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Changes in the Structure of Wages in the 1980's: An Evaluation of Alternative Explanations

By JOHN BOUND AND GEORGE JOHNSON*

During the 1980's, a period in which the average level of real wage rates was roughly stagnant, there were large changes in the structure of relative wages, most notably a huge increase in the relative wages of highly educated workers. This paper attempts to assess the power of several alternative explanations of the observed relative wage changes in the context of a theoretical framework that nests all of these explanations. Our conclusion is that their major cause was a shift in the skill structure of labor demand brought about by biased technological change. (*JEL* J31)

During the 1980's, there were three major changes in the structure of wages in the United States. First, there was a precipitous rise in the relative wages and earnings of workers with high levels of education. The average wage of a college graduate increased relative to the average wage of a high school graduate by over 15 percentage points from 1979 to 1988. The high-school/elementary-school wage differential also increased substantially. Second, for those in the labor force who had not completed college there was a large increase in the average wage of older workers relative to younger workers. Third, the average wage of women relative to the average wage of men increased by about 8 percent, resulting in a fall in the average wage disadvantage of women from 30 percent of men's wages in 1979 to 24 percent in 1988.

The fact that there have been large changes in the distribution of wages and income has been widely documented (see e.g., Kevin Murphy and Finis Welch, 1991; McKinley Blackburn et al., 1990/91; Marvin Kosters, 1989). There is, however, less consensus concerning the *causes* of these phenomena than about the changes themselves. There are four—not necessarily exclusive—explanations that have received recent attention. The first attributes the wage-structure changes to the decline in manufacturing employment, in large part associated with the increase in the trade deficit during the 1980's, which may have increased the relative demand for better-educated workers and female workers (Murphy and Welch, 1991). The second explanation concentrates on the loss of wage premia paid to blue-collar males in certain industries because of the declines of manufacturing employment and the power of unions (Barry Bluestone and Bennett Harrison, 1988). The third attributes the wage-structure changes to changes in technology, brought on in large part by the computer revolution (Jacob Mincer, 1991). The fourth attributes the rise in the relative wages of college graduates to a slowdown in the rate of growth of the college-educated population, caused in turn by the drop in the size of the cohort entering the labor market during the 1980's (Murphy and Welch, 1989).

While some work has been done evaluating each of the above explanations, such work has usually involved assessing the merits of each in isolation. The purpose of this paper is to evaluate comprehensively the explanatory power of each of the explana-

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tions. Our major conclusion is that, while each of the other three contributed slightly to the explanation of observed relative wage movements, their primary cause was technical change.

The remainder of the paper is organized as follows. Section I describes the major changes in the wage structure during the 1970's and 1980's. Section II sets out a simple model from which the testable implications of the alternative explanations are obtained. Section III reports the results of the tests, and Section IV summarizes our major conclusions.

I. Changes in the Structure of Wages

Our first task is to document the changes in the wage structure that occurred during the 1980's as well as, for comparative purposes, the changes in the structure of wages that occurred in the 1970's. The analysis is based on imputed wage rates from questions on usual weekly earnings and hours from the Current Population Survey (CPS) for 1973–1974, 1979, and 1988. Each sample eliminates all workers in agriculture, forestry, and fisheries, as well as private household service and individuals with imputed hourly wages less than \$1.00 or greater than \$100 in 1979 dollars.[1] Each sample included only persons between the ages of 18 and 64 who reported employment as their major normal weekly activity.

Each of the resultant samples (66,808 for 1973–1974, 145,744 for 1979, and 149,011 for 1988) was then split into 32 subsamples based on four values of completed years of schooling, S (dropouts: $S<12$; high school: $S=12$; some college: $12<S<16$; and college: $S\geq 16$), four levels of potential labor-market experience, X (0–9, 10–19, 20–29, and 30+ years), and two sexes. For each subsample in each of the three periods, the logarithm of the wage rate for each individual was regressed on X and dummy variables for educational attainment (where relevant), nonwhite, part-time employment, residence in an SMSA, four major regions, and employment in 17 major industries (a list of which are given in Table 2). From the estimated regression coefficients we then calculated the estimated mean log wage in each period of workers in the ith education/experience/sex group, Y_i, evaluated for whites working full time in SMSA's in the mean region and industry for the group at particular years of schooling (8 for dropouts, 12 for high school, 14 for some college, and 16 for college) and at the midpoints of the experience ranges ($X=5$, 15, 25, and 35).[2]

The estimated values of the real average hourly wage rates of the 32 groups [$\exp(Y_i/P)$, where P is the value of the CPI relative to its value in 1988] are reported in columns (i)–(iii) of Table 1. A striking feature of these results is the downward trend in most real wage rates over the entire period. The average per annum growth of real wages between 1973 and 1979 was -0.010,[3] and the equivalent rate of growth between 1979 and 1988 is -0.008. Fringe benefits are not included in CPS wages, but from aggregate data on the ratio of supplements to wages and salaries to total compensation (*Economic Report of the President*, 1990 table C-24), our estimates should be adjusted upward by 0.006 and 0.001 for, respectively, the 1973–1979 and 1979–1988 periods to reflect total per annum compen-

[1]A problem with the CPS data is that weekly earnings are capped at \$999, which introduces a downward bias in wage rates, especially for prime-age men with high levels of education in the 1988 sample. To deal with this, we used the unedited weekly-earnings measure, which is "top-coded" at \$1,923, for observations that were at the \$999 cap. Very few respondents, even among the critical groups, exceeded this higher cap. For the approximately 20 percent of the relevant observations for whom earnings were not available in the unedited file, we assigned the geometric mean among the top-capped individuals.

[2]A dummy variable for 1974 was included in the 1973–1974 regressions (the two years having been merged in order to yield an adequate sample of employment by industry), and this was set equal to zero for the computation of the Y_i's.

[3]This is equal to $\sum_i[Y_i(1979)-Y_i(1973)]k_i(1973)/6-\ln[P(1979)/P(1973)]/6$, where $k_i(1973)$ is group i's share of total employment in 1973–1974.

TABLE 1—ESTIMATED AVERAGE REAL HOURLY WAGE RATES (IN 1988 DOLLARS), RELATIVE WAGE CHANGES, AND EMPLOYMENT DISTRIBUTIONS BY EXPERIENCE, EDUCATION, AND SEX FOR 1973, 1979, AND 1988

		Real wage levels			Fixed-weight relative wage changes		Employment distributions		
Experience (years)	Education	1973 (i)	1979 (ii)	1988 (iii)	1973–1979 (iv)	1979–1988 (v)	1973 (vi)	1979 (vii)	1988 (viii)
Men:									
0–9	dropouts	7.52	7.20	5.54	0.020	−0.192	0.027	0.023	0.015
	high school	9.69	8.96	7.31	−0.015	−0.134	0.077	0.079	0.060
	some college	10.61	9.89	8.51	−0.008	−0.080	0.041	0.043	0.034
	college	12.69	11.38	12.16	−0.046	0.136	0.043	0.048	0.041
10–19	dropouts	9.96	9.61	7.45	0.027	−0.185	0.033	0.021	0.018
	high school	12.69	12.09	10.31	0.014	−0.089	0.062	0.057	0.067
	some college	14.60	13.43	12.06	−0.021	−0.037	0.023	0.031	0.036
	college	16.95	15.29	14.81	−0.040	0.038	0.028	0.036	0.050
20–29	dropouts	11.37	10.25	8.53	−0.041	−0.113	0.037	0.024	0.014
	high school	13.92	12.81	11.91	−0.020	−0.003	0.046	0.040	0.045
	some college	15.33	14.37	13.93	−0.002	0.039	0.015	0.016	0.022
	college	18.62	17.10	17.08	−0.022	0.069	0.020	0.022	0.028
30+	dropouts	11.30	10.74	10.17	0.012	0.015	0.078	0.054	0.029
	high school	13.65	13.02	12.05	0.015	−0.007	0.051	0.051	0.042
	some college	15.39	14.60	14.27	0.010	0.047	0.014	0.015	0.014
	college	18.26	16.88	17.64	−0.016	0.114	0.011	0.015	0.016
Women:									
0–9	dropouts	5.80	5.48	4.82	0.005	−0.058	0.014	0.012	0.008
	high school	7.14	6.87	6.18	0.024	−0.035	0.066	0.069	0.055
	some college	8.91	7.79	7.52	−0.071	0.034	0.028	0.038	0.038
	college	10.42	9.29	10.00	−0.052	0.144	0.027	0.036	0.040
10–19	dropouts	6.68	5.96	5.11	−0.051	−0.084	0.016	0.013	0.011
	high school	8.21	7.74	7.60	0.004	0.052	0.040	0.049	0.058
	some college	10.11	9.21	9.29	−0.052	0.079	0.011	0.019	0.034
	college	11.29	10.64	11.38	0.003	0.138	0.011	0.017	0.036
20–29	dropouts	6.17	6.31	5.81	0.085	−0.013	0.022	0.015	0.011
	high school	8.22	7.96	7.74	0.030	0.042	0.040	0.038	0.049
	some college	9.23	8.90	9.64	0.027	0.150	0.009	0.012	0.022
	college	12.04	10.54	11.25	−0.070	0.135	0.010	0.011	0.019
30+	dropouts	6.38	6.59	6.20	0.095	0.009	0.040	0.029	0.019
	high school	8.39	8.07	7.96	0.024	0.056	0.045	0.048	0.046
	some college	9.59	9.12	9.59	0.012	0.121	0.009	0.012	0.013
	college	12.50	10.52	11.15	−0.110	0.128	0.008	0.008	0.009

sation growth. This yields annualized (fixed-weight) rates of growth of real wages of −0.4 percent and −0.7 percent over the two periods.

A graphical illustration of the three major stylized facts about the wage structure that are the focus of this paper is provided in Figures 1 and 2. Here, the estimated wage rates in 1979 and 1988 for the four educational categories (dropouts, high school graduates, those with some college, and college graduates) are plotted against the midpoints of the four labor-market experience intervals, separately for men and women. Columns (iv) and (v) of Table 1 report the fixed-weight proportional relative wage

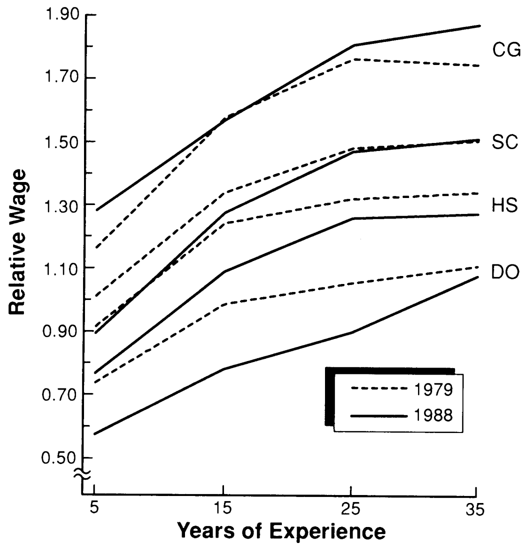

FIGURE 1. ESTIMATED RELATIVE WAGE RATES BY EDUCATION FOR MEN BY YEARS OF EXPERIENCE, 1979 AND 1988

Notes: CG = college graduates, SC = some college, HS = high school, and DO = dropouts.

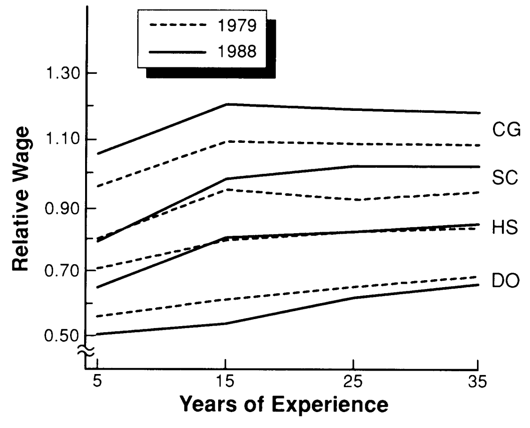

FIGURE 2. ESTIMATED RELATIVE WAGE RATES BY EDUCATION FOR WOMEN BY YEARS OF EXPERIENCE, 1979 AND 1988

Notes: CG = college graduates, SC = some college, HS = high school, and DO = dropouts.

changes for each of the 32 demographic groups in the two time periods.[4] [For example, the wage of male college graduates relative to high school graduates with five years of experience increased from $11.38/8.96 = 1.27$ in 1979 to $12.16/7.31 = 1.66$ in 1988, and the change in the logarithm of this relative wage is the difference in the two groups' relative wage changes in column (v), $0.136-(-0.134)=0.270$.]

The first stylized fact, the increase in the relative wages of more-educated workers during the 1980's, is very clearly seen from inspection of Figures 1 and 2. With very few exceptions, the change in the average relative wage position of more-educated workers, experience and sex held constant, was higher than that for less-educated workers. The average proportionate change in the wages of college graduates relative to those of high school graduates was 0.163 for men and 0.118 for women. Over the 1973–1979 period, on the other hand, the college relative wage fell (by 0.035 for men and 0.073 for women).[5] The increase in the wages of high school graduates relative to those of dropouts during the 1980's was also very large except for the highest experience cate-

[4]The value of each of the relative wages in Figures 1 and 2 is $\text{rel}_i = \exp(Y_i - \Sigma_i Y_i k_i)$, and the fixed-weight proportional wage change of each group is $\Delta Y_i - \Sigma_i \Delta Y_i k_i$. The change in rel_i(as well as the average real wage change with variable weights) also depends on the value of $\Sigma_i (Y_i + \Delta Y_i)\Delta k_i$, the change in the average wage in the economy due to changes in the demographic composition of the work force. The value of this term was 0.004 over the 1973–1979 interval (the mean worker was more educated but was younger and more likely to be female), but over the 1979–1988 interval its value was 0.042 (because the labor force got older).

[5]The social rate of return to a four-year college program for men behaved in a similar manner from 1973 to 1988. Assuming a 2,000-hour work year, average retirement at age 61, and that the resource cost of each year of college equals average tuition at private universities, the real internal rate of return fell from 4.8 percent to 4.1 percent from 1973 to 1979 and then increased to 7.3 percent in 1988. If tuition had not risen during the 1980's at 5.1 percent per annum in real terms, the rate of return in 1988 would have been 8.4 percent, or double its 1979 value. Interestingly, the steep decline in the real wages of young high school graduates in concert with the rise in real tuition costs caused the average ratio of opportunity costs (lost earnings) to total costs to fall from 0.73 in 1973 to 0.58 in 1988.

gory; for $X \leq 30$, the proportionate change in the average high-school/dropout relative wage was 0.072 for men and 0.060 for women. These differentials changed very little during the 1970's.

The second fact about the wage structure in the 1980's is that for those who did not complete college there was an increase in the relative earnings of older workers, especially for men. This is reflected in an increase in the slopes of the relevant relative-wage profiles in Figures 1 and 2. For noncollege workers (all except college graduates) the average proportionate wage change of those with more than 19 years of experience exceeded that of younger workers by 0.107 for men and by 0.043 for women. During the 1973–1979 period, this relative wage was constant for men, but it increased by 0.058 for women.

The third major wage-structure development is reflected in the fact that the 1988 relative wage profiles for women in Figure 2 tend to be higher relative to their 1979 values than is true for men in Figure 1. The average fixed-weight proportional wage change for women was 0.076 greater than that for men, which represented an acceleration of the 1973–1979 difference in wage changes by gender of 0.016. Indeed, the average logarithmic wage advantage of men (at the demographic weights for each year) declined from 0.392 in 1973 to 0.363 in 1979 and to 0.280 in 1988, which means that the relative wages of women (based on geometric means) were 0.675, 0.696, and 0.756 over the three periods.[6]

A major problem in the interpretation of the wage-structure developments of the 1980's, which has been noted in previous work on this topic (e.g., Murphy and Welch, 1992), is that there were large increases in the relative supplies of most of the demographic groups whose relative wages increased. (The correlation between the proportional change in relative wages across the 32 demographic groups, [column (v) in Table 1] and the proportional change in relative supply [the logarithm of the ratio of column (viii) to column (vii)] is +0.51.) Other things equal, there should have been a *decrease* in the relative wages of these groups, for the work force got more educated and more female during the 1980's. The obvious strategy for explaining the wage-structure developments of the 1980's is to look for the set of demand-shift factors that were sufficiently powerful to overcome the effects of demographic changes that would have caused the wage structure to move in the opposite direction.

II. A Conceptual Framework for Evaluating Alternative Explanations

In order to perform an empirical analysis of the reasons for the observed relative changes of the 1980's, it is useful to set out a simple theoretical model that incorporates all of the major explanations. The aggregate work force is composed of I demographic groups (defined by age, education, and sex). The wage rate of group-i workers in industry j is W_{ij}, and this is conveniently defined as the product of the "competitive" wage, W_{ic}, and a relative rent, μ_{ij}. If the nonpecuniary attributes of employment in all industries were identical and there were no unions or other factors causing wage rates to deviate from their competitive norm, the μ_{ij}'s would be identically equal to 1. Whatever the reasons, however, there is substantial evidence that quality-adjusted wage levels vary across industries (Sumner Slichter, 1950; Alan Krueger and Lawrence Summers, 1988).

Defining Y_{ij} and M_{ij} as the logarithms of W_{ij} and μ_{ij}, the geometric mean of the wage rate for group-i workers is

$$(1) \qquad Y_i = Y_{ic} + \sum_j M_{ij}\phi_{ij}$$

[6]The value of the gross average log wage of women relative to men increased by slightly more than did their fixed-weight relative wages relative to those of men primarily because of an increase in their average educational attainment relative to men in the work force. An additional source of the wage gap between men and women, which is netted out of our figures, is the fact that women are more likely to be employed part-time (15.5 percent versus 5.5 percent in 1988). Since there is a negative effect on individual wages of being employed part-time (15 percent for women and 22 percent for men in 1988), this adds approximately two percentage points to the gender differential in each of the years.

where $Y_{ic} = \ln(W_{ic})$ and $\phi_{ij} = N_{ij}/N_i$ is the proportion of group-i workers who are employed in industry j.

This suggests an initial classification of explanations of relative wage movements into those that focus on market factors (changes in relative demand or supply that affect the W_{ic}'s) and those that focus on institutional factors (changes in the values of the μ_{ij}'s or their incidence), for the change in the relative average log wage of each group i is

$$(2)\quad dY_i = dY_{ic} + \sum_j (\phi_{ij} dM_{ij} + M_{ij} d\phi_{ij}).$$

The change in the relative average wage for group-i workers is (2) less its weighted average across all I groups. This can change for a particular group either because its competitive wage grows faster or slower than average or because of changes in the average level or average incidence of industry wage premia, represented by the two parts of the second term on the right-hand side of (2).

Estimation of the initial relative wage impact of changes in average group premia is straightforward. The task of devising tests of alternative explanations of changes in competitive wage levels is more challenging. To do this, we utilize a conventional model of the determination of competitive wages for each of the I demographic groups and the employment level for each group in each of J industries (N_{ij}). There are five (admittedly simplifying) assumptions made in the model (the equations of which are set out in the Appendix) so that it can be used for purposes of testing. First, output in each industry is a function of efficiency units of employment, $b_{ij}N_{ij}$, of each of the demographic groups, where b_{ij} is an index of the technical efficiency of group-i workers in industry j. The single intrafactor elasticity of substitution, σ, is assumed to be constant and equal across industries (with, following Daniel Hamermesh [1986], $1 < \sigma < \infty$), but different industries may employ the different groups in different proportions.[7] Second, the demand for the output of each industry is a function of its relative price and an exogenous shift parameter. Third, the employment levels of all groups in each industry (the N_{ij}'s) are determined by equations setting the marginal revenue products of the I labor inputs equal to their competitive wage rates.[8] Fourth, the economy is at full employment in the sense that the total effective aggregate labor supply (i.e., measured labor force minus frictional unemployment) of each labor group (N_i) is employed in the J industries in the economy. The final assumption is that the N_i's are exogenous. In particular, the aggregate supply of each demographic group does not depend on its relative average wage.

The model leads to the conclusion that the change in the competitive wage of group-i workers (relative to the change in the aggregate wage) depends positively on their average rate of technical change (relative to all groups) $d(\ln b_i)$, negatively on their relative supply change $d(\ln N_i)$, and positively on the change in their relative product-demand-shift index $d(\ln D_j)$. Sub-

[7]A more general specification, which would allow partial elasticities of complementarity among labor groups to vary (rather than assume that all cross-elasticities are equal to $1/\sigma$) can be set out theoretically but is intractable empirically. It turns out, however, that the proportional changes in the supplies of groups that might be considered a priori to be substitutable for each other to some extent (e.g., men and women dropouts in the youngest experience interval) are highly correlated. Thus, it does not make much difference exactly how the labor groups are aggregated.

[8]This assumption can be justified theoretically on the grounds that (a) the μ_{ij}'s result from the effects of trade unions on relative wages and that (b) unions and management bargain over both wages and employment. These assumptions imply that the competitive wage (W_{ic}), rather than the negotiated wage ($\mu_{ij}W_{ic}$), figures in the determination of employment levels by industry (see Henry Farber [1987] for an extensive discussion of this); in other terms, changes in the μ_{ij}'s yield solutions that are off the conventional demand curves. The alternative assumption is that employment is set in each industry such that marginal revenue products are equal to negotiated wages. In this case, an increase in a particular μ_{ij} will lower W_{ic} because of employment effects, and the sign of the effect of a change in μ_{ij} on the average wage for that group is ambiguous. Our simplifying assumption thus implies a possible bias in our empirical results toward acceptance of the explanation involving changes in the $\Sigma_j \mu_{ij}\phi_{ij}$'s.

stituting Appendix equation (A9) for dY_{ic} in (2), we have

$$(3)\quad dY_i = (1-1/\sigma)d(\ln b_i) - (1/\sigma) \times d(\ln N_i) + (1/\sigma)d(\ln D_i) + \sum_j (\phi_{ij} dM_{ij} + M_{ij} d\phi_{ij}).$$

The four alternative explanations of the wage-structure changes of the 1980's are nested in this equation, and three of them (all but that dealing with technical change) can be directly confronted by the data.

To illustrate the operation of the model, consider the possible explanations of the increase in the relative wage of college graduates during the 1980's. Figure 3 shows the relative-demand and relative-supply functions for college-educated labor, with w the wage and n the supply of college-educated workers relative to other labor. The initial values of w and n are w_0 and n_0. We know that w increased to w_1'' in the face of an increase in n to n_1 [in fact, from the data in Table 1, $\Delta(\ln w) = 0.111$ and $\Delta(\ln n) = 0.139$ for the 1979–1988 period]. If the initial equilibrium were at point a in Figure 3 with the demand curve n_d (which means that average relative rents $[\Sigma_j \mu_{ij} \phi_{ij}]$ for college and noncollege workers were equal), we would expect w to fall to w_1' through the operation of the $-(1/\sigma)d(\ln N_i)$ term in (3). That w rose to w_1'' implies, in the absence of a change in relative rents, that the relative-demand function shifted to n_d'. This, according to (3), could have been caused by shifts in product demand [the $d(\ln D_i)$ term in (3)] or changes in technology [the $d(\ln b_i)$ term in (3)] that were relatively favorable to college-educated workers.

The other possible explanation of the increase in w is that the initial relative-demand function was n_d' but that noncollege workers received higher rents (through the effects of unionism) such that the equilibrium of the economy was point a in Figure 3. The increased relative supply of college graduates would have driven w down to w_1', but the rents disappeared during the

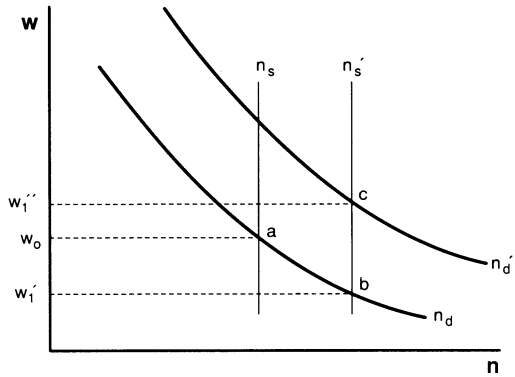

FIGURE 3. SHIFTS IN THE RELATIVE DEMAND FOR (n_d) AND RELATIVE SUPPLY OF (n_s) COLLEGE-EDUCATED LABOR)

1980's, resulting in an increase in the college relative wage to w_1'', an amount equal to the distance between points b and c in Figure 3 above what it would have been in the absence of whatever institutional changes occurred.

III. Evidence on Alternative Explanations

We now apply the model developed in Section II to the question of the causes of the wage structure that were described in Section I, examining in turn explanations that focused on changes in average rents, changes in the structure of product demand, and technical change. These explanations are, of course, not mutually exclusive.

A. *Changes in the Industrial Wage Structure*

In order to discern how much of the observed wage-structure changes was due to changes in the demographic composition of employment between high- and low-wage industries and how much was due to changes in industry wage differentials, it is necessary to estimate group-average wage rates by industry, the Y_{ij}'s. These can be obtained from the estimated parameters of our original 32 regressions for each of the three years in which log CPS wages were regressed on experience and dummy variables for education, part-time status, nonwhite,

TABLE 2—ESTIMATED AGGREGATE WEIGHTS (ϕ_j) AND WAGE EFFECTS (M_j) FOR 17 CPS INDUSTRIES IN 1973, 1979, AND 1988

	Weight (ϕ_j)			Wage effect (M_j)		
Industry	1973 (i)	1979 (ii)	1988 (iii)	1973 (iv)	1979 (v)	1988 (vi)
Construction	0.067	0.059	0.059	0.206	0.137	0.112
Durables/mining	0.190	0.179	0.141	0.073	0.089	0.099
Nondurables	0.119	0.107	0.088	0.018	0.020	0.025
Transport	0.041	0.042	0.039	0.112	0.122	0.069
Utilities	0.035	0.035	0.034	0.151	0.147	0.166
Wholesale trade	0.041	0.041	0.041	0.014	0.000	−0.004
Retail trade	0.137	0.141	0.147	−0.202	−0.150	−0.175
Finance	0.054	0.065	0.075	0.029	0.012	0.079
Business services	0.028	0.032	0.054	−0.079	−0.072	−0.051
Personal services	0.019	0.019	0.024	−0.282	−0.179	−0.218
Entertainment	0.007	0.008	0.010	−0.133	−0.098	−0.120
Medical	0.025	0.031	0.039	−0.068	−0.063	−0.058
Hospitals	0.044	0.048	0.048	−0.006	0.015	−0.049
Welfare	0.014	0.017	0.024	−0.282	−0.222	−0.252
Education	0.096	0.090	0.084	−0.103	−0.107	−0.055
Professional services	0.019	0.023	0.028	0.084	0.044	0.104
Public administration	0.065	0.064	0.068	0.126	0.070	0.093

area size, region, and major industry of employment. The average log wage for group-i workers in industry j is the predicted value for the particular educational level (8, 12, 14, or 16) and experience level (5, 15, 25, or 35) for white, full-time workers residing in SMSA's in the average region for that group with the relevant industry coefficient in effect.[9]

The next step is to estimate average industry wage effects across all groups in each period. For each of the three years these were obtained from a weighted (by $[k_i\phi_{ij}]^{0.5}$) regression of the 32×17 Y_{ij}'s on dummy variables for (all but one of) the 32 groups and 17 industries. This decomposes each Y_{ij} into a group effect and a common-industry effect, and the deviations of the estimated value of the latter from its mean in the three years are reported as M_j (the estimated value of the log of μ_j) in columns (iv)–(vi) of Table 2. For example, in 1973 the estimated average hourly wage in construction relative to education, other things held constant, was $\exp[0.206-(-0.103)]=1.36$.[10]

The estimated contribution of changes in average industry wage effects to the relative wage change of each demographic group can then be calculated as $\Sigma_j M_j\Delta\phi_{ij}+\Sigma_j(\phi_{ij}+\Delta\phi_{ij})\Delta M_j$. The first term is the part due to changes in the industry weights, and the second term represents the parts

[9]Tables reporting the relevant Y_{ij}'s as well as the industry distribution of employment, the ϕ_{ij}'s, are available from the authors upon request.

[10]The estimates of these industry wage effects are based on the specification that the proportionate effect of working in industry j is the same for all I groups (i.e., $\mu_{ij}=\mu_j$). This is, of course, testable, and it turns out that there are a few exceptions to the assumption. For example, noncollege men earned about 3-percent more than other groups in the first five industries listed in Table 2 (the location of the majority of blue-collar unionism). College women earned 20-percent more than other groups in 1973 in the four "public service" industries (hospitals, welfare and religious, education, and public administration), possibly reflecting a combination of discrimination in other industries and self-selection, but this coefficient fell to insignificance in the later years. Although these exceptions to the assumption of identical industry effects are interesting, they have virtually no effect on the quantitative contribution of industry wage effects to the explanation of wage-structure changes and, therefore, are ignored.

TABLE 3—ESTIMATED EFFECTS OF CHANGES IN INDUSTRY WAGE EFFECTS AND IN UNION MEMBERSHIP ON RELATIVE WAGE CHANGES DURING 1973–1979 AND 1979–1988

Comparison groups	Sex	Relative wage change (i)	Industry effects: Weights (ii)	Industry effects: Wages (iii)	Union effect (iv)
A. 1973–1979:					
College/high school	men	−0.035	0.009	−0.002	0.012
	women	−0.073	0.005	−0.007	0.016
High school/dropout ($X < 30$)	men	−0.006	0.007	−0.001	0.004
	women	−0.002	−0.003	−0.010	0.004
Old/young (noncollege)	men	−0.004	0.007	−0.004	0.001
	women	0.058	0.012	−0.001	0.003
Women/men		0.011	0.002	0.012	0.004
B. 1979–1988:		(v)	(vi)	(vii)	(viii)
College/high school	men	0.163	0.016	0.020	0.013
	women	0.118	0.010	0.005	−0.006
High school/dropout ($X < 30$)	men	0.072	−0.001	0.003	0.002
	women	0.060	0.002	0.009	0.006
Old/young (noncollege)	men	0.107	0.010	0.005	0.010
	women	0.043	0.009	0.004	0.002
Women/men		0.076	−0.001	0.006	0.007

due to changes in industry wage effects (evaluated at the weights at the end of the period). The values of the estimated contribution of the two sources of change in average industry wage effects to the summary relative wage changes for 1973–1979 and 1979–1988 discussed in Section I are reported in Table 3. Columns (i) and (v) give the relative wage changes in the 1970's and 1980's, (ii) and (vi) give the effects of changes in industry weights, and (iii) and (vii) give the effects of industry wage effects. For example, of the 0.163 proportional increase in the wage of male college graduates relative to high school graduates during the 1980's, 0.016 can be attributed to differential movements between high- and low-wage industries and 0.022 can be attributed to changes in industry wage effects, or a total of $0.038/0.163 = 22$ percent of the change. All of the estimated changes in net average industry effects for the 1980's are in the right direction, but they do not explain a very large part of any of the summary relative wage changes.[11]

An alternative way of looking specifically at the question of whether workers' relative rents changed is to examine changes in the

[11]The use of only 17 industries in our analysis, which is done for purposes of consistency with the rest of the paper, raises the possibility of downward aggregation bias. To check for this, we ran regressions for all men and all women (instead of separate regressions for 16 education/experience groups) in each period, using 45 detailed industries instead of the 17 major industries in the above analysis. The results implied that most of the total industry wage-structure effects (the sum of the compositional and wage-change effects) were picked up by the dummy variables for 17 industries. The major exceptions to this were confined to noncollege men during the 1980's, for whom the use of detailed industry dummies increased the total industry wage effects by up to one-half. However, this still falls far short of a full explanation of the relevant wage-structure changes.

incidence of unionism. Between 1979 and 1988, the fraction of all workers in our sample who were union members fell from 27 percent to 19 percent, a slight acceleration of the decline in union membership that has been occurring since the mid-1950's.[12] Assume that the wage of each worker is equal to the competitive wage for the relevant demographic group times the union/nonunion relative wage if the worker is a union member ($U=1$), that is, $W_{ic}\exp(\lambda U)$. Then the geometric average of the wages of demographic group i is $Y_i = Y_{ic} + \lambda U_i$, where λ is the logarithmic union/nonunion wage differential and U_i is the proportion of group-i workers who are unionized. Holding λ constant and assuming no feedback from the extent of unionization to the competitive wage, the contribution of changes in unionization to the wage of each group is $\lambda \Delta U_i$.

Using the upper-bound estimate of λ by H. Gregg Lewis (1986), 0.15, estimates of the contribution of changes in unionism to the relative wage change of each demographic group in each of the two time periods of this analysis can then be calculated. These calculations for the summary relative wage changes are reported in columns (iv) and (viii) of Table 3. For example, the fraction of male high school graduates (with no college) who were union members fell from 0.385 in 1979 to 0.270 in 1988, and the corresponding figures for male college graduates were 0.174 and 0.146. Thus, the estimated contribution of changes in union incidence to the male college/high-school relative wage for the 1980's was $0.15\times[-0.028-(0.115)]=0.013$. It is clear from inspection of column (viii) that the decline of unionism in the United States during the 1980's had *at most* a small effect on overall relative wage changes.

[12] Our union-membership data were taken from the May version of the CPS for 1973, 1979, and 1988. For an analysis of the causes of the overall decline in unionization in the United States see Henry Farber (1987).

B. *Changes in the Structure of Product Demand*

It is clear that not very much of the wage changes of the 1980's can be explained—even with perhaps unrealistically favorable assumptions—by changes in the industrial wage structure or in the incidence of unionism. It is necessary to focus (as other investigators, such as Murphy and Welch [1991], have concluded) on changes in relative competitive wage levels.

Changes in the structure of product demand can in principle shift the relative labor-demand functions for different groups, and the presumption of this explanation is that in the 1980's these demand shifts were both in the right direction and large. As can be seen from inspection of columns (ii) and (iii) of Table 2, there was a significant shift during the 1980's in the fraction of total employment ($\phi_j = \sum_i \phi_{ij} k_i$) from industries that were the traditional employers of male blue-collar labor (like manufacturing) toward industries that employ larger fractions of women and highly educated labor (like finance and professional services). In addition to lowering the average wages of blue-collar males because wage levels are higher in manufacturing than in the expanding industries (the weights effect, $\sum_j M_j \Delta\phi_{ij}$, in Table 3), such a shift would also have lowered their wage levels in, say, hardware stores because they are less scarce.

One index of the influence of product demand shifts on relative labor-demand functions, which has been used by Richard Freeman (1975), Murphy and Welch (1991), and Lawrence Katz and Murphy (1992), is the average employment growth by industry weighted by the initial employment distribution of each demographic group. This is

$$\text{(4)} \qquad \mathrm{EMP}_i = \sum_j \Delta(\ln\phi_j)\phi_{ij}$$

where $\Delta(\ln\phi_j)$ is the proportionate change in the logarithm of industry j's share of aggregate employment in each period. The

TABLE 4—PROPORTIONATE EMPLOYMENT CHANGES ($\Delta[\ln \phi_j]$) AND DERIVED DEMAND INDEXES ($\Delta[\ln x_j]$) BY INDUSTRY, FOR 1973–1979 AND 1979–1988

	1973–1979			1979–1988		
Industry	$\Delta\phi$ (i)	Δx_1 (ii)	Δx_2 (iii)	$\Delta\phi$ (iv)	Δx_1 (v)	Δx_2 (vi)
Construction	−0.126	0.000	0.037	0.014	0.148	0.182
Durables/mining	−0.057	0.034	−0.041	−0.224	−0.171	−0.270
Nondurables	−0.096	−0.034	−0.109	−0.179	−0.125	−0.221
Transport	0.035	0.096	0.021	−0.070	−0.027	−0.135
Utilities	−0.004	−0.019	0.092	−0.007	−0.079	−0.201
Wholesale trade	−0.004	0.002	0.045	−0.023	0.044	0.099
Retail trade	0.036	0.050	0.087	0.061	0.167	0.203
Finance	0.178	0.075	0.120	0.156	0.112	0.117
Business services	0.128	0.133	0.173	0.534	0.573	0.617
Personal services	−0.023	−0.013	0.005	0.252	0.365	0.399
Entertainment	0.210	0.132	0.173	0.211	0.265	0.310
Medical	0.215	0.105	0.152	0.241	0.195	0.252
Hospitals	0.080	−0.036	0.013	0.018	−0.123	−0.060
Welfare	0.230	0.055	0.117	0.355	0.237	0.320
Education	−0.057	−0.215	−0.146	−0.059	−0.286	−0.198
Professional services	0.242	0.087	0.140	0.199	0.123	0.200
Public administration	−0.008	−0.056	−0.006	0.072	−0.027	0.048

calculated values of $\Delta(\ln \phi_{ij})$ for each of 17 major industries for 1973–1979 and 1979–1988 are reported under $\Delta\phi$ in columns (i) and (iv) of Table 4. EMP_i can be considered as a rough proxy for the discrete version of $d(\ln D_i)$ in (3). The presumption of the demand-shift explanation of the wage-structure changes of the 1980's is that the EMP_i's were both in the right direction (i.e., toward more-educated, older, and female labor) and sufficiently large to overwhelm the "perverse" effect of increased relative supplies of most of the demographic groups whose relative wages increased.

In fact, the values of EMP_i are positively correlated with changes in relative supply for the 1979–1988 period (the slope coefficient of a weighted regression of EMP_i on the change in the log change in supply across the 32 demographic groups was 0.049 with a standard error of 0.020). Although the size of these shifts is not sufficiently large to constitute the whole explanation, they at least have the right sign, so previous studies have concluded that part of the explanation of the wage-structure phenomena is found in product demand shifts.

A problem with this conclusion is that it is necessary to net out the effect of changes in the relative supply of the different groups. If, for example, there were a very large increase in the relative supply of college-educated labor over some time interval, we would expect, other things equal, a tendency for industries that are the most skill-intensive to grow relative to other industries. The use of total relative-employment changes by industry as a proxy for product demand shifts, therefore, may confound product demand shifts with relative-supply changes.

An alternative approach that gets around this possible bias is to estimate a discrete version of Appendix equation (A10), and the resultant estimated demand-shift indexes by industry for the two periods are reported under Δx_1 in columns (ii) and (v) of Table 4. Because the estimated value of $\Delta(\ln x_j)$ picks up employment growth in each industry as a deviation from the weighted rates of growth across demographic groups of its initial employment distribution, $\Delta(\ln x_j)$ is greater than $\Delta(\ln \phi_j)$ in industries (like construction) that tended to hire

TABLE 5—PROPORTIONATE SUPPLY CHANGES, ALTERNATIVE PRODUCT-DEMAND-SHIFT INDEXES, AND SPECIFIC-INDUSTRY TECHNICAL CHANGE BY AGGREGATED GROUPS FOR 1973–1979 AND 1979–1988

Comparison groups	Sex	Supply (i)	Demand-change indexes: EMP (ii)	DEM$_1$ (iii)	DEM$_2$ (iv)	Specific technical change (v)
A. 1973–1979:						
College/high school	men	0.204	0.033	−0.051	−0.014	0.016
	women	0.172	−0.039	−0.132	−0.091	0.004
High school/dropout ($X<30$)	men	0.304	0.013	−0.002	−0.001	−0.012
	women	0.316	0.042	0.005	0.028	−0.006
Old/young (noncollege)	men	−0.154	−0.002	−0.006	−0.010	0.028
	women	−0.212	−0.014	−0.012	−0.012	0.043
Women/men		0.119	0.038	−0.021	0.005	−0.023
B. 1979–1988:		(vi)	(vii)	(viii)	(ix)	(x)
College/high school	men	0.176	0.048	−0.032	0.022	0.033
	women	0.334	0.001	−0.127	−0.082	0.016
High school/dropout ($X<30$)	men	0.338	0.006	−0.007	−0.006	0.031
	women	0.310	0.042	0.021	0.048	0.061
Old/young (noncollege)	men	−0.109	−0.019	−0.035	−0.039	0.109
	women	0.012	−0.008	−0.029	−0.029	0.052
Women/men		0.164	0.053	0.011	0.046	−0.037

low-educated and male labor and smaller in industries (like education) that have the opposite demographic composition.

The next step is to calculate a derived demand-shift index, $\mathrm{DEM}_{1i}=\Sigma_j\Delta(\ln x_j)\phi_{ij}$, which is analogous to the calculation of the EMP_i index in (4). For the 1979–1988 period, however, this index exhibits a small *positive* correlation with relative-supply changes, which is not favorable to acceptance of the product-demand-shift explanation as an important source of the wage-structure developments of the 1980's. What is in fact happening is that, while the traditional employers of males with low education were declining during the 1980's, so were some industries that traditionally employ a large fraction of college graduates, notably education and public administration.

The average values of the alternative demand-shift indexes for the summary demographic groups are reported in Table 5, in columns (ii) and (vii) under EMP for the index based on industry employment changes and in columns (iii) and (viii) under DEM_1 for the index based on estimated derived demand changes. For example, the proportional change in the relative supply of male college graduates relative to high school graduates [from column (vi)] was 0.176, which, by (3), implies that their relative wages should have decreased by $(1/\sigma)\times0.176$ (instead of rising by 0.163). Use of the EMP index of relative-demand changes suggests that $(1/\sigma)\times0.048$ of the failure of this relative wage to fall could be accounted for by product demand shifts. Use of the DEM_1 index, however, only deepens the mystery, for that index suggests that demand shifts were on balance slightly unfavorable to highly educated labor.

C. *Intra-industry Employment Shifts*

It is clear from the results thus far that the major wage-change phenomena of the 1980's are *not* adequately accounted for by explanations based on institutional factors or changes in the structure of product demand. This leads us to the consideration of the remaining possibility, that the 1980's were characterized by major changes in technology that were nonneutral with respect to different types of labor. Variations across demographic groups in the ceteris paribus effects on wages of technical change or changes in average group quality are reflected in the $(1-1/\sigma)d(\ln b_i)$ term in (3). Given the maintained assumption that $\sigma > 1$, the wage-structure facts are attributable to this set of explanations if the relative values of the b_i's for the more educated, older, and female demographic groups increased during the 1980's. The major difficulty with this explanation, unlike the explanations involving industry wage effects, supply, and product demand, is (as in the analysis of the sources of economic growth) that it involves the residuals of the intrafactor demand function rather than directly observable phenomena.

It has long been argued that in periods of rapid technical change the relative demand for highly educated workers may increase because of their superior ability to adapt to and refine new methods of production (see Richard Nelson and Edmund Phelps, 1966).[13] The 1980's, as well as the 1970's to some extent, have been characterized popularly as a period in which computer technology was adopted throughout most of the U.S. economy, and there have been several case studies that suggest that changes in production methods have been favorable to professional and technical workers relative to blue-collar workers (see e.g., Jerome Mark, 1987). To the extent that this technical change was common across most industries in the economy, we would expect that the $d(\ln b_i)$'s for certain demographic groups would have risen relative to others.[14]

If it were true, however, that the rate of growth of the technical-efficiency parameter for group-i workers in industry j, $d(\ln b_{ij})$, were equal to the weighted mean for that group, $d(\ln b_i)$, plus a random error, all that is predicted by the model in Section II is that the relative wages of those groups most favored by the technical change would rise. Since changes in the b_i's are not observed, there is no way to test this. The thrust of the recent literature on technical change, however, is that the effects of spurts of innovation on the relative demand for skilled labor may vary across industries. If this is so, some of the variation in the values of $d(\ln b_i)$ can be identified.

To make this concrete, suppose that there is a subset of industries, say J′, in which the rate of growth of the efficiency parameters for a subset of the demographic groups, say I′, differs from their average growth in other industries. Specifically,

$$(5)\quad d(\ln b_{ij}) = \begin{cases} c_{0i} + c_{1i} & i \text{ in I}' \text{ and } j \text{ in J}' \\ c_{0i} & \text{otherwise.} \end{cases}$$

This implies that the average change across all industries in the efficiency parameters

[13]Several implications of the complementarity between human capital and the rate of technical change were tested by Welch (1970) and more recently by Ann Bartel and Frank Lichtenberg (1987) and Jacob Mincer (1989).

[14]It is difficult to argue that the 1970's and 1980's were a period of rapid growth in overall labor productivity, for, as pointed out in Section I, the average real wage has been essentially stagnant since 1973 as compared to its 1.5–2-percent annual growth over the preceding 150 years. An alternative interpretation of recent technical developments, in terms of the CES production function given by (A1), is that the new computer technology caused a relative shift in the δ_{ij}'s toward highly skilled demographic groups. Given $\sigma > 1$, this is observationally equivalent to relative increases in the b_i's for highly skilled demographic groups. In a time-series analysis of CPS earnings differentials, Mincer (1991) provides some direct evidence that the partial productivity rebound of the 1980's was skill-biased.

for a group in I′ is $d(\ln b_i) = c_{0i} + c_{1i}T_i$, where T_i is the proportion of group i's employment that is in the J′ industries. To estimate the extent of this group/industry specific technical change, note that

$$(6)\quad d\left[\ln(b_{ij}/b_i)\right] = c_{1i}D_{\mathrm{I}'}(D_{\mathrm{J}'} - T_i)$$

where $D_{\mathrm{I}'}$ and $D_{\mathrm{J}'}$ are dummy variables for the relevant groups and industries. Following Appendix equation (A12), equation (6) should be substituted for $d[\ln(b_{ij}/b_i)]$ in the regression equation to estimate industry demand shifts, and the coefficient on this variable is an estimate of $(\sigma - 1)c_{1i}$. The resultant proportional change in the average value of b_i for each group can then be calculated as

$$(7)\qquad d(\ln b_i) = c_{0i} + c_{1i}T_i.$$

In other words, the change in the average-efficiency parameter for group-i workers equals a *general* component c_{0i}, which applies to all industries (and is unobservable), plus a specific component $c_{1i}T_i$, which applies only to certain industries.

Looking at the underlying data on the industry distributions of employment for the 32 demographic groups (the ϕ_{ij}'s), it is clear that in both the 1973–1979 and 1979–1988 periods there was a greater shift out of manufacturing and similar industries for younger and less-educated workers than for those with the opposite characteristics. For example, the fraction of male dropouts with $X < 10$ who were employed in durable goods and mining fell from 0.252 in 1979 to 0.157 in 1988, but the fraction of male college graduates with $X < 10$ employed in this sector increased from 0.148 to 0.150 during this period. Further analysis of the data suggested a similar pattern in four of the five traditional blue-collar industries (durables/mining, nondurables, transportation, and public utilities, but *not* in construction). These four industries were aggregated into the J′ sector (that in which there was differential technical change by demographic group). The group characteristics that were selected to be included as dummy variables included those that were seen in Section I to have had an important influence on relative wage changes in the 1979–1988 period: the four educational groups and the four experience groups separately for men and women, the four experience groups for those who had not completed college separately for men and women, and gender by itself.

The test of the null hypothesis that the 13 dummy variables that represent the above characteristics did not add to the explanation of $\Delta(\ln \phi_{ij})$ was rejected in both periods at better than the 0.001 level. The results show that there was a major shift in the employment structure in this blue-collar sector toward more-educated workers and, for those who had not completed college, older workers. Further, these shifts were slightly stronger in the 1980's than in the 1970's.

The summary values of $(\sigma - 1)c_{1i}T_i$ for the aggregated comparison groups are reported under SPEC in columns (v) and (x) of Table 5. It is clear that the intra-industry shifts represented by the variable were an important source of the relative demand shifts for high-school/dropout and the non-college old/young comparisons, but they were unfavorable for women relative to men in the aggregate. Inclusion of the dummy variables to capture intra-industry employment shifts also changes the estimated derived demand indexes, the $\Delta(\ln x_j)$'s, which are reported under Δx_2 in columns (iii) and (vi) of Table 4. These estimates are generally between the $\Delta(\ln \phi_j)$'s and the original $\Delta(\ln x_j)$'s, and the resultant indexes of the estimated effect of product demand shifts on labor demand, which are reported for the summary comparison groups under DEM_2 in columns (iv) and (ix) of Table 5, are usually between EMP and DEM_1.

An alternative interpretation of the positive effect of experience on the employment shares of workers who had not finished college in the blue-collar sector concerns the effect of seniority systems in the face of declining employment in these industries. Since these industries pay relatively high wages (see Table 2) and thus have low rates of labor turnover, a reduction in their relative importance in the economy (from a 0.385 share of nonagricultural employment in 1973 to 0.363 in 1979 and then to just 0.302 in 1988) must have caused a reduction in their hiring of young workers. Thus, the employment share of younger noncollege workers in the blue-collar sector fell, not because of a decline in the productivity of less-experienced relative to more-experienced workers, but because older workers had the right to retain their "good" jobs.[15]

D. *General Technical Change*

Variation in the other component of the proportionate change in the average-efficiency parameter for each group, c_{0i}, is not directly observable. It is, however, clear from inspection of the summary measures in Table 5 that the addition of relative specific technical change to the other explanations does not add enough to outweigh the perverse effects of relative supply changes for the 1980's. Obviously, something else is going on, and the only remaining candidate within our structure is variation in the c_{0i}'s. Fortunately, it appears that other things appear also to have been affecting wage changes in the 1970's, and this permits us to estimate the c_{0i}'s indirectly.

Following (3), the per annum growth of the relative wage of group-i workers over each period ($t=1$ for 1973–1979 and $t=2$ for 1979–1988) may be written as

$$(8)\quad dY_{ai}(t) = -(1/\sigma)dN_{ai}(t) + (1-1/\sigma)c_{0i}(t)' + u_i(t)$$

where $dY_{ai}(t)$ is the annualized proportionate change in the relative wage of group-i workers adjusted for the change in total average industry wage effects, $dN_{ai}(t)$ is the per annum proportionate change in relative supply adjusted for product demand shifts and industry-specific technical change, $c_{0i}(t)'$ is the per annum value of general technical change, and $u_i(t)$ is a random error term. It then follows that the difference between the rates of growth of adjusted relative wages in the two periods is

$$(9)\quad d^2Y_{ai} = -(1/\sigma)d^2N_a + (1-1/\sigma)[c_{0i}(2)' - c_{0i}(1)'] + u_{i'}$$

where $d^2Y_{ai} = dY_{ai}(2) - dY_{ai}(1)$, $d^2N_{ai} = dN_{ai}(2) - dN_{ai}(1)$, and $u_i' = u_i(2) - u_i(1)$.

We now specify that the growth in each group's efficiency parameter relating to all industries in the 1979–1988 period equals its value in the 1973–1979 period plus a difference A_i, that is,

$$(10)\qquad c_{0i}(2)' = c_{0i}(1)' + A_i.$$

It is assumed initially that A_i is uncorrelated with either $c_{0i}(1)'$ or d^2N_{ai}, which is equivalent to assuming that the pattern of general technical change in the two periods was (more or less) identical. If these assumptions are correct, the reciprocal of the elasticity of intrafactor substitution can be estimated by regressing d^2Y_{ai} on d^2N_{ai}, for the influence of general technical change disappears as a fixed effect. The estimated slope coefficient of a regression (weighted by $[k_i(1979)]^{0.5}$) of d^2Y_{ai} on d^2N_{ai} is -0.588 (SE $= 0.127$), which implies a value of σ of 1.70. This is approximately the midpoint of past estimates of that parameter (see the surveys by Hamermesh and James Grant [1979] and Richard Freeman [1986]).[16] It is then possible to obtain an estimate of the effect (common to both periods) of general technical change on the per annum growth rate of group-i workers, $(1-1/\sigma)c_{01}'$, by computing the average of the residuals,

[15]The fact that there was no perceptible change in the age composition of employment in the construction industry is consistent with this seniority interpretation of the results. Construction is characterized by very high rates of labor turnover, so seniority is much less important than in the other blue-collar industries.

[16]The estimated coefficient of dY_{ai} on dN_{ai} for the 1973–1979 and 1979–1988 periods are, respectively, -0.077 (SE $= 0.023$) and $+0.094$ (SE $= 0.044$). The first of these yields an estimate of the intrafactor substitution elasticity that is implausibly large (the implicit σ is $1/0.077 \approx 13$), and the second has the wrong sign. Obviously the dN_{ai}'s in both periods are positively correlated with some omitted variable.

$dY_{ai}(t)+(1/\sigma)dN_{ai}(t)$, over the two periods.[17]

To test for the possibility that the pace of general technical change for some groups may have risen or fallen from the 1970's to the 1980's (i.e., that certain A_i's were not zero) we added several dummy variables for sets of demographic groups to the right-hand side of (9). The only one that yielded a statistically significant result was that for five young, low-education groups.[18] Inclusion of this dummy variable, YNGLO, yielded an estimated coefficient on d^2N_{ai} of -0.571 (SE = 0.128) and a coefficient on YNGLO of -0.026 (SE = 0.012). This result is consistent with $\sigma = 1.75$, and it implies that the relative wages of young workers with low levels of education fell by 2.6 percent per annum faster than they would have in the absence of the acceleration of technical change against them. With this modification, the estimated common values of $(1-1/\sigma)c_{0i'}$ were recalculated. The estimated effects of general technical change on relative wage changes are $\text{GEN}_i(1) = 6\times(1-1/\sigma)c_{0i'}$ for the 1973–1979 period and $\text{GEN}_i(2) = 9\times[(1-1/\sigma)c_{0i'} - 0.026\ \text{YNGLO}_i]$ for the 1979–1988 period. The average relative values of these estimates for the summary comparison groups are reported under GEN in columns (vi) and (xiii) of Table 6.

It is apparent from inspection of the estimated values of GEN for the 1980's that our major conclusion, which will be discussed more completely below, is that the principal cause of the significant wage-structure changes of the past decade was a shift in the structure of the b_i's that were extremely favorable to certain groups, especially women and the highly educated. We have interpreted the source of this shift as an exogenous change in technology, but the basic result has other interpretations.

First, the average value of GEN of women relative to men during the 1980's was 0.145, which means that, relative rents, supply, product demand, and intraindustry composition held constant, women's wages grew 1.6 percent per year faster than men's wages because of relative proportional changes in average b_i's. (The corresponding value for the 1973–1979 period was 1.4 percent.) Some of this may have been attributable to changes in production technology that were relatively favorable to women,[19] but much of the decline in the gender gap may reflect an improvement in the unobserved labor quality of women (see James Smith and Michael Ward, 1984; June O'Neill, 1985). Since our results are based on CPS data, which do not measure actual as opposed to potential labor-market experience, we cannot identify how much of the decline in the gender differential was due to technical change versus an increase in the average extent of labor-market attachment on the part of women (or other explanations, such as a gradual decline in labor-market discrimination against women).[20]

A second problem of interpretation involves the large negative values of GEN for

[17]A lower-bound estimate of the standard error of $(1-1/\sigma)c'_{0i}$ is the standard error of estimate of the second-difference demand function, which equals 0.0284 divided by two. This implies that the standard errors of the estimated effects of general technical change on relative wage changes are (at least) 0.085 and 0.128 for the 1973–1979 and 1979–1988 periods, respectively.

[18]The five groups with unit values of YNGLO include both men and women dropouts and high school graduates in the lowest experience interval and male dropouts with $X = 10$–19. Several other dummy variables were added to this regression (all women, all college graduates, and so forth), but the estimated coefficients on these variables were small and statistically insignificant, indicating that their per annum c_{0i}'s did not change between the 1970's and 1980's.

[19]Compared to men, women tend to work in occupations that on average impose higher intellectual, as opposed to physical, demands (see Johnson and Gary Solon, 1986). Changes in production practices brought about by the introduction of computers would thus tend to favor "women's jobs" vis-à-vis "men's jobs."

[20]Using data from the Panel Study on Income Dynamics, which has detailed information on respondents' work histories, Allison Wellington (1991) reports that nearly half of the observed increase in wages of women relative men from 1976 to 1985 can be attributed to changes in job- and labor-market-attachment variables. This still leaves a great deal of women's improvement unexplained.

TABLE 6—DECOMPOSITION OF ESTIMATED SOURCES OF 1973–1979 AND 1979–1988 RELATIVE WAGE CHANGES

			Source of relative wage change					
						Technical change		
Comparison groups		Relative wage change (i)	Rents (ii)	Supply (iii)	Demand (iv)	Specific (v)	General (vi)	Unexplained (vii)
A. 1973–1979:								
College/ high school	men	−0.035	0.007	−0.117	−0.008	0.009	0.073	0.001
	women	−0.073	−0.002	−0.098	−0.052	0.002	0.120	−0.043
High school/ dropout ($X < 30$)	men	−0.006	0.006	−0.174	0.000	−0.007	0.153	0.016
	women	−0.002	−0.013	−0.181	0.016	−0.004	0.158	0.022
Old/young (noncollege)	men	−0.004	0.003	0.088	−0.006	0.016	−0.104	−0.001
	women	0.058	0.011	0.121	−0.007	0.024	−0.076	−0.015
Women/men		0.016	0.013	−0.066	0.003	−0.013	0.086	−0.007
B. 1979–1988:		(viii)	(ix)	(x)	(xi)	(xii)	(xiii)	(xiv)
College/ high school	men	0.163	0.036	−0.100	0.013	0.019	0.196	−0.001
	women	0.118	0.015	−0.191	−0.047	0.009	0.270	0.062
High school/ dropout ($X < 30$)	men	0.072	0.002	−0.193	−0.003	0.018	0.267	−0.019
	women	0.060	0.011	−0.177	0.027	0.035	0.202	−0.038
Old/young (noncollege)	men	0.107	0.015	0.062	−0.022	0.062	−0.023	0.013
	women	0.043	0.013	−0.007	−0.016	0.030	−0.006	0.029
Women/men		0.076	0.005	−0.094	0.026	−0.021	0.145	0.015

younger workers with low levels of education, especially in the 1980's. The discussion of this section has been in terms of an exogenous change in technology that lowered the b_i's of this group, but this is also subject to other explanations. The first of these is the possibility that workers with low levels of education who entered the labor market in the 1980's had a much lower level of innate ability than their older counterparts who entered the labor market in the 1970's. If this were true, the low relative values of the $d(\ln b_i)$'s would be a reflection of a recent decline in the effectiveness of precollege education rather than an exogenous change in technology.[21] Another possible explanation of the decline in the relative wages of young workers with low skill levels is that they were the most susceptible to competition from undocumented immigrants. In this case, the low $d(\ln b_i)$'s for these groups would be reflecting the underestimation of the "true" $d(\ln N_i)$'s applying to this type of labor.[22]

[21]John Bishop (1991) reports evidence of a widening in academic achievement scores between college and noncollege youth during the 1980's. Perhaps more importantly, Bishop also reports a sharp increase in the fraction of college students majoring in relatively renumerative fields such as business. It is therefore possible that the more-educated members of the youngest cohort in our analysis are relatively more motivated (in the acquisitive sense) than was true of the older cohorts.

[22]For evidence that the recent wave of immigration has had a negative effect on the wages of "native"

There is, on the other hand, some direct evidence in favor of the interpretation of variation in the $d(\ln b_i)$'s across demographic groups as reflecting changes in technology. The first concerns the effect on individual wages of the use of computers. Respondents of supplementary surveys of the CPS in 1984 and 1989 were asked whether they used computers on their jobs, and Krueger (1991) has estimated that the proportional ceteris paribus effect on wages of computer use was 0.170 in 1984 and 0.188 in 1989. The fraction of all workers who reported using computers on the job rose from 25 percent to 37 percent in 1989, other determinants held constant, and both the incidence of computer use and its absolute increase over the five-year period were greater for women than for men (43 percent versus 32 percent in 1989) and greater for more-educated workers (8, 29, and 59 percent in 1989 for dropouts, high school graduates, and college graduates, respectively). Further, Krueger estimates that from one-third to two-thirds (depending on model specification) of the 1984–1989 increase in the estimated effect of education is directly attributable to the use of computers.[23]

A second piece of direct evidence in favor of the technical-change interpretation of variation in the $d(\ln b_i)$'s concerns the effect of "high-tech" capital on the structure of labor demand within manufacturing industry. Using Bureau of Economic Analysis data, Ernst Berndt and Catherine Morrison (1991) report estimates of the effects of this type of capital, which includes computer, communications, and photocopy equipment and instruments, as distinct from other producers' durable equipment and structures. They report a dramatic increase in the ratio of high-tech to total capital stock in manufacturing from 0.095 in 1976 to 0.257 in 1986. Further, within two-digit manufacturing industries, increases in the high-tech intensity of capital are associated with both shifts in labor demand from production toward nonproduction workers and increases in the average educational attainment of production workers.[24]

E. *Decomposition of Relative Wage Changes*

We have now accumulated evidence on each of the potential explanations of the dramatic changes in the wage structure in the 1980's and can summarize the results. Following (3), the proportionate relative wage change of each demographic group equals the change in total industry wage effects plus the effects of relative supply changes, product demand shifts, and average technical change, which we have separated into that arising in specific industries and in general. Estimates of the contribution of each of these effects for the comparison groups are reported for 1973–1979 in columns (ii)–(vi) of Table 6 and for 1979–1988 in columns (ix)–(xiii). Columns (vii) and (xiv) are the amounts of these relative wage changes that remain unexplained.

The decompositions for the 1980's suggest a fairly consistent story. First, total changes in average industry wage effects

workers with low skill levels, see Joseph Altonji and David Card (1991). The potential impact of undocumented immigration on relative wages in the context of our model can be assessed by assuming that (a) these immigrants are perfect substitutes for the demographic groups in YNGLO and (b) their labor force was equal to 3 million in 1979 (but not enumerated in official statistics). In order to have been responsible for all of the relative specific-industry and general technological change against the YNGLO groups from 1979 to 1988, the employment of undocumented immigrants would have to have increased over this period by 26.6 million (compared to the native YNGLO work force of 19.2 million).

[23]Krueger's estimates of the fraction of the increase in the educational differential explained by computers refer only to their *direct* effect on wages. There was also an *indirect* effect, due to the reduced supply for other functions (e.g., managing) of those groups that had the largest increases in use of computers. There is no way to resolve with current data sets how large these indirect effects might have been, but our results suggest that it is possible that they were *very* large.

[24]Berndt and Morrison (1991) also find a complementarity between skill and the other forms of capital, which is consistent with the hypothesis of Zvi Griliches (1969). They find, however, that high-tech capital is more complementary with skill than are the other types.

[column (ix)] were in the right direction but accounted for a small fraction of relative wage changes. Second, for the comparisons involving education and gender, relative supply changes [column (x)] were large and in the wrong direction. The relative decrease in the supply of older relative to younger noncollege males, however, does account for a large proportion of the increase in the slope of the age/earnings profile for that group. Third, our estimates of the effects of product demand shifts (based on DEM_2 in Table 5) on relative wages are small and of uneven direction. Fourth, the two forms of technical change, SPEC and GEN, comprise the principle source of the increase in educational differentials and the decrease in the gender differential, and the large positive values of SPEC for older noncollege workers account for a large amount of the increase in their relative wages.

An alternative way to explain the wage changes of the 1980's, an approach followed most recently by Katz and Murphy (1992), is to focus on the deceleration of changes in the demographic composition of the labor force from the 1970's to the 1980's. For example, during the 1970's the per annum rate of growth of college graduates relative to high school graduates for males was $0.204/6 = 0.034$, but this fell to $0.176/9 = 0.019$ in the 1980's. If the 1973–1979 trend had continued during 1979–1988, other things held constant, the male college/high-school proportional relative wage would have increased by $9 \times (1/\sigma) \times (0.034 - 0.019) = 0.077$ less than the actual increase of 0.163. Similar results apply to the high-school/dropout and old/young differentials for both men and women, but not to the college/high-school differential for women or to the gender differential.

The shortcoming of the approach dealing with the deceleration of relative-supply changes is that it does not explain why the structure of relative demand is changing. The approach does, however, suggest that the 1970's would have been characterized by the sorts of wage-structure changes that prevailed in the 1980's if there had not been large increases in enrollment rates starting in the 1960's.

IV. Conclusions

We have attempted to evaluate the evidence concerning several alternative explanations of the dramatic wage-structure developments in the United States during the 1980's. Our analysis points strongly to the conclusion that the principal reason for the increases in wage differentials by educational attainment and the decrease in the gender differential is a combination of skilled-labor-biased technical change and changes in unmeasured labor quality. Interestingly, these sources of wage change applied to the 1970's as well as the 1980's, but they did not cause major changes in the wage structure in the 1970's because of the abnormally large increases in the relative supply of educated labor during that time. The extremely large relative-wage decrease of young workers with low educational attainment during the 1980's is more difficult to explain, because a large part of the source of this decrease did not apply in the 1970's.

It is interesting to speculate about what the results imply about the course of relative wages in the future. Given a continuation of the increase in the relative demand within industries for highly educated labor, wage differentials by education are likely to continue to increase unless there is a sharp rise in college attendance and completion rates. Such an increase does not appear to be likely in the near future (see Bishop and Shani Carter, 1990) in the absence of drastic changes in educational policy at all levels.

Appendix

The Determination of Competitive Wage Rates

Assume that output of each of J industries (Q_j) depends on employment of each of the I demographic groups (N_{ij}) according to the CES (constant elasticity of substitution) function

$$\text{(A1)} \quad Q_j = {}_{aj}\left[\sum_i \delta_{ij}(b_{ij}N_{ij})^{(\sigma-1)/\sigma}\right]^{\sigma/(\sigma-1)}$$

where b_{ij} is an index of the technological efficiency of group-i workers in industry j, a_j is a parameter representing the (neutral) technological efficiency of the industry and the effect of capital intensity, and σ is the elasticity of intrafactor substitution, which is assumed to be equal across industries. The relative demand for the output of industry j relative to some reference industry r is assumed to be

$$(A2) \qquad Q_j/Q_r = \theta_j P_j^{-\varepsilon} \qquad j \neq r$$

where P_j is the price of Q_j relative to Q_r, θ_j is an exogenous parameter reflecting consumer tastes and other factors (such as foreign competition) relative to good r, and ε is the absolute price elasticity of product demand for each industry. The marginal conditions for each industry are given by

$$(A3) \qquad P_j \partial Q_j/\partial N_{ij} = P_j a_j \delta_{ij} b_{ij}^{1-1/\sigma}(Q_j/N_{ij})^{1/\sigma} = W_{ic}.$$

Finally, it is assumed that the economy is at full employment, such that the effective (fixed) labor force of each group is allocated among the J industries, that is,

$$(A4) \qquad N_i = \sum_j N_{ij}.$$

The $2J-1+I\times(J-1)$ equations represented by (A1)–(A4) comprise a model in which the J Q_j's, I W_{ic}'s, $J-1$ P_j's, and $I\times J$ N_{ij}'s are determined as functions of the $J-1$ θ_j's, $I\times J$ b_{ij}'s, J a_j's, and I N_i's. The model is easily manipulated to obtain a few useful results. First, the share of total group-i employment in industry j is

$$(A5) \qquad \phi_{ij} = \delta_{ij}^{\sigma}(b_{ij}/b_i)^{\sigma-1}x_j/D_i$$

where b_i is the average value of the technological-efficiency parameter for group-i workers across industries and

$$(A6) \qquad D_i = \sum_j \delta_{ij}^{\sigma}(b_{ij}/b_i)^{\sigma-1}x_j$$

$$(A7) \qquad x_j = a_j^{\sigma-1}\theta_j^{\sigma/\varepsilon}Q_j^{1-\sigma/\varepsilon}.$$

Second, the ratio of the competitive wage for group-i workers to that of some other group s is

$$(A8) \qquad W_{ic}/W_{sc} = (b_i/b_s)^{1-1/\sigma}(D_i/D_s)^{1/\sigma}(N_i/N_s)^{-1/\sigma}$$

where D_i is an index of the effects of the θ_j's, a_j's, and Q_j's, and proportional changes in its values are referred to as a "product-demand-shift index." Holding constant the variables that affect W_{sc}, the total logarithmic derivative of (A8) is

$$(A9) \qquad d(\ln W_{ic}) = (1-1/\sigma)d(\ln b_i) + (1/\sigma)d[\ln(D_i/N_i)].$$

The third useful result from the model concerns the estimation in the demand-shift variable in (A9), $d(\ln D_i)$, which reflects changes in the θ_j's and a_j's. Total differentiation of (A6) yields the product-demand-shift index

$$(A10) \qquad d(\ln D_i) = \sum_j \phi_{ij} d(\ln x_j)$$

for $\sum_j \phi_{ij} d[\ln(b_{ij}/b_i)] = 0$. The $d(\ln x_j)$'s are not directly observed, but the total derivative of (A5) is

$$(A11) \qquad d(\ln \phi_{ij}) = (1-\phi_{ij})d(\ln x_j) - \sum_{k\neq j}\phi_{ik}d(\ln x_k) + (\sigma-1)\left[d(\ln(b_{ij}/b_i))\right]$$

which may be rewritten in matrix form as equation (A12), below.

$$(A12) \qquad \begin{bmatrix} d(\ln\phi_{11}) \\ d(\ln\phi_{12}) \\ \vdots \\ d(\ln\phi_{IJ}) \end{bmatrix} = \begin{bmatrix} 1-\phi_{11} & -\phi_{12} & \cdots & -\phi_{1J} \\ -\phi_{11} & 1-\phi_{12} & \cdots & -\phi_{1J} \\ \vdots & \vdots & & \vdots \\ -\phi_{I1} & -\phi_{I2} & \cdots & 1-\phi_{IJ} \end{bmatrix} \begin{bmatrix} d(\ln x_1) \\ d(\ln x_2) \\ \vdots \\ d(\ln x_J) \end{bmatrix} + \begin{bmatrix} d[\ln(b_{11}/b_1)] \\ d[\ln(b_{12}/b_1)] \\ \vdots \\ d[\ln(b_{IJ}/b_I)] \end{bmatrix}$$

In the absence of any information about the pattern of industry/group-specific technical change, the $d[\ln(b_{ij}/b_i)]$'s are treated as an error term, and the $d(\ln x_j)$'s may be estimated by ordinary least squares and then substituted back into (A10) to obtain estimates of the product-demand-shift index.

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