described in Appendix 1.

Table 1. Composition of the Samples in the United States and Canada. (Standard Errors in Parentheses)

	United States				Canada			
	Sample Proportions		Unionization Rates		Sample Proportions		Unionization Rates	
Category	1981 (1)	1988 (2)	1981 (3)	1988 (4)	1981 (5)	1988 (6)	1981 (7)	1988 (8)
Age Categories								
Age 17–19	0.062	0.052	0.055	0.054	0.062	0.056	0.183	0.142
	(0.001)	(0.001)	(0.003)	(0.003)	(0.002)	(0.001)	(0.010)	(0.009)
Age 20–24	0.148	0.124	0.195	0.095	0.155	0.128	0.351	0.254
	(0.001)	(0.001)	(0.003)	(0.003)	(0.002)	(0.002)	(0.008)	(0.007)
Age 25–34	0.308	0.321	0.288	0.185	0.301	0.315	0.436	0.383
	(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.003)	(0.006)	(0.005)
Age 35–44	0.204	0.248	0.325	0.263	0.211	0.246	0.464	0.467
	(0.001)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.007)	(0.006)
Age 45–54	0.163	0.158	0.355	0.297	0.160	0.162	0.479	0.496
	(0.001)	(0.001)	(0.004)	(0.004)	(0.002)	(0.002)	(0.008)	(0.008)
Age 55–64	0.114	0.096	0.346	0.286	0.112	0.094	0.492	0.491
	(0.001)	(0.001)	(0.005)	(0.005)	(0.002)	(0.002)	(0.009)	(0.010)
Education Categories								
8 Years and Less	0.074	0.052	0.334	0.215	0.154	0.098	0.524	0.526
	(0.001)	(0.001)	(0.006)	(0.006)	(0.002)	(0.002)	(0.008)	(0.010)
9–12 Years	0.520	0.499	0.334	0.256	0.504	0.470	0.431	0.416
	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.004)
13–15 Years	0.190	0.203	0.254	0.200	0.209	0.271	0.383	0.361
	(0.001)	(0.001)	(0.003)	(0.003)	(0.003)	(0.003)	(0.007)	(0.006)
16 Years and More	0.216	0.246	0.178	0.139	0.133	0.161	0.363	0.357
	(0.001)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)	(0.008)	(0.007)
All Workers	—	—	0.285	0.214	—	—	0.426	0.402
			(0.002)	(0.001)			(0.003)	(0.003)
Sample Size	95,149	83,245	7,676	83,245	25,751	25,855	25,751	26,855

Note: Samples include individuals age 17–64 who report average hourly wages above $1.75 (in 1981 national currency). See text for further details.

Sources: Public use files of the May 1981 Current Population Survey (CPS); 1981 and 1988 Outgoing Rotation Group Supplements of the CPS; 1981 Canadian Survey of Work History; and 1988 Canadian Labour Market Activity Survey.

There are also striking differences in unionization rates by industry in the United States and Canada. Columns 3 and 4 of Table 2 show that, in some industries, Canadian rates of unionization are as high as 68% (Public Administration) or 71% (Education). Another difference is that unionism uniformly declined in all U.S. industries except public administration. Such a systemic decline did not occur in Canada, although unionization shifted from traditional blue-collar industries such as natural resources, manufacturing, and construction toward service-based industries.

Supply, Demand, and Overall Economic Performance

A primary conclusion of the recent U.S. literature on widening wage inequality is that the decline in relative wages of less-skilled workers was driven by relative shifts in labor demand. The two leading explanations for these shifts are skill-biased technical change and international trade. Other factors, such as changes in supply, have also been implicated. Before turning to institutional factors, we discuss possible similarities and differences between the two countries in demand and supply forces.

Table 1 shows that the supply of workers by age and education categories changed at comparable rates in the United States (columns 1 and 2) and Canada (columns 5 and 6) during the 1980s. The age distributions of both countries were extremely similar. In both countries, the middle of the age distribution (ages 25–44) showed some growth over the period. The tails of the distribution (younger and older workers) showed significant declines in both countries.

In both countries, average levels of schooling increased, although U.S. men had more years of schooling on average. In particular, the proportion of men holding a university degree increased by three percentage points in both Canada (from 13% to 16%) and the United States (from 22% to 25%). On the other hand, the supply of workers with a post-secondary education but no college degree increased faster in Canada (community colleges, CEGEPs in Quebec, one- or two-year certificates in universities) than in the United States.

Overall, the supply of workers with more than a high school degree has thus increased faster in Canada than in the United States. The difference in supply shocks is even more pronounced in relative terms, since these workers represented a smaller fraction of the work force in Canada than in the United States in 1981. In a simple supply and demand setting, one would thus expect education wage differentials to decrease in Canada relative to the United States. Freeman and Needels (1993) indeed argued that supply factors account for half of the differential growth in education wage differentials in the two countries.

Turning to demand, columns 1 and 2 of Table 2 show the industrial distribution of employment in the United States and Canada in 1981 and 1988. Given the difficulty of measuring directly the underlying demand shocks due to trade and technology, several authors—for example, Bound and Johnson (1992) and Katz and Murphy (1992)—have proxied these shocks using "fixed manpower requirement indices." The key variable in these demand indices is the change in employment by industry that induces different changes in the demand for different types of labor. Table 2 shows that the differences in the industrial distribution of employment between the countries are rather minor. The share of employment in natural resources and manufacturing industries decreased in both countries. That decline was offset by the growth of retail trade, business services, and personal and other services in both countries. Since the patterns of industrial restructuring are similar in the two countries, though more pronounced in the United States, they are unlikely to play an important role in differential changes in wage inequality.

Alternative demand measures have been proposed by looking more directly at trade flows and technological innovation. For example, Card, Kramarz, and Lemieux (1996) compared relative demand shocks in the United States and Canada using changes in the import penetration rate and the utilization rate of personal computers. They found very similar patterns between the two countries in these two measures of demand. We conclude that changes in labor demand do not play a major role in differential changes in wage inequality.[11]

Finally, it is worth noting that the two countries had similar overall economic performance. For instance, the employment/population ratios followed similar trends in the two countries.[12] Likewise, as measured by growth in real GNP, the two countries were very similar. Although 1981 and 1988 both preceded major recessions, real GNP grew at an average annual rate of

[11]In a previous version of this paper, we did try to assess the role of changes in supply and demand using the methodology of Bound and Johnson (1992). We were unable to explain any systematic divergence in Canada/U.S. wage inequality using this approach.

[12]See Card and Riddell (1993) for a study of the relative employment performance of the two countries.

Table 2. Industrial Distribution of Employment and Unionization in the United States and Canada, 1981 and 1988.

Industry	Employment Share 1981 (1)	Employment Share 1988 (2)	Unionization Rate 1981 (3)	Unionization Rate 1988 (4)
United States				
1. Natural Resources	2.12	1.29	22.16	20.90
2. Durable Manufacturing	20.97	17.38	37.90	26.61
3. Non-Durable Manuf.	10.41	9.30	33.16	24.07
4. Construction	8.75	9.90	37.90	24.05
5. Transportation	5.56	5.45	52.76	39.09
6. Communication & Utilities	5.26	5.25	49.81	45.12
7. Wholesale Trade	5.36	5.24	14.28	9.41
8. Retail Trade	13.17	14.18	12.40	7.97
9. Finance, Ins., Real Est.	4.21	4.86	5.42	4.88
10. Education	5.86	5.65	39.59	35.30
11. Health and Welfare	3.53	3.69	19.25	14.59
12. Business Services	2.28	3.82	10.72	6.06
13. Personal and Other Serv.	6.91	7.97	10.59	7.61
14. Public Administration	6.10	6.04	29.77	35.63
Canada				
1. Natural Resources	4.71	4.01	47.03	39.73
2. Durable Manufacturing	15.84	14.77	49.93	47.19
3. Non-Durable Manuf.	12.09	11.72	45.55	44.55
4. Construction	7.95	8.05	44.94	38.56
5. Transportation	7.57	5.92	59.36	57.88
6. Communication & Utilities	4.81	5.13	66.92	58.14
7. Wholesale Trade	5.85	5.97	14.78	14.35
8. Retail Trade	9.91	10.98	15.88	17.03
9. Finance, Ins., Real Est.	3.70	3.80	5.40	11.95
10. Education	5.71	5.86	72.73	71.42
11. Health and Welfare	3.13	3.35	52.05	59.02
12. Business Services	3.31	3.85	7.13	11.74
13. Personal and Other Serv.	6.36	7.91	16.17	15.47
14. Public Administration	9.05	8.67	66.39	67.83

Note: Samples include individuals age 17–64 who report average hourly wages above \$1.75 (in 1981 national currency). See text for further details.

3% in the United States and at a rate of 3.4% in Canada between 1981 and 1988. This suggests that the same underlying economic forces were at play in the United States and Canada during the 1980s.

Changes in Summary Measures of Wage Inequality in the United States and Canada

Before looking at the effect of institutions on changes in wage inequality, it is useful to summarize these changes using comparable data for the United States and Canada. In Table 3 we summarize various measures of wage inequality for the two countries. In the top panel of the table the wage measure is log hourly earnings; in the bottom panel we use the log weekly wage for the sake of comparison with other studies.

In the United States real hourly wages fell and income inequality increased by all of our measures. In Canada, the increase in hourly wage inequality was more modest and real wages grew slightly. The standard deviation of log hourly wages in Canada grew by only 0.010 compared to 0.056 in the United States. The differential between the 10th and 90th percentiles of log

Table 3. Distribution of Hourly and Weekly Wages in the United States and Canada, 1981 and 1988.

	United States			Canada		
Description	1981 (1)	1988 (2)	Change (3)	1981 (4)	1988 (5)	Change (6)
			Log Hourly Wage			
1. Mean Real Wage (geometric mean)	7.658	7.566	−0.093	8.578 [7.154]	8.915 [8.007]	0.337 [0.853]
2. Standard Deviation	0.504	0.560	0.056	0.476	0.486	0.010
3. 25–75 Differential	0.721	0.808	0.086	0.667	0.736	0.069
4. 10–90 Differential	1.330	1.492	0.162	1.280	1.365	0.085
5. 10–50 Differential	0.740	0.799	0.058	0.708	0.783	0.075
6. 50–90 Differential	0.590	0.693	0.103	0.572	0.582	0.010
			Log Weekly Wage			
1. Mean Real Wage (geometric mean)	301.59	301.95	0.35	332.34 [277.18]	348.52 [313.02]	16.18 [35.84]
2. Standard Deviation	0.627	0.701	0.074	0.588	0.642	0.054
3. 25–75 Differential	0.725	0.811	0.086	0.694	0.782	0.088
4. 10–90 Differential	1.455	1.710	0.255	1.384	1.555	0.171
5. 10–50 Differential	0.842	0.965	0.123	0.818	0.942	0.124
6. 50–90 Differential	0.613	0.745	0.132	0.565	0.613	0.048
Sample Size	95,149	83,245		25,751	26,855	

Note: Samples include individuals age 17–64 who report average hourly wages above $1.75 (in 1981 national currency). Mean wages are expressed in constant 1981 units of the national currency, except for the numbers in square brackets, which represent Canadian dollars converted into constant 1981 U.S. dollars using the prevailing exchange rates in 1981 and 1988.

wages grew half as fast in Canada (8.5%) as in the United States (16%). Interestingly, the greatest difference between the two countries seems to be at the top end of the distribution. Growth in the 10th–50th percentile differential was comparable in the two countries, while the 50th–90th percentile differential grew ten times faster in the United States than in Canada.

Table 3 highlights the fact that in Canada hourly wage inequality grew much more slowly than weekly earnings inequality.[13] For instance, the standard deviation of log weekly earnings increased by 0.054 while the standard deviation of hourly wages increased by only 0.010. Corresponding figures for the United States are 0.074 and 0.056, respectively. Changes in weekly earnings inequality thus seem to be a good proxy for changes in inequality in the price of labor in the United States but not in Canada.[14] This suggests that weekly hours of work are more responsive than wages to changes in labor market conditions in

[13]This fact was pointed out by Doiron and Barrett (1996) using the same data but differently constructed samples. They also reported similar findings for women in Canada.

[14]Using data from the Canadian Census and the Survey of Consumer Finances, authors such as Davis (1992) and Gottschalk and Joyce (1992) have found that weekly earnings inequality increased almost as fast in Canada as in the United States. We obtain similar results using the 1981 SWH and the 1988 LMAS but find that these results give little indication of changes in hourly wage inequality. Note that Picot, Myles, and Wannell (1990) also reported that data from the 1981 and 1986 Canadian censuses show a rate of growth in earnings inequality comparable to that found using data from the 1981 SWH and 1986 LMAS.

Canada, while the opposite is true in the United States. Identifying the source of this difference would be an interesting topic for future research.

Changes in the Overall Density of Wages

One drawback of summary measures of wage inequality is that they give little indication of what happens where in the distribution of wages. To get a more complete picture of the changes in the distribution of wages, we present kernel density estimates of the density of log wages in Figures la (United States) and lb (Canada).[15] It is clear from the figures that, in the United States, there was a substantial flattening of the wage distribution between 1981 and 1988. One striking feature of these graphs is the sharp decline in the density near the minimum wage. Even more striking is the fact that all of the "fattening" of the lower tail of the distribution of wages occurred below the 1981 minimum wage of $3.35 (in real terms). These graphs thus suggest that *the 23% decline in the real minimum wage from 1981 to 1988 accounts for most of the increase in U.S. wage inequality in the left tail of the wage distribution.*

Compared with the U.S. wage distribution, the Canadian wage distribution reported in Figure lb is much less skewed to the left and looks more log-normal. In addition, there is almost no noticeable effect of the minimum wage on the distribution of wages in 1981. This may partly be explained by the large inter-province variation in the minimum wage in 1981: the minimum wage ranged from $3.00 in Ontario in the first months of 1981 to $4.00 in Quebec in the last months of 1981. By contrast, the minimum wage was the same in Ontario and Quebec in 1988 ($4.55 in the first six months of 1988, $4.75 later). This may explain why a concentration of workers just above the minimum wage can be observed in 1988 but not in 1981. Note that this concentration of workers above the minimum wage in 1988 is the only substantial difference between two otherwise similar wage distributions.

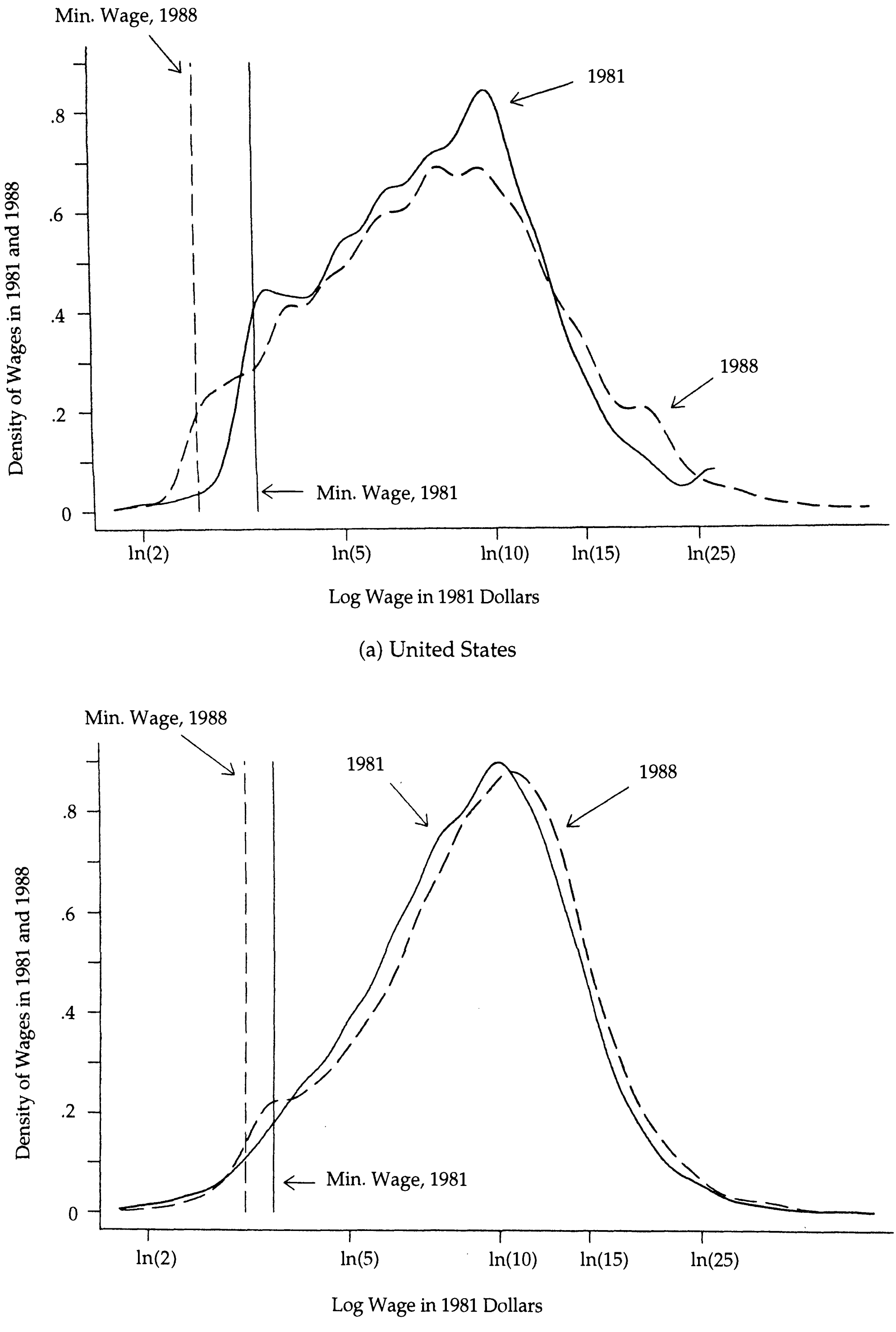
Changes in Wage Inequality Between and Within Groups of Workers

It is well known that both between- and within-group wage inequality increased in the United States in the 1980s. Rows 1 and 2 of Table 4 compare the changes in U.S. and Canadian between-group inequality by presenting a selected number of education and age wage differentials. Changes in within-group inequality are considered in row 3 by computing the standard deviation of log wages among the 23 age and education groups available.

First consider the results for all workers (union and nonunion) in columns 1 and 2. In Canada, the most striking source of increasing inequality is in the differential between older and younger men, especially at lower levels of schooling. The differential between men aged 45–54 and those aged 25–34 with 9–12 years of education increased by 13 percentage points over the period 1981–88. In the United States, however, the increased inequality seems to be more closely tied to increases in returns to schooling. As is evident from rows 2a and 2b of Table 4, education differentials grew more quickly in the United States than in Canada, where they remained roughly constant. In the United States, there was a 16 point increase in the college–high school differential (16 and more versus 9–12 years of education) for the 25–34 age group, and an 11 point increase for the 45–54 age group. As mentioned above, differences in supply are a natural explanation for the differential growth in education wage differentials in the two countries.

The changes in within-group standard deviations summarized in row 3 of Table 4 show that this measure of dispersion grew somewhat larger over the 1981–88 period in the United States and generally fell in Canada. When the between- and within-

[15]All estimates are obtained using a Gaussian kernel. After some experimentation, we decided to use a bandwidth of 0.075 for all the samples. See DFL for more detail.



(a) United States

(b) Canada

Figure 1. Distribution of Real Wages, 1981–1988.

Table 4. Changes in Wage Inequality in the United States and Canada Between 1981 and 1988.

	All Workers		*Nonunion Only*		*Union Only*	
Measure	*U.S.* *(1)*	*Canada* *(2)*	*U.S.* *(3)*	*Canada* *(4)*	*U.S.* *(5)*	*Canada* *(6)*
1. *Age Differentials* (Age 45–54 - Age 25–34)						
a. 9–12 Years Education	0.075	0.128	0.043	0.113	0.047	0.119
b. 16 or More Years' Educ.	0.030	0.120	0.041	0.128	0.077	0.061
2. *Education Differentials* (16 or More - 9–12 Years)						
a. Age 25–34	0.159	0.007	0.195	0.037	0.072	–0.034
b. Age 45–54	0.113	–0.001	0.193	0.052	0.102	–0.091
3. *Within-Cell Standard Deviations*						
9–12 Years' Education:						
a. Age 25–34	0.036	–0.017	–0.006	–0.021	0.021	–0.039
b. Age 45–54	0.030	–0.053	0.029	–0.023	–0.004	–0.106
16 or More Years' Education:						
c. Age 25–34	0.026	0.005	0.064	0.008	0.033	–0.037
d. Age 45–54	0.001	–0.060	–0.010	0.004	0.036	–0.095
e. Average over All Cells	0.031	–0.033	0.025	–0.031	0.010	–0.062
Overall Standard Deviation of Log Hourly Wage:	0.056	0.010	0.056	0.027	0.020	–0.050

Note: Entries are difference in mean log wages between indicated groups, or standard deviations of mean log wages within indicated groups. Samples include individuals age 17–64 who report average hourly wages above $1.75 (in 1981 national currency). See text for further details.

group components are taken together, the standard deviation of log hourly wages grew 11% in the United States compared to only 2% in Canada.

We document changes in wage inequality among nonunion workers in columns 3 and 4 of Table 4. Since nonunion workers are a majority in both countries, it is not surprising that changes in the structure of wages resemble the general patterns we observe for all workers. Some of these patterns are more striking, however, for the nonunion sector:

1. In Canada, increases in wage inequality in the nonunion sector are mostly attributable to a massive increase in wage differentials by age, although there was also a small increase in the returns to education. The differential between workers aged 45–54 and workers aged 25–34 with at least some high school increased by 11 percentage points over the period 1981 to 1988. For workers with 16 and more years of education, this age differential increased by 13 points. Wage differentials by education grew only by 4 to 5 percentage points.

2. In the United States, the story is just the opposite: wage differentials by age in the nonunion sector grew by only 4 percentage points, while differentials by schooling grew dramatically. The college/high school wage differential both for younger workers and for older workers increased by 20 percentage points.

3. Changes in within-group standard deviations follow opposite patterns in the two countries. In Canada, within-group standard deviations either changed little over the period or declined slightly. In the United States, they either increased or remained approximately constant.

Some interesting differences emerge when we turn to the experience of union workers in the two countries (columns 5 and 6). Among Canadian union workers, education wage differentials and within-group standard deviations decreased substantially during the 1980s. On the other hand, the increase in age differentials in the union sector was almost as large as the increase in age differentials in the nonunion sector. In the United States, education wage differentials and the within-group standard deviation of log wages increased, albeit a bit more slowly than in the nonunion sector; age differentials increased faster in the union sector than they did in the nonunion sector.

Except for age differentials, wage inequality thus increased faster in the nonunion sector than in the union sector in both Canada and the United States. This supports the view that inequality increased faster in the United States than in other industrialized countries because of the small size of its union sector. This hypothesis and others will be examined in more detail in the next section.

Effect of Unions and Minimum Wages on Changes in the Overall Wage Distribution

We now turn to the following question: What was the overall effect of unions on wage inequality in Canada and the United States? We look at this question in two ways.

First, for each country we ask the question: What would be the distribution of wages if union workers were paid according to the wage schedule prevailing in the nonunion sector? In other words, how would the wage distribution be different in the "absence of unions"? We develop a simple semiparametric procedure to answer this question.

Second, we repeat that comparison but focus on particular aspects of the distribution such as the variance of log wages and standard age and education wage differentials. This enables us to separate easily the effect of unions into between- and within-group components.

We also look at the effect of another labor market institution, the minimum wage, on the distribution of wages. One reason we focus on the effects of unions and the minimum wage is that these two factors have clear and testable implications for both the mean and the variance of log wages (or other measures of inequality) for any given group of workers. The impact of unions on the variance is due the "leveling" effect of unions, while the impact of the minimum wage is due to the censoring of the lower tail of the wage distribution. By contrast, the standard supply and demand factors discussed above have implications only for the mean wages among age and education groups.

Semiparametric Estimates of the Union Effect on the Distribution of Wages

We use a simple semiparametric method to estimate the distribution of wages that would prevail if all workers were paid according to the observed nonunion wage schedule. Note that this is a more modest task than trying to estimate the distribution that would prevail in the absence of unions. The point is that the observed wage schedule in the nonunion sector may itself depend on the unionization rate because of general equilibrium effects (for example, nonunion companies pay workers more because of the threat of unionization) or selection biases (nonunion workers are a non-random sample of the population). For the sake of expositional brevity, however, we will occasionally use the shorthand "absence of unions" to refer to the state that would result if all workers were paid according to the observed nonunion wage schedule.

This semiparametric scheme is based on a simple reweighting of the distribution of wages of nonunion workers. The unweighted distribution of wages in the nonunion sector is an inappropriate estimate of the distribution of wages in the absence of unions because the distribution of characteristics among nonunion workers is not the same as the distribution of characteristics among all (union and nonunion) workers. For example, blue-collar workers are underrepresented in the sample of nonunion workers. More "weight" has thus to be put on blue-collar workers in the nonunion sample to get the same proportion of blue-collar workers as in the overall sample. We describe the procedure in detail in Appendix 2.

Semiparametric estimates of the wage distribution that would prevail if all workers were paid according to the nonunion wage schedule are compared to the actual wage distribution in Figures 2 (United States) and 3 (Canada). The figures indicate qualitatively similar effects of unions on the distribution of wages. Unions tend to "move" workers from the middle-lower tail of the distribution to a peak slightly above the median of the distribution. This effect is most pronounced for Canadian workers in 1988 (Figure 3b) and least so for U.S. workers in 1988 (Figure 2b). Not surprisingly, the figures indicate that unions have little effect on the distribution of wages either near and below the minimum wage or in the upper tail of the distribution.

Returning to Figure la, a rough summary of the changes in the distribution of wages in the United States is that the decline in unionism leveled the peak slightly above the median of the distribution, while decreases in the minimum wage created a concentration of low-wage jobs at the bottom end of the distribution. Neither unions nor the minimum wage can explain, however, why the upper tail of the distribution became "fatter" in the United States.

The Effect of Unions on the Variance of Log Wages

We next focus on a specific measure of wage inequality, the variance of log wages. Unlike other measures of wage inequality, the variance can easily be decomposed into between- and within-group components, and into components attributable to unions, minimum wages, and so on.

As mentioned previously, the distribution of wages that would prevail if all workers were paid according to the nonunion wage schedule is obtained by weighting each observation i by a weighting factor $\theta_i = \theta(x_i)$ (see Appendix 2). Note that these weighting factors are normalized to have a mean of one. It is straightforward to incorporate standard CPS (or LFS) weights in this setting by multiplying them by θ_i.

As shown in DFL, these weights can than be used to compute the density of wages or other statistics using standard estimation methods over the reweighted sample of nonunion workers. In what follows, we show how a modification of the procedure described in DFL can be used to decompose the effect of unions on the variance of log wages into several components of interest.[16]

Consider the variance of log wages for all workers

$$(1) \qquad Var(w) = \frac{1}{N}\sum_i (W_i - E(w))^2,$$

and the variance for nonunion workers

$$(2) \qquad Var(w|U=0) = \frac{1}{N_n}\sum_{U_i=0} (W_i - E(w|U=0))^2.$$

The variance of log wages that would prevail if everybody were paid according to the nonunion wage schedule is simply

$$(3) \qquad Var^n(w) = \frac{1}{N_n}\sum_{U_i=0} \theta_i (W_i - E^n(w))^2,$$

where

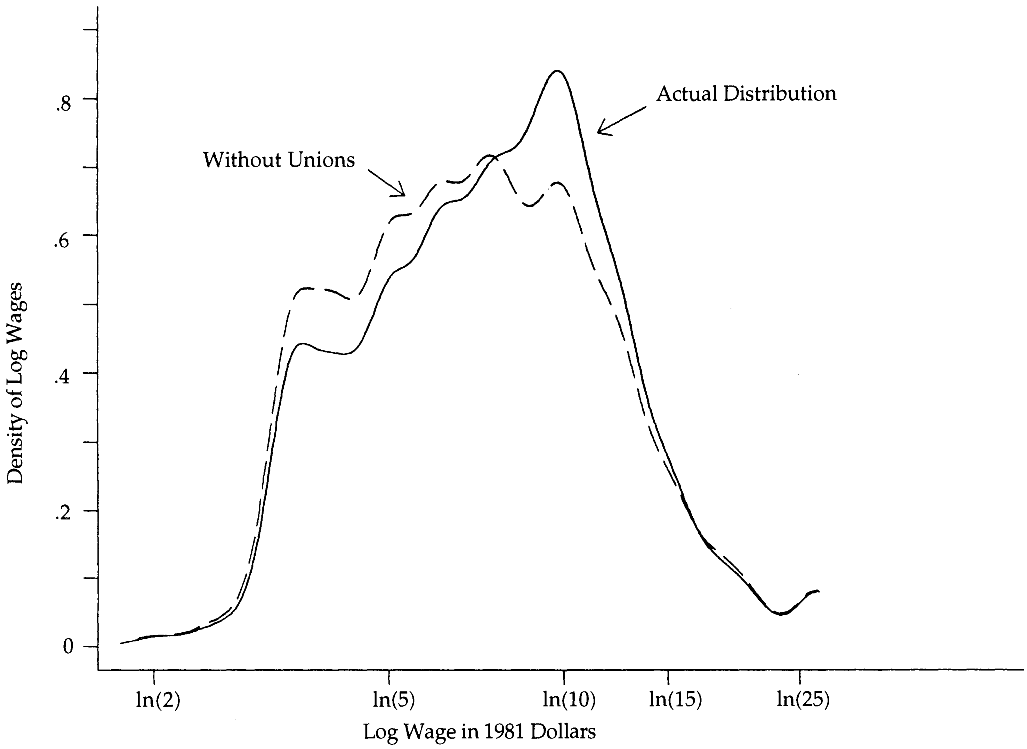
$$(4) \qquad E^n(w) = \frac{1}{N_n}\sum_{U_i=0} \theta_i W_i.$$

The estimated effect of unions on the variance of log wages is

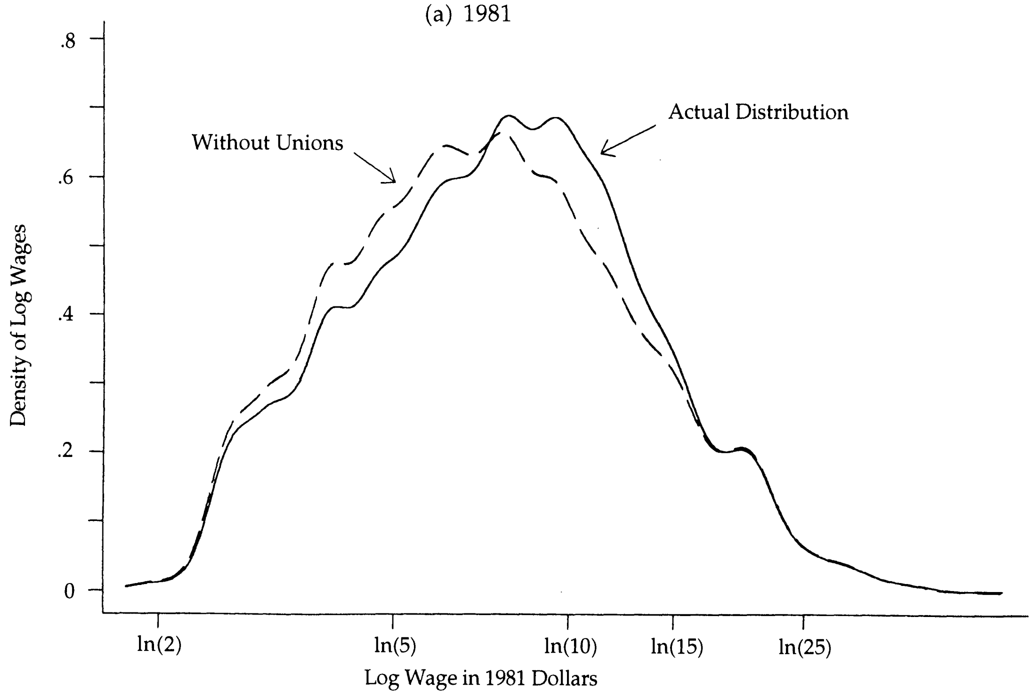
$$(5) \qquad \mathrm{Var}(w) - \mathrm{Var}^n(w).$$

It is interesting to decompose the effect of unions into a wage compression effect between different groups of workers, a wage compression effect within groups of workers, and a wage gap effect. The decomposition is based on two standard wage equations, one for the union sector, another for the nonunion sector:

[16]The decomposition is similar to the one in Freeman (1980), although the reweighting procedure simplifies the computations.



(a) 1981



(b) 1988

Figure 2. Effect of Unions on the U.S. Distribution of Real Wages.

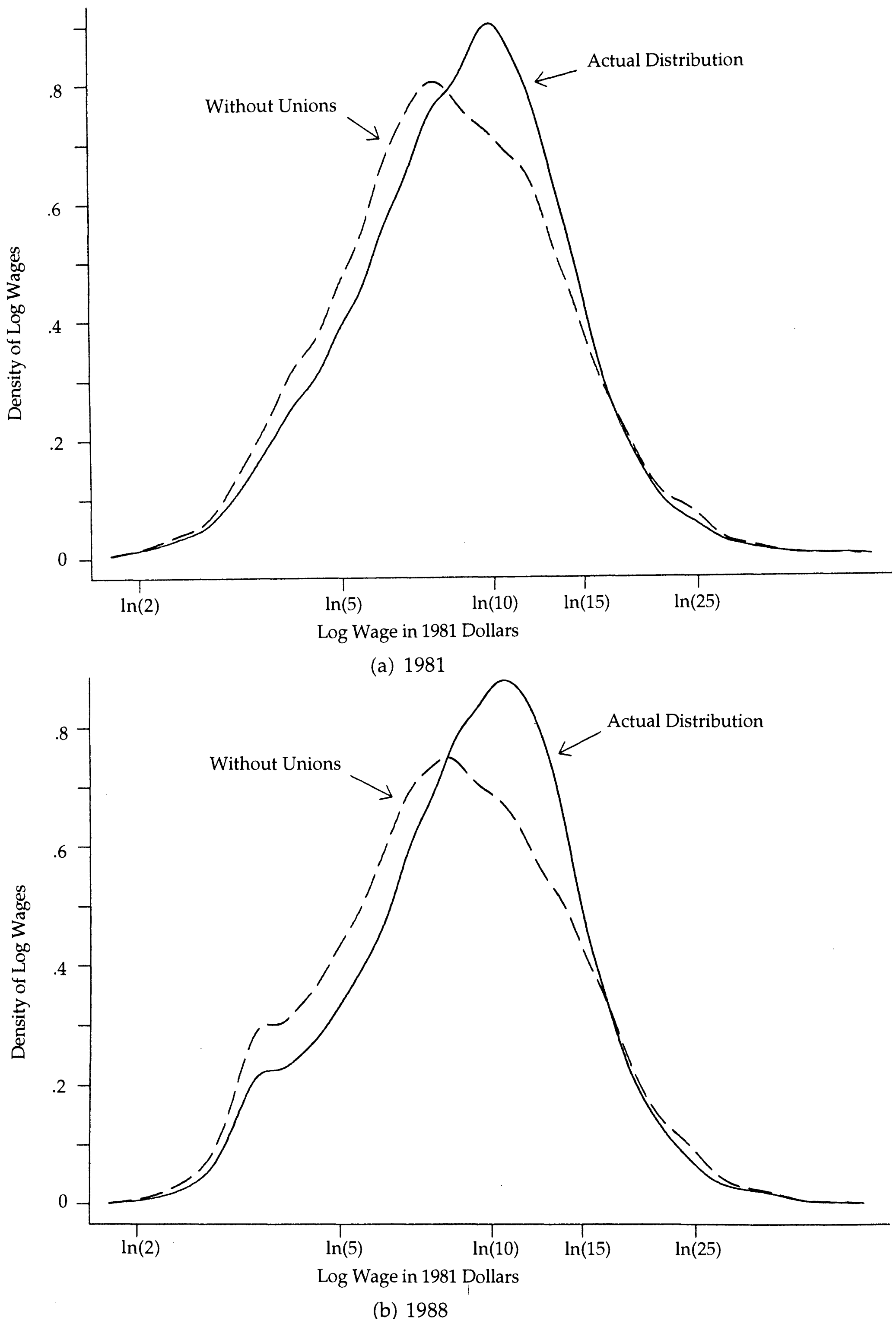


Figure 3. Effect of Unions on the Canadian Distribution of Real Wages.

$$(6) \qquad w_i^S = X_i\beta^S + \varepsilon_i^S, \; S = U, N,$$

where S is an indicator whose value is U if we are describing union wage-setting, and N if we are describing wage-setting in the nonunion sector. The vector X_i consists of a vector of observable characteristics such as dummy variables for the age and education categories, for part-time or full-time worker, for marital status, and for industry and occupation. The unobservable characteristics (or error terms) ε_i^U and ε_i^N are assumed to have a zero conditional mean.

Using simple calculations, the union effect of equation (5) can be rewritten as the sum of three separate components. The wage compression effect between groups of workers is equal to

$$(7) \qquad \bar{U} \times [\mathrm{Var}(X\beta^U | U = 1) - \mathrm{Var}(X\beta^N | U = 1)].$$

This effect is typically believed to be negative because of the "leveling" effect of unions. In other words, unions tend to standardize wages by reducing the returns to various skills such as education ($\beta_U < \beta_N$). Similarly, the wage compression effect of unions within groups of workers is equal to

$$(8) \qquad \bar{U} \times [\mathrm{Var}(\varepsilon^U | U = 1) - \mathrm{Var}(\varepsilon^N | U = 1)].$$

This effect is also typically believed to be negative because of the "leveling" effect. If the error term ε_i is interpreted as unobservable skills, the "leveling" effect simply means that unions reduce the returns to these unobservable skills. Finally, the wage gap effect is given by

$$(9) \qquad \bar{U} \times (1 - \bar{U}) \times [\bar{\Delta}^2 - (\bar{\Delta} - \Delta)^2],$$

where $\bar{\Delta}$ is the difference in mean wages between the union and the nonunion sector and Δ is the union wage gap (for union workers).[17] Unlike the wage compression effects discussed above, this effect is typically believed to be positive. As stressed by several authors (for example, Friedman 1962), this effect is the consequence of the wage premium received by union workers. The point is that unions may increase inequality by advantaging some (union) workers at the expense of other (nonunion) workers.

Table 5 presents the results of this decomposition for both countries in both years. We also add two numbers to the display to calculate the magnitude of the effect of unions. In columns 1 and 4 we present the actual components of variance that prevailed in the particular country and year. Next we recompute this component, applying the nonunion wage-setting equation to the distribution of union characteristics (both observed and unobserved). We label this "Variance Without Unions" and display the calculations in columns 2 and 5. Given the actual variance, and the variance without unions, the "Union Effect" reported in columns 3 and 6 is merely the difference between these two quantities. The union effect in rows 3, 6, and 7 corresponds to the terms in equations (7), (8), and (9).

Note that before computing the effect of unions on the variance of wages, we control for changes in the real value of the minimum wage using a simple imputation procedure described in Appendix 2. The idea is to impute the part of the 1981 distribution (in real terms) that lies at or below the 1981 minimum wage to workers earning this 1981 real minimum wage or less in 1988. This yields a simulated 1988 distribution that would have prevailed if the minimum wage had remained at its (higher) 1981 level.

As expected, in both countries and years, the union wage compression effects between and within groups of workers tend to reduce the overall variance of log wages. Also as expected, the wage gap effect is positive. This effect is small, however, because men with characteristics highly valued in the nonunion sector are less likely to be unionized in both countries ($\bar{\Delta}$ is smaller than Δ).

The most interesting comparison between the two countries is how the equalizing influence of unions has *changed* over

[17]Formally, we have $\bar{\Delta} = \bar{X}^U\beta^U - \bar{X}^N\beta^N$, where $\bar{X}^U$ and $\bar{X}^N$ are the sample averages of x_{it} in the union and nonunion sectors, respectively, and $\Delta = \bar{X}^U(\beta^U - \beta^N)$.

Table 5. Overall Effect of Unions on the Variance of Log Wages in the United States and Canada.

Measure	Year 1981: Actual Variance (1)	Year 1981: Variance w/o Unions (2)	Year 1981: Union Effect (3)	Year 1988: Actual Variance (4)	Year 1988: Variance w/o Unions (5)	Year 1988: Union Effect (6)
United States						
Wage Compression Effect Between Groups of Workers						
1. Var($X\beta$ | union)	0.0383	0.0647	−0.0263	0.0456	0.0867	−0.0411
2. Var($X\beta$ | nonunion)	0.1268	0.1268	0.0000	0.1617	0.1617	0.0000
3. Average Between-Group Variance ($\bar{u}$*row 1 + (1− $\bar{u}$) *row 2)	0.1012	0.1088	−0.0076	0.1368	0.1456	−0.0088
Wage Compression Effect Within Groups of Workers						
4. Var(ε | union)	0.1059	0.1521	−0.0462	0.1188	0.1529	−0.0341
5. Var(ε | nonunion)	0.1545	0.1545	0.0000	0.1670	0.1670	0.0000
Total Within-Group Variance						
6. Var(ε) ($\bar{u}$*row 4 + (1− $\bar{u}$) *row 5)	0.1405	0.1538	−0.0133	0.1567	0.1640	−0.0073
Wage Gap Effect						
7. $\bar{u}*(1-\bar{u})*\Delta^2$	0.0069	0.0000	0.0068	0.0076	0.0000	0.0076
Overall Variance of Wages						
8. Var($X\beta + \varepsilon$) (row 3 + row 6 + row 7)	0.2486	0.2627	−0.0141	0.3011	0.3096	−0.0085
Canada						
Wage Compression Effect Between Groups of Workers						
1. Var($X\beta$ | union)	0.0279	0.0447	−0.0168	0.0326	0.0808	−0.0482
2. Var($X\beta$ | nonunion)	0.0762	0.0762	0.0000	0.1307	0.1307	0.0000
3. Average Between-Group Variance ($\bar{u}$*row 1 + (1− $\bar{u}$) *row 2)	0.0556	0.0627	−0.0072	0.0913	0.1107	−0.0194
Wage Compression Effect Within Groups of Workers						
4. Var(ε | union)	0.1248	0.1764	−0.0517	0.0979	0.1573	−0.0594
5. Var(ε | nonunion)	0.1913	0.1913	0.0000	0.1641	0.1641	0.0000
Total Within-Group Variance						
6. Var(ε) ($\bar{u}$*row 4 + (1− $\bar{u}$) *row 5)	0.1629	0.1850	−0.0220	0.1374	0.1613	−0.0239
Wage Gap Effect						
7. $\bar{u}*(1-\bar{u})*\Delta^2$	0.0075	0.0007	0.0068	0.0121	0.0010	0.0111
Overall Variance of Wages						
8. Var($X\beta + \varepsilon$) (row 3 + row 6 + row 7)	0.2261	0.2484	−0.0224	0.2408	0.2730	−0.0322

Note: $\bar{u}$ is the unionization rate and Δ is the difference between average union and average nonunion wages. Samples include individuals age 17–64 who report average hourly wages above \$1.75 (in 1981 national currency). See text for further details.

the period. Consider the wage compression effect of unions on the variance between groups of workers in row 3 of both panels. In Canada, this union effect grew from −.007 in 1981 to −.019 in 1988. Over the same period, the increase in this wage compression effect was much smaller in the United States (from −.008 in 1981 to −.009

in 1988). Similarly, the wage compression effect of unions on the variance within groups of workers (row 6) became smaller in absolute terms in the United States but remained relatively constant in Canada.

We draw two conclusions from Table 5:

—*Canadian unions have a more equalizing influence on male wages.* This pattern stems from two conditions: more men are unionized in Canada, and within the unionized sector Canadian wages are more equalized.

—*The equalizing influence of unions has grown larger in Canada over the period, whereas this influence has diminished in the United States.* Largely accounting for this pattern, again, is the fact that within the union sector, U.S. unions have been less successful over time in maintaining their equalizing influence, while Canadian unions have grown somewhat more successful. The sharp decrease in union density in the United States has also played a role.

The Contribution of Unions and of Minimum Wage Laws to Changes in the Variance of Log Wages

In Table 6, we decompose relative changes in the variance of log wages into the effect of the minimum wage (column 2) and the effect of unions (column 5). We further decompose the effect of unions into a component attributable to changes in the equalizing effects of unions (column 3) and a component attributable to changes in the unionization rate (column 4).[18] The effect of changes in the minimum wage is obtained by contrasting the 1988 actual variance of log wages to the variance that would have prevailed if the minimum wage had remained at its 1981 level (in real terms). This hypothetical variance is obtained by using the imputation procedure described at the end of Appendix 2.

The key results in Table 6 can be summarized as follows:

—In the absence of unions, the variance of log wages would have increased by 9.6% less in the United States than it actually did. Most of this effect is attributable to the decline in unionization as opposed to changes in the equalizing effects of unions. This effect is also smaller than those found by Card (1992) and Freeman (1993), who used data for longer time periods than 1981–88.

—In the absence of unions, the variance of log wages would have increased by 101% more in Canada than it actually did. Contrary to the case in the United States, this effect is almost solely due to the fact that Canadian unions became relatively "more equalizing" in 1988 than in 1981.

—The decline in the minimum wage in the United States accounted for 22% of the increase in the variance of log wages from 1981 to 1988. This finding is consistent with the semiparametric evidence reported in Figure 1a and with the results reported for the period 1979 to 1988 in DFL.

—*A striking two-thirds of the U.S./Canada difference in changes in the variance of log wages from 1981 to 1988 is explained by unions and by minimum wage laws.* Each of these labor market institutions accounts for about a third of the U.S./Canada difference. Most of the effect of unions is attributable to the fact that Canadian unions, unlike U.S. unions, became more equalizing in 1988 than in 1981.

This last finding points to the large, though neglected, role of labor market institutions in explaining the very different changes in wage inequality across countries. It is interesting to note that Katz, Loveman, and Blanchflower (1995) conjectured that unions and minimum wage legislation played a major role in explaining the very different experiences of France and the United Kingdom with regard to changes in inequality over the 1980s. By comparing Canada and the United States, we have been able to directly test these

[18]The contribution of wage equalizing effects of unions is obtained by computing the effect of unions on the variance of log wages in 1981 and 1988 that would have prevailed if the unionization rate had remained constant over this period. The effect of changes in the unionization rate is obtained by doing just the opposite (holding equalizing effects constant but changing the unionization rate).

Table 6. Decomposition of Cross-Country Differences and Temporal Changes in the Variance of Log Wages.

Description	Actual Difference (or Change) in the Variance (1)	Effect of Difference (or Change) in: Minimum Wage (2)	Union Equalizing Effects (3)	Unionization Rates (4)	Total Effect of Unions (5)
(U.S. 1981)-(Canada 1981)	0.0278	—	0.0024 (8.6)	0.0056 (20.3)	0.0080 (28.9)
(U.S. 1988)-(Canada 1988)	0.0773	0.0169 (21.9)	0.0126 (16.3)	0.0114 (14.7)	0.0240 (31.0)
(U.S. 1988)-(U.S. 1981)	0.0596	0.0129 (21.7)	0.0018 (3.0)	0.0040 (6.6)	0.0057 (9.6)
(Canada 1988)-(Canada 1981)	0.0101	−0.0040 (−39.3)	−0.0093 (−92.0)	−0.0009 (−9.1)	−0.0102 (−101.1)
(Change in U.S.)-(Change in Canada)	0.0495	0.0169 (34.2)	0.0111 (22.4)	0.0049 (9.8)	0.0160 (32.2)

Note: The numbers in parentheses represent the effect on the variance of wages as a percentage of total changes in the variance. The effect of the minimum wage is computed by comparing actual changes in the variance of wages to changes that would have prevailed if the distribution of wages at or below the 1981 real minimum wage in 1988 was as in 1981. The total effect of unions on changes in the variance of wages (column 5) is calculated using the estimated effect of unions reported in Table 5 (columns 3 and 6 of row 8). Columns (3) and (4) show this total union effect decomposed into the impact of changes in the effect of unions on the variance of wages, holding constant unionization rates (column 3), and changes in the unionization rates, holding constant the effect of unions on the variance of wages (column 4). See the text for further details.

hypotheses, and we find that these factors do in fact play a major role.

Union and Minimum Wage Effects on Other Measures of Wage Inequality

While other papers have studied the changes in wage inequality in Canada and the United States, our finding on the importance of institutions is new. For example, Freeman and Needels (1993) recognized the importance of institutions but concluded that differential supply changes are the most important source of differences between the two countries.

Our findings are easily reconciled with those of Freeman and Needels (1993) by noting that the two studies focus on different measures of inequality. Freeman and Needels (1993) focused on the evolution of the college–high school differential, whereas we look at more global measures of wage dispersion (entire distribution and variance of log wages). Table 7 shows how this difference in focus can yield very different conclusions on the role of institutions. The table presents the effect of unionization and of the minimum wage on changes in the various measures of wage inequality reported in Table 4. The numbers reported in the table are computed from the reweighted samples used in Table 6. This illustrates once again the advantage of the DFL method. Once proper weighting factors have been estimated, any measure of wage inequality can be computed using standard methods by properly weighting the sample.

Rows 1 and 2 of Table 7 show that institutional factors do not explain much of the differential Canada/U.S. changes in age and education differentials. Explanations for these changes must thus be found elsewhere. On the other hand, institutional factors explain more systematically the evolution of within-group (row 3.e) and overall wage inequality.

In other words, the findings of Freeman

Table 7. Effect of Institutional Changes on Changes in Various Measures of Wage Inequality.

Measure	*United States* Total Change (1)	Min. Wage Effect (2)	Union Effect (3)	*Canada* Total Change (4)	Min. Wage Effect (5)	Union Effect (6)
1. *Age Differentials* (Age 45–54 - Age 25–34)						
a. 9–12 Years Education	0.075	−0.012	0.034	0.128	0.007	−0.018
b. 16 or more Years Educ.	0.030	−0.002	0.008	0.120	0.001	−0.017
2. *Education Differentials* (16 or More - 9–12 Years)						
a. Age 25–34	0.159	−0.004	−0.001	0.007	−0.000	−0.029
b. Age 45–54	0.113	0.006	−0.027	−0.001	−0.006	−0.028
3. *Within-Cell Standard Deviations*						
9–12 Years' Education:						
a. Age 25–34	0.036	−0.001	0.028	−0.017	0.000	0.016
b. Age 45–54	0.030	0.016	0.001	−0.053	−0.011	−0.015
16 or More Years' Education:						
c. Age 25–34	0.026	0.003	0.006	0.005	−0.000	0.009
d. Age 45–54	0.001	0.008	0.010	−0.060	0.003	−0.053
e. Average over All Cells	0.031	0.006	0.007	−0.033	−0.003	−0.004
Overall Standard Deviation of Log Hourly Wage	0.056	0.011	0.006	0.010	−0.004	−0.009

Note: See the note to Table 4 for a description of the samples and of the measures of wage inequality used.

and Needels (1993) on the role of supply shocks in the evolution of education wage differentials can be perfectly consistent with our general conclusion that institutions explain most of the Canada/U.S. differences in changes in wage inequality. Education wage differentials are one important dimension of wage inequality but not the only one. Our findings do not suggest that factors like supply and demand do not explain any of the difference between the United States and Canada. They simply suggest that these factors play a small role compared to the one played by institutions when broadly defined measures of wage inequality are considered.

Conclusion

In the 1980s wage inequality grew larger in the United States while it remained roughly constant in Canada. Returns to education increased much faster in the United States than in Canada, but returns to experience increased more in Canada than in the United States. During the same period, unionization rates fell precipitously in the United States but declined very little in Canada. Similarly, the real minimum wage declined by 23% in the United States but by only 12% in Canada.

Using large computerized data files for men in the United States and Canada in 1981 and 1988, we have found that unions and the minimum wage accounted for two-thirds of the differential growth of wage inequality between the two countries. These results suggest that labor market institutions are an important explanation for the difference in the evolution of overall wage inequality in the two countries.

Appendix 1
Canadian Data

In this Appendix, we describe how the Canadian samples were edited to make them comparable to the U.S. samples. For each month of the year (each week in 1988) we know whether the individual was working on each job. We determined which job (if any) was the main job by comparing usual hours of work on the different jobs. Once it was established that a job was a main job during a month, a weight of 1/12 was assigned to this job. The weight was augmented by 1/12 for each additional month in which the job was determined to be a main job. Our final sample is thus a sample of jobs, and not a sample of individuals. When weighted using the above procedure, however, this sample of jobs becomes equivalent to a sample of main jobs that would be obtained at a given point of time. This "weighted sample" is thus comparable to the sample obtained in the CPS. One feature of both our reweighted Canadian sample and the CPS sample is that short jobs tend to be undersampled relative to an annual sample of jobs. Our measures of the distribution of earnings will thus tend to put "less weight" on people weakly attached to the labor force than measures based on annual earnings.

Appendix 2
Semi-Parametric Estimates of the Effects of Unions on the Minimum Wage

In this appendix we describe briefly our semiparametric density estimation method. Full details can be found in DFL.

Let w refer to wages, x to characteristics other than union status. The definition of conditional probability yields the following representation of the overall distribution of wages.

$$f(w) = \int f(w|x)f(x)\,dx. \tag{10}$$

First consider the effect of unions. It is useful to define two other densities. First, the observed density of wages in the nonunion sector is given by

$$f(w|u=0) = \int f^n(w|x)f(x|u=0)\,dx, \tag{11}$$

where $f^n(w|x) \equiv f(w|x, u=0)$. Likewise, the observed density of wages in the union sector is given by

$$f(w|u=1) = \int f^u(w|x)f(x|u=1)\,dx, \tag{12}$$

where $f^u(w|x) \equiv f(w|x,u=1)$.

By analogy to the Oaxaca decomposition, we are interested in what distribution would prevail if all workers (not just nonunion workers) were paid under the wage structure in the nonunion sector, or, more formally:

$$f^n(w) = \int f^n(w|x)f(x)\,dx. \tag{13}$$

This equation is similar to equation (11) except that we integrate over the distribution of characteristics x for all workers ($f(x)$) instead of nonunion workers only ($f(x|u=0)$). Estimation of the above density can be made simple by noting that Bayes' Law implies

$$f(x) = \frac{f(x|u=0)\Pr(u=0)}{\Pr(u=0|x)}. \tag{14}$$

By substituting (14) into (13) we get

$$f^n(w) = \int f^n(w\mid x)\frac{f(x\mid u=0)\Pr(u=0)}{\Pr(u=0\mid x)}dx, \tag{15}$$

$$= \int \theta(x)f^n(w|x)f(x|u=0)\,dx, \tag{16}$$

where $\theta(x) = Pr(u=0) / Pr(u=0|x)$. Notice that equation (16) is identical to equation (11) except for the weighting factor $\theta(x)$, which is estimated by noting that $Pr(u=0)$ is merely the proportion of nonunion members in the sample, while $Pr(u=0|x)$ can be estimated by a discrete choice model like a Probit.[19] The weighting factors used in this paper were calculated using conventional "human capital" covariates entered in a fairly unrestricted way in a probit model.

Kernel density estimation with weights $\theta_i = \theta(x_i)$ is a straightforward modification of the usual Rosenblatt-Parzen Density estimator

$$\hat{f}(x) = \frac{1}{N}\sum_{i=1}^{N}\frac{\theta_i}{h}K\left(\frac{x-X_i}{h}\right),$$

where K refers to the kernel or weighting function. It is also easy to incorporate standard CPS (or LFS) weights in this setting by multiplying them by θ_i. Finally, note that our estimates were not sensitive to choice of reasonable variations in bandwidth or kernel.

[19] $Pr(u=0|x)$ can also be estimated nonparametrically by dividing up the sample by the characteristics x and calculating the proportion of individuals in each cell.

In the case of the minimum wage, a similar procedure can be used under the assumption that minimum wages have no spillover or employment effects, and that the shape of the wage distribution at or below the minimum wage does not depend on its level (see DFL for more details). Consider what would happen to the 1988 wage distribution if the minimum wage were raised to its higher 1981 level. Under the three above assumptions, a simulated distribution is obtained by replacing the 1988 observations at or below the real 1981 minimum wage by the corresponding 1981 observations weighted by the factor

$$\theta(z) = \frac{\Pr(t=88|z, w <= m_{81})\Pr(t=81)}{\Pr(t=81|z, w <= m_{81})\Pr(t=88)},$$

where t is a dating variable, z consists of the characteristics x and the union status variable u, and m_{81} is the 1981 value of the minimum wage. This weighting factor has to be used to account for possible changes in the distribution of characteristics of minimum wage and subminimum wage workers (first term) and for differences in the size of the 1981 and 1988 samples (second term). See DFL for more details.

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