1. Changes over time in union relative wage effects in Great Britain and the United States

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> As all economics data files have weaknesses – measurement error, unmeasured variables, sample survey quirks – and all model specifications are questionable, contaminated by data mining, any 'finding' ought to be replicated on several data sets and under 'plausible' model specifications before one accepts it as valid. Replication with additional data and specifications contrasts sharply with the practice of econometricians who postulate a 'true' model, use maximum likelihood search procedures to extract its parameters from data, and stop, as if technical prowess rather than robustness of results was the key to credibility. In fact, as all practitioners know, any single piece of complex econometric analysis rarely convinces anyone, for the more sophisticated the econometrics, the greater the danger the results derive from the model than from the world. In economics it is the accumulation of disparate lines of evidence, not the elegance of the statistical technology for a single estimator, that is compelling. (Freeman 1989, p. xi)

1 INTRODUCTION

It is a well-known result that unions raise wages, holding constant characteristics of the individual as well as of the industry, area and workplace. Of particular interest is by how much does this vary by country, across groups and through time? In this chapter these questions are examined using broadly comparable data for Great Britain and the United States. I first started work on this issue as a graduate student at Queen Mary College in 1983 after being persuaded to do so by Bernard Corry. He told me that (a) this was a great topic and (b) there was nowhere else as wonderful as QMC to do this work. Probably because I was young and inexperienced at the time I believed him on both counts! I eventually wrote my PhD thesis[1] on the issue, with lots of input from both Maurice and

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Bernard. Over the subsequent dozen years or so I have revisited the question several times, both on my own and with a number of co-authors.[2] Hence it seemed to be a good topic to return to for their Festschrift. Having spent the last eight years in the United States it seemed to make sense to compare what unions do on the wage front in the two countries. So here are a few thoughts from the other side of the pond.

What is estimated in this literature is the difference between *ceteris paribus* earnings of union and non-union workers. That is, how much would wages change if an individual changed from non-union to union status, holding constant their individual and workplace characteristics. The vast majority of work estimating the effects of unions on relative wages has been based on US data. The definitive works in this area are by Lewis (1963, 1986). The first of his two books measured the effects of unions using relatively aggregated data at the industry level backed up by case study evidence. In the 1986 volume, Lewis examined approximately 200 studies that had used micro-data to estimate the effect of unions. He concluded that it was not possible to use 'macro' data to estimate the union wage gap and that methodologically estimating an ordinary least squares (OLS) equation with wages on the left, union status on the right with a group of controls, was probably the best way to estimate the size of the effect. Panel estimates had problems of misclassification and measurement error while simultaneous equation methods suffered from poor identification due to a lack of suitable instruments. Lewis (1986) argued that estimates obtained using OLS were likely to be upper bounds of the true effect because of the omission of controls correlated with the union status variable.

After an examination of the results of the US studies, many of which he re-estimated himself, Lewis concluded that during the period 1967–79, the US mean wage gap was approximately 15 per cent. He found that the gap was greater for blacks than whites; in services than in manufacturing; for construction than for other non-manufacturing; for blue-collar workers than for white-collar; for private than for public sector workers. The estimates for men and women were approximately the same. The wage gap falls as years of schooling, establishment or firm size and industry unemployment rates rise. For age, years of experience and years of seniority the gap at first falls and then rises. The robustness of Lewis's results were broadly confirmed by Jarrell and Stanley (1990) using meta-analysis, although their mean estimate of the wage gap for the period was a little lower than that obtained by Lewis.

Over the past couple of decades there has been a growing body of literature estimating the size of the union wage gap outside the USA. There are a few studies for Canada which suggest that the union wage gap is in the range 10–15 per cent. This estimate appears to have remained fairly constant over time.[3] In Australia the estimated range is between 7 and 17 per cent with most estimates at the lower

end of the range.[4] Moll (1993) obtained estimates for South Africa in 1985 of 24 per cent for black blue-collar workers (19 per cent for black men and 31 per cent for black women) and 13 per cent for whites in 1985. For South Korea, Park (1991) obtained estimates of 4.2 per cent for men and 5 per cent for women. Wagner (1991) found significant positive union effects for blue-collar workers in Germany while Schmidt (1995) found small but significant wage differentials of under 6 per cent. Neither Schmidt (1995) nor Schmidt and Zimmermann (1991) were able to find evidence of significant union wage gaps for men.[5]

For Great Britain there have been approximately 20 studies some based on establishment data,[6] and some on individual data.[7] The consensus seems to be (see Booth 1995 for a discussion) that the mean union wage gap is approximately 10 per cent. Despite the rapid decline in union density experienced in the UK since 1979, there is some evidence to suggest that the gap has remained roughly constant since 1970 (see Blanchflower 1991; Stewart 1995), the year for which the earliest estimate is available (Shah 1984), although there is some dispute on this question (see Lanot and Walker 1998). The disaggregated pattern of results reported by Lewis (1986) for the USA appear to be broadly repeated for the UK. The main exception is that the wage gap in the UK appears to be larger for females than it is for males (see Blanchflower 1991; Main 1996). Some care has to be taken in comparing the estimates from the British studies because they often relate to quite disparate groups of workers – usually manuals (Stewart 1983), sometimes male manuals only (for example, Shah 1984) and occasionally full-time male manuals (Stewart 1983). Moreover, all of the estimates obtained from establishment-level data (from the Workplace Industrial Relations Survey series) exclude workers employed at small workplaces (Blanchflower 1984; Stewart 1987). As we show below, there is considerable variation in the differential across groups (for example, by establishment size, occupation, industry sector and so on) which means that sample exclusions will potentially result in biases in the 'overall' result.

In what follows, a series of estimates for the union wage gap in the 1980s and 1990s are presented for both Great Britain and the USA. Comparable micro-data files at the level of the individual are used to estimate log hourly earnings equations which contain similar groups of control variables for the two countries. The British data files are a lot smaller than those for the US, which necessitates making use of data from a number of different sources. Clearly one would wish also to examine the extent to which unions are able to influence the total compensation package including fringe benefits. Unfortunately relatively little is known about the extent to which unions are able to influence fringe benefits, primarily because of a lack of suitable data. Such literature as does exist – virtually all of which is for the USA – suggests that these effects are large (see Freeman and Medoff 1984). None of our data files contain adequate information that would allow us to examine this issue.

Before moving to estimating union wage gaps it is appropriate to place these results in the wider context of the changes in the labour market experience of the two countries over the last couple of decades.

There are five basic facts to be kept in mind.

First, unemployment was generally higher in the USA than it was in the UK from 1965–1980. The picture reversed itself in the later period, 1980–95 (Figure 1.1).

Second, both employment and the size of the labour force increased rapidly over the period 1980–94 in the USA as is illustrated in Table 1.1. The UK experienced only small growth on both of these dimensions over this period. Consequently, the gap in the employment/population ratios between the two countries widened over the period.

Third, there was substantial growth in earnings inequality in the 1970s and 1980s in the USA. Earnings inequality declined in the UK in the 1970s but increased in the 1980s. This is illustrated in Figures 1.2a and 1.2b which are taken from Katz et al. (1995) and which plot the time series of overall wage inequality of men and women as measured by the log hourly wage differential between the ninetieth and the tenth percentiles of the wage distribution. There is much

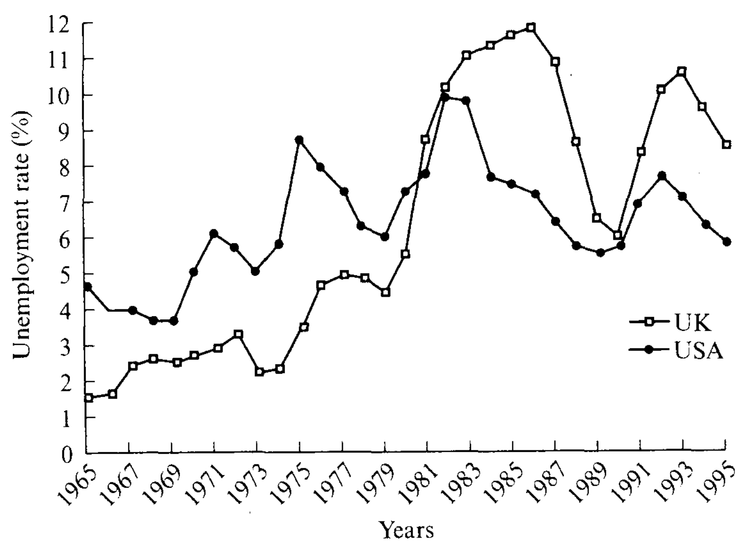


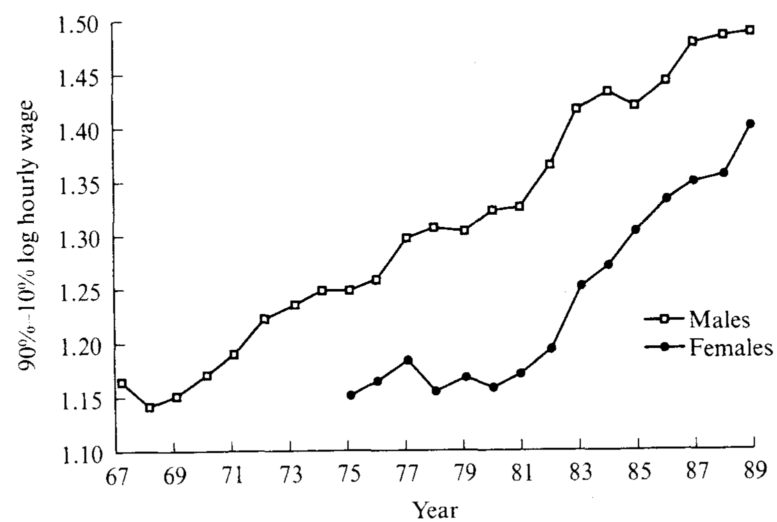
Figure 1.1 Unemployment rates (%) in the UK and the USA, 1965–95

less evidence of rising wage inequality in other countries over the period (see the various papers in Freeman and Katz 1995).

Table 1.1 Employment and size of labour force, USA and UK, 1980–94

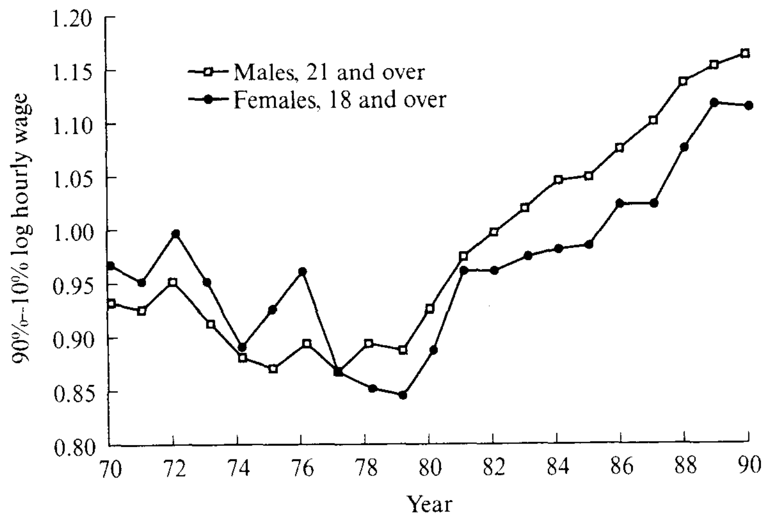
	Employment (millions)		Employment/population rate (%)		Labour force (millions)	
	USA	UK	USA	UK	USA	UK
1980	99.3	24.7	59.2	58.1	106.9	26.5
1985	107.1	24.2	60.1	55.1	115.4	27.2
1990	118.7	26.6	62.8	59.2	125.8	28.5
1992	118.4	25.5	61.5	56.5	128.1	28.4
1993	120.2	25.3	61.7	56.2	129.2	28.3
1994	123.0	25.6	62.5	56.5	131.0	28.3

Source: *Statistical Abstract of the US, 1996–7.*



Source: Katz et al. (1995).

Figure 1.2a Changes in earnings inequality, United States, 1967–89

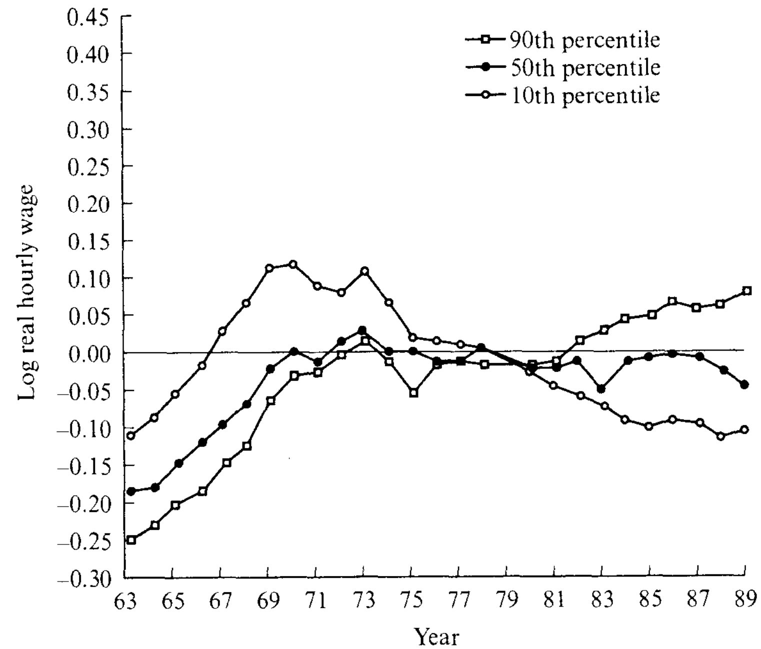


Source: Katz et al. (1995).

Figure 1.2b Earnings inequality in Great Britain, 1970–90

Fourth, real wage growth has been much higher in the UK than in the USA, and especially so at the low end of the distribution. Figures 1.3a and 1.3b, which are also taken from Katz et al. (1995) illustrate this by plotting the real earnings of the ninetieth, fiftieth and tenth percentiles of the wage distributions for men in the two countries. More precisely the figure displays the log ratio of each group's real earnings in each year relative to the group's real earnings in 1979 (the base year). The two panels show fairly similar increases in the 90–10 differential in the 1980s in Britain and the USA, but indicate that these increases implied a 0.12 decline in log real hourly earnings from 1979 to 1989 at the tenth percentile in the US wage distribution and a 0.12 increase in real log earnings at the same point of the British distribution. The figure indicates that only in the USA was rising wage inequality in the 1980s accompanied by declining real wages for low wage males. Even the median US male employee experienced a modest decline in log real hourly earnings from 1979–89; his UK counterpart experienced a 0.24 increase in log real hourly earnings over the same period.

Fifth, union density rates declined steadily in the USA from 1970.[8] In the UK density increased in the 1970s and then declined dramatically thereafter (Figure

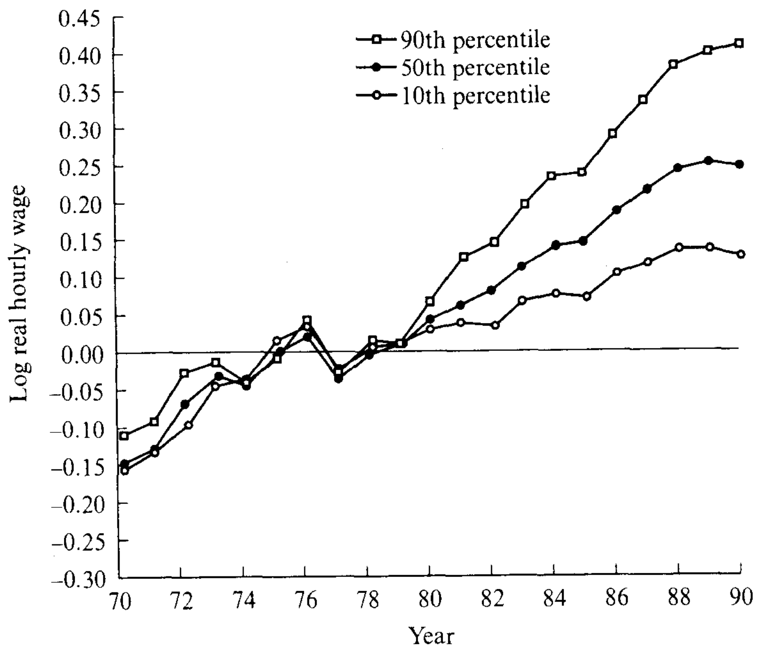


Source: Katz et al. (1995).

Figure 1.3a United States, 18–65 years old

1.4). The decline in density was also pronounced in Japan and Austria. Some countries experienced increases in density over the period (for example, Denmark, Finland and Sweden). For a discussion, see Blanchflower and Freeman (1992) and Blanchflower (1996).

In the next section, data from the 1983–95 Outgoing Rotation Group files of the Current Population Survey (CPS) are used to obtain estimates of the impact of trade unions on hourly earnings for the USA. Separate estimates are obtained from each of these cross-sections which allow us to examine movements in the union wage gap over time. In the following section, cross-section data from various individual level datasets for Great Britain are used to perform a similar exercise for Great Britain. It is found that the level of the union wage gap is untrended over time, but positively correlated with the unemployment rate in both countries.

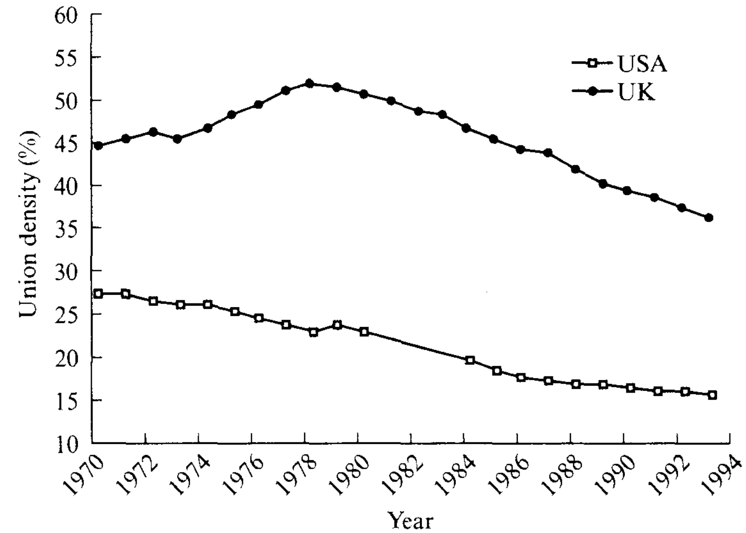


Source: Katz et al. (1995).

Figure 1.3b Great Britain, 21 years old and older

2 UNION WAGE DIFFERENTIALS IN THE USA

In Table 1.2 the results of estimating log hourly earnings equations for 1983 using data from the Outgoing Rotation Group files of the CPS are reported. Control variables are a union status dummy plus age and its square, a gender dummy, years of schooling, a part-time dummy, two race dummies, three sector of work dummies, two self-employment dummies plus 50 state and 50 industry dummies. In total there are just over 170 000 observations. The dependent variable is defined as the log of usual hourly earnings for hourly paid workers and for the remainder as the log of usual weekly earnings/usual weekly hours.[9] The overall union wage effect is estimated at 15.5 per cent (antilog of 0.1445 from column 1 minus one because the dependent variable is in logarithms). Columns 2–8 of the table report disaggregated estimates. The union wage gap is higher in the private sector (16.9 per cent) than it is in the public sector (8.8 per cent). Results by gender and race are all close to 15 per cent.



Source: Visser (1996).

Figure 1.4 Union density rates, UK and USA, 1970–93

Table 1.3 reports the results of estimating a very similar specification to that in Table 1.2, using the same data source, but now for 1993. The main exception is that years of schooling are replaced by 15 schooling dummies to distinguish highest level of schooling attended. This change is necessary because of changes in the CPS survey design. Results are very similar to those for 1983 discussed above. The union wage differential remains unchanged at 15.5 per cent – remarkably the estimates in columns 1 of both Tables 1.2 and 1.3 vary only at the fourth place of decimals. Once again there is little difference by gender or race. However, the differential in the public sector in 1993 is slightly higher than it was in 1983 (8.8 per cent and 11.8 per cent, respectively). Overall the wage gap in the USA is very close to that estimated in Lewis (1986), despite both a dramatic decline in density along with a large increase in earnings inequality that has occurred since then.

Is the high differential in the USA an artefact of sample selectivity? In Blanchflower and Freeman (1992) it was argued that this is not the correct way to interpret the data and this is still my view. The reasons given, which are still relevant, were as follows.

1. Evidence within the USA tends to reject the notion that union wage effects are large when union density is small. Union wage differentials tend to be

Table 1.2 Log hourly earnings equations, USA, 1983

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Union	0.1445	0.1564	0.0841	0.1312	0.1469	0.1469	0.1359	0.1440
	(53.39)	(48.11)	(17.01)	(37.46)	(57.21)	(50.32)	(16.69)	(9.67)
Age	0.0480	0.0482	0.0483	0.0613	0.0339	0.0490	0.0370	0.0424
	(109.29)	(99.94)	(44.97)	(94.71)	(57.21)	(105.18)	(24.72)	(15.64)
Age^2	−0.0005	−0.0005	−0.0005	−0.0006	−0.0004	−0.0005	−0.0004	−0.0004
	(91.34)	(83.59)	(37.44)	(79.08)	(48.49)	(88.31)	(20.23)	(12.57)
Male	0.2237	0.2382	0.1697	—	—	0.2324	0.1428	0.1829
	(103.27)	(97.42)	(37.38)			(100.37)	(20.08)	(15.42)
Years schooling	0.0581	0.0555	0.0683	0.0567	0.0571	0.0592	0.0496	0.0473
	(144.28)	(120.54)	(82.97)	(107.22)	(92.00)	(135.89)	(37.73)	(23.90)
Black	−0.0966	−0.1059	−0.0719	−0.1390	−0.0511	—	—	–
	(26.66)	(25.19)	(10.41)	(26.21)	(10.61)			
Other non-white	−0.0689	−0.0733	−0.0565	−0.0859	−0.0452	—	—	–
	(11.21)	(10.42)	(4.60)	(9.97)	(5.32)			
Part-time	−0.1658	−0.1527	−0.2302	−0.2000	−0.1373	−0.1651	−0.1609	−0.1672
	(60.71)	(50.87)	(35.04)	(42.70)	(42.04)	(56.83)	(17.53)	(10.82)
Federal government	0.0483	—	—	0.0545	0.0474	0.0373	0.0509	0.0949
	(5.16)			(4.24)	(3.55)	(3.40)	(2.26)	(3.03)
State government	−0.0257	—	−0.0828	−0.0577	0.0078	−0.0327	0.0206	0.0100
	(4.12)		(7.00)	(6.23)	(0.94)	(4.82)	(1.11)	(0.32)
Local government	−0.0339	—	−0.0947	−0.1028	0.0213	−0.0378	0.0198	−0.6807
	(6.87)		(8.25)	(13.94)	(3.30)	(7.03)	(1.45)	(2.43)
Constant	−0.4789	−0.4564	−0.4076	−0.5455	−0.0196	−0.5187	0.0134	−0.1427
N	173 404	140 854	32 298	92 756	80 648	152 668	15 204	5 532
R^2	0.4928	0.4956	0.4585	0.4750	0.4255	0.4996	0.4444	0.4770
$\bar{R}^2$	0.4925	0.4952	0.4568	0.4744	0.4247	0.4992	0.4404	0.4666

Note: Private sector subsample excludes self-employed. All equations include 50 state and 50 industry dummies + 2 self-employment dummies. *t*-statistics in parentheses.

Source: CPS Outgoing Rotation Group File, 1993: NBER 50 Variable Uniform Extract, 1979–1993.

Table 1.3 Log hourly earnings equations, USA, 1993

	(1) All	(2) Private sector	(3) Public sector	(4) Men	(5) Women	(6) Whites	(7) Blacks	(8) Other non-whites
Union	0.1440	0.1501	0.1113	0.1424	0.1306	0.1444	0.1413	0.1455
	(45.99)	(38.25)	(20.63)	(33.68)	(27.91)	(42.05)	(16.01)	(9.94)
Age	0.0430	0.0433	0.0444	0.0507	0.0473	0.0444	0.0340	0.0357
	(89.34)	(82.05)	(36.59)	(69.33)	(54.74)	(85.85)	(21.85)	(14.76)
Age^2	−0.0004	−0.0004	−0.0004	−0.0005	−0.0004	−0.0004	−0.0003	−0.0004
	(74.89)	(69.01)	(30.66)	(56.17)	(47.44)	(72.11)	(18.17)	(12.54)
Male	0.1423	0.1524	0.1032	—	—	0.1537	0.0576	0.1093
	(59.33)	(57.08)	(19.02)			(58.93)	(7.79)	(10.32)
Veteran	0.0139	0.0159	0.0161	−0.0201	−0.0048	0.0093	0.0341	0.0325
	(4.01)	(4.04)	(2.19)	(5.13)	(0.36)	(2.50)	(3.04)	(1.54)
Federal government	0.0975	—	—	0.0859	0.1187	0.1002	0.1012	0.0873
	(13.67)			(8.66)	(11.65)	(12.30)	(5.81)	(3.21)
State government	−0.0194	—	−0.1070	−0.0411	0.0128	−0.0208	0.0173	−0.0469
	(3.28)		(12.27)	(4.59)	(1.64)	(3.15)	(1.08)	(1.80)
Local government	−0.0288	—	−0.1278	−0.0621	0.0062	−0.0381	0.0183	0.0281
	(5.73)		(14.75)	(8.09)	(0.95)	(6.79)	(1.39)	(1.22)
Black	0.1237	−0.1343	−0.0896	−0.1674	−0.0805	—	—	–
	(33.63)	(31.70)	(12.36)	(33.01)	(16.80)			
American Indian	−0.0471	−0.0683	−0.0026	−0.0459	−0.0421	—	—	–
	(4.34)	(5.08)	(0.14)	(2.95)	(2.83)			
Asian or Pacific Islander	−0.0919	−0.0947	−0.0738	−0.1064	−0.0731	—	—	−0.0202
	(14.60)	(13.60)	(5.06)	(11.85)	(8.43)			(1.19)
Other non-white	−0.0900	−0.1003	−0.0247	−0.0915	−0.0797	—	—	−0.0441
	(6.13)	(6.28)	(0.66)	(4.58)	(3.73)			(2.06)
Part-time	−0.1772	−0.1674	−0.2227	−0.2280	−0.1473	−0.1755	−0.1908	−0.1550
	(61.29)	(52.62)	(32.04)	(44.58)	(42.97)	(56.28)	(20.20)	(11.46)
Constant	0.3428	0.4444	−0.5819	0.2608	0.6319	0.2274	0.5972	0.7909
N	171439	140323	147479	87257	84182	147479	15969	7991
R^2	0.4172	0.4637	0.4748	0.4682	0.4480	0.4748	0.4396	0.4706
$\bar{R}^2$	0.4708	0.4632	0.4743	0.4675	0.4471	0.4743	0.4352	0.4622

Note: All equations also include 15 schooling dummies, 50 state dummies and 50 industry dummies and 2 self-employment dummies. Excluded categories are private sector and white. Private sector excludes the self-employed. *t*-statistics in parentheses.

Source: CPS Outgoing Rotation Group File, 1993: NBER 50 Variable Uniform Extract, 1979–1993.

greater the greater the extent of unionization in the sector (see Lewis 1986; Freeman and Medoff 1984), presumably because this gives unions greater bargaining power.

2. If selectivity were the major cause of the estimated large effects of unionism on wages in the USA, similar differences in other market outcomes should be expected, which is not found.
3. Third, the fact that employers as well as workers affect union density makes the direction of the selectivity effect uncertain. One might well argue that selectivity operates to bias down union wage effects in the USA as employers fight hardest against unions that have the most potential for raising wages and accept unions when they have the least potential.
4. Massive employer opposition to unions in the USA but not elsewhere is consistent with the greater demand by unions for higher wages there than in other countries.

All of this does not deny the possibility that our estimates may be contaminated by the reverse effects of density on wage differentials, rather it is probable that this potential contamination is unlikely to reverse the finding that union wage differentials are relatively high for similar workers in the USA.

Table 1.4 Estimates of the union wage gap, USA, 1983–95

Year	Union coefficient	Union wage gap	Number of observations
1983	0.1445	15.6	173 404
1984	0.1519	16.4	172 970
1985	0.1428	15.4	179 710
1986	0.1435	15.4	178 969
1987	0.1366	14.6	180 165
1988	0.1360	14.6	172 813
1989	0.1375	14.7	176 158
1990	0.1300	13.9	184 731
1991	0.1222	13.0	179 261
1992	0.1330	14.2	176 492
1993	0.1440	15.5	171 439
1994	0.1470	15.8	149 819
1995	0.1390	14.9	152 274
Average		14.9	172 939
Total observations			2 248 205

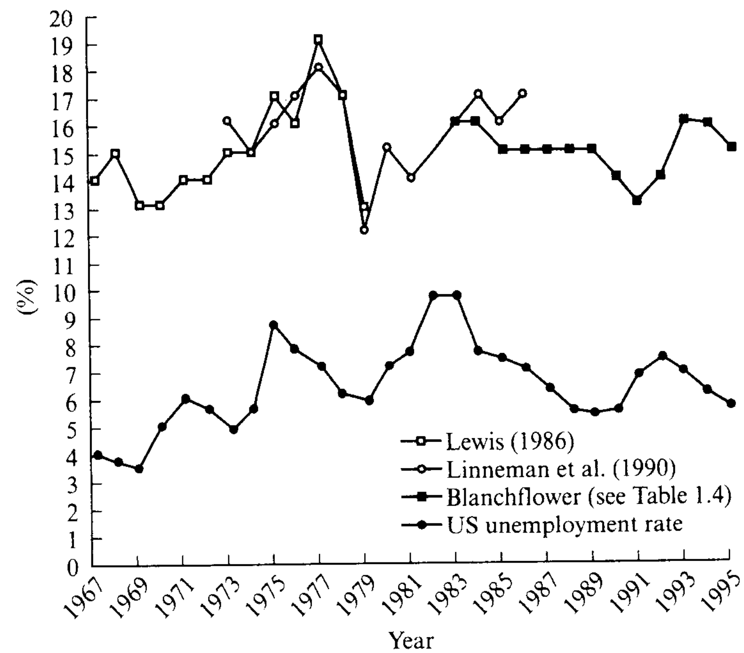


Figure 1.5 Union wage premia, USA, 1967–95

A number of earlier studies have examined the extent to which the mean union wage gap has varied over time in the USA. As discussed earlier, Lewis (1986) reported an average union wage gap for the period 1967–79 of 15 per cent. He further reported on time-series movements in the differential, by taking the mean estimate by year from each of the 150 studies he surveyed. Subsequent work by Linneman et al. (1990), extending earlier work on the subject presented in Linneman and Wachter (1986), used the CPS to estimate wage gaps for full-time non-agricultural workers for the years 1973–86. The two sets of results are plotted in Figure 1.5. In addition the figure also plots new estimates for the intervening years between 1983 and 1995 that were obtained from estimating eleven further earnings equations of the form presented in column 1 of both Tables 1.2 and 1.3. The data are drawn from the Outgoing Rotation Group files of the CPS for each year in turn.[10] In total there are two and a quarter million observations, with an average of about 170 000 observations per year. The union dummy is always highly significant with t-statistics everywhere greater than 40. The full results are shown in Table 1.4. The union wage gap is calculated as the natural antilog of the union coefficient minus one.

Where the three series in Figure 1.5 overlap, there is considerable agreement on both the size and movements in the mean wage gap – the estimates are never more than one percentage point apart. The average union wage premium over the period 1967–93 is estimated at approximately 15 per cent. There is a clear cyclical pattern: there are obvious peaks in 1977, 1983/84 and 1993 and troughs in 1979 and 1991. These cyclical patterns approximate fairly closely movements in the aggregate unemployment rate, which is also plotted in Figure 1.5. Unemployment was low in 1979 and high in 1983 and 1992, for example. When unemployment is low the union wage premia appears to be low and vice versa. Despite some evidence of cyclicality the dominant impression from the figure is the relative constancy of the differential over this long time period, even though the labour market has, along other dimensions, experienced so much turbulence over the same time period.

Table 1.5 Time-series union wage gap regressions, USA, 1967–95

	(1)	(2)	(3)	(4)	(5)	(6)
Unemployment rate $_t$	0.5500		0.0021			
	(2.78)		(0.01)			
Unemployment rate $_{t-1}$		0.8090	0.8076	0.6946	0.5491	0.7408
		(4.80)	(3.46)	(3.63)	(3.10)	(4.26)
Union wage gap $_{t-1}$				0.2008		
				(1.22)		
CPI						0.0013
						(0.02)
Time trend	–0.0201	–0.0539	–0.0539	–0.0487	0.2413	–0.0587
	(0.62)	(1.87)	(1.83)	(1.69)	(1.43)	(1.84)
Constant	11.9905	10.9027	10.8984	8.5025	8.0924	11.2538
N	28	27	27	27	27	27
R^2	0.2421	0.4901	0.4901	0.5211	0.4619	0.4410
Adjusted R^2	0.1814	0.4476	0.4235	0.4587	0.4170	0.3681

Note: t-statistics in parentheses.

A series of time-series equations were estimated to explain movements in the union differential over time (Table 1.5). This is simply meant to be illustrative: no attempt is made to correct for serial correlation or to estimate any fully specified model. Where there was more than one estimated differential for a single year I simply took the average. It appears that

1. unemployment lagged one period enters positively and significantly;
2. contemporaneous unemployment is insignificantly different from zero;
3. there is statistically weak evidence of a significant time trend in the data;
4. the lagged dependent variable is everywhere insignificant; and

5. the consumer price index (CPI) (and lags) were also always insignificant.

Hence we conclude that the union wage differential is untrended but positively correlated with the unemployment rate and uncorrelated with the inflation rate.[11] This presumably arises because union wages are less responsive to changes in the unemployment rate than are non-union wages, confirming wage curve results in Blanchflower and Oswald (1994). The unemployment elasticity of pay is higher in the non-union sector than it is in the union sector (ibid., Table 4.19, p. 159).

An obvious question to ask is why have union membership and union employment been in decline given the relative constancy of the union wage premium? As was noted above, the level of the differential – at about 15 per cent – is still very high by international standards. The United States decides union membership through an adversarial electoral process at plant level which has evolved into a system where management has a greater say in unionization outcomes than it does in other countries. The benefits to employers in removing unions from the workplace often outweigh the costs of doing so. The cost to unions in organizing recruitment drives is high.[12] Bender (1997) has argued that the loss of economies of scale in union organizing is an important factor in explaining union decline. It is much harder for employers in other countries to get rid of unions than it is in the USA. Even in the UK there are only a very few examples of union de-recognition. Employers are unable to hide from a union; they have no place to go.

The decline in US unionism seems to have been driven by employer opposition, fuelled by more competitive product markets, increased international trade and a favourable legal environment, which has meant that there have been smaller economic rents to be shared with workers than was true in the past.[13] It is unlikely to be a coincidence that the generally lower union–non-union wage differentials that operated in the late 1980s and 1990s, as compared with those that existed in the 1970s, were associated with a marked slowing in the rate of decline in US union density (see Figure 1.4). Similarly the 1970s, which was a period of high and growing union wage differentials, saw rapid declines in private sector union density.

Linneman et al. (1990) have gone even further and suggested that the evidence of a relatively constant aggregate union wage premium is a 'statistical artifact' (1990, p. 51). High premium industries, they show, have been increasing their union wage premia and losing employment shares and hence membership of trade unions. Union wage premia in private services, they argue, have held constant or fallen. They argue that even though unions have been hurt by exogenous factors which have created shifts in demand from goods to service-producing industries, unions have been hurt most by the rising wage premia. Supporting evidence for this view is presented by Freeman (1986) who found a positive correlation between the union wage gap and a proxy for managerial opposition to unions – the number of unfair labour practices per worker in National Labor Relations Board elections. Farber (1990) also concludes that the decline is principally a result of increased employer opposition to unions along with lower demand for union services by workers.

3 UNION WAGE DIFFERENTIALS IN GREAT BRITAIN

There is no single long time-series of large cross-sections for Great Britain that contain data on both wages and union status. The Labour Force Surveys (LFSs), which are large and contain a long time-series of cross-sections, have only recently started asking questions on wages. Although there are quite a number of British individual level micro-data files, only occasionally do they contain data on these two crucial variables. We make use of data from the 1983 General Household Survey (GHS), the 1993 and 1994 Labour Force Surveys, the British Social Attitudes Survey (BSA), 1983–94 and the British Household Panel Study (BHPS), 1991–93. Where possible we try to replicate results by year on more than one data file. The number of observations is everywhere much less than was available for the USA, ranging from a high of about 8000 per year in the GHS and LFS to a few hundred observations a year for the BSA surveys.

Table 1.6 reports the results of estimating a log hourly earnings equation for Great Britain using data from the General Household Survey for 1983. This is the same source used by Green (1988). When missing values are deleted, data are available on only about 8000 individuals. A group of control variables similar to those used for the USA (age and its square, gender, race, highest qualification, size of establishment, region and industry and month of interview) were included. Consistent with earlier studies which found estimates of about 10 per cent, the estimated differential is approximately 11.2 per cent. Results in the public and private sectors are very similar. The differential for females is higher than for males (12.5 per cent and 8.6 per cent, respectively) confirming earlier work by Blanchflower (1991) and Main (1996).

Table 1.7 performs a similar exercise but now for 1993/94 using pooled data from the 1993 and 1994 Labour Force Surveys. (Union membership data is not available in any subsequent GHS or in earlier LFSs.) Even though these surveys are large, only the Outgoing Rotation groups – a fifth of the sample – are asked to report wages, which means a total sample size of just over 16 000 when the two years are pooled. The purpose behind pooling is to raise sample size. The female differential is now more than double that of males (12.5 per cent and 6.0 per cent, respectively). Separate estimates are reported for the two years. The picture is very similar to that of the USA, where despite large declines in density, the union wage differential remained largely unchanged between 1983 and 1993/94. Our estimate of the overall wage gap is little changed, and not significantly different, from the 1983 result; it is now estimated at 9.8 per cent.

Table 1.6 Log hourly earnings equations, Great Britain, 1983

	(1) All	(2) Private sector	(3) Public sector	(4) Males	(5) Females
Union	0.1064	0.0973	0.0960	0.0821	0.1176
	(10.23)	(7.20)	(5.86)	(5.78)	(7.89)
Age	0.0646	0.0670	0.0543	0.0842	0.0449
	(30.83)	(26.07)	(14.81)	(28.15)	(14.94)
Age^2	–0.0007	–0.0007	–0.0006	–0.0009	–0.0005
	(27.34)	(23.03)	(13.05)	(25.28)	(13.20)
Male	0.2987	0.3140	0.2865	—	—
	(27.63)	(22.43)	(17.28)		
Part-time	–0.0795	–0.0981	–0.0405	–0.0480	–0.0379
	(6.30)	(5.87)	(2.18)	(1.54)	(2.71)
Black	–0.1030	–0.1387	–0.0406	–0.1199	–0.0692
	(3.91)	(4.01)	(1.04)	(3.62)	(1.63)
Public sector	0.0509	—	—	0.0219	0.1046
	(4.79)			(1.18)	(5.00)
Constant	–0.3995	–0.7158	–0.5495	–0.7316	–0.0306
N	7951	5106	2845	4442	3509
R^2	0.5260	0.4992	0.5493	0.4586	0.4617
$\bar{R}^2$	0.5226	0.4938	0.5404	0.4518	0.4532

Note: Equations include 17 highest qualification dummies, 4 size of establishment dummies, 11 month dummies, 10 region dummies and 10 industry dummies. *t*-statistics in parentheses.

Source: General Household Survey, 1983.

The evidence of a relatively constant differential over time in the UK is consistent with earlier work presented in Blanchflower (1991), where union wage effects were estimated using data from the 1983–87 and 1989 British Social Attitude Surveys. I obtained estimates of 10 per cent with no significant variation over the years.[14] In Table 1.8 the equations reported in the earlier paper are re-estimated adding the 1990, 1991, 1992 and 1994 surveys. Results are reported for both annual earnings and for hourly earnings. The reason for the shorter time run when hourly earnings are used arises because of the lack of suitable hours data in the earliest years. The small differences from the earlier results arise from variation in the group of right-hand side variables. The

Table 1.7 Log hourly earnings equations, Great Britain, 1993–94

	(1) All	(2) Private sector	(3) Public sector	(4) Male	(5) Female	(6) 1993	(7) 1994
Union	0.0934	0.0990	0.0806	0.0584	0.1177	0.0861	0.0997
	(11.31)	(9.36)	(6.15)	(5.14)	(9.78)	(7.48)	(8.37)
Age	0.0618	0.0656	0.0449	0.0800	0.0484	0.0662	0.0576
	(31.22)	(28.40)	(10.91)	(28.17)	(16.33)	(23.72)	(20.39)
Age^2	−0.0007	−0.0007	−0.0005	−0.0009	−0.0006	−0.0007	−0.0006
	(26.69)	(24.11)	(9.48)	(24.20)	(14.48)	(20.68)	(17.06)
Male	0.1812	0.1814	0.1782	—	—	0.1984	0.1698
	(20.65)	(17.66)	(12.32)			(14.74)	(14.17)
Part-time	−0.0974	−0.1062	−0.0979	−0.0873	−0.0728	−0.0917	−0.1032
	(10.40)	(6.88)	(6.28)	(3.89)	(6.82)	(6.97)	(7.72)
Black	−0.1006	−0.0983	−0.0948	−0.1701	−0.0695	−0.1687	−0.0328
	(2.94)	(2.08)	(2.00)	(3.28)	(1.54)	(3.40)	(0.69)
Asian	−0.1065	0.0142	0.0121	−0.1810	0.0042	−0.0899	−0.1232
	(3.77)	(0.26)	(0.22)	(4.98)	(0.10)	(2.22)	(3.11)
Other	−0.1151	−0.0765	−0.0803	−0.1231	−0.1170	−0.1637	−0.0382
	(2.94)	(1.26)	(1.32)	(2.19)	(2.18)	(3.26)	(0.61)
Public sector – type nk	0.0983	—	—	0.0667	−0.1862	0.1023	n/a
	(5.48)			(3.18)	(0.75)	(5.05)	
Nationalized industry	0.0900	—	0.0509	0.0948	0.0890	0.0554	0.1017
	(2.72)		(1.24)	(2.32)	(1.61)	(0.71)	(2.66)
Central government	0.1607	—	0.0246	0.0787	0.2276	0.2360	0.1262
	(7.30)		(0.86)	(2.33)	(7.83)	(6.07)	(4.57)
Local government	0.0902	—	−0.0618	0.0427	0.1207	0.1049	0.0758
	(6.17)		(2.39)	(1.52)	(6.91)	(4.77)	(3.78)
University	0.0511	—	−0.0902	−0.0227	0.0986	0.0721	0.0323
	(1.85)		(2.70)	(0.46)	(2.98)	(1.55)	(0.92)
Health authority	0.1616	—	−0.0303	0.0188	0.2026	0.1907	0.1347
	(8.31)		(1.04)	(0.41)	(9.16)	(6.67)	(5.02)
Other	0.0496	—	−0.0771	−0.1413	0.1174	0.0376	0.0525
	(1.03)		(1.50)	(1.51)	(2.09)	(0.53)	(0.79)
N	16159	11352	4807	8014	8145	8131	8028
R^2	0.4108	0.4077	0.3669	0.4055	0.3656	0.4217	0.4072
$\bar{R}^2$	0.4078	0.4038	0.3561	0.3995	0.3593	0.4160	0.4013

Note: Equations also include 11 region dummies, 31 qualification dummies, 11 industry dummies, 6 size of establishment dummies and a year dummy. *t*-statistics in parentheses.

Source: Labour Force Surveys, 1993/1994.

Table 1.8 British Social Attitudes Survey earnings equations

	(1) All	(2) 1983	(3) 1984	(4) 1985	(5) 1986	(6) 1987	(7) 1989	(8) 1990	(9) 1991	(10) 1993	(11) 1994
a) Annual (83–94)											
Union	0.1174	0.0995	0.1187	0.0489	0.1516	0.1114	0.1005	0.0890	0.0929	0.1941	0.1319
	(11.92)	(2.62)	(3.20)	(1.38)	(5.80)	(4.04)	(3.70)	(3.51)	(3.74)	(5.04)	(3.69)
N	10 850	691	699	781	1418	1282	1352	1177	1104	1019	1327
R^2	0.6698	0.6628	0.6605	0.6838	0.6952	0.6835	0.6771	0.6623	0.6366	0.5539	0.5021
Adjusted R^2	0.6687	0.6496	0.6474	0.6729	0.6895	0.6770	0.6707	0.6546	0.6278	0.5422	0.4921
b) Hourly (85–94)											
Union	0.1170	n/a	n/a	0.0118	0.1580	0.1101	0.1075	0.0869	0.0697	0.1557	0.1341
	(10.94)			(0.33)	(6.20)	(4.12)	(3.93)	(3.28)	(2.42)	(3.86)	(3.91)
N	9443			764	1418	1282	1352	1177	1104	1019	1327
R^2	0.4563			0.4624	0.4594	0.4745	0.4258	0.3872	0.2980	0.2915	0.2680
Adjusted R^2	0.4544			0.4435	0.4493	0.4636	0.4145	0.3734	0.2810	0.2729	0.2534

Note: Unless otherwise stated controls also included in each equation were age and its square, male dummy, part-time dummy, years of schooling, 10 region dummies, 9 industry dummies, a manual dummy and, in the overall equations, 9 year dummies. t-statistics in parentheses.

Source: British Social Attitudes Surveys, 1983–1994, current employees only.

Table 1.9 British Household Panel Study hourly earnings equations

	OLS							Fixed effects			
	(1) All	(2) 1991	(3) 1992	(4) 1993	(5) All	(6) Males	(7) Females	(8) All	(9) Males	(10) Females	(11) All
Union	0.1057	0.0832	0.1178	0.1192	0.0887	0.0618	0.1457	0.0373	0.0317	0.0359	0.0293
	(12.57)	(5.92)	(8.07)	(7.82)	(6.80)	(5.37)	(11.77)	(3.56)	(2.09)	(2.47)	(2.41)
Union*1992					0.0302						0.0172
					(1.67)						(1.60)
Union*1993					0.0228						0.0078
					(1.24)						(0.69)
N	13 434	4 794	4 413	4 227	13 434	6 586	6 848	13 434	6 586	6 848	13 434
R^2	0.4939	0.5026	0.4939	0.4838	0.4940	0.4638	0.4437				
Adjusted R^2	0.4918	0.4971	0.4878	0.4773	0.4919	0.4594	0.4394				
N (individuals)								6 119	2 999	3 120	6 119
R^2 within								0.1169	0.1462	0.1171	0.1172
R^2 between								0.0100	0.0145	0.0065	0.0103
R^2 overall								0.0162	0.0195	0.0109	0.0166

Note: Unless otherwise stated controls also included in each equation were age and its square, male dummy, 2-year dummies, part-time dummy, 11 highest qualification dummies, 17 region dummies, 9 industry dummies, and 10 size of establishment dummies. N (individuals) refers to the number of people specific fixed effects in the regression.

Source: British Household Panel Study, 1991–1993 (waves 1–3), current employees only.

finding of the constancy of the differential at about 10 per cent is unchanged. There is no significant variation across the years.

Table 1.9 provides further validation using data of the constancy result using the first three sweeps of the British Household Panel Study for the years 1991–93. The first column pools all three waves of the survey. A union wage gap of 11.1 per cent is estimated. The next three columns report results separately by year. Estimates vary from 9.3 per cent in 1991 to 12.7 per cent in 1993. Interaction terms between the year dummies and the union status dummy were insignificantly different from zero, suggesting no significant variation in the differential over time. Differentials for women are more than double those of men, confirming results reported above. Finally, in columns 8 to 11 the first panel estimates of the union wage gap for Great Britain are reported. In this specification the three years are pooled and a full set of people fixed effects. As was found in the US literature, panel estimates are generally lower than the OLS estimates, principally because of problems of misclassification and measurement error (see Freeman 1984). Now the wage gap has fallen to 3.8 per cent, with only a small difference between the findings for men and women. As far as I am aware this is the first panel estimate of the wage gap that has been reported for Great Britain. It would be interesting to trace how this estimate moves through time as more sweeps of this survey become available.

Table 1.10 summarizes the findings of the extent of variation in the wage gap in Great Britain from 1983–94. Other estimates are available but they are for specific subgroups (for example, males, manuals, or more usually, male manuals); we report here only estimates using individual level data for all groups of employees. The average is 10.7 per cent.

Table 1.10 Hourly earnings union wage premia

Year	BSA	GHS	LFS	BHPS	Average
1983		11.2			11.2
1985	1.2*				1.2
1986	17.1				17.1
1987	11.6				11.6
1989	11.3				11.3
1990	9.1				9.1
1991	7.2			8.7	8.0
1992				12.5	12.5
1993	16.8		9.0	12.7	12.8
1994	14.4		10.5		12.5
Average					10.7

Note: * = Not significantly different from zero.

Figure 1.6 plots the average for the years from the final column against the unemployment rate. As was found for the USA, the wage gap appears to vary positively with the unemployment rate, although the evidence is considerably weaker given the small number of time points. Table 1.11 reports a number of time-series regressions to examine the relationship between the wage gap and the unemployment rate. Column 1 reports a very poor equation for Great Britain alone. Because of the small sample size we pool the British and US data (columns 2 and 3). We confirm the earlier result that the gap is positively correlated with the unemployment rate: we find it now for the two countries. The British dummy is significantly negative in columns 2 and 3. Once again there is no significant time trend in these data, even though the unemployment rate itself is positively trended. Indeed, the finding of relative constancy of the differential through time in Great Britain seems quite robust.[15] In assessing the impact of unions at the macroeconomic level, it should be kept in mind that, even though the union wage differential appears to have remained roughly constant, it applies to a considerably smaller fraction of the workforce in 1993 than it did

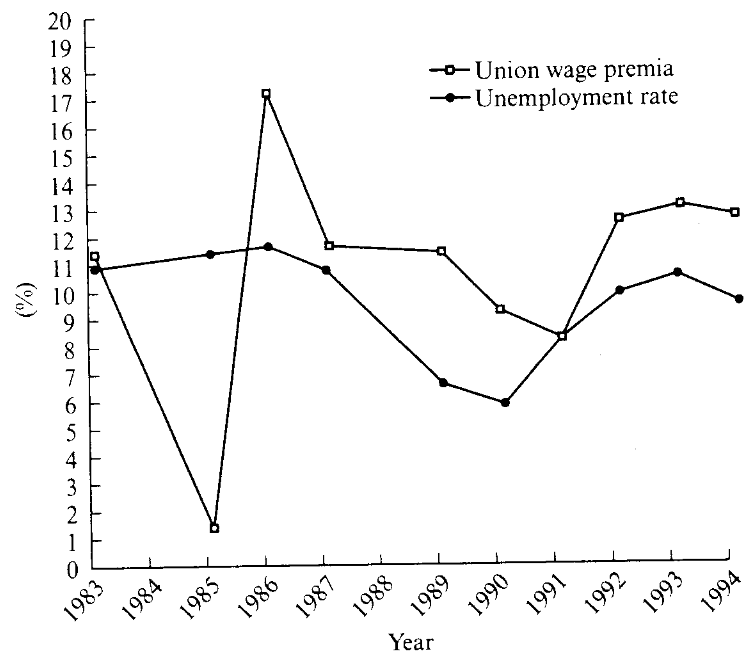


Figure 1.6 Union wage premia, UK, 1983–94

ten or twenty years ago. There are now fewer workers getting the 10 per cent union wage premium than there were.

Table 1.11 Time-series union wage gap regressions, USA, 1967–95 and Great Britain, 1983–94

	(1)	(2)	(3)
Unemployment rate$_{t-1}$	0.8732	0.5542	0.6337
	(1.08)	(2.13)	(2.40)
Time trend	0.6685	–0.0189	–0.0200
	(1.22)	(0.33)	(0.36)
GB dummy		–6.2492	–7.5024
		(4.53)	(4.51)
Union wage gap$_{t-1}$			–0.2224
			(1.32)
Constant	–1.9770	11.8497	14.7071
N	9	36	36
R^2	0.2366	0.4603	0.4888
Adjusted R^2	–0.0179	0.4097	0.4229
Sample	GB only	GB + USA	GB + USA

Note: t-statistics in parentheses.

Table 1.12 Disaggregated union wage gap estimates, USA and UK (%)

	UK	USA
All	9.8	15.5
Male	6.0	15.3
Female	12.5	14.0
Age < 30	12.6	18.8
Age ≥ 30 and < 50	8.0	14.5
Age ≥ 50	7.1	12.4
Public sector	8.4	11.8
Private sector	10.4	16.2
Least educated	11.1	25.4
Middle educated	11.1	17.1
Most educated	4.3	9.0

continued

Table 1.12 continued

	UK	USA
Non-whites	16.1	15.7
Whites	9.6	15.5
Part-timers	14.8	30.3
Full-timers	7.8	13.2
Manufacturing	8.3	7.0
Non-manufacturing	17.5	10.2
Manual	18.8	24.2
Non-manual	6.7	12.2

Notes: 'Least educated' is defined as at least 11 years of schooling in the UK and not graduated from high school in the USA; 'middle educated' is defined as 12–15 years of schooling in the UK and graduated from high school/some college in the USA; 'most educated' is defined as at least 16 years of schooling in the UK and completed college in the USA; 'manual' in the USA includes forestry and logging occupations.

It is of interest to examine the extent to which the union wage gap varies across groups in the two countries. We have already presented a number of disaggregated results by gender, race and the public/private sector in Tables 1.2–3 and 1.6–7 above. Table 1.12 reports estimates for a number of other categories each derived from separate wage regressions for 1993 for the USA and 1993/94 for Great Britain. Wage gaps are the same by race and gender in the USA; in contrast, in Great Britain non-whites and women have higher differentials. In all other respects the results appear similar across the two countries; the gap is higher among the young, the least educated, part-timers, non-manuals and for workers in the non-manufacturing sector.[16]

4 CONCLUSIONS

The main findings of this chapter may be summarized as follows.

1. The union wage gap averages 15 per cent in the USA and 10 per cent in Great Britain.
2. The gap moves pro-cyclically but appears to be untrended in both countries. Union wages are sticky.
3. The size of the wage gap varies across groups. In both the USA and Great Britain the differential is relatively high in the private sector, in non-manufacturing, for manuals, the young and the least educated.

4. In the US there are no differences by race or gender in the size of the differential. In Great Britain it is higher both for women and for non-whites.

In comparison with other countries it does appear that the size of these estimated wage gaps are quite high. Union wage differentials in other countries are generally less than 10 per cent (see Blanchflower 1996). It does appear that countries that have experienced rapid declines in union membership do have the highest wage differentials. The fact that the differential has remained more or less constant in both Great Britain and the USA is a puzzle, particularly given the rapid declines in union membership in both countries. The evidence is not consistent with the widely held view that union power has been emasculated. We still have a great deal to learn about the time-series properties of the union wage differential and its correlates.

NOTES

1. See Blanchflower (1985).
2. See Blanchflower (1984, 1986a, 1986b, 1991, 1996); Blanchflower and Freeman (1992, 1994); Blanchflower and Machin (1996); Blanchflower and Oswald (1990, 1994); and Blanchflower et al. (1990).
3. Examples of studies for Canada are Simpson (1985) who found 11 per cent for 1974; Grant et al., (1987) with 12–14 per cent for 1969 and 13–16 per cent in 1970 and Green (1991) who obtained an estimate of 15 per cent for 1986.
4. Australian studies include Kornfeld (1993) who found 7–10 per cent for young people between 1984 and 1987. Mulvey (1986) obtained 7 per cent for women and 10 per cent for men using a 1982 sample. Christie (1992) using 1984 data obtained an estimate of 16.6 per cent using OLS and 17.2 per cent using simultaneous equation methods. Blanchflower and Freeman (1992) found 8 per cent for the period 1985–87.
5. In Blanchflower (1996) I obtained union wage gap estimates based on similar data and specifications for the following countries – * implies not significantly different from zero.

	%		%
Australia	9.2	Japan	47.8
Austria	14.6	Netherlands	3.7*
Canada	4.8*	New Zealand	8.4
Germany	3.4	Norway	7.7
Ireland	30.5	Spain	0.3*
Israel	7.0*	Switzerland	0.8*
Italy	7.2		

The very large estimate for Japan appears to arise because of the lack of controls for workplace/firm size. Some of the estimates are based on only a few hundred observations so care has to be taken in interpreting these results.

6. Blanchflower (1984); Stewart (1987, 1990, 1991 and 1995); Blanchflower and Oswald (1990); Blanchflower et al. (1990); Metcalf and Stewart (1992); Machin et al. (1993); and Blanchflower and Machin (1996).
7. Stewart (1983); Shah (1984); Green (1988); Symons and Walker (1988); Blackaby et al. (1991); Blanchflower (1991); Main and Reilly (1992); Murphy et al. (1992); Main and Reilly (1993); and Main (1996).

8. Interestingly enough Canada, which has many of the same firms and trade unions that exist in the US, has not seen declines in density – 1970 = 31 per cent; 1980 = 36.1 per cent; 1990 = 35.8 per cent; 1993 = 37.4 per cent (source: Visser 1996).
9. In both cases we use edited or computed data, including allocated values.
10. The dependent variable is usual hourly earnings as defined above. All equations include age and its square, a gender dummy, three public sector dummies, two self-employment dummies (incorporated and unincorporated), a part-time dummy, 50 state dummies and 49 industry dummies. For the years 1983–88 two race dummies were included. After 1989 this was increased to four. From 1983–91 a years of schooling variable was used. For 1992 and 1993, because of the change in the sample to a credential-based schooling measure, 15 qualification dummies were used. For 1994 and 1995, six full-time/part-time dummies were also included.
11. For a discussion on these issues and rather different results for the earlier period 1920–80, see Pencavel and Hartsog (1984).
12. Chaison and Dhavale (1990) estimate that maintenance of current density levels will require unions to make, over and above current organizing budgets, an expenditure of $300 million annually.
13. For further discussion on this point, see Blanchflower and Freeman (1992).
14. All years pooled 10 per cent; 1983, 11 per cent; 1984, 11 per cent; 1985, 1 per cent (not significant); 1986, 13 per cent; 1987, 9 per cent; 1989, 11 per cent. Dependent variable in Blanchflower (1991) was log of annual earnings. Note that the dependent variable is banded with open-ends. The banding changes over the years. Mid-points are used and the ends are closed in an inevitably ad hoc way.
15. It should be noted, however, that Gregg and Machin (1992) have found evidence in a sample of large quoted UK firms that union wages increased less than non-union wages. while Ingram (1991) found similar evidence for some sectors. Differences in wage *growth* is likely to work to reduce union differentials moderately. Lanot and Walker (1998) provide simultaneous equation estimates for an earlier period 1978–85 and show a marked *increase* in the size of the wage gap for married male manuals, although they do observe some decline in the differential at the end of the period (see their Figure 3).
16 Due to small sample sizes in the UK it is not appropriate to disaggregate further by occupation or industry. The large sample sizes for the USA allow further disaggregation. The following estimated wage gaps for 1993 were obtained from separate wage regressions. Sample sizes are also reported in the second column.

	%	
Occupation		
Executive, administrative and managerial	6.0	20 349
Professional specialty occupations	13.7	24 524
Technicians and related support occupations	12.3	54 720
Sales occupations	14.5	18 645
Administrative support occupations, including clerical	12.0	29 609
Service occupations	20.1	24 999
Farming, forestry and fishing	30.4	2 848
Precision production, craft and repair	19.6	17 691
Machine operators, assembler etc.	23.6	11 775
Transportation and material moving equipment occupations	32.8	7 428
Handlers, equipment cleaners, helpers and labourers	31.6	7 105
Industry		
Agriculture	13.9	2 534
Mining	7.9	1 197
Construction	30.0	8 136
Manufacturing	8.3	30 400
Transportation and public utilities	21.1	12 752
Wholesale trade	5.9	6 336
Retail trade	20.4	29 324
Finance, insurance and real estate	2.1	11 318
Business and repair services	16.9	8 006

Personal services	8.8	5 710
Entertainment and recreation services	32.7	769
Professsional and related services	12.4	43 265
Public administration	8.3	9 486

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