What Effect Do Unions Have on Wages Now and Would Freeman and Medoff Be Surprised?*

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I. *Introduction*

> Everyone "knows" that unions raise wages. The questions are how much, under what conditions, and with what effects on the overall performance of the economy" (Freeman and Medoff, 1984, p. 43).

Richard Freeman and James Medoff's (F&M) pathbreaking 1984 book *What Do Unions Do?* has had an enormous impact. According to Orley Ashenfelter, one of the commentators in a review symposium on the book published in January 1985 in the *Industrial and Labor Relations Review*, the response of the popular press to the book "has only been short of breathtaking" (p. 245).[1] It received rave reviews at the time it was written and unlike most books has withstood the test of time. It is certainly the most famous book in labor economics and industrial relations. One of the other reviewers in the symposium, Dan Mitchell called it "a landmark in social science research" and so it has proved (p. 253). We went to the *Social Science Citations Index* and typed in "What do unions do" (hereinafter WDUD) and found that it had been cited by other academics more than one thousand times.[2] Herein we show that the vast majority of their commentary written in the early 1980s is still highly applicable despite the fact that private sector unionization has been in precipitous decline. An old adage is that a classic book is one that everyone talks about but nobody reads. F&M's work is not one of those. It is a true classic because it continues to be a book that anyone — scholar or layman — interested in labor unions needs to read!

Central to the thesis propounded by F&M is that there are two faces to unions — the undesirable monopoly face — which enables unions to raise wages above the competitive level which results in a loss of economic efficiency. This inefficiency arises because employers adjust to the higher union wage by hiring too few workers in the union sector. The second, more desirable face to be examined in detail by others in this symposium, is the collective voice face which enables unions to channel worker discontent into improved workplace conditions and productivity. Our study concentrates on the monopoly face of unions and its impact on relative wages. We explore the various claims made by F&M about how unions affect wages and update them with new and better data.

We examine in some detail the role of the public sector, which was largely ignored by F&M. This was a perfectly understandable omission at the time but is less appropriate today given the importance of public sector unionism in the United States.[3] In Section I we report F&M's main findings. In Section II we discuss the main labor market changes that have occurred since WDUD was written. Section III reports our estimates of wage gaps disaggregated by various characteristics used by F&M. We also examine wage gaps that F&M did not examine, namely those in the public sector and for immigrants. Section IV examines time series changes in the union wage gap. Section V models the determinants of changes in the union wage premium at the level of the industry, occupation, and state. Section VI outlines our main findings and discusses whether F&M would have been surprised about these findings when they wrote WDUD.

II. *Summary of F&M's Findings on Union Wage Effects*

F&M reported that early work on union wage effects used aggregate data on different industries, occupations, and areas. Much of this work was summarized in Lewis (1963). The reason that such aggregated data were used was that "data on the wages of union versus nonunion individuals or establishments was neither available nor, given the state of technology, readily amenable to statistical analysis" (1984, p. 44). These studies found a union wage effect on average of 10–15 percent. The more recent studies F&M examined, including a number of their own, used micro data at the establishment level but more usually at the individual level. In Table 1 F&M showed that the union differential in the 1970s was 20–30 percent using cross-sectional data (the seven numbers in the table averaged out at 25.3 percent). Such estimates may still suffer from bias because differences due to the skills and abilities of workers are wrongly attributed to unions. F&M also considered "before and after" comparisons and argued that, although they represent a way to eliminate ability bias they also suffer from measurement error problems derived from mismeasurement of the union status measure (Hirsch, 2003). F&M reported 12 estimates using panel data in their Table 2 for the 1970s: these are sizable but smaller than the cross-section estimates they examined, averaging out at 15.7 percent.[4]

F&M used data from the May 1979 Current Population Survey (CPS) to obtain a series of disaggregated estimates using a sample of nonagricultural, private sector, blue-collar workers aged 20–65. They reported that unions raise wages most for the young, the least tenured, whites, men, the least educated, blue-collar workers and in the largely unorganized South and West.[5] Furthermore, F&M found, using data for 62 industries from the 1973–1975 May CPS, that there was considerable variation in the size of the differential.[6]

F&M argued that the amount of union monopoly power is related to the wage sensitivity of the demand for organized labor. The smaller response of employment to wages the greater, they argued, is the ability of unions to raise wages without significant employment loss. Areas where employment is less responsive to wage changes, such as air transport, they argued should be where one would expect to find sizable wage gains.

F&M then argued that the differential likely depends on the extent to which the union is able to organize a big percentage of workers — the higher the percentage the higher the differential (p. 51). F&M found that for blue-collar workers in manufacturing a 10 percent increase in organizing generates a 1.5 percent increase in union wages. In contrast, they argued that the wages of nonunion workers do not appear to be influenced by the percentage of workers organized. In terms of the characteristics of firms and plants F&M obtained the following results: (a) union differentials depend on the extent to which the firm bargains for an entire sector rather than for individual plants within a sector; (b) wage differentials tend to fall with size of firm/plant/workplace; and (c) there was no clear empirical evidence on the relationship between product market power and differentials primarily as it is so difficult to measure power.

In terms of macro changes in differentials, F&M found that the 1970s were a period of increases in the union wage premium. F&M conjectured that a possible explanation was the sluggish labor market conditions then prevailing. Wages of union workers, they argued, tend to be less sensitive to business cycle ups and downs — particularly due to three-year contracts. This implies the union wage premium moves counter-cyclically — high in slumps when the unemployment rate is high and low in booms when the unemployment rate is low. However, F&M found that inflation and unemployment explained less than 50 percent of the rising union differentials in the 1970s. Nor did the rising wage differentials of the 1970s represent an historical increase in union power. The early 1980s, according to F&M, were a period of "give-backs" where unions agreed to wage cuts. Union wage gains were not a major cause of inflation.

F&M ended chapter three by estimating the social cost of monopoly power of unions. Loss of output due to unions they found to be "quite modest," accounting for between 0.2 percent and 0.4 percent of GNP or between $5 billion and $10 billion.

F&M drew six conclusions on the union wage effect: (a) The common sense view that there is a union wage effect is correct; (b) the magnitude of the differential varies across workers, markets, and time periods; (c) variation in the union wage gap across workers is best understood by union standard rate policies arising from voice; (d) variation in the union wage gap across markets is best understood by union monopoly power and employer product market power; and (e) wage premia in the 1970s were substantial but they returned to more "normal" levels in the 1980s; and (f) social loss due to unions is small.

III. *Changes in the Labor Market since WDUD*

Union density rates in the United States have fallen rapidly from 24 percent in 1977 to 13 percent in 2002 (Hirsch and Macpherson, 2002).[7] The decline was most dramatic in the private sector where in 2002 fewer than one in ten workers were union members. Density remains higher in manufacturing than in services. However, Table 1 suggests that union membership has roughly the same disaggregated pattern in 2001 as it did in 1977 — union density is higher among men than women; for older versus younger workers; in regions outside the South; and in transportation, communication, and construction. The exceptions are by race, where in 1977 rates were higher among nonwhites, but there is little difference by 2001, and by schooling. In 1977 membership rates for those with below high school education were nearly double those with above high school education. In 2001 they were approximately the same. So the highly qualified have increased their share of union employment.

The number of private sector union members declined between 1983 and 2002 from 11.9 million to 8.7 million while the number of public sector union members actually increased from 5.7 million to 7.3 million (Hirsch and Macpherson, 2003, Table 1c). Due to the growth in total employment in the public sector, however, the proportion of public sector workers who were union members was exactly the same in 2001 and 1983 (37 percent).[8] By 2002, 46 percent of all union members were in the public sector compared with 32.5 percent in 1983.

IV. *Union Wage Gaps since WDUD*

What has happened to the union wage differential between 1979 and 2001? Table 2 presents union wage gaps obtained from estimating a series of equations for each of the major sub-groups examined by F&M who used the 1979 May CPS file on a sample of nonagricultural about private sector, blue-collar workers aged 20–65. Their sample was very small, 6,000 observations. Rather than use the estimates reported by F&M to ensure large sample sizes we decided to pool together six successive May CPS files from 1974–1979 and compare those to wage gaps estimated for the years 1996–2001 using data from the Matched Outgoing Rotation Group (MORG) files of the CPS. Columns 1 and 2 estimate wage gaps for the private sector for 1996–2001 and 1974–1979, respectively. Columns 3 and 4 present equivalent estimates for the sample used by F&M of nonagricultural, private sector, blue-collar workers aged 20–65.

Hirsch and Schumacher (2002) show a "match bias" in union wage gap estimates due to earnings imputations.[9] This bias arises because workers in the CPS have earnings imputed using a "cell hot deck" method so wage gap estimates are biased *downward* when the attribute being studied (e.g., union status) is not a criterion used in the imputation. By construction, the individuals with imputed earnings have a union wage gap of about zero; hence omitting them raises the size of the union wage gap. They show that standard union wage gap estimates such as reported in Blanchflower (1999) are understated by about three to five percentage points as a result of including individuals with imputed earnings.

Table 1

Disaggregated Union Membership Rates, 1977 and 2001, in Percent

	1977	2001
All	24	14
Private Sector	22	9
Public Sector	33	37
Private Sector Employees		
Men	27	12
Women	11	6
Whites	20	9
Nonwhite	27	10
Ages 16-24	12	4
Ages 25-44	23	9
Ages 45-54	27	13
Ages >=55	22	10
< High School	23	7
High School	25	12
> High School	13	8
North East	24	12
Central	25	12
South	13	5
West	22	9
Agriculture, Forestry & Fisheries	3	2
Mining	47	12
Construction	36	19
Manufacturing	34	15
Transportation, Communication, and Other Public Utilities	48	24
Wholesale & Retail Trade	10	4
FIRE	4	3
Services	7	6

Source: 1977, *What Do Unions Do?* 2001 authors' calculations from the ORG file of the CPS.

Unfortunately, consistently excluding those individuals with imputed earnings over time is not a simple matter.[10] Herein we follow the procedure suggested by Hirsch and Schumacher (2002) and that we used in Blanchflower and Bryson (2003).[11] All allocated earners are identified and excluded for the years 1996–2001 in the MORG files. Because the May CPS sample files do not report allocated earnings in 1979–1981, the series are adjusted upward by the average bias of .033 found by Hirsch and Schumacher using these May CPS data for 1979–1981. Earnings were not allocated in the years 1973–1978. For the period 1973–1979 total sample size was approximately 184,000 compared with 547,000 for the later period. In each year from 1996–2001

Table 2

Private Sector Union/Nonunion Log Hourly Wage Differentials, 1974–1979 and 1996–2001, in Percent

	Private Sector		Freeman & Medoff's Sample	
	1974-1979	1996-2001	1974-1979	1996-2001
Men	19	17	27	28
Women	22	13	27	24
Ages 16-24	32	19	35	23
Ages 25-44	17	16	26	28
Ages 45-54	13	14	22	27
Ages >=55	19	16	29	28
Northeast	14	11	21	22
Central	20	15	27	27
South	24	19	29	26
West	23	22	31	34
< High school	33	26	31	29
High school	19	21	25	28
College 1-3 years	17	15	28	28
College >=4 years	4	3	17	14
Whites	21	16	28	27
Non-white	22	19	28	30
Tenure 0-3 years	20	20	28	n/a
Tenure 4-10	16	15	19	n/a
Tenure 11-15	10	11	12	n/a
Tenure 16+	17	8	28	n/a
Manual	30	21	n/a	n/a
Non-manual	15	4	n/a	n/a
Manufacturing	16	10	19	19
Construction	49	39	55	45
Services (excl. construction)	34	16	43	29
Private sector	21	17	28	28

Notes: 1996-2001 data files exclude individuals with imputed hourly earnings. Controls for 1996–2001 are 50 state dummies, 46 industry dummies, gender, 15 highest qualification dummies, private nonprofit dummy, age, age squared, log of weekly hours, four race dummies, four marital status, year dummies + union membership dummy (n=546,823). Estimates for 1974–1979 are adjusted upwards by the average bias found during 1979–1981 of .033. Controls for 1974–1979 are nine census division dummies, 46 industry dummies, years of education, age, age squared, log of weekly hours, four race dummies, four marital status dummies, five year dummies + union membership dummy (n=183,881). Tenure estimates for 1974–1979 obtained from the May 1979 CPS and for 1996–2002 files and February 1996 and 1998 Displaced Worker and Employee Tenure Supplements and January 2002 and February 2000: Displaced Workers, Employee Tenure, and Occupational Mobility Supplements. Freeman/Medoff's sample consists of non-agricultural private sector blue-collar workers aged 20-65 (n=64,034 for 1974-1979 and 142,024 for 1996-2001).

there are approximately 130,000 observations for the private sector in the MORG; in the May files, sample sizes are approximately 31,000.

Comparing F&M's sample and the wider private sector sample for the 1970s (columns 4 and 2, respectively, in Table 2), F&M's sample generates a larger wage gap for all, with the exception of the least educated.[12] The difference between the two samples is large, with the F&M sample generating a premium for the entire private sector, which is a third larger (28 percent as opposed to 21 percent) than in the wider sample. However, patterns in the wage gaps across workers are similar. (a) By *sex,* there is little difference in the size of the gap. (b) By *age,* the union effect is U-shaped in age and largest among the youngest who tend to be the lowest paid. (c) By *tenure,* the pattern is also U-shaped. (d) By *education,* unions raise wages most for the least educated, with the most highly educated having the lowest premium. (e) By *race,* unions raise wages by a similar amount for whites and nonwhites. (f) By *occupation,* although not reported in column 4, F&M (1984, pp. 49–50) report larger gains for blue-collar than for white-collar workers. The manual/non-manual gap in column 2 bears this out. (g) By *region,* unions had the largest effects in the relatively unorganized South and West, with more modest effects in the relatively well-organized Northeast. (h) By *industry,* construction and services have the highest premia.

What has happened since the 1970s?

(i) By *sex,* the wage gap has declined for women, but remained roughly stable for men so that, by the late 1990s, the union wage gap was higher for men than for women. The rate of decline in women's union premium is underestimated in F&M's restricted sample, but is still apparent. (j) By *age,* the U-shaped relationship apparent in the 1970s has disappeared because there has been a precipitous decline in the premium for the youngest workers, while the older workers' wage gap has remained roughly constant. In the full private sector sample, young workers still benefit most from unionization, though this is not apparent in F&M's restricted sample. (k) By *tenure,* as in the case of age, the U-shaped relationship between tenure and the union premium apparent in the 1970s has disappeared, because low- and high-tenured workers have seen their wage gap fall substantially while middle-tenure workers have experienced a stable union wage gap. Now, it seems the premium declines with tenure. (l) By *education,* the lowest educated continue to benefit most from union wage bargaining, but not to the same degree as in the 1970s. Although the trend is not so apparent in F&M's sample, the wage gap has fallen most for high school dropouts. (m) By *race,* a three-percentage point gap has opened up between the union premium commanded by nonwhites and the lower premium for whites. (n) By *occupation,* the union premium has collapsed for non-manual workers. Despite some decline in the premium for manuals, their wage gap was 17 percentage points larger than that for non-manuals by the late 1990s (compared with only five percentage points in the 1970s). (o) By *region,* the wage gap remains largest in the West and the South though, in F&M's sample, there is no difference in the premium in the South and Central regions. The wage gap remains smallest in the Northeast. (p) By *industry,* the wage gap remains

Table 3

Union Wage Differentials in the Public Sector, in Percent

	1983–1988		1996–2001	
	Wage Gap	Sample Size	Wage Gap	Sample Size
Private	22	(754,056)	17	(567,627)
Public	13	(165,276)	15	(110,833)
Federal	2	(33,633)	8	(20,938)
State	9	(42,942)	10	(34,919)
Local	16	(88,642)	20	(60,981)
Male	8	(77,528)	10	(48,298)
Female	17	(87,748)	16	(62,534)
Age <25	28	(15,603)	23	(7,771)
Age 25-44	13	(93,676)	15	(53,798)
Age 45-54	8	(32,127)	11	(33,830)
Age >=55	13	(23,870)	14	(15,433)
New England	17	(33,540)	17	(20,148)
Central	16	(38,863)	16	(25,930)
South	10	(51,785)	12	(33,522)
West	10	(41,088)	13	(31,232)
<High School	26	(13,217)	18	(29,775)
High School	15	(48,037)	13	(21,536)
College 1-3	13	(35,097)	11	(9,672)
College >= 4 Years	8	(68,925)	11	(50,029)
Whites	13	(131,676)	14	(85,893)
Nonwhites	15	(33,600)	16	(24,939)
Manual	18	(17,874)	18	(9,679)
Non-manual	13	(147,402)	14	(101,150)
Registered nurses (95)	5	(2,945)	6	(1,854)
Teachers (156-8)	15	(25,147)	21	(19,484)
Social workers (174)	12	(2,870)	12	(2,716)
Lawyers (178)	5	(1,014)	17	(1,184)
Firefighters (416-7)	15	(1,866)	19	(1,227)
Police & correction officers (418-424).	16	(6,068)	18	(5,503)

Notes: Sample excludes individuals with allocated earnings. Controls and data as in Table 2.

largest in construction and smallest in manufacturing. The decline in the differential was particularly marked in services. We return to industry differentials later.

Two points stand out from these analyses. First, *no group of workers in the broader private sector sample has experienced a substantial increase in its union premium.* Indeed, the only group recording any increase at all is those aged 45–54 whose premium rose from 13 percent to 14 percent. Clearly, unions have found it harder to maintain a wage gap since F&M wrote. Second, with the exception of the manual/non-manual gap, those with the highest premiums in the 1970s saw the biggest falls, so there has been some convergence in the wage gaps. This finding is apparent whether we compare trends using F&M's sample (columns 3 and 4) or the broader private sector sample (columns 1 and 2). This trend may be due to an increasingly competitive U.S. economy, where workers commanding wages well above the market rate are subject to intense competition from nonunion workers. Nevertheless, with the exception of the most highly educated and non-manual workers, the wage premium remains around 10 percent or more.

Public Sector. F&M said little or nothing on the role of unions in the public sector, although, as noted above, Freeman has subsequently written voluminously on the issue. Given that the remaining bastion of U.S. unionism is now the public sector, if F&M were writing today they would likely have devoted a considerable amount of space in a twenty-first century edition of WDUD to the public sector. More evidence on how the role of unions in the public sector has changed since WDUD was written is reported by Gunderson elsewhere in this symposium.

The size of the public sector grew (from 15.6 million to 19.1 million or 22.4 percent) between 1983 and 2001, but as a proportion of total employment it fell from 18.0 percent to 16.1 percent. Union membership in the public sector grew even more rapidly (from 5.7 million to 7.1 million or by 24.6 percent). Furthermore, by 2001 public sector unions accounted for 44 percent of all union members compared with 32.5 percent in 1983.

Table 3 is comparable to Table 2 for the private sector in that it presents disaggregated union wage gap estimates. Because sample sizes in the public sector are small using the May CPS files we again decided to use data from the ORG files of the CPS for the years 1983–1988 for comparison purposes with the 1996–2001 data. Data for the years 1979–1982 could not be used, as no union data are available. A further advantage of the 1983–1988 data is that information is available on individuals whose earnings were allocated who were then excluded from the analysis.

The main findings are as follows: (1) The private sector union wage gap has fallen over the two periods (21.5 percent to 17.0 percent) whereas a slight *increase* was observed in the public sector (13.3 percent to 14.5 percent, respectively); (2) the majority of the worker groups in Table 3 experienced *increases* in their union wage premium over the two periods, but wage gaps declined markedly for those under 25 and with less than a high school education; (3) there was little change in public sector union wage gaps for men or women. In marked contrast to the private sector where men had higher differentials than women, wage gaps in both periods in the public sector were *higher* for women than for men; (4) unions benefit workers most in local government and least in the federal government, although the differential for federal workers increased over time;[13] (5) just as for the private sector, the wage benefits of union membership are greatest for manual workers, the young, and the least educated; (6) there are only small differences in union wage gaps for nonwhites compared to whites in both the public and the private sectors; (7) in contrast to the private sector where wage differentials were greatest in the South and West, in the public sector exactly the opposite is found. Differentials are higher in the public sector in New England and the Central region in both time periods whereas the reverse was the case in the private sector; and (8) wage gaps increased over time for teachers, lawyers, firefighters and police.

Immigrants.[14] F&M also said nothing about the extent to which U.S. labor unions are able to sign up immigrants as members and by how much they are able to raise their wages. Using the data available in the CPS files since the mid–1990s we calculate wage gaps for the period 1996–2001. We find little variation in union wage gaps by length of time the immigrant had been in the United States, holding characteristics constant as well as wage gaps for the U.S.-born. However, differentials by source country are large. Differentials for Europeans (11.6 percent for Western Europe and 12.7 percent for Eastern Europe) are well below those of the native born (16.8 percent). Estimates are also in low double digits for Asians, Africans, and South Americans (13.3 percent, 11 percent, and 12.2 percent, respectively). In contrast the wage gap for Mexicans is 28 percent.

V. *Time Series Changes in the Union Wage Gap*

F&M reported that the 1970s was a period of rising differentials for unions, although they did not separately estimate year-by-year results themselves. Table 4, which is taken from Blanchflower and Bryson (2003),[15] reports adjusted estimates of the wage gap using separate log hourly earnings equations for each of the years from 1973 to 1981 using the National Bureau of Economic Research's (NBER) May Earnings Supplements to the CPS (CPS)[16] and for the years since then using data from the NBER's (MORG) files of the CPS.[17] The MORG data for the years 1983–1995 were previously used in Blanchflower (1999).[18] For both the May and the MORG files a broadly similar, but not identical, list of control variables is used, including a union status dummy, age and its square, a gender dummy, education, race, and hours controls plus state and industry dummies.[19]

The first column of Table 4 reports time-consistent estimates of union wage gaps for the total sample whereas the second and third columns report them for the private sector. To solve the match bias problem discussed above, as in Tables 2 and 3 we followed the procedure suggested by Hirsch and Schumacher (2002). Results obtained by Hirsch and Schumacher (2002) with a somewhat different set of controls are reported in the final column of the table. For a discussion of the reason for these differences, see Blanchflower and Bryson (2003). The time series properties of all three of the series are essentially the same.

The wage gap averages between 17 and 18 percent over the period, and is similar in size in the private sector as it is in the economy as a whole. The table confirms F&M's comment (1984, p. 53) that "the late 1970s appear to have been a period of substantial increase in the union wage premium." What is notable is the high differential in the early-to-mid 1980s and a slight decline thereafter, which gathers pace after 1995, with the series picking up again as the economy started to turn down in 2000.

Table 5 presents estimates of both the unadjusted and adjusted union wage gaps for the private sector. The sample excludes individuals with imputed earnings. In col-

Table 4

Union Wage Gap Estimates for the United States, 1973–2002, (%)
(excludes workers with imputed earnings)

Year	All Sectors Blanchflower/Bryson	Private Sector Blanchflower/Bryson	Private Sector Hirsh/Schumacher
1973	14.1	12.7	17.5
1974	14.6	13.8	17.5
1975	15.1	14.3	19.2
1976	15.5	14.6	20.4
1977	19.0	18.3	23.9
1978	18.8	18.6	22.8
1979	16.6	16.3	19.7
1980	17.7	17.0	21.3
1981	16.1	16.3	20.4
1983	19.5	21.2	25.5
1984	20.4	22.4	26.2
1985	19.2	21.0	26.0
1986	18.8	20.1	23.9
1987	18.5	20.0	24.0
1988	18.4	19.1	22.6
1989	17.8	19.2	24.5
1990	17.1	17.6	22.5
1991	16.1	16.6	22.0
1992	17.9	19.2	22.5
1993	18.5	19.6	23.5
1994	18.5	18.2	25.2
1995	17.4	18.0	24.5
1996	17.4	18.4	23.5
1997	17.4	17.7	23.2
1998	15.8	16.1	22.4
1999	16.0	16.9	22.0
2000	13.4	14.3	20.4
2001	14.1	15.1	20.0
2002	16.5	18.6	
1973-2001 average	17.1	17.6	22.4

Notes: Wage gap estimates calculated taking anti-logs and deducting 1. Columns 1 and 2 are taken from Table 3 of Blanchflower and Bryson (2003). Column 3 is taken from column 5 of Table 4 of Hirsch and Schumacher (2002). Data for 1973-1981 are from the May CPS Earnings Supplements. a) 1973-1981 May CPS, n=38,000 for all sectors, and n=31,000 for the private sector. Controls comprise age, age^2, male, union, years of education, 2 race dummies, 28 state dummies, usual hours, private sector and 50 industry dummies. For 1980 and 1981 sample sizes fall to approximately 16,000 because from 1980 only respondents in months 4 and 8 in the outgoing rotation groups report a wage. Since the May CPS sample files available to us do not include allocated earnings in 1979–1981, the series in columns 2 and 4 are adjusted upward by the average bias of 0.033 found by Hirsch and Schumacher (2002) using these May CPS data for 1979–1981. The data for 1973–1978 do not include individuals with allocated earnings and hence no adjustment is made in those years. b) Data for 1983–2002 are taken from the MORG files of the CPS. Controls comprise usual hours, age, age squared, four race dummies, 15 highest qualifications dummies, male, union, 46 industry dummies, four organizational status dummies, and 50 state dummies. Sample is employed private sector nonagricultural wage and salary workers aged 16 years and above with positive weekly earnings and non-missing data for control variables (few observations are lost). All allocated earners were identified and excluded for the years 1983–1988 and 1996–2001 from the MORG files. For 1989–1995, allocation flags are either unreliable (in 1989–1993) or not available (1994 through August 1995). For 1989–1993, the gaps are adjusted upward by the average imputation bias during 1983–1988. For 1994–1995, the gap is adjusted upward by the bias during 1996-1998. In each year there are approximately 160,000 observations for the U.S. economy and 130,000 for the private sector in the MORG; in the May files, sample sizes are approximately 38,000 and 31,000 respectively until 1980 and 1981 when sample sizes fall to approximately 16,000 and 13,000, respectively, as from that date on only respondents in months four and eight in the outgoing rotation groups report a wage. The Hirsch and Schumacher (2002) wage gap reported in column 3 is the coefficient on a dummy variable for union membership in a regression where the log of hourly earnings is the dependent variable. The control variables included are years of schooling, experience and its square (allowed to vary by gender), and dummy variables for gender, race and ethnicity (3), marital status (2), part-time status, region (8), large metropolitan area, industry (8), and occupation (12).

Table 5

The Ratio between Unadjusted and Adjusted Union Wage Gap Estimates for the United States, 1983–2002 (%)
(excludes workers with imputed earnings)

Year	Unadjusted	Adjusted	Unadjusted/Adjusted
1983	48.3	21.2	2.28
1984	48.3	22.4	2.16
1985	47.0	21.0	2.24
1986	44.8	20.1	2.23
1987	45.2	20.0	2.26
1988	44.6	19.1	2.34
1989	38.0	19.2	1.98
1990	34.3	17.6	1.95
1991	32.8	16.6	1.98
1992	32.5	19.2	1.69
1993	34.0	19.6	1.74
1995	34.6	18.0	1.92
1996	35.8	18.4	1.95
1997	36.1	17.7	2.04
1998	33.2	16.1	2.07
1999	32.5	16.9	1.92
2000	29.4	14.3	2.06
2001	29.8	15.1	1.98
2002	35.6	18.6	1.91

Notes: Column 1 obtained from a series of private sector log hourly wage equations that only contained a union membership dummy and a constant. Reported here is the antilog of the coefficient minus 1. Column 2 is from Table 4. Column 3 is column 1/column 2. Sample is employed private sector nonagricultural wage and salary workers aged 16 years and above with positive weekly earnings and non-missing data for control variables (few observations are lost)
Source: ORG files of the CPS, 1983–2001.

umn 1 of Table 5 we report the results of estimating a series of wage equations by year that only include a union dummy as a control. These numbers are consequently different from those reported by Hirsch and Macpherson (2002, Table 2a) who report raw unadjusted wage differences between the union and nonunion sectors but do not exclude individuals with imputed earnings.[20] Throughout the unadjusted wage gap is higher than the adjusted wage gap, implying a positive association between union membership and wage-enhancing employee or employer characteristics. However, the unadjusted gap has declined more rapidly than the regression-adjusted gap since 1983. In 1983 the unadjusted estimate was 128 percent higher than the adjusted estimate. In 2002 the difference had fallen to 91.5 percent higher.

To establish what is driving this effect, Hirsch et al. (2002) decompose the unadjusted wage gap into its three components — employment shifts, changes in worker characteristics, and changes in the residual union wage premium. Using CPS data for the private sector only, they find almost half (46 percent) of the decline in the union-nonunion log wage gap over the period 1986–2001 is accounted for by a decline in the regression-adjusted wage gap. Sixteen percent of the decline is accounted for by changes in worker characteristics and payoffs to those characteristics, chief among these is the increase in the union relative to nonunion percentage of female workers. The remaining 38 percent of the decline in Hirsch et al.'s unadjusted wage gap was due to sectoral shifts and payoffs to the occupational sectors of workers. The sectoral changes that stand out are the substantial decline in union relative to nonunion employment in durable manufacturing, and the decline in relative pay (that is, the industry coefficient) in transportation, communications and utilities, a sector with a large share of total union employment.

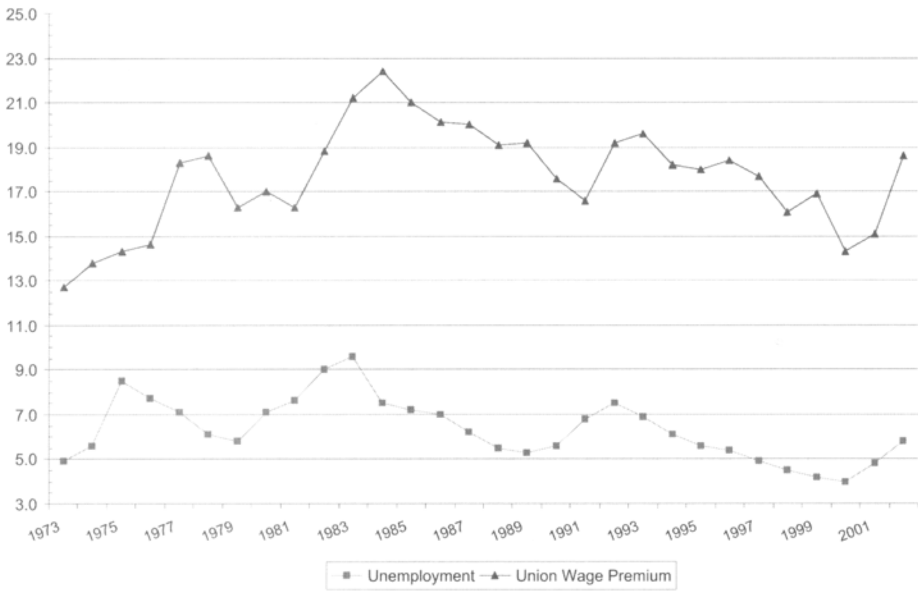
The results reported in Table 4 are broadly comparable to the estimates obtained by Lewis (1986) in his Table 9.7, which summarized the findings of 165 studies for the period 1967–1979. Lewis concluded that during this period the U.S. mean wage gap was approximately 15 percent. His results are reported in Table 6.[21] The left panel contains estimates for the six years prior to our starting point in Table 4. It appears that the unweighted average for this first period, 1967–1972, of 14 percent is slightly below the 16 percent for the second interval, 1973–1979. The estimates for the later period are very similar to those shown in Table 4 — which also averaged 16 percent — and have the same time-series pattern. In part, Lewis's low number for 1979 is explained by the fact that the 1979 May CPS file included allocated earners and hence the estimates were not adjusted for the *downward* bias caused by the imputation of the earnings data.[22]

Figure 1 plots the point estimates of the U.S. union wage premium, taken from the first column of Table 5, against unemployment for 1973–2002. The premium moves counter-cyclically. There are three main factors likely influencing the degree of counter-cyclical movement in the wage gap. The first, cited by F&M (1984, pp. 52–53) as the reason for the widening wage gap during the Depression of the 1920s and 1930s, is the greater capacity for union workers to "fight employer efforts to reduce wages" when market conditions are unfavorable. Conversely, when demand for labor is strong, employees rely less on unions to bargain for better wages because market rates rise anyway. The second factor is that union contracts are more long term than nonunion ones and, as such, less responsive to the economic cycle, so union wages respond to economic conditions with a lag.

When inflation is higher than expected, a greater contraction in the premium can occur because nonunion wages respond more to higher inflation. However, the third factor, which should reduce the cyclical sensitivity of the union wage premium, is the cost-of-living-adjustment (COLA) clauses in union contracts that increase union wages in response to increases in the consumer price level. According to F&M (1984, p. 54) the percentage of union workers covered by these agreements rose dramatically in the 1970s, from 25 percent at the beginning of the decade to 60 percent at the end of the decade. However, F&M's estimates for manufacturing suggest that COLA provisions "contributed only a modest amount to the rising union advantage" in the 1970s. Bratsberg and Ragan (2002) revisit this issue and find the increased sensitivity of the pre-

Figure 1

Movements in the U.S. Private Sector Wage Premium, 1973-2002



mium to the cycle is due in part to reduced COLA coverage from the late 1980s, but we find no such evidence (see below).

Commenting on the growth of the union wage premium during the 1970s, F&M (1984, p. 54) suggested that "at least in several major sectors the union/nonunion differential reached levels inconsistent with the survival of many union jobs." They were right. In the 1970s and early 1980s, the wage gap in the private sector rose while union density fell, as predicted in the standard textbook model of how employment responds to wages where the union has monopoly power over labor supply. In the classic monopoly model, demand for labor is given, so a rise in the union premium results in a decline in union membership since the premium hits employment. The fact that unions pushed for, and got, an increasing wage premium over this period, implies that they were willing to sustain membership losses to maintain real wages, or that unions were simply unaware of the consequences of their actions.

From the mid-1990s, the continued decline in union density was accompanied by a *falling* union wage premium because demand for union labor fell as a result of two pressures. The first was increasing competitiveness throughout the U.S. economy: Increasing price competition in markets generally meant employers were less able to pass the costs of the premium onto the consumer, so that pressures for wages to conform to the market rate grew. Second, union companies faced greater nonunion competition. Declining union density, by increasing employers' opportunities to sub-

Table 6

U.S. Mean Wage Gap: 1967–1979

Year	# Studies	Mean Estimate	Year	# Studies	Mean Estimate
1967	20	14%	1973	24	15%
1968	4	15%	1974	7	15%
1969	20	13%	1975	11	17%
1970	8	13%	1976	7	16%
1971	20	14%	1977	10	19%
1972	7	14%	1978	7	17%
			1979	3	13%

stitute nonunion products for union products, fueled this process. So too did rising import penetration: If imports are nonunion goods, regardless of U.S. union density, they increase the opportunity for nonunion substitution. These same pressures also increased the employment price of any union wage gap (the elasticity of demand for union labor).

VI. *Industry, Occupation, and State-Level Wage Premia*

So far, we have focused primarily on union wage effects at the level of the individual and the whole economy. However, the literature on the origins of the union wage premium focuses largely on firms and industries because the conventional assumption is that unions can procure a wage premium by capturing quasi-rents from the employer (Blanchflower et al., 1996). If this is so, there must be rents available to the firm arising from its position in the market place, and unions must have the ability to capture some of these rents through their ability to monopolize the firm's labor supply. Individual-level data can tell us little about these processes. Instead, the literature has concentrated on industry-level wage gaps. In this section we model the change in the union wage premium at three different units of observation — industry, state, and occupation.

Industries. As we noted above, F&M reported wage gap estimates by the extent of industry unionism. They (1984, p. 50) comment on substantial variation in the union wage effect by industry, with gaps ranging between 5 percent and 35 percent in the CPS data for 1973–1975. F&M's results are reported in the first column of Table 7. We used our data to estimate separate results by two-digit industry for 1983–1988 and 1996–2001. We chose these years as it was possible to define industries identically using the 1980 industry classification. Using these data we also found considerable variation by the size of the wage gap by industry as shown in Table 7. There is less variation in the wage gap by industry in the later period than in the earlier period with only three industries, construction (41 percent), transport (36 percent), and repair services (37 percent) having a differential of over 35 percent, compared with six in the earlier period which includes the same three — construction (52 percent); transport (44

Table 7

Union Wage Effects by Industry Using CPS Data

Estimates by Industry	FM 1973–1975	1983–1988	1996–2001
<5%	13	11	10
5–15%	17	15	19
15–35%	24	12	12
>=35%	8	6	3
# Industries	62	44	44

percent) and repair services (37 percent) — plus agricultural services (41 percent); other agriculture (56 percent) and entertainment (47 percent).[23]

Where is the union wage premium rising, and where is it falling? We estimated the regression-adjusted wage gaps in 44 industries during the 1980s (1983–1988) and then in the late 1990s (1996–2001). In contrast to the analysis by worker characteristics, which reveal near universal decline in the premium — at least in the private sector — we found that the wage gap rose in 17 industries and declined in 27 — results are presented in an appendix available on request from the authors. The gap rose by more than ten percentage points in autos (+12 percent) and leather (+19 percent). It declined by more than 20 percentage points in other agriculture (–33 percent) retail trade (–20 percent) and private households (–29 percent). Many of the industries experiencing a rise in the union premium between 1983 and 2001 would have been subject to intensifying international trade (machinery, electrical equipment, paper, rubber and plastics, leather) but this is equally true for those experiencing declining premiums (such as textiles, apparel, and furniture). Horn (1998) found that increases in import competition increased union density and decreased in the wage premium within manufacturing industries. This occurred because union density fell slower than overall employment when faced with import competition. Horn also found that imports from OECD countries decreased union density; imports from non-OECD countries tended to raise union density within an industry.

There is a negative correlation between change in union density and change in the premium (correlation coefficient –0.39). Some of the biggest declines in the premium have been concentrated in sectors where the bulk of private sector union members are concentrated, as Table 8 indicates. It shows the three industries with more than a 10 percent share in private sector union membership in 2002. In construction and transport, which both make up an increasing proportion of all private sector union members, the premium fell by around 10 percentage points. In retail trade, where the share of private sector union membership has remained roughly constant at 10 percent, the premium fell 20 percentage points. The decline in the wage gap for the whole economy, presented earlier, is due to the fact that the industries experiencing a decline in their wage gap make up a higher percentage of all employees than those experiencing a widening gap. The results are similar to those presented by Bratsberg and Ragan (2002) who found that, over the period 1971–1999, the regression-adjusted wage gap

Table 8

All Private Sector Union Members: Share of Membership

	Share of Membership, 1983	Share of Membership, 2002	Change in Premium, 1983–2001
Construction	9.3	13.5	-10.7
Retail Trade	10.2	10.5	-20.3
Transport	9.7	12.3	-8.0

closed in 16 industries and increased in 16 others. Their analysis is not directly comparable to ours, but where industry-level changes are presented in both studies, they tend to trend in the same direction. Only in one industry (transport equipment) do Bratsberg and Ragan report a significant increase in the wage gap where we find a decline in the wage gap.

These changes in the union wage premium by *industry* over time are worth detailed investigation, even though F&M did not present such analyses. Our first step was to estimate 855 separate first-stage regressions, one for each of our 45 industries in each year from 1983–2001 with the dependent variable the log hourly wage along with controls for union membership, age, age squared, male, four race dummies, the log of hours, and 50 state dummies. The sample was restricted to the private sector and excluded all individuals with allocated earnings. Three sectors with very small sample sizes (toys, tobacco, and forestry and fisheries) were deleted. We extracted the coefficient on the union variable, giving us 19 years * 42 industries or 798 observations in all. The adjustments discussed earlier were made to deal with imputed earnings. The coefficient on the union variable was then turned into a wage gap taking anti-logs, deducting 1 and multiplying by 100 to turn the figure into a percentage. We used the ORG files to estimate the proportion of workers in the industry who were union members both in the private sector and overall and mapped that onto the file. Unemployment rates at the level of the economy are used as industry-specific rates are not meaningful: Workers move a great deal between industries and considerably more than they do between states. (A table providing information on the classification of industries used and the average number of observations each year is included in the data appendix available on request from the authors.) Regression results, reported in Table 9, columns 1 and 2, estimate the impact of the lagged premium, lagged unemployment, and a time trend on the level of the industry-level wage premium. The number of observations is 756 as we lose 42 observations in generating the lag on the wage premium and the union density variables.

In the unweighted equation in column (1) the lagged premium is positively and significantly associated with the level of the premium the following year indicating regression to the mean. Unemployment and the time trend are not significant. However, once the regression is weighted by the number of observations in the industry in the first-stage regression, (column (2)) lagged unemployment is positive and signifi-

Table 9

Industry, State, and Occupation-Level Analysis of the Private Sector Union Wage Premium, 1983–2001

	(1)	(2)	(3)	(4)	(5)	(6)
Level of Analysis	Industry	Industry	State	State	Occupation	Occupation
Premium$_{t-1}$	.2584*	.3453*	.2051*	.2366*	.0907*	.1746*
	(.0367)	(.0350)	(.0337)	(.0333)	(.0379)	(.0374)
Unemployment rate$_{t-1}$	.6333	.5866*	.4373*	.5366*	.3799	.5823*
	(.4035)	(.2821)	(.1449)	(.1175)	(.5084)	(.2900)
Time	-.0463	-.2344*	-.1547*	-.0651	-.3419*	-.2416*
	(.1056)	(.0762)	(.0468)	(.0379)	(.1343)	(.0788)
State/industry/ occupation dummies	50	50	41	41	41	41
Weighted by # obs at 1st stage	No	Yes	No	Yes	No	Yes
R^2	.6187	.7749	.5071	.5861	.7345	.8453
N	756	756	918	918	756	756

Source: Outgoing Rotation Groups of the CPS, 1984-2001. Samples exclude individuals with imputed earnings.

cant, indicating counter-cyclical movement in the premium, while the negative time trend indicates secular decline in the premium.

Bratsberg and Ragan (2002) reported that the industry-level premium was influenced by a number of other variables.[24] In particular they found that COLA clauses reduced the cyclicality of the union premium and that increases in import penetration were strongly associated with rising union premiums.[25] They also found some evidence that industry deregulation had mixed effects. Their main equations (their Table 2) did not include a lagged dependent variable. Table 10 reports results using their data for the years 1973–1999 using their method and computer programs that they kindly provided to us. Column 1 of Table 10 reports the results they reported in column 2 of their Table 2. Column 2 reports our attempt to replicate their findings. We are unable to do so exactly — the problem appears to arise from the use of the xtgls routine in STATA which gives different results on our two machines.[26] There are several similarities — we find import penetration both in durables and nondurables, COLA clauses, deregulations in communications, and the unemployment rate all have positive and significant effects. We also found, as they did, that deregulation in finance lowered the premium. In contrast to Bratsberg and Ragan, however, the inflation rate and the two interaction terms with the unemployment rate were insignificant. The model is rerun in column 3, but without the insignificant interaction term. A linear time trend is added in column 4: this is negative and significant, and eliminates the COLA effect and the

Table 10

Industry Level Analysis of the Union Wage Premium in the Private Sector, 1973–1999

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Premium$_{t-1}$					.6030*	.2759*	.6001*	.2468*	.3196*
					(.0274)	(.0350)	(.0284)	(.0361)	(.0333)
Time				-.0019*	-.0012*	.0002	-.0011*	-.0001	-.0009*
				(.0004)	(.0003)	(.0004)	(.0003)	(.0004)	(.0003)
Unemployment rate	.0187*	.0131*	.0108*	.0083*	.0064*	.0064*	.0061*	.0070*	.0052*
	(.0017)	(.0017)	(.0011)	(.0014)	(.0010)	(.0021)	(.0011)	(.0022)	(.0010)
COLA	.0763*	.0767*	.0403*	.0155	-.0065	.0139	.0041	.0156	.0141*
	(.0313)	(.0303)	(.0126)	(.0134)	(.0090)	(.0140)	(.0108)	(.0144)	(.0096)
Inflation	-.0182*	-.0077	.0012	.0006	.0024*	.0026	.0020*	.0032	.0002
	(.0065)	(.0069)	(.0008)	(.0008)	(.0007)	(.0015)	(.0008)	(.0016)	(.0008)
*Unempt rate*COLA*	-.0092*	-.0047							
	(.0038)	(.0036)							
*Unempt rate*Inflation*	.0026*	.0012							
	(.0009)	(.0009)							
Import penetration	.2048*	.2201*	.2362*	.3090*	.1688*	.1234*	.1738*	.1668*	.1811*
Durables	(.0427)	(.0414)	(.0441)	(.0424)	(.0326)	(.0416)	(.0461)	(.0549)	(.0302)
Import penetration	.1655*	.1459*	.1491*	.1698*	.0939*	.0880*	.0914*	.0945*	.1043*
Nondurables	(.0513)	(.0525)	(.0509)	(.0488)	(.0302)	(.0208)	(.0419)	(.0265)	(.0314)
Dereg.Communications	.0752*	.0609*	.0589*	.0612*	.0451*	.0625*	.0506*	.0734*	.0532*
	(.0316)	(.0244)	(.0246)	(.0248)	(.0200)	(.0307)	(.0234)	(.0261)	(.0193)
Deregulation Rail	.0329	.0400	.0394	.0580	.0200				.0333
	(.0905)	(.0844)	(.0855)	(.0839)	(.0616)				(.0606)
Deregulation Trucking	-.0716	-.0617	-.0630	-.0394	-.0139				-.0332
	(.0560)	(.0570)	(.0565)	(.0518)	(.0429)				(.0398)
Deregulation Air	.0554	.0684	.0661	.0815	.0087				.0214
	(.1262)	(.1217)	(.1190)	(.1161)	(.0852)				(.0804)
Deregulation Finance	-.0614*	-.0599*	-.0587*	-.0329	.0179				-.0174
	(.0191)	(.0188)	(.0195)	(.0203)	(.0160)				(.0150)
Weighted	Yes	Yes	Yes	Yes	Yes	No	Yes	No	Yes
Method	GLS	GLS	GLS	GLS	GLS	GLS	OLS	OLS	GLS
Wald Chi^2/ R^2	2,325.01	2,781.32	2,686.37	3,190.74	10,961.71	1,220.21	.8973	.6516	6,189.4
N	832	832	832	832	832	832	832	832	806

Notes: All equations also include a full set of 31 industry dummies. Data are taken from Bratsberg and Ragan 2002. GLS regression estimated with industry specific AR(1) process in error term. Where indicated ach observation in the GLS regressions is weighted by the industry observation count of the first step following Berntsberg and Ragan (2002). Column 9 excludes Retail Trade. Standard errors in parentheses.

negative effect of deregulation in the finance sector. Column 5 adds the lagged union wage premium, which is positive and significant. Its introduction makes inflation positive and significant. In columns (6) to (8) models are run without the four insignificant deregulation dummies. Column (6) indicates that using an unweighted regression, the size of the lagged premium effect drops markedly and the time trend and inflation lose significance, showing these results are sensitive to the weighting of the regression. The smaller coefficient on the lagged dependent variable is unsurprising given that there is much less likely to be variation in the union wage gap estimates in industries with large sample sizes that have higher weights in the former case. We are able to confirm Bratsberg and Ragan's finding that the unemployment rate, deregulation in communications, and import penetration in both durables and nondurables have positive impacts on the premium but not the findings on COLA, inflation, or any of the other deregulations identified.

That import penetration in durable and nondurable goods sectors increases the premium suggests that union wages are more resilient than nonunion wages to foreign competition. Import penetration is likely correlated with unmeasured industry characteristics that depress the premium inducing a negative bias that is removed once industry characteristics are controlled for. Import penetration has likely reduced demand for union and nonunion labor, with union wages holding up better than nonunion wages, but at the expense of reduced union employment. There are theoretical and empirical reasons as to why this might occur. For instance, since union wages tend to be less responsive to market conditions generally, union wages may be sluggish in responding to increased import competition. Alternatively, industries characterized by "end-game" bargaining may witness perverse union responses to shifts in product demand as the union tries to extract maximum rents in declining industries (Lawrence and Lawrence, 1985). Another possibility is that increased import penetration reduces the share of union employment in labor-intensive firms and increases it in capital-intensive firms. Greater capital intensity reduces elasticity of demand for union labor, allowing rent-maximizing unions to raise the premium (Staiger, 1988).

It isn't obvious that weights should be used if we regard each industry as a separate observation. In cross-country comparisons which, say, contrasted outcomes for Switzerland, the United Kingdom, and the United States, it wouldn't make a lot of sense to weight by population and thereby make the observation from the United States 4.67 times more important than that of the United Kingdom and 39.3 times more important than Switzerland.[27] Columns (1) to (6) are GLS estimates accounting for potential correlation in error terms. Column (7) switches to a weighted OLS and shows that results are not sensitive to the switch. The unweighted OLS in column (8) gives broadly the same results as the unweighted GLS in column (6). Taking off the weights has a much bigger effect than switching from GLS to OLS.

Furthermore, the industries defined by Bratsberg and Ragan are *very* different in size. Some industries are very broadly defined — for example industry 32 Services covers SIC codes 721–900 whereas tobacco, for example, covers one SIC code (130). Retail trade averaged 19,075 observations. Column 9 of Table 10 illustrates the sen-

Table 11

Union Wage Effects by State Using CPS Data

Estimates by State	1983–1988	1996–2001
<5%	0	0
5–15%	6	21
15–35%	43	30
>=35%	2	0
# States (+D.C.)	51	51

sitivity of the results to industry exclusions. It is exactly equivalent in all respects to column 5 of Table 10 except that it drops the 32 observations from retail trade. The lagged dependent variable falls dramatically from .60 to .32. The COLA variable is now significantly *positive* while the inflation variable moves from being significantly positive to insignificant. The unweighted results (not reported) are little changed. Bratsberg and Ragan's results appear to be sensitive to both the use of weights and the sample of industries used.

States. In the United States unions are geographically concentrated by town, county, district, and state. Often towns next to each other differ — one is a union town, the other is nonunion. Waddoups (2000) used this interesting juxtaposition of union and nonunion zones to estimate the impact of unions on wages in Nevada's hotel-casino industry.[28] Although they share many features and are subject to broadly similar business cycles, most of the 50 states in the United States are comparable in size and economic significance to many countries. They also differ markedly in their industrial structures and unionization rates. Assuming union density proxies union bargaining power, this implies different premiums across states. However, as noted earlier, F&M found the union premium at the regional level was inversely correlated with union density, with the premium highest in the relatively unorganized South and West. To explore this issue further, and to assess changes over time, we estimated separate wage gaps for two time periods at the level of the 50 states plus Washington, D.C. Results at the level of the state were also estimated and are summarized in Table 11 below — full state-level results are available in the appendix available on request from the authors. The data used are from the Outgoing Rotation Group files of the CPS. It was not possible to identify each state separately in the May CPS, so F&M did not report such results. Hence, we compared results from a merged sample of the 1983–1988 with those obtained from our 1996–2001 files.[29] The correlation between changes in state density and state premia is negative but small (–0.10).

First we notice that *the variation in the union wage premium is much less by state than it is by industry.* Only two states in the earlier period had gaps of at least 35 percent — North Dakota (35 percent) and Nebraska (37 percent) and none in the later period. There has been a downward shift in the premium generally, as indicated by the movement from the 15–35 percent category to the 5–15 percent category. The mean state union wage gap was 23.4 percent between 1983 and 1988, falling by 6.2 percentage points to 17.2 percent in 1996–2001. The premium fell in all but five states, with South Dakota recording the biggest decline (16.8 percentage points). In four of the five states where the premium rose, it only increased by a percentage point or two (Vermont, Massachusetts, Wyoming, and Hawaii). The premium only rose markedly in Maine, where it increased 9 percentage points (from 7 percent to 16.1 percent). Since the early 1980s, union density fell by an average of 5.7 percentage points, with Pennsylvania (–10.6 percent) and West Virginia experiencing the biggest decline (–11 percentage points). The premium appears to have declined more in smaller states than it has in bigger states. The five biggest states of California, Texas, Florida, New York, and Illinois had small changes in their wage gaps (–1.4 percent; –6.7 percent; –10.1 percent; –0.6 percent and –4.5 percent, respectively). The five smallest states measured by employment tended to have big declines in the differentials: New Mexico (–14.10 percent); Alabama (–14.20 percent); Nebraska (–15.00 percent); Arkansas (–15.20 percent); South Dakota (–16.80 percent).[30]

We then ran 969 separate first-stage regressions, one for each state in each year from 1983–2001 with the dependent variable the log hourly wage along with controls for union membership, age, age squared, male, four race dummies, the log of hours, and 44 industry dummies. The sample was restricted to the private sector, and allocated earnings were dealt with as described earlier. We extracted the coefficient on the union variable, giving us 19 years * 51 states (including D.C.), 969 observations in all. We then mapped to that file the unemployment rate in the state-year cell.[31] Once again we ran a series of second-stage regressions where the dependent variable is the one-year level of the premium (obtained by taking anti-logs of the union coefficient and deducting one) on a series of RHS variables including the lagged premium and lagged unemployment and union density rates.[32] Results are reported in columns (3) and (4) of Table 9. The number of observations is 918 — we lose 51 observations in generating the lag on the wage premium and the union density variables. Both unweighted and weighted results are presented where the weights are total employment in the state by year. Controlling for state fixed effects with 50 state dummies we find that with an unweighted regression (column (3)), the lagged premium is positive and significant, as it was at industry level. Again, as in the case of industry-level analysis, the effect is apparent when weighting the regression (column (4)). The positive, significant effect of lagged state-level unemployment confirms the counter-cyclical nature of the premium: The effect is apparent whether the regression is weighted or not. There is also evidence of a secular decline in the state-level premium, but only where the regression is unweighted.

State fixed effects account for state-level variance in union density where the effect is fixed over time. However, Farber (2003) argued that there remain potential unobserved variables which simultaneously determine density and wages, but which are time-varying, and thus not picked up in fixed effects, which might bias our results.

Occupations. Finally, we moved on to estimate wage gaps at the level of the occupation pooling six years of data for each of the time periods 1983–1988 and 1996–2001.

Table 12

Union Wage Effects by Occupation Using CPS Data

Estimates by Occupation	1983–1988	1996–2001
<5%	10	10
5–15%	9	9
15–35%	14	17
>=35%	11	8
# Occupations	44	44

In each case we used files from the Outgoing Rotation Group files of the CPS. As with our estimates by industry, there is considerable variation by occupation both in the first period and the second. The variation is greater than was found when the analysis was conducted at the level of states — once again results by occupation are reported in an appendix available on request from the authors. The results are summarized in Table 12. In the first period 11 occupations had wage gaps over 35 percent — primarily manual occupations. In the second seven of these occupations still had gaps of over 35 percent. Out of the 44 groups, 13 showed increases in the size of the differential over time while the remainder had decreases. We used the same method described above for industries and states, with occupations defined in a comparable way through time. Columns (5) and (6) of Table 6 show that whether the occupation-level analysis is weighted or not, there is clear evidence of regression to the mean, with the lagged premium positive and significant, as well as evidence of a secular decline in the premium. A significant counter-cyclical effect is evident when the regression is weighted, but not in the unweighted regression.

In all three units of observation we have used — industry, state, and occupation — there is evidence that the private sector premium moves counter-cyclically and that it has been declining over time. In all three cases the lagged level of the premium entered significantly positively and was larger when the weights were used than when they were not. The size of the lag was greatest when industries were used as the unit of observation and least when occupations were used. Translating the results from levels into changes — that is by deducting t-1 from both sides — leaves all of the other coefficients unchanged. Using the weighted results in Table 6 the results reported below imply mean convergence.

State level	$\Delta\text{Premium}_{t\text{-}t\text{-}1}$	=	$\text{-.7949Premium}_{t\text{-}1}$
Industry level	$\Delta\text{Premium}_{t\text{-}t\text{-}1}$	=	$\text{-.6457Premium}_{t\text{-}1}$
Occupation level	$\Delta\text{Premium}_{t\text{-}t\text{-}1}$	=	$\text{-.8254Premium}_{t\text{-}1}$

The *higher* the level of the premium in the previous period the lower the change in the next period.

VII. *What Have We Learned and Would F&M Have Been Surprised about These Results When They Wrote WDUD?*

The Private Sector Wage Premium is Lower Today Than It Was in the 1970s. This would not have surprised F&M. Indeed, they predicted that the premiums of the 1970s were unsustainable due to their impact on union density (F&M, 1984, p. 54). Perhaps it is surprising that the premium remains as high as it has. One possibility is that, even though union bargaining power has declined, union density continues to decline, implying that there is some employment spillover into the nonunion sector. If wage setting in the nonunion sector is more flexible than it used to be, this additional supply of labor to the nonunion sector may depress nonunion wages more so than in the past, keeping the premium higher than anticipated.

The Union Wage Premium is Counter-Cyclical. The decline in the premium in good times is what seems to explain much of the decline in the premium since the mid–1990s. Far from being a surprise to F&M, they identified the counter-cyclical nature of the premium. We show the premium is counter-cyclical at the state, occupation, and industry levels. F&M said COLA's could dampen counter-cyclical movements in the premium, but thought their significance had been overplayed. This is confirmed: we find no COLA effect, though this finding is contested by others.

There Is Evidence of a Secular Decline in the Private Sector Union Wage Premium. There is evidence at state, industry, and occupation level of a downward trend in the private sector union wage premium accompanying the marked decline in union presence in the private sector. The effect is sensitive to weighting in the case of the state-level and industry-level premia, but not in the case of the occupational premia. Interestingly, the wage gap appears to have declined most in the smallest states (e.g., New Mexico, Alabama, Nebraska, Arkansas, and South Dakota) and declined least in the bigger states (e.g., California, Texas, New York, and Illinois). It would not have been a surprise to F&M that there had been some reduction in the ability of unions over time to raise wages as the proportion of the work force they bargain for has declined.

There Remains Big Variation in the Premium across Workers. Patterns in the premium across worker types resemble those found by F&M. The F&M sub-sample generally overstated the size of the premium in the population as a whole, but we suspect this would not have surprised F&M. A decline in the premium over time seems to have occurred across all demographic characteristics in the private sector, but there is regression to the mean, the biggest losers being among the most vulnerable and certainly the lowest paid workers (the young, women, and high school dropouts). But perhaps the real surprise is just how large the premium still is for some of these workers (26 percent for high school dropouts, 19 percent for under 25 years old). One puzzle is the remarkable rise in the share of union employment taken by the highly qualified, yet they continue to receive the lowest union premium. Why? Do they look for something else from their unions, e.g. professional indemnity insurance or voice, or are their unions less effective? In contrast to the private sector, the public sector has experienced a small increase in the premium, an increase apparent for most public sector employees. In the public sector industry-level bargaining remains the norm, maintaining union bargaining power. So, perhaps F&M would not have been surprised by this result. We look at one group of workers that F&M did not consider: immigrants. The premium varies little by year of entry to the United States, but does depend on the country of birth, with Mexicans benefiting from the highest premium. Whether this reflects the human capital, occupational mix, or costs of immigration faced by different groups is a matter for further research, but we suspect the results would not have surprised F&M.

There Is Big Variation in Industry-Level Union Wage Premia. F&M also found wide variation in industry-level premiums and might have expected this to persist because unions' ability to push for a premium, and employers' ability to pay, is determined by industry-specific factors (such as union organization and the availability of nonunion labor, regulatory regimes, bargaining, and product market rents). However, we find convergence in industry-level premiums since F&M wrote, that is, a falling premium where it was once large, and a rising premium where it was once small. (Overall decline at the economy level is due to the fact that the former constitutes a larger share of employment than the latter.) F&M may well have been surprised by this regression to the mean, because in 40 percent of the industries we examined there was a rise in the premium.

State-Level Union Wage Premia Vary Less Than Occupation- and Industry-Level Premia. F&M did not explicitly compare variations in the premia at the state and industry levels. Freeman expressed the view to us that it made sense to him that there would be more variation at the occupation and industry levels as they are more closely approximated to markets. No surprise here.

Union Workers Remain Better Able Than Nonunion Workers to Resist Employer Efforts to Reduce Wages When Market Conditions Are Unfavorable. Import penetration is a good proxy for competition in the traded goods sector. Although the impact of imports on the U.S. wage distribution is often overstated (Blanchflower, 2000a), it may be expected to play a role where imports permit substitution of union products for nonunion products. If imports reduce demand for domestic output and, in turn, demand for labor, this should reduce union and nonunion wages (assuming the supply of nonunion labor is not perfectly elastic). Whether the premium rises or falls with increased import penetration depends on the relative responsiveness of union and nonunion wages to demand shifts resulting from foreign competition. Unions' ability to resist employer efforts to reduce wages when market conditions are unfavorable was cited as one reason for the counter-cyclical premium by F&M (1984, pp. 52–53).

There Has Been a Decline in the Unadjusted Wage Gap Relative to the Regression-Adjusted Wage Gap. Union members do have wage-enhancing advantages over nonmembers, but these have diminished in recent years, implying changes in the selection of employees into membership. It is unlikely that F&M would have predicted this.

Public Sector Wage Effects Are Large and Similar to Those in the Private Sector. F&M did not examine public sector effects and we suspect F&M would not have predicted this.

VIII. *Policy Implications*

In spite of — or perhaps because of — the inexorable decline in union membership since F&M wrote, and despite some evidence of a recent secular decline in the premium, the union wage premium in the United States remains substantial and is substantial by international standards (Blanchflower and Bryson, 2003). How should U.S. policy analysts treat this piece of information: Should policy support unions or make life more difficult for them?

From a standard economic perspective a substantial union premium has to be bad news: In bargaining on their members' behalf unions distort market wage setting, resulting in a loss of economic efficiency. This may have consequences for jobs and investment for the firms involved and potential impacts for the economy as a whole in terms of inflation and output. However, it is not obvious that markets do operate in the textbook competitive fashion, so there may indeed be rents available that employers are at liberty to share with workers. Furthermore, there is evidence — confirmed herein — that unions are particularly good at protecting the wages of the most vulnerable workers. If the most vulnerable are receiving less than their marginal product — and, perhaps, even if they are receiving it — there may be moral and ethical grounds for supporting unions. Finally, whether analysts like it or not, there is substantial unmet demand for union representation in the United States.[33] People are not getting what they want, raising the question of whether policy should be deployed to assist unions to organize.

F&M found the union premium effect on output was modest, and inflation effects were negligible. There is no reason to alter this judgment. However, the size of the social costs of unionization requires an answer to a fundamental, unanswered question, namely: Where does the premium originate? F&M assume that it necessarily originates in the monopoly rents of employers in privileged market positions, although there are at least three possibilities: (1) unions increase the size of the pie because union workers are more productive than like workers in a nonunion environment; (2) unions operate in firms with excess profits arising from a privileged market position (this may arise either because unions only seek to organize where there are excess profits available, or because these sorts of employers have something particular to gain in contracting with unions); and (3) the premium is simply a tax on normal profits.

It really matters which of these three options it is. If (1) is true, there's no implication for workplace survival or union density, indeed, if this were the case, one might expect a growth in unionization as firms recognize the advantages in unionizing. If it is (2) there is limited damage, but with (3) there are real problems for firms, with potential effects for investment, jobs, and prices. Of course, different employers may be in different positions at any point in time and the weight attached to the three options may differ over time with the business cycle, structural change in the economy, and so on. What evidence is there on the above? Hirsch, in his contribution to this symposium, covers these issues in some detail.

Herein we focus on some key points. (i) F&M (1984, p. 54) speculate that at least in some areas, the wage gap has "reached levels inconsistent with the survival of many union jobs." Blanchflower and Freeman (1992) take this issue further and argue that the levels of the union wage differential were so high this gave incentives to employers to remove the union — the benefits of removing the unions appeared to outweigh the costs. Well, what has happened to union jobs? There is a growing body of evidence that employment growth rates are lower in the union sector, suggesting (2) or (3). The evidence is for the United States (Leonard, 1992), Canada (Long, 1993), and Australia (Wooden and Hawke, 2000) and Britain (Blanchflower et al., 1991; Bryson, 2001). However, Freeman and Kleiner (1999) and DiNardo and Lee (2001) find no clear link between unionization and closure. Indeed, F&M record instances in which union workers accepted cuts in normal wages ("givebacks"), sometimes to keep employers in business. (ii) Unionization is more common where employers operate in monopolistic or oligopolistic product markets, suggesting that unions do try to extract surplus rents. (iii) The premium falls when one accounts for workplace heterogeneity. The implication is that some of the premium usually attached to membership is, in fact, due to union workplaces paying higher wages than nonunion workplaces. If this is so, why? Perhaps unions target organization efforts on workplaces that have rents to share; as noted above, they would be foolish not to. Or employers use unions as agents to deliver lower quit rates to recoup investment in human capital (which makes them more productive/profitable). Abowd et al. (1999) stated that higher paying firms are more profitable or more productive than other firms. If this is so for union firms, they are operating at a higher level of performance than other firms. It makes sense that unions are located in higher performing workplaces since, despite higher wages and a negative union effect on performance, they continue to survive, albeit with slower employment growth rates. A central thesis of WDUD was that unions were more productive so this would not surprise F&M. (iv) The evidence that unions have a substantial negative impact on employment growth suggests that the social cost of unions may be larger than F&M calculated. If the premium reduces the competitiveness of union firms, they will lose employees and, as a consequence, union organizing will get tougher for unions. This is exactly what has happened. On the other hand, if unions do not command a premium, they lose their best selling point for prospective customers. It's Catch–22.

From a public policy perspective, it is not obvious from this evidence that unions, taken as a whole, operate to the detriment of the economy or, if they do, that the magnitude of the problem is really that great. Even the evidence on slower employment growth rates is open to the criticism that the link is not causal but arises because unions are often located in declining sectors. What is lacking from the discussion above is a realization that unions and their effects are heterogeneous: Taking wages alone, unions appear to operate very differently across individuals, states, and industries. Before we can make any clear policy prescriptions about what governments should do to, or for, unions we need to know more about the nature of this heterogeneity to establish under what conditions employers and unions can increase the size of the pie.

And then finally we asked Richard Freeman whether he was surprised by any of these findings. He kindly read the paper and gave us three responses. First, he was most surprised by the fact that the public sector wage effects are so large and so similar to those in the private sector. Second, he said he was surprised how little we know about the social costs of unions in the twenty-first century. Consequently he was unsure about the magnitude of the social costs, but would like to see empirical work on the issue although he said would be "stunned" if there were large effects, but as any good empiricist he would let the data speak. Third, he said was not surprised by any of our other findings.

NOTES

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[1]The symposium included an introduction by the editor John Burton along with reviews by Orley Ashenfelter, Barry Hirsch, David Lipsky, Dan Mitchell, and Mel Reder, plus a reply by Richard Freeman and Jim Medoff.

[2]It has been cited an astonishing 1,024 times over its near 20-year life. In 2002 alone it had 34 cites. It is clear that the book continues to be relevant.

[3]Richard Freeman has devoted a lot of his subsequent writings to an examination of unionism in the public sector including Freeman (1986, 1988), four chapters written with co-authors in Freeman and Ichniowski (1988), and a co-authored paper on union wage gaps for police (Freeman et al., 1989). Freeman's other post-F&M work on union wage effects includes an international comparative paper with one of the authors (Blanchflower and Freeman, 1992), and the impact of union decline on rising wage inequality in the United States — see for example Freeman, (1993, 1995, 1999); Freeman and Katz (1995); Freeman and Needels (1993); and Freeman and Revenga (1998). For a discussion of the issues involved see Blanchflower (2000a).

[4]For further discussions on these issues, see Lewis (1986), Freeman (1984), and Blanchflower and Bryson (2003).

[5]In contrast Lewis (1986), who did not restrict his analysis to nonagricultural private sector blue-collar workers aged 20–65 as F&M did, found no differences by either gender or color, although he confirmed F&M's other disaggregated results. Lewis reported a number of additional disaggregated results that were not examined by F&M. Lewis found the wage gap was greater for married workers; U-shaped in age/experience; U-shaped in tenure/seniority minimizing at 22–24 years of seniority and was higher the higher is the unemployment rate. He found mixed results regarding any relationship with the industry concentration ratio.

[6]Sample sizes in many cases were likely very small as is made clear from their footnote 11 which says that they limited their sample of industries to ones containing at least five union and nonunion members. The rule resulted in only four industries being dropped.

[7]F&M use 1977 union density rates in Table 2.

[8]These estimates are obtained from the ORG files. Although numbers for the public and private sectors as a whole are available since 1973, the breakdown by federal, state, and local employee only begins in 1983.

[9]We do not deal here with a further problem identified by Card (1996) of misclassification of self-reported union status in the CPS, first identified by Mellow and Sider (1983). Card concludes that about 2.7 percent are false positives and 2.7 percent are false negatives. Given that there are more nonunion workers than union workers, this means the union density rate is biased upwards. See Farber (2001) for a discussion and a procedure to adjust the union density rate for error. In 1998, the observed private sector rate of 9.7 per-

cent translates to an adjusted rate of 7.4 percent (the figures for 1973 were 25.9 percent and 24.5 percent, respectively).

[10]The number of wage observations followed by the percentage imputed in parentheses (hourly + non-hourly paid) in the NBER MORG are given below. Note in 1995 allocation information is only available on one-third of the wage observations, hence the small sample.

1979	171,745 (16.5%)	1986	179,147 (10.7%)	1993	174,595 (4.6%)	2000	161,126 (29.8%)
1980	199,469 (15.8%)	1987	180,434 (13.5%)	1994	170,865 (0%)	2001	171,533 (30.9%)
1981	186,923 (15.2%)	1988	173,118 (14.4%)	1995	55,967 (23.3%)	2002	184,137 (30.4%)
1982	175,797 (13.7%)	1989	176,411 (3.7%)	1996	152,190 (22.2%)		
1983	173,932 (13.8%)	1990	185,030 (3.9%)	1997	154,955 (22.2%)		
1984	177,248 (14.7%)	1991	179,560 (4.4%)	1998	156,990 (23.6%)		
1985	180,232 (14.3%)	1992	176,848 (4.2%)	1999	159,362 (27.6%)		

[11]A revised version of their paper, due for publication in the *Journal of Labor Economics* in 2004, exploits the unedited earnings data for those years.

[12]Although there are some differences in the levels of the wage differences reported in columns 3 and 4 of Table 3, the majority of these results are consistent with the findings reported by F&M in their Figure 3. Major exceptions are F&M's finding that wage gaps were higher for men than women and for nonwhites compared with whites. We suspect such differences arise because of the small sample size in the May 1979 CPS of 6,018 used by F&M.

[13]Nearly all federal workers have wages set by civil service pay schedules, but these are not set by collective bargaining in any meaningful sense. Even with workers in the exact same federal job, one will say they are a union member and the other will say they are not. Thus the low union premium for federal workers may not be very meaningful. For state and local workers (and some groups of federal workers like postal workers), the union status variable provides meaningful information. We thank Barry Hirsch for this point.

[14]Tables for the analyses presented in this section are available from the authors.

[15]We have added data for 2002 as the 2002 ORG has recently become available.

[16]The May extracts of the CPS extracts in Stata format from 1969–1987 are available from the NBER at <http://www.nber.org/data/cps_may.html>.

[17]Hirsch et al. (2002) have compared union wage gap estimates obtained from the BLS quarterly Employment Cost Index (ECI) constructed from establishment surveys and from the annual Employer Costs for Employee Compensation (ECEC) with those obtained using the CPS. They find that union/nonunion wage trends in the three series "are consistent neither with each other nor with the CPS," and ultimately conclude that "we find ourselves relying most heavily on results drawn from the CPS" (Hirsch et al., 2002, p. 23).

[18]There was no CPS survey with wages and union status in 1982.

[19]Following Mincer, it is more usual to include a term in potential experience rather than a direct measure of age. We use education, however, for reasons of comparability as the CPS Outgoing Rotation Group files from 1993 report qualifications rather than years of schooling.

[20]Similarly in the 2003 edition of Hirsch and Macpherson's *Union Membership and Earnings Data Book,* recently received.

[21]There is a dissonance between the estimates Lewis offers by way of summary in his introductory chapter and those given in his Table 9.7 produced here (Lewis, 1986, p. 9).

[22]Lewis (1986) had 35 studies using the CPS, 1970–1979; 16 studies using the 1967 Survey of Economic Opportunity; 25 studies using the Panel Study of Income Dynamics, 1967–1978; 15 studies using Michigan Survey Research Center survey data other than the PSID, including the 1972–1973 Quality of Employment Survey; 22 studies using the National Longitudinal Surveys of 1969–1972; and eight studies exploiting other sources.

[23]F&M's smaller sample sizes by industry could account for some of the greater variation in their estimates.

[24]Bratsberg and Ragan (2002) also use CPS data. But their analysis differs in several ways. First, they assess trends over the period 1971–1999 whereas we present trends over the period 1983–2001. Second, we adjust for wage imputation as recommended by Hirsch and Schumacher (2002) whereas Bratsberg and Ragan do not. Third, specifications producing the regression-adjusted estimates differ somewhat. Fourth, the samples differ. In particular, Bratsberg and Ragan exclude government workers, and they present results for some different industries. Fifth, their wage premium relates to weekly wages whereas all of our estimates are derived from (log) hourly wages.

[25]The import penetration variables are calculated as the ratio of imports to industry shipments. Bratsberg and Ragan (2002) in their footnote 19 report that they tabulated shipments through 1994 from Feenstra (1996) and thereafter from the U.S. Bureau of the Census, U.S. Merchandise Trade, series FT900 (December) and Manufactures' Shipments, Inventories and Orders (<http://www.census.gov>).

[26]When the equations are run on the two machines using OLS they are identical. The problem appears to arise from the different tolerances used across computers and not from differences in the STATA programs.

[27]According to the 2002 Human Development Report Table 5 (<http://hdr.undp.org/>) the population of the United States in 2000 was 283.2 million compared with 60.6 million in the UK and 7.2 million in Switzerland.

[28]He finds wages of highly unionized occupations in Las Vegas's hotel and gaming industry are significantly higher than wages of identical occupations in less unionized Reno.

[29]The source of the data is the *Union Membership and Coverage Database* which is an Internet data resource providing private and public sector union membership, coverage, and density estimates compiled from the Current Population Survey (CPS) using BLS methods. Economy-wide estimates are provided beginning in 1973; estimates by state, detailed industry, and detailed occupation begin in 1983; and estimates by metropolitan area begin in 1986. The *Database*, constructed by Barry Hirsch (Trinity University) and David Macpherson (Florida State University), is updated annually and can be accessed at <http://www.unionstats.com>.

[30]The main exceptions are Maine, Hawaii, and Vermont which are small and which had increases in the wage gap. Florida is the fourth largest state after California, Texas, and New York but its differential declined by 10 percentage points.

[31]Source: <http://data.bls.gov/labjava/outside.jsp?survey=la>.

[32]We experimented with both the level of the unemployment rate and the log and the latter always worked best.

[33]According to Peter Hart Associates, the percentage of nonmembers saying they would vote for a union hit 50 percent in 2002, the highest percentage since their figures began in 1984 (when the figure was only 30 percent). See <http://www.aflcio.org/mediacenter/resources/upload/LaborDay2002Poll.ppt>.

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