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Can Falling Supply Explain the Rising Return to College for Younger Men? A Cohort-Based Analysis

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Abstract

The college-high school wage gap for younger men has more than doubled since the mid-1970s. For older male workers, however, the gap is about the same today as it was 25 years ago. We show that the rise in relative returns to college for younger workers is attributable to permanent "cohort effects" for groups that entered the labor force after 1975. We argue that these cohort effects are caused by a shifts in the relative supply of college workers, associated with a dramatic trend break in educational attainment that occurred for men born after 1950. We estimate models in which the college wage gap for a particular age group depends on the relative supply of college-educated workers in the group and on underlying trends in the relative productivity of highly-educated workers. Trends in cohort-specific supplies of college-educated workers explain the relative rise in returns to college for younger men, and a substantial fraction of the overall increase in returns to college for all age groups. Controlling for cohort-specific supply effects, we find little evidence that relative demand shifts associated with the diffusion of computer technology, or changing unionization, have had much effect on the returns to college.

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One of the most remarkable trends in the U.S. labor market is the recent rise in education-related wage differentials.[1] Among men, for example, the gap in average hourly earnings between workers with a college degree and those with only a high school diploma rose from about 25 percent in the mid-1970s to 40 percent in 1998. A less-known fact is that virtually all of this rise is attributable to gains in the relative earnings of younger college-educated workers. Figure 1, for example, plots the college-high school wage gap for younger (ages 26-35) and older (ages 46-60) male workers. While the gap for young workers has more than doubled since 1975, the wage gap for older workers is about the same today as it was 20 years ago. As a consequence of these divergent trends, the age structure of the college wage premium is much different today than in the past. In the 1960s and 1970s the college premium rose with age -- as predicted by Mincer's (1974) "human capital earnings function".[2] Currently, however, the college-high school wage gap is roughly constant for men after age 30.

In this paper we explore a simple explanation for the trends in Figure 1. We argue that the shifting structure of the returns to college over the past 20 years is a reflection of permanent "cohort effects" for recent labor force entrants that explain both the rise in the average return to education and the twisting of the age profile of returns. Consistent with the standard human capital earnings function, we find no evidence of differential cohort effects among men born in the first half of the century. Starting with cohorts who began entering the labor market in the 1970s, however, successive generations of young men have had systematically higher returns to college throughout their labor market careers. The entry of these new cohorts has driven up the average college-high

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[1]See Katz and Autor (1999) for an exhaustive review.

[2]Mincer (1974) posited that the log of earnings depends on a linear education term and a concave quadratic function of labor market experience (age minus education minus 5). This formulation implies that the difference in log wages between college and high school workers of the same age rises linearly with age. Recent research (e.g. Murphy and Welch, 1990) has allowed more flexible functions of experience, for example, $\log w = \beta S + g(A\text{-}S\text{-}5)$, where S=education and A=age. As long as g is concave and increasing, the implied college high school wage gap is increasing with age.

school wage differential among all male workers, and also shifted the age structure of the college-high school wage gap.

We then go on to analyze alternative explanations for the emergence of cohort effects in the return to education. A key observation is that for most of this century, successive cohorts of men had rising college completion rates. This trend stopped -- and abruptly reversed -- with cohorts born in the early 1950s. The post-1950 break in the trend toward higher educational attainment precisely mirrors the emergence of cohort effects in the returns to education. Thus, a potential explanation for the rise in returns to college for cohorts born after 1950 is the failure of the relative supply of college educated workers in successive cohorts to keep pace with a steadily increasing demand for educated workers.[3]

We evaluate this cohort-specific relative supply explanation against an alternative hypothesis that focuses on the demand for computer skills. Some analysts (e.g. Krueger, 1993) have argued that part of the increase in returns to education over the 1980s is essentially an omitted variables bias, attributable to the rising demand for computer-literate workers and the positive correlation between higher education and computer skills. To the extent that younger cohorts of college-educated workers have higher levels of computer skills, this argument suggests that younger cohorts should exhibit a bigger rise in the return to college. We also consider a third institutional factor -- declining unionization -- which may have contributed to a decline in the relative wages of younger high school-educated men.

Our findings suggest that shifts in cohort-specific supplies of highly-educated workers, coupled with an hypothesis of steadily increasing demand for educated workers, can explain *all* of the relative rise in the returns to college for younger men over the past two decades. Once supply

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[3]This hypothesis -- although ignoring differences across cohorts -- is proposed by Katz and Murphy (1992) as an explanation for the rise in average returns to education.

factors are taken into account, we find no evidence that computer skills have had much effect on returns to college. Controlling for cohort-specific supplies of educated labor, shifts in unionization exert only a small effect on the relative return to education for younger cohorts -- in line with the magnitude suggested by earlier studies (e.g. DiNardo, Fortin, and Lemieux, 1996).

II. Returns to College: Age, Year, and Cohort Effects

Throughout most of this paper we focus on a particular measure of the "return" to college education -- the difference in mean log earnings between individuals *of the same age* who have a college degree and those with a high school diploma. The advantage of this measure is that it compares individuals who attended elementary and secondary schooling together, and were subject to the same influences on their decision as to whether to attend college (such as college deferments for the draft). A potential disadvantage is that it ignores any differences in labor market experience between people of the same age who have different levels of schooling. Under the traditional human capital earnings function, for example, one would expect the college-high school earnings gap for people of the same age to rise over the lifecycle. Since we account for age effects in our models we do not regard this as a problem, and in view of our focus on "cohorts" we believe it is most natural to compare college and high school earnings for men of the same age. As a check on the validity of our conclusions, however, we present some parallel results from an analysis based on "experience cohorts".

Table 1 presents age-specific estimates of the college-high school wage gap, derived from data for male workers age 26-60 in the 1976-1997 March Current Population Surveys (CPS). Each entry in the table represents the estimated college gap in log hourly earnings from a pooled set of CPS samples. For example, the 1979-81 entries are based on data from the 1980-82 March CPS files (which report annual earnings and hours for 1979-1981).[4] The wage gaps are estimated from regression models fit separately by cohort to samples of men with either a high school or college education.[5] In addition to a dummy for college education (whose coefficient is reported in Table 1), the models include controls for age and race, and indicators for the year from which the observation was drawn.

We stress that the entries in Table 1 are estimates of the wage premium for those with *exactly* a college degree relative to those with a high school diploma. As discussed in Appendix A, the wage gap between those with a college degree and those with post-graduate schooling expanded over the 1980s and early 1990s, and was substantially affected by the change in wording of the CPS education question in the early 1990s. Thus, the wage gap between high school graduates and workers with a college degree or more does not precisely parallel the gaps in Table 1.

The entries in Table 1 provide a variety of information on the evolution of the college-high school wage gap. Comparisons down a column of the table show the changing college premium for a specific age group (as in Figure 1). Among 26-30 year olds, for example, the college wage gap roughly tripled from 1975 to 1995. For slightly older men (ages 30-45) the gap also rose, although less than for the youngest age group. Finally, for men over age 45, the gap dipped between 1975 and 1980, then returned to its earlier level, with little net change from 1975 to 1995.

Comparisons across the rows of Table 1 reveal the age profile of the college-high school wage gap at a point in time. These profiles are graphed in Figure 2. In the mid-1970s the wage gap was an increasing and slightly concave function of age. Between 1975 and 1980 the entire profile shifted down, with the exception of the youngest age group, whose gap remained constant. By the mid-

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[4]Appendix A presents more information on the CPS samples and on our procedures to eliminate outliers, etc.

[5]We pool the samples from adjacent CPS samples by cohort, rather than age. Thus, we pool 25-29 year olds in the 1980 CPS with 26-30 year olds in the 1981 CPS and 27-31 year olds in the 1982 CPS.

1980s the gaps for older workers were back to their levels in the mid-1970s, but the gaps for the *two* youngest age groups were much higher. Moving to 1989-91, the gaps for the *three* youngest age groups were substantially higher than those in the mid-1970s, while those for the older cohorts were not too different. Finally, in 1994-96, the gaps for the *four* youngest age groups were well above the levels of the mid-1970s, but the gaps for older age groups were still comparable to those 20 years earlier.

The shifting age profiles in Figure 2 suggest that there are at least two separate forces underlying the evolution of the college wage premium over time. First, overall returns to college dipped from 1975 to 1980, then returned to their previous level by 1985. Second, starting with men who were age 26-30 in 1980 (i.e. those born from 1950 to 1954), later cohorts have had higher returns to college than cohorts born before 1950. Cohort-specific college-high school age profiles can be read along the diagonals of Table 1. These profiles suggest a fairly consistent 10 percentage-point rise in the gap from age 26-30 to age 31-35, another 8-9 percentage point rise from age 31-35 to age 36-40, roughly 5-7 percentage point gains from age 36-40 to 41-45 and from 41-45 to 46-50, and 3-5 percentage point gains per 5-year interval thereafter. Relative to the profile for any given cohort, however, the cross-sectional profiles are distorted by cohort effects, which have pushed up the relative return to college for those born after 1950.

Table 2 presents a series of models, fit to the college-high school wage gaps presented in Table 1, that provide formal tests of this story. These models decompose the observed wage gaps for different age groups in different years into an additive set of age, year, and cohort effects. Formally, let r_{at} represent the observed mean college-high school wage gap for individuals of age group a in year t. Note that people who are age a in year t were born in t-a: hence t-a is a convenient index of their cohort. The models in Table 2 assume that

$$(1) \qquad r_{at} = b_a + c_{t-a} + d_t + e_{at}$$

where b_a are age effects (for 5-year age bands), c_{t-a} are cohort effects (for 5-year birth cohorts), d_t are year effects (for time periods 5 years apart), and e_{at} represents a combination of sampling error (arising from the estimation of r_{at}) and specification error. Since the sampling variances of the estimated r's are known, it is easy to construct goodness-of-fit tests for the null hypothesis of no specification error.[6]

The model in column 1 of Table 2 includes a full set of unrestricted cohort, year, and age effects. Since age, year, and cohort are linearly dependent, the individual effects are not identified in this specification, although it provides a benchmark for the entire class of models described by equation (1). As shown in the second row of the table, the unrestricted model provides a statistically acceptable fit.

Columns 2 and 3 report estimated models that exclude any cohort effects. The model in column 2 is fit to the entire set of returns in Table 1 while the model in column 3 is fit to data for the 7 oldest cohorts only (i.e. for men born before 1950). The contrast between these two models is informative. The "no cohort effects" assumption is clearly rejected by the overall data, but is acceptable for the pre-1950 cohorts. The estimated year effects from these two specification are also quite different. In column 2, the year effects show a decline in returns from 1975 to 1980, then steadily rising returns from 1980 to 1995. In column 3, however, the year effects show a decline in returns from 1975 to 1980 that is reversed by 1985, with relative stability thereafter. Focusing only on older cohorts, one would conclude that returns to college have not increased from the mid-1970s to the mid-1990s.

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[6]Specifically, let r represent the vector of estimated gaps, and let μ represent the true gaps. Given our estimation procedures, r-μ is normally distributed with mean 0 and a diagonal covariance matrix Σ. Let S represent a consistent estimate of Σ. Under the null that $\mu=f(\pi)$, where π is a vector of parameters (e.g. cohort, age, and year effects), $(r-f(\hat{\pi}))'S^{-1}(r-f(\hat{\pi}))$ is asymptotically distributed as chi-square with degrees of freedom equal to the number of elements of r minus the number of linearly independent columns of the matrix of derivatives of $f(\pi)$. This quadratic form is simply the sum of squared residuals from a weighted regression, where the weights are the inverse sampling variances of the gaps.

Finally, columns 4 and 5 report models that include a restricted set of cohort effects, as well as unrestricted year and age effects. The specification in column 4 assumes that the four oldest age cohorts (born 1915-1934) have the same cohort effect, while the specification in column 5 assumes that the seven oldest cohorts (born 1915-1949) have the same cohort effect. Both specifications provide acceptable fits, and have similar interpretations. Relative to the model with no cohort effects (column 2), models that include permanent cohort effects for younger cohorts imply that most of the rise in average returns to college from 1975 to 1995 is attributable to inter-cohort differences. Indeed, the specification in column 5 shows essentially negligible year effects for 1985, 1990, or 1995 relative to 1975. Virtually all of the rise in the average college-high school wage gap from 1975 to 1995 is attributed to permanently higher returns to college for cohorts who entered the labor market in the 1970s. While not reported in the table, a more parsimonious model that includes age effects, a single year effect for 1980, and cohort effects for cohorts born after 1950 provides a statistically acceptable description of the data.[7]

One potential criticism of the models in Table 2 is the assumption that the cohort, age, and year effects are additive. Any aggregate shocks (captured by the year effects) must uniformly raise or lower the college wage premiums for all age groups. A number of authors (notably Freeman and Katz, 1995) have argued against this assumption, suggesting instead that aggregate shocks tend to have a bigger effect on younger workers -- at least in the short run.[8] One way to test this idea is to allow for age-specific "loading factors" that translate a given aggregate shock into differential effects on different age groups. Formally, equation (1) is replaced by

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[7]The goodness of fit statistic for this model is 18.07 with 23 degrees of freedom. The estimated cohort effects are almost the same as the ones reported in column 5 of Table 3.

[8]Freeman and Katz (1995, page 5) call this the "active labor market hypothesis". They argue that older workers are less likely to renegotiate their wages in any time interval: thus, it may take a long time for wage structure of older workers to adjust to a new equilibrium level.

(2) $r_{at} = b_a + c_{t\text{-}a} + h_a \cdot d_t + e_{at}$,

where h_a is the loading factor on the aggregate year effect for age group a. Normalizing h_a=1 for older age groups, Freeman and Katz's hypothesis implies that $h_a > 1$ for younger groups.

Table 3 presents estimation results for a series of variants of this model. All of the specifications include age and year effects, as well as cohort effects for groups born after 1950. In columns 1 and 2, the year effects are allowed to have a different loading factor for the very youngest group (ages 26-30) relative to all older groups; in columns 3 and 4 there are separate loading factors for 26-30 year olds and 31-35 year olds. We report models with a full set of year effects (columns 1 and 3) as well as models that only include a 1980 year effect (columns 2 and 4). The four specifications lead to similar conclusions. Although younger workers are slightly more responsive to aggregate shocks, the difference is not statistically significant.[9] Moreover, the patterns of the estimated year and cohort effects are hardly affected by allowing for a greater responsiveness among younger workers. We continue to find that the rise in average returns to college between the mid 1970s and the mid 1990s is largely attributable to permanently higher returns (i.e. cohort effects) for men born after 1950.

We have investigated the robustness of the findings in Tables 1 and 2 to two other specification choices. One is the use of post-1975 data, rather than a longer sample period. This choice was dictated by our preference for using hourly earnings as a measure of relative pay, and the fact that data on weeks of work and hours per week were only collected in categorical form prior to the March 1976 CPS. To evaluate the effect of extending our sample period, we constructed a consistent series of college-high school wage gaps in average *weekly* earnings using data from the

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[9]One piece of evidence against the "active labor market hypothesis" is the fact that between 1975 and 1980 (a period of declining college wage premiums for all age groups) the college premium fell least for the youngest age group. If younger workers are more sensitive to aggregate shifts, one might have expected a bigger decline for the youngest group.

1960 Census, the March 1970-72 CPS, and the post-1976 Current Population Surveys.[10] These gaps are presented in Appendix Table B-1. In general, the estimated wage gaps for the 1975-1995 period are similar using hourly or weekly earnings data, although the gaps in *weekly* earnings for older workers rose somewhat between 1975 and 1995 whereas the gaps in *hourly* earnings were constant.[11] The pre-1975 data show that returns to college were about 3 percentage points higher in 1970 than 1975, and about 2 points higher in 1959 than 1975. Most importantly for our purposes, the age profile of the college premium was very stable from 1959 to 1975: in fact a model with only age and year effects fits the returns data for 1959, 1970, and 1975 very well.[12] Unlike the later period, there is no indication of any "twisting" in the age profile of returns prior to 1980.

Estimation results for models based on equation (1) over the 1959-95 period are presented in Appendix Table B-2, and are very similar to the results in Table 2. In particular, a model with age, year, and cohort effects adequately describes the observed wage gaps, whereas a model with only year and age effects is soundly rejected. On the other hand, a model with year and age effects provides a reasonably good fit to data for cohorts born before 1950. As in column 5 of Table 2, a model that includes year and time effects, and cohort effects for men born after 1950, also provides a statistically acceptable fit to the observed data. In this model, the estimated year effect for 1995 (relative to 1975) is about one-third as large as the corresponding coefficient from a model with no cohort effects. Thus, cohort effects "explain" about 60 percent of the rise in returns to college (based on weekly wages) from 1975 to 1995. The remaining increase in returns (which is attributable to a relative rise in the hours of college workers) is concentrated in the 1980-85 period. Once cohort

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[10]We used the same categorical codes for weeks in all years, and assigned midpoints to each category.

[11]This divergence reflects a rise in the relative hours of college workers over the 1980s.

[12]This model has 8 parameters for 21 observed gaps: the chi-squared statistic is 9.88 with 12 degrees of freedom (p-value 0.62). The estimated year effects (relative to 1975) are 0.018 (0.008) for 1959 and 0.033 (0.010) for 1970.

effects are taken into account, the college premium was fairly stable from 1985 to 1995, whether earnings are measured hourly or weekly.

A second specification issue is the use of age, rather than potential labor market experience, to define cohorts. To investigate the effect of this choice, we re-estimated the college-high school wage gaps by experience cohort, assuming that college-educated men enter the labor market on average about 5 years later than men with only a high school degree.[13] The resulting wage gaps are presented in Appendix Table C-1. In the mid-1970s, the college wage premium was relatively constant across experience cohorts, as is assumed in the specification of a conventional human capital earnings function. Over the 1980s, however, the premium for newly entering workers rose quickly, while the premium for older cohorts remained relatively constant. By the mid-1990s, the college premium for men with 1-10 years of experience was about 53 percent, while the premium for men with 26-35 years of experience was about 38 percent (about what it was in 1975).

Models like equation (1), fit by experience group and year, are presented in Appendix Table C-2. The estimation results closely parallel the findings in Table 2. In particular, a model with cohort, experience, and time effects fits the data well, whereas a model with only experience and year effects is easily rejected. As with age-based cohorts, the source of the poor fit is the rising return to college among younger experience cohorts. A model fit to the oldest experience cohorts provides an acceptable fit, as does a model that includes cohort effects for the four most recent experience cohorts (those who "entered" the labor market after the mid-1970s). Finally, the addition of cohort effects for the most recent entrant groups substantially reduces the measured increase in overall returns to college: about two-thirds of the rise in average returns is "explained" by cohort effects of the post-1975 entrants. Moreover, all of the unexplained rise that remains is concentrated in the

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[13]Thus, college educated men age 26-30 are in the same experience cohort as high school educated men age 21-25. The assumption of a 5-year gap was made mainly for convenience, to correspond to the 5-year observation intervals we use throughout this paper.

1980-85 period: between 1985 and 1995 there is no evidence of a further rise in aggregate returns to education.[14]

The results of these two specification checks give us reasonable confidence in two key conclusions: (1) recent cohorts of workers have permanently higher returns to college than earlier cohorts; (2) cohort effects are an important part of the explanation for the apparent rise in overall returns to college that occurred between the mid-1970s and the mid-1990s.

III. Cohort-Specific Supply and the College Wage Premium

a. Relative Supply of College Education by Cohort

Having found that returns to college contain important inter-cohort differences, it is natural to ask what causes these differences. In this paper we focus on a very simple answer: cohort-specific supplies of college-educated labor. Before presenting a formal model that incorporates this idea, it may be helpful to look at some descriptive evidence on inter-cohort differences in the relative supply of college-educated labor. Table 4 presents the college completion rates of different age groups in different years, along with information on the fraction of college completers with some post-graduate education.[15] Inspection of the data in Panel A of this table reveals an interesting fact: since the mid-1970s there has been a noticeable drop in the college completion rate among cohorts who are just entering the labor market. Similarly, the data in panel B show a decline in the fraction of college graduates with any post graduate education, beginning about the same time.

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[14]Murphy and Welch (1993) present an analysis of the evolution of the college premium from 1968 to 1988 in which they control for cohort effects in experience (see their Table 3.1, page 108). They conclude that cohort effects play a minor role in the increase in the college wage premium between 1978 and 1988. This result is consistent with our finding that most of the increase in the college premium in the first half of the 1980s is due to unexplained time effects as opposed to cohort effects. Cohort effects become much more important after the end of the sample period considered by Murphy and Welch.

[15]The entries in Table 4 pertain to all men, not just workers.

Precise inter-cohort comparisons of college completion rates are somewhat difficult using the raw data in Table 4 since college completion tends to rise with age, and not all cohorts are observed at each age. To facilitate cohort comparisons, we fit a model with cohort and age effects to the data in panel A of Table 4 (augmented with earlier data for 1970 and 1960).[16] The resulting cohort effects (standardized to age 26-30) are plotted in Figure 3. This graph illustrates a dramatic "trend break" in educational attainment that occurred with cohorts born after 1950. Prior to 1950, each successive cohort had a higher college completion rate. After the peak 1945-49 cohort, however, educational attainment actually fell slightly, and subsequently stagnated.[17] [18]

A striking feature of Figure 3 is that the timing of the break in the inter-cohort trend in college completion rates coincides with the emergence of rising cohort effects in the college wage premium. Among pre-1950 cohorts, college completion rates were rising and the wage premium was stable. After the peak 1945-49 cohort, college completion rates stagnated and the wage premium began to rise. The coincidence of timing suggests a possible causal link between cohort-specific relative supplies of college educated labor and cohort-specific returns to college. We turn to a formal model that elucidates such a link.

b. Cohort-Specific Supply and Demand

Although existing research on the sources of the rising return to higher education has emphasized the role of supply variation, most studies have focused on aggregate returns rather than

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[16]We used March 1969-71 CPS files, and the 1960 Census, to obtain this additional data. The earlier data yield more observations, at a younger age, for the older cohorts.

[17]Freeman and Needels (1993) also notice this fall in the educational attainment of young men born in the 1950s (age 25-34 in the mid-1980s) relative to slightly older cohorts.

[18]Women born in the 1920s-1940's had lower levels of college completion than men, and slightly faster inter-cohort growth rates. Women's college rates continued to rise among the post-1950 cohorts, and surpassed men's by the 1965-69 cohort.

differences by age or cohort (e.g. Freeman, 1976; Freeman and Needels, 1993; Katz and Murphy, 1992). In order to incorporate these differences, we begin with an aggregate production function that depends on age-group-specific inputs of low-education and high-education labor.[19] Based on the evidence in Table 1 that college wage premiums for different age groups follow different paths over time, we assume that production is *separable* in age-specific labor inputs:

$$(3) \qquad y_t = f(\, g^1(C_{1t},H_{1t};\theta_{1t}),\, g^2(C_{2t},H_{2t};\theta_{2t}),\, \ldots\, g^A(C_{At},H_{At};\theta_{At})\,)$$

where $g^a(C_{at},H_{at};\theta_{at})$ is a sub-production function for inputs of age group a in year t (a=1,2,...A), C_{at} represents inputs of "college-equivalent" labor of age a in period t, H_{at} represents inputs of "high school-equivalent" labor of age a in period t, and θ_{at} is a relative productivity shock that affects workers of age a in year t. Separability of the production function implies that the relative demand for high and low education workers in any age group depends only the relative wages of workers in that age group. We view this as an extreme assumption that is unlikely to be literally true: nevertheless, the data seem consistent with it. We further assume that the sub-production functions are CES:

$$(4) \qquad g^a(C_{at},H_{at};\theta_{at}) = [\, (\theta_{at}C_{at})^\rho + H_{at}{}^\rho \,]^{1/\rho},$$

where $-\infty < \rho \leq 1$ is the same for all age groups. Efficient utilization of labor inputs requires that the relative employment of high and low education workers of age a varies with their relative wage:

$$(5) \qquad C_{at}/H_{at} = \theta_{at}{}^{\sigma-1}\,(w_{cat}/w_{hat})^{-\sigma},$$

where $\sigma = 1/(1-\rho) \geq 0$ is the elasticity of substitution between high and low education workers, and w_{cat} and w_{hat} are the wage rates of high and low education workers of age a in period t.

If relative employment ratios are taken as exogenous, equation (5) leads to a model for the college-high school wage gap $r_{at} \equiv \log(w_{cat}/w_{hat})$ of the form:

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[19]We follow Johnson (1997), Katz and Murphy (1992), and many other researchers in assuming that within age groups there are two types of labor.

(6) $r_{at} = (\sigma\text{-}1)/\sigma \log \theta_{at} - 1/\sigma \log (C_{at}/H_{at}) + e_{at}$,

where e_{at} reflects sampling variation in the measured gap and any other sources of variation in age-specific wage premiums. The empirical implications of this equation depend on what assumptions are made about the structure of the relative productivity shock affecting workers of age a in year t. A variety of arguments suggest that the relative productivity of college equivalent workers varies with age: we therefore include age effects in all our specifications. We consider two alternative assumptions about relative productivity over time. One alternative is to simply include year effects. A more restrictive alternative is to include a linear trend.[20] Combining these assumptions with equation (6) leads to a model for the observed college wage premiums that includes a relative supply term, age effects, and some combination of year effects and/or a linear trend ($d_L t$):

(7) $r_{at} = b_a + d_t + d_L t - 1/\sigma \log (C_{at}/H_{at}) + e_{at}$.

Note that the trend coefficient d_L is not identifiable if a full set of year effects are included.

It is worth noting that when unrestricted age effects are included in equation (7), an aggregate linear trend (which raises the relative productivity of college workers in all age groups) is statistically indistinguishable from a *cohort-specific trend* which raises the relative productivity of college workers in successive cohorts. While the two alternatives provide identical fits, and yield identical estimates of the relative supply effect, they lead to a different set of relative age effects, and have rather different economic implications. We return to this issue in the final section of the paper.

In general the model described by equation (7) is not strictly nested in the simple age-year-cohort effects framework of equation (1), since C_{at}/H_{at} can vary arbitrarily by age and cohort. If the relative supply of college-equivalents is fixed in a given cohort, or only varies within cohorts because of age effects, however, then (7) is nested in (1). For example, suppose that the log supply ratio

[20] Katz and Murphy (1992) use a linear trend to proxy increasing relative demand for highly educated workers at the aggregate level.

among workers who are age a in year t consists of a cohort effect for that group, λ_{t-a}, and an age effect ϕ_a that is constant across cohorts:

$$(8) \qquad \log\left(C_{at}/H_{at}\right) = \lambda_{t-a} + \phi_a \ .$$

In this case equation (7) can be re-written as a model with age effects, year effects, and cohort effects (analogous to equation (1)):

$$r_{at} = b_a - (1/\sigma)\cdot\phi_a + d_t + \eta_{t-a} + e_{at} \ ,$$

where the cohort effects depend on the cohort effects in the relative supply of college equivalents:

$$(9) \qquad \eta_{t-a} = - \ 1/\sigma \ \lambda_{t-a} \ .$$

This discussion suggests that there at least three approaches to evaluating the cohort relative supply hypothesis as an explanation for the shifting age pattern of the college wage premiums in Table 1. One is to fit a model like equation (7) directly to the wage gaps, using the observed relative supply ratios for each age group in each year as exogenous determinants of the wage gap. A minor drawback to this approach is that (7) is not nested in a general age-year-cohort model. A second approach is to estimate (7) using a standardized relative supply index for each cohort -- e.g. the relative fraction of college graduates at ages 26-30. Since each cohort is assigned a single supply variable under this scheme, the resulting model is strictly nested in a general age-year-cohort model, facilitating comparisons to the models in Table 2. A third approach is to estimate the cohort effects in the college wage premiums from equation (1), and the cohort effects in the relative supply ratios from equation (8), and then estimate equation (9) in a second step using the two sets of estimated cohort effects. Standard arguments can be used to show that the third approach is equivalent to the second, provided that the second-step estimation is appropriately weighted. As a practical matter the first approach yields very similar results too, since cohort-specific relative supplies of college education are well-described by a model like equation (8). For simplicity, we therefore present results based on the direct estimation of (7).[21]

Table 5 presents estimates of several specifications of this model. For reference, the first column of the table reproduces the model from column 2 of Table 2, which includes only year and age effects. Column 2 adds the cohort-specific relative supply variable log (C_{at}/H_{at}).[22] The estimated coefficient shows a strong negative effect of relative supply on the college wage premium (with a t-ratio over 10), and implies a relatively high elasticity of substitution between high school and college labor ($1/\sigma \approx 0.2 \Rightarrow \sigma \approx 5$). The addition of the supply variable substantially improves the fit of the model: relative to column 1, the chi-squared statistic drops by almost 90 percent.[23]

The model in column 3 adds a linear trend and removes the year effects. Although the estimated relative supply coefficient is hardly affected, the fit of the model is substantially worse. Inspection of the year effects in column 2 hints at the source of the difficulty: controlling for relative supply and age effects, aggregate returns to college dipped between 1975 and 1980, then rose over the 1980s and 1990s. A simple way to allow for this phenomenon is to retain the linear trend term and add a single year effect for 1980. This specification -- shown in column 4 -- fits about as well as the unrestricted year effects model, and provides an interesting perspective on the aggregate trend in observed college wage premiums. According to the estimates, the relative demand for college-

[21]Results from the other approaches are available on request.

[22]Following other researchers (e.g. Johnson, 1997) we measure the number of college equivalents in an age group as the number of those with exactly 16 years of education, plus 1.2 times the number with post-graduate education, plus 0.5 times the number with 13-15 years of education. We measure the number of high school equivalents as the number with 12 years of education, plus 0.5 times the number with 13-15 years of education, plus 0.83 times the number with less than a high school degree. We experimented with using the relative supply of college equivalents in the population as a whole versus the relative supply among workers only. The estimates are very similar for either variable, suggesting that relative supply can be taken as exogenous. Throughout the paper, we use a relative supply index that sums workers and nonworkers.

[23]The goodness of fit statistics reported in Table 5 (and subsequent tables in this paper) make no allowance for the fact that the relative supply variable is estimated, rather than known. This adjustment will lower the fit statistic, and raise the associated p-values.

educated labor has risen fairly steadily at about 1 percentage point per year, apart from a 6 percentage point drop in the late 1970s that was reversed in the early 1980s.

Columns 5 and 6 present two expanded models that provide specification tests against the simple model in column 4. In column 5 we include a year effect for 1995. This addition has almost no effect on the estimates of the other parameters. Moreover, the estimated coefficient of the 1995 dummy is very close to zero, confirming that apart from the dip in aggregate returns from 1975 to 1980, the trend in relative productivity of college workers has been fairly steady. In column 6 we include dummies for the two youngest cohorts in our analysis. Again, the addition of these variables has very little effect on the other estimates, and the estimated cohort effects are themselves close to zero, providing no evidence against the simpler model.

We have also investigated the robustness of the results when experience, rather than age, is used to define cohorts. The estimates obtained using the wage gaps by experience cohorts (Appendix Table C-1) are similar to those reported in Table 5. For example, in a model comparable to column 4 of Table 5 (experience dummies, trend, relative supply, and 1980 dummy as explanatory variables), the estimated effect of relative supply is -0.208 (standard error of 0.064) and the estimated trend is 0.012 (standard error of 0.002). Unlike the models where age is used to define cohorts, however, the fit of models where experience is used to define cohorts is not statistically acceptable at standard confidence levels.[24] This suggests that the cohort-specific supply and demand model presented above is consistent with the data when age is used to define cohorts (the fit of the model is acceptable, see Table 5), but is inconsistent with the data when experience is used to define cohorts.[25] This provides

[24]For example, the chi-squared statistic of the experience model corresponding to the specification reported in column 4 of Table 5 is 198.1 (p-value of 0.00).

[25]Another specification test of the model consists of adding the relative supply of workers who are 5 years older (or 5 years younger) to the regression model. Under the assumption of strict separability across age groups (equation 3), the relative supply of other cohorts (older or younger) should have no effect on wages once the cohort's own relative supply has been controlled for. While we cannot reject this

an additional reason to focus our analysis on cohorts defined on the basis of age instead of experience.

IV. Other Sources of Inter-cohort Differences in the College Wage Premium

Although changes in cohort-specific relative supplies of college-educated labor seem to go a long way toward explaining the presence of cohort effects in the college wage premium, it is worth considering other possible explanations for the data in Table 1. In this section we focus on two alternatives: the rising demand for computer skills; and changes in unionization. We also briefly consider the role of immigration.

a. Computer Skills

It is widely believed that the computer revolution has increased the demand for workers who know how to use computers, and for those whose talents are complementary with computer technologies (see e.g. Krueger, 1993; Bound and Johnson, 1992; Katz and Autor, 1999). This "skill-biased" demand shift is hypothesized to have raised the relative wages of better-educated workers throughout the economy. Krueger (1993) considers a model in which individuals who use a computer on the job receive a positive wage premium. Since college-educated workers are more likely to use computers on the job, the computer premium is effectively an omitted variable in a standard earnings model that leads to an upward bias in the measured college-high school wage gap.[26] A slightly different mechanism is posited by Autor, Katz, and Krueger (1997), who interpret

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hypothesis when age is used to define cohorts, the same hypothesis is soundly rejected when experience is used to define cohorts. The results from these tests are available upon request.

[26]Krueger's assumption that the wage premium for workers who use a computer is a "return" to a specific skill is criticized by DiNardo and Pischke (1997), who show that workers who use other tools at work also receive higher wages.

computer usage rates of different skill groups as indicators of their complementarity with new technologies. They argue that the relative computer usage rates of different education groups are a good proxy for the relative demand shifts induced by the diffusion of computer-related technologies through the economy, and show that groups that use computers more have had rising relative wages.

To the extent that computer skills are concentrated among younger cohorts of highly-educated workers, either of these mechanisms might be expected to influence the age structure of the college-high school wage gap and lead to higher returns to college for recent cohorts. Figure 4 shows the differentials in computer use rates between college and high school graduates by age in 1984, 1989, and 1993. These data are derived from supplements to the October Current Population Surveys, which asked workers whether they use a computer on the job (for work-related tasks). The graph confirms that the college-high school computer use gap is higher among younger workers. Interestingly, however, the gaps for different age groups have risen more or less uniformly since 1984. Thus, in contrast to Figure 2, there is no indication of a "twisting" age profile of the computer use gap after 1984.

Prior to 1984 there are no comparable data on computer use rates, so it is difficult to judge whether the age profile of the college-high school gap shifted during the late 1970s and early 1980s. One possible assumption is that computer use was negligible before 1981, the year the IBM-PC was introduced to the market. Under this (admittedly extreme) assumption, the college-high school gap was 0 for all age groups in 1975-76 and 1979-81. This would imply that the gap rose more for young workers in the early 1980s, potentially explaining the relative rise in the college wage premium for younger age groups during this period.

Table 6 presents a series of regression models that attempt to test this hypothesis more directly. These models express the college-high school wage premium as a function of age and year effects, the relative supply of college-equivalents, and the college-high school gap in computer use.

The specification can be interpreted in two ways: either as an aggregated version of the earnings model fit by Krueger (1993); or along the lines of the analysis in Autor, Katz and Krueger (1997). In the former case, the coefficient on the computer gap variable is an estimate of the individual wage premium earned by workers who use a computer.[27] In the latter case, the coefficient measures the relative demand shock affecting an age/education group with a particular computer use rate.[28] In the absence of other data, the computer gaps in 1975-76 and 1979-81 are assumed to be zero for all age groups.

For reference, column 1 includes the "base" specification from column 4 of Table 5, which excludes any computer use variables. Column 2 presents a model that drops the relative supply and trend variables, and includes only age and year effects and the college-high school computer gap. Taken at face value, the large positive computer gap coefficient in this specification suggests an important role for technology in explaining the evolution of the college wage premium. However, once relative supply is added to the model (column 3) the computer gap coefficient falls to about 10 percent. This estimate is roughly comparable to Krueger's (1993) estimate of the individual wage premium for using a computer at work, although it is relatively imprecise (with a t-ratio under 1.0).

A potential limitation of these results is the assumption that the college-high school gap in computer usage was zero before 1982. As a check, we re-estimated the models using only data for

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[27]To be concrete, suppose that $\log w_{iat} = S_{iat}r_{at} + C_{iat}p + e_{iat}$, where w_{iat} is the wage of individual i in age group a and year t, S_{iat} is a dummy equal to 1 if i has a college degree (the sample includes only high school graduates or college degree holders), r_{at} is the "true" college wage premium for age group a in year t, C_{iat} is a dummy equal to 1 if i uses a computer at work, and p is the computer use premium. This model implies that the estimated college wage premium in year t for age group a is $r_{at} + \Delta C_{at}p$, where ΔC_{at} is the college-high school computer use gap for workers in age group and year t.

[28]Specifically, suppose that workers in age group a with education level j face a relative demand shock D_{ajt} in period t. The college high school wage gap for age group a in year t depends on $D_{act} - D_{aht}$. Assume that $D_{ajt} = b_a + d_t + gC_{ajt}$, where C_{ajt} is the fraction of individuals in the group who use a computer on the job, b_a are age effects and d_t is a year effect. This model implies that the college wage premium in year t for age group a depends on age and year effects, other supply-side factors, and $g\Delta C_{at}$, where ΔC_{at} is the college-high school computer use gap.

1984-86, 1989-91 and 1994-96. The results are presented in Table 7. The base specification (column 1) yields estimates of the relative supply and trend effects that are virtually identical to those in the larger sample. The computer gap coefficients are less stable across samples, however, and are actually negative in the restricted sample.[29] There is little evidence of a systematic linkage between the computer use patterns of different age/education groups and the evolution of age-specific college wage premiums over the 1980s and early 1990s.

b. Changing Patterns of Unionization

In addition to supply and demand factors, a third potential explanation for the changing relative wage structure is the effect of labor market institutions (see Fortin and Lemieux, 1997 for a review). Historically, trade unions exerted an important influence on the wage structure of adult male workers.[30] Over the past two decades, union coverage rates -- especially for younger, less well-educated men -- have fallen sharply (see e.g. Card, 1998). Since union coverage is typically associated with a 10-15 percent wage premium, recent shifts in relative unionization may have raised the college-high school wage gap among younger workers. Figure 5 shows the differences in union coverage rates between high school and college workers by age group in 1975, 1980, 1985, 1990, and 1995.[31] In the mid-1970s high school workers were about 30 percent more likely to be covered by union contracts than college workers of the same age. This pattern shifted in the early 1980s, as the union rates of young high school graduates dropped and unions spread to the public sector. Since

[29]We also tried a model that included the computer use gap with different coefficients for the three periods 1984-86, 1989-91, and 1994-96. The results were very similar to those in Tables 6 and 7.

[30]Another potentially important labor market institution -- the minimum wage -- has relatively little impact on male workers over age 25 with at least 12 years of education.

[31]The 1975 data are taken from pooled May 1974, 1975, and 1976 CPS files. The 1979 data are taken from the May 1979 CPS files. The later data are from the merged outgoing rotation group files of the monthly CPS.

1985 the relative unionization rate of high school workers has continued to fall, but more uniformly across age groups.

A simple way to incorporate the effect of changing unionization on the college-high school relative wage gap is to include the difference in union coverage between college and high school workers as an additional explanatory variable in equation (7). Assuming that unions raise wages by a constant percentage (with no "threat" or "spillover" effects) the coefficient of the union gap will equal to union wage premium.[32] Estimation results for this exercise are presented in columns 4-6 of Tables 6 and 7.

When the union coverage gap is entered without controlling for the relative supply of college-equivalent workers (as in column 4 of Table 6 or 7), the estimated coefficient is large and positive. The estimates suggest that each percentage point reduction in the relative union coverage of high school workers raises the college wage premium by 0.8 to 1 percent. Once the relative supply variable is added to the specifications, however, the union coefficient drops substantially. In Table 6 the coefficient is in the 15-20 percent range, consistent with conventional estimates of the union wage effect. In Table 7, which uses only 1985-95 data, the coefficient is negative, but very imprecise. An examination of the age profiles in Figure 5 suggests an explanation for the weaker findings in the post-1985 subsample. From 1985 to 1995 the age profiles of the union gap shift down more-or-less in parallel, leaving little independent variation in the union gap once age effects and trends are taken into consideration.

Focusing on the results for the full sample, the estimates in columns 5 or 6 of Table 6 imply

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[32]Using the notation of footnote 23, let $\log w_{iat} = S_{iat}r_{at} + U_{iat}m + e_{iat}$, where U_{iat} is a dummy equal to 1 if individual i in age group a and year t has his wages set by a union, and m is the union wage premium. This model implies that the estimated college wage premium in year t for age group a is $r_{at} + \Delta U_{at}m$, where ΔU_{at} is the difference in coverage between college and high school workers in age group and year t. If the union wage premium varies with the extent of union coverage (because of threat or spillover effects, for example) then the union wage effect recovered from an aggregate specification may be higher or lower than the relative union wage effect estimated at the micro level.

that the relative decline of unions among younger high school men has contributed modestly to the rise in the college wage premium for these workers. For example, between 1975 and 1995 the union gap for 26-30 year olds fell by about 0.25. Assuming a coefficient of 0.16, the implied effect on the college-high school wage gap is 0.04 -- or about one-fifth of the actual rise in the wage gap for this age group. The implied effects on 31-35 year olds are comparable. For older workers, the declines in the union gap are smaller (more like 10 percentage points) suggesting very small effects on the college wage gap.

The final right-hand column of Tables 6 and 7 includes all four potential factors explaining the variation in the college wage premium: the relative supply of college-equivalent workers; a linear trend; the gap in computer usage rates between college and high school workers; and the gap in union coverage rates. In addition, the models include unrestricted age effects and (in Table 6) a single dummy for 1980. In both the overall sample and the post-1985 subsample, this model provides an acceptable fit to the observed college wage premiums. In neither case, however, is there any significant improvement over a model that ignores computer use and unionization.

We have also examined a fourth potential explanation for the rise in the college wage premium among recent cohorts: increases in the relative fraction of immigrants in recent cohorts. Over the past 30 years U.S. immigration rates have risen sharply at the same time as the source-country composition of immigrant inflows has shifted toward Mexico, Central America, and Asia (see Card, DiNardo and Estes, 1998). These trends have been associated with a rise in the fraction of immigrants, particularly among the less-skilled segment of the workforce. One might conjecture that the relative fraction of immigrants with a college versus high school education has shifted between recent and older cohorts, leading to a shift in the measured college wage premium.[33]

[33]A more subtle possibility is that the source country composition of immigrant groups with 12 and 16 years of education has shifted. Since the immigrant-native wage gap varies by source country (as well as other factors), such shifts could differentially affect the measured college wage premiums for different age

Our CPS data files for the 1970s and 1980s lack any information on immigrant status. Starting in 1994, however, immigrant information is available in the CPS. Using this data, a simple way to assess the role of immigration is to re-estimate the college-high school wage gaps excluding all immigrants. The results, which are presented in Appendix Table A-3, show that the cohort-specific college-high school wage gaps for 1994-96 are very similar whether immigrants are included or excluded. A comparison of the fractions of immigrants in different education groups helps explain why this is true (see Appendix Table A-3). Although immigrants are over-represented at very low and very high levels of education, the fraction of immigrants among high school- educated men is about equal to the fraction among college graduates. Moreover, immigrant-native wage differentials are not too different for college-educated and high school-educated workers. Thus, immigrant composition has only a small impact on the estimated college-high school wage premiums of different age groups.

V. Discussion

This paper is motivated by the observation that the college-high school wage gaps of different age groups have not moved together over the past two decades. In particular, the college premium for younger workers has risen substantially, while the premium for older workers is about the same today as it was in the mid-1970s. We have argued that this shifting structure can be largely explained by a combination of cohort-specific relative supplies of college-equivalent labor, and steady rises in the relative productivity of college workers. Within this framework, the increase in the college-high school wage gap over the past two decades is attributable to steadily rising relative demand for educated labor, coupled with the dramatic trend-break in the relative supply of college educated workers that occurred for cohorts born after 1950.

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groups over time.

These conclusions raise a number of interesting questions for further research and analysis. A central question is what has caused the trend in relative demand for college-educated workers that plays such an important part in the models in Tables 5-7? While many observers attribute the trend to skill-biased technological change, we have found little connection to one readily available indicator of technological change -- the use of computers in the workplace. A more sophisticated view is that computers are just the latest in a steady stream of technological innovations that have raised the relative productivity of more highly-educated workers over the past 40 or 50 years. It remains an open question whether this hypothesis can be rigorously tested.

In searching for explanations for the trend in relative demand it is important to keep in mind that our models cannot distinguish between an aggregate trend that affects the relative demand for better-educated workers of all age groups, and a cohort-specific trend that shifts the relative demand for college graduates in newly-entering cohorts, leaving relative demand for older workers fixed.[34] One way of choosing between these alternatives is to consider their implications for the changing age structure of the college wage premium. The <u>aggregate trend</u> model implies that the steeply-rising age profile of college gaps in the mid-1970s was largely an artefact of the relative shortage of older college-educated workers at that time. Over the next two decades, as the relative supply of college workers increased among older workers, the college premiums for these workers fell -- counteracting the aggregate trend that raised returns for younger workers, and leading to a roughly constant premium for older men between 1975 and 1995. An <u>inter-cohort</u> trend model implies that each successive cohort must have a higher relative supply of college graduates to keep up with a rising cohort-specific relative demand. Under this interpretation, the rise in relative returns for younger workers over the past two decades reflects the failure of post-1950 cohorts to keep up with the

[34]If one replaces the aggregate trend in the models in Tables 5-7 with a cohort-specific trend (that increments by 5 between each 5-year birth cohort) all the coefficient estimates in these tables -- including the trend coefficients themselves -- are identical. Only the age effects change.

earlier trend toward rising college completion rates, while the stability of the premium for older workers reflects the aging of the pre-1950 cohorts.

A second question is raised by the observation that even controlling for cohort-specific relative supplies and long-run trends, returns to college fell between 1975 and 1980, and subsequently rose between 1980 and 1985. The existing literature on aggregate returns to education has explained this dip as a response to the upsurge in supply of educated labor for the economy as a whole (e.g. Katz and Murphy, 1992). While such a reaction is inconsistent with an age-separable production function such as equation (3), a more general model would allow the relative supplies of educated labor in all age groups to affect the college premium for any specific age group.

Aggregate trends in the relative supply of college-educated labor are illustrated in Figure 6. The figure plots the log of the relative supply of college-equivalents among male workers and male and female workers between 1960 and 1995.[35] The rate of growth of supply of college workers was somewhat faster in the 1970s than the 1960s. Comparing 1975 to 1980, however, it appears that relative supply was further "above trend" in the mid-1970s than in 1980.[36] Thus, it seems unlikely that shifts in the aggregate supply of educated labor can explain much of the dip in overall returns to education in the 1975-80 period.

Table 8 presents a series of models that evaluate the evidence more formally. These models include the aggregate supply index for male and female workers shown in Figure 6 as an added covariate. (Results using an aggregate index based on male workers are very similar). For reference,

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[35]We use the 1960 Census and 1970-97 CPS files to construct this graph. We define workers as those with positive wages and hours in the previous year. We use simple counts of workers in each education range (rather than hours-weighted counts) to construct high school and college equivalents.

[36]Between 1970 and 1975 the cohort born from 1945 to 1949 entered the labor market. These men had the highest rates of college completion of any earlier (or later) cohort. Between 1975 and 1980 the cohort born in the early 1950s entered the labor market. This cohort had much lower college completion rates than the preceding cohort, accounting for the slowdown in relative supply.

column 1 reproduces the "base" specification from column 4 of Table 5: this model includes an age-group-specific relative supply variable, a trend, and a dummy for returns in 1980. In column 2 we drop the 1980 dummy and include instead the aggregate supply index. The estimated coefficient of the aggregate supply variable is negative and significant, and remarkably similar to the estimate reported by Katz and Murphy (1992) based on a longer time series of aggregate returns.[37] However, as shown by the goodness-of-fit statistics at the bottom of the table, the relative supply index does not perform as well as a simple dummy for 1980. Moreover, when we include both an aggregate supply index and the 1980 dummy (column 3), the estimated coefficient of the aggregate index becomes very small, and insignificantly different from 0.

The relatively short time period of our base sample (only 5 observations over 20 years) may make it difficult to identify the effect of the aggregate supply index. To investigate this possibility we estimated similar models using the longer series of college wage gaps shown in Appendix Table B-1, based on *weekly* earnings from the 1960 Census, the 1970-72 CPS, and our 1975-95 CPS samples. Column 4 shows the results of re-estimating the basic specification (column 1) for the 1975-95 period, but using the college-high school gaps in weekly earnings. The results are quite similar to those based on the hourly wage gaps. Column 5 shows the results for the same specification, but with additional data for 1959 and 1970. The estimates of the relative supply effect, the trend, and the 1980 year effect are very similar, although the goodness of fit is worse, reflecting (in part) the rise and fall in returns from 1959 to 1975. The specification in column 6 adds the aggregate supply index: the results are fairly similar to those in column 2 over the shorter sample period. Finally, in column 7 we include the relative and aggregate supply indexes, a trend, and a 1980 dummy. In this specification the effect of the aggregate supply variable is substantially attenuated, but is marginally

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[37]Katz and Murphy (1992, page 69) report an estimated inverse elasticity of 0.709, using a combined supply index for men and women. The trend coefficient in column 4 is also virtually identical to the one they report.

significant. More importantly, however, the 1980 dummy still shows a dip in returns to college, relative to earlier or later periods. These results demonstrate the robustness of our findings with respect to the role of the relative supply variable, but leave open the question of whether there is a supply-based explanation for the decline and subsequent rise in aggregate returns to college in the early 1980s.

A third important question raised by our findings is what caused the break in the inter-cohort trend toward rising college completion rates after 1950? Some insight into this question is provided by an intergenerational perspective. Much existing research has shown that children's education outcomes are heavily affected by their parent's education (see e.g. Solon, 1999). For example, consider a simple inter-generational model of educational attainment:

(10) $S_{ic} = \alpha_c + \beta P_{ic} + u_{ic}$,

where S_{ic} is the education level of a son in cohort c, and P_{ic} is the average education of his parents. A typical estimate of β is in the range of 0.4 to 0.5 (Card, 1999, Table 2). Using such a model it is possible to decompose the rise in mean education between cohorts 1 and 2 as:

$$S_2 - S_2 = \alpha_2 - \alpha_1 + \beta (P_2 - P_1),$$

where S_c is the mean level of education for sons in cohort c, and P_c is the mean level of education of parents for that cohort.[38] Likewise, one can decompose the change in the rate of growth of educational attainment across cohorts into a component due to the change in the rate of growth of parental education, and an "unexplained" component.

We used data from the pooled 1977-1996 General Social Survey (GSS) to examine the sources of the slow-down in the rate of growth of educational attainment beginning with cohorts born in the early 1950s. The mean levels of sons' and parents' education for cohorts born from 1925 to

[38] If the coefficient β varies by cohort the decomposition includes another term. In practice the intergenerational coefficient is very stable across cohorts in the data set we used, and we could not reject the hypothesis of a constant coefficient. We therefore ignore variation in β.

1969 are plotted against each other in Figure 7.[39] Consistent with the data in Figure 3, the GSS samples show steadily rising educational attainment for cohorts born until the late 1940s, then an abrupt stagnation. Looking at the figure it is clear that some of the slowdown in the growth of education across cohorts is attributable to a slowdown in the rate of growth of parents' education: comparing the 1945-49 cohort to the 1925-29, for example, mean parents' education rose by 2.45 years, while between the 1945-49 and 1965-69 cohorts, mean parents' education rose only 1.65 years. Using an estimate of the intergenerational coefficient of β=0.45 this accounts for about 20 percent of the relative slowdown in the rate of growth of sons' education.[40] The remainder is attributable to changes in the cohort-specific constants in equation (10).

A broader view suggests that some of the change in the cohort-specific constants in equation (10) may also be attributable to intergenerational effects, via peer-group effects or similar mechanisms.[41] For example, suppose that (10) is expanded to

$$(11) \quad S_{ic} = \alpha_c' + \beta P_{ic} + \gamma P_c + u_{ic},$$

where P_c is the mean level of parental education for sons in cohort c, and γ represents an externality effect. The combination of $\beta+\gamma$ can be estimated by a cohort-level model:

$$S_c = \alpha_c' + (\beta+\gamma)P_c + u_c ,$$

assuming that variation in the remaining cohort effect, α_c', is orthogonal to mean parental education.

[39]Our sample contains 5384 men born from 1920 to 1969 who were ages 26-60 at the time of a survey, and who reported valid data on their mother's and father's education. The mean levels of sons' education in Figure 7 are regression estimates by cohort, adjusting for the age at which education is observed.

[40]The estimate of β is obtained from a regression that includes a quadratic in age and cohort effects. The standard error is 0.02; the R-squared is 0.25. The mean son's education rose by 2.62 years between the 1925-29 cohort and the 1945-49 cohort, and by 0.34 years between the 1945-49 cohort and the 1965-69 cohort.

[41]See Borjas (1992) for an application of this idea to differences in education across ethnic groups.

The data for the pre-1950 cohorts in Figure 7 imply an estimate of $(\beta+\gamma)$ that is very close to 1.[42] Assuming that $(\beta+\gamma)=1$, about 40 percent of the slow-down in the growth of educational attainment across cohorts after 1950 can be explained by the slow-down in the rate of growth of parental education. This is presumably an upper-bound estimate of any intergenerational effects, and still leaves a substantial share of the slow-down unexplained.

One potential explanation for the slow-down in college completion for post-1950 cohorts is that cohorts at the peak of the baby boom were "crowded out" of college. The 1950-54 cohort was about 13 percent bigger than the 1945-49 cohort, while the 1955-59 cohort was 27 percent bigger.[43] These comparisons suggest that the U.S. college system would have had to continue expanding in the early 1970s if the peak baby boom cohorts were to attend college at <u>higher</u> rates than the 1945-49 cohort and maintain the historic inter-cohort trend. Moreover, female college attendance rates were rising, leading to even greater competition for college slots in the 1970s. Total enrollment in higher education rose 70 percent from 1965 to 1975, but grew only 12 percent from 1975 to 1985, and 11 percent from 1985 to 1995. Total *male* enrollment rose 51 percent from 1965 to 1975, but fell about 3 percent from 1975 to 1985 and ended up only slightly higher in 1995 than in 1975.[44] While these data are potentially consistent with "supply constraints" arising from a slow-down in investment in the infrastructure of higher education, alternatively they may simply reflect a reduction in the rate of growth of demand for college education among the post-1950 cohorts of men. Further research on the demand for college education among the post-war cohorts is needed to identify the potential

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[42]The estimated coefficient is 0.97 (standard error 0.04). However, this estimate is only based on 5 observations.

[43]These figures are based on the number of people in the cohort, when the cohort was age 20-24 (taken from the 1997 *Digest of Education Statistics*, Table 15). Due to immigration, these figures probably exaggerate the relative size of more recent cohorts at a slightly younger age.

[44]These numbers are based on total enrollment of people age 14-34 in institutions of higher education, measured in the October CPS: see 1997 *Digest of Education Statistics*, Table 212.

contributions of the various factors.

VI. Conclusions

Contrary to the impression conveyed by some of the existing literature, not all education-based wage differentials in the U.S. economy have expanded over the past two decades. In particular, the college-high school wage differential for men over age 45 is no higher today than it was in the mid-1970s. By comparison, the college premium for younger men has risen by over 100 percent. We argue that these changes can be explained by a model in which the college wage gap for a particular age group depends on the relative supply of college educated workers in that group, and on underlying trends in the relative productivity of better-educated workers. Within this framework, the rising relative wage of younger college educated workers is explained by the relative stagnation in college completion rates for cohorts born after 1950, who began to enter the labor force in the 1970s. Once relative supply and underlying trends are taken into account, we find little evidence that the diffusion of computers, or shifting union coverage, have had much affect on the returns to college education.

Our findings point to a number of interesting issues for further research. One is to better understand the sources of the persistent trend in the relative productivity of more highly educated workers. A second is to account for the dip and subsequent rebound in returns to college that occurred for all age groups between 1975 and 1985. A third important issue for public policy is to explain the sharp break in the trend of rising educational attainment that occurred in the 1970s. If cohorts born after 1950 had been able to maintain the trend set by earlier cohorts, education levels would be substantially higher in the U.S. today, and the college wage premium would be lower.

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Appendix A

a. Sample Description

The wage gaps in Table 1 are based on samples of men in the March CPS. The 1975-76 data are drawn from the March 1976 and 1977 CPS; the 1979-81 data are drawn from the March 1980-82 CPS; the 1984-86 data and drawn from the 1985-87 CPS; the 1989-91 data are drawn from the March 1990-92 CPS; and the 1994-96 data are drawn from the 1995-97 CPS. The samples for "year t" (t=1975, 1980, 1985, 1990, 1995) include men who were age 26-60 in t, plus men who were 25 to 59 in t-1; plus men who were 27 to 61 in t+1. Hourly wages are formed by dividing annual wage and salary earnings by the product of weeks worked during the year and usual hours per week. We use the CPI to deflate all wages to 1989 dollars. Individuals whose wages are less than $2.00 per hour or more than $150 per hour in 1989 dollars are excluded. The wage gaps are estimated in separate regression models for each cohort in each "year", using samples of men with exactly 12 or exactly 16 years of education. Each model includes a dummy for 16 years of education, a linear age term, an indicator for nonwhite race, and dummies for which CPS sample the observation was drawn from.

The education distributions in Table 4 are based on similar samples, except that all men (workers and nonworkers) are included.

b. Wages of Workers with Post-Graduate Education

The college wage premiums in this paper pertain to individuals with a college degree and no post-graduate education. Appendix Table A-1 presents the wage gaps (by cohort and "year") between men with any post-graduate education, and those with exactly a college degree. These gaps have been rising over time, and their age structure has changed. Comparisons over time are potentially affected by the introduction of a new education question in the 1992 CPS. Under the old (pre-1992) wording, individuals were asked their number of years of completed education. Individuals who report more than 16 years of completed education are considered to have some post-graduate education. Under the new wording, individuals are asked their highest degree: individuals who report holding a masters degree, Ph.D., or professional degree are considered to have some post-graduate education. An analysis of the February 1990 CPS file, which asked individuals both education questions, suggests that the change in wording leads to a reduction in the fraction of people with post-graduate education, and a corresponding increase in the fraction with exactly a college degree. As shown in Appendix Table A-2 (row 3), the fraction of male college graduates age 26-60 who are coded as having more than a college degree falls from 44.8 percent under the old question to 36.6 percent under the new question. This change also raises the estimated college-high school wage gap slightly (see row 1 of Appendix Table 2), and raises the wage differential between post-graduates and graduates.

Appendix Table A-1: Wage Premiums for Post-College Schooling by Age and Year

	Age Range						
	26-30	31-35	36-40	41-45	46-50	51-55	56-60
1975-76	0.011	0.013	0.028	0.032	-0.027	-0.021	-0.102
	(0.017)	(0.020)	(0.025)	(0.026)	(0.030)	(0.032)	(0.042)
1979-81	0.020	0.059	0.065	0.028	-0.008	0.058	0.025
	(0.014)	(0.013)	(0.017)	(0.019)	(0.020)	(0.022)	(0.026)
1984-86	-0.019	0.065	0.106	0.055	0.071	0.029	-0.023
	(0.016)	(0.015)	(0.015)	(0.018)	(0.022)	(0.025)	(0.028)
1989-91	0.035	0.063	0.143	0.118	0.116	0.118	0.040
	(0.017)	(0.016)	(0.016)	(0.016)	(0.019)	(0.024)	(0.028)
1994-96	0.064	0.154	0.239	0.265	0.257	0.224	0.137
	(0.021)	(0.019)	(0.019)	(0.019)	(0.019)	(0.024)	(0.032)

Notes: Standard errors in parentheses. The elements of the table are estimates of the difference in mean log wages between individuals with more than 16 years of education and those with exactly 16 years of education in the indicated years and age range.

Appendix Table A-2: Effects of the Change in CPS Question Wording on the Measured Returns to College and Post-College Education

	Old Wording	New Wording
1. Estimated College-High School Wage Gap (excluding people with post-graduate education)	0.360 (0.017)	0.377 (0.017)
2. Estimated College-High School Wage Gap (including people with post-graduate education)	0.413 (0.015)	0.434 (0.015)
3. Fraction of College Graduates with Post-graduate schooling	0.448	0.366
4. Estimated Return to Post-college Education	0.124 (0.022)	0.159 (0.023)

Notes: Standard errors in parentheses. Sample includes 5,876 men age 26-60 in the outgoing rotation groups of the February 1990 CPS who report a weekly or hourly wage between $3.00 and $200 per hour. The samples in row 1 are further restricted to men who report either exactly high school or exactly college education (3,167 observations in column 1; 3,070 in column 2). The samples in rows 2 and 4 are further restricted to men who report either exactly high school or college-or-higher education (3,923 observations in column 1; 3,688 in column 2). Estimates in rows 1-2 are coefficients of an indicator for college degree in models that also include a cubic in age. Estimates in row 4 are coefficients of an indicator for post-college education in models that also include an indicator for college degree and a cubic in age.

Appendix Table A-3: College-High School Wage Gaps Including and Excluding Immigrant Workers, 1994-96 Sample

	Percent Immigrants:			College-High School Wage Gap:	
Age Group	All (1)	High School (2)	College Degree (3)	Overall (4)	No Immigrants (5)
26-30	15.3	11.3	10.4	0.317 (0.013)	0.297 (0.013)
31-35	13.7	9.4	11.3	0.419 (0.013)	0.416 (0.013)
36-40	12.0	7.3	10.7	0.436 (0.013)	0.447 (0.014)
41-45	10.9	7.3	10.8	0.417 (0.015)	0.415 (0.016)
46-50	11.4	9.5	10.8	0.363 (0.016)	0.355 (0.017)
51-55	10.1	6.8	10.9	0.349 (0.020)	0.354 (0.021)
56-60	10.1	7.1	11.1	0.396 (0.025)	0.396 (0.028)

Note: Based on March 1995-97 CPS files. Entries in columns 1-3 are percentages of men who are immigrants: column 1 includes all education groups; column 2 is for men with a high school diploma; column 3 is for men with a college degree (and no post-graduate education). Entries in columns 4-5 are log wage premiums for a college degree -- see Table 1 for details. Estimates in column 4 are for all men in the age range; estimates in column 5 are for native-born men only. Standard errors in parentheses.

Appendix Table B-1: College High School Wage Differentials in Average Weekly Wages by Age and Year

	Age Range:						
	26-30	31-35	36-40	41-45	46-50	51-55	56-60
1959	0.148 (0.008)	0.296 (0.008)	0.350 (0.009)	0.365 (0.012)	0.360 (0.014)	0.378 (0.017)	0.374 (0.023)
1969-71	0.181 (0.014)	0.272 (0.016)	0.375 (0.017)	0.394 (0.018)	0.364 (0.020)	0.403 (0.024)	0.403 (0.031)
1975-76	0.121 (0.016)	0.262 (0.018)	0.326 (0.021)	0.354 (0.022)	0.365 (0.024)	0.376 (0.025)	0.394 (0.035)
1979-81	0.112 (0.012)	0.181 (0.013)	0.253 (0.016)	0.302 (0.018)	0.355 (0.019)	0.354 (0.019)	0.352 (0.023)
1984-86	0.300 (0.013)	0.333 (0.014)	0.347 (0.015)	0.415 (0.019)	0.412 (0.021)	0.441 (0.023)	0.455 (0.027)
1989-91	0.330 (0.013)	0.420 (0.014)	0.422 (0.015)	0.442 (0.016)	0.413 (0.019)	0.412 (0.023)	0.467 (0.027)
1994-96	0.344 (0.015)	0.484 (0.015)	0.498 (0.015)	0.447 (0.017)	0.417 (0.018)	0.388 (0.024)	0.431 (0.031)

Notes: Standard errors in parentheses. The elements of the table are estimates of the difference in mean log wages between individuals with 16 and 12 years of education in the indicated years and age range. Data for 1959 are taken from the 1960 Census; later data are taken from the March CPS. Average weekly earnings are estimated by dividing annual earnings by an estimate of weeks worked, based on midpoints of 7 intervals. See note to Table 1.

Appendix Table B-2: Decompositions of College-High School Wage Differentials by Age and Year into Cohort, Age, and Time Effects

	Unrestricted Model	No Cohort Effects	10 Oldest Cohorts Only	Models with Restricted Cohort Effects: 7 Oldest Same	10 Oldest Same
Degrees of Freedom	24	36	26	29	32
Chi-squared (p-value)	27.97 (0.26)	213.88 (0.00)	42.94 (0.02)	32.04 (0.32)	47.91 (0.04)
R-squared	0.99	0.89	0.97	0.98	0.98
Year Effects (relative to 1975-76):					
1959	--	0.018 (0.022)	0.021 (0.012)	-0.006 (0.012)	0.021 (0.011)
1969-71	--	0.035 (0.026)	0.033 (0.014)	0.023 (0.012)	0.033 (0.013)
1979-81	--	-0.043 (0.025)	-0.051 (0.014)	-0.044 (0.011)	-0.051 (0.013)
1984-86	--	0.083 (0.025)	0.049 (0.016)	0.063 (0.012)	0.049 (0.014)
1989-91	--	0.123 (0.025)	0.067 (0.017)	0.080 (0.014)	0.059 (0.014)
1994-96	--	0.148 (0.026)	0.044 (0.021)	0.085 (0.016)	0.057 (0.016)
Fraction of 1975-96 Change Explained[a/]	--	0.00	0.70	0.43	0.61
Cohort Effects:					
1935-39	--	--	--	-0.035 (0.019)	--
1940-44	--	--	--	-0.034 (0.012)	--
1945-49	--	--	--	-0.050 (0.014)	--
1950-54	--	--	--	-0.014 (0.016)	0.033 (0.011)
1955-59	--	--	--	0.061 (0.019)	0.114 (0.014)
1960-64	--	--	--	0.094 (0.022)	0.153 (0.017)
1965-69	--	--	--	0.093 (0.028)	0.157 (0.025)

Notes: Standard errors in parentheses. Models are fit by weighted least squares to the 49 age-group by year college-high school wage gaps shown in Table B-1. Weights are inverse sampling variances of the estimated wage gaps. See note to Table 2.

[a/] This row shows the proportional reduction in the 1994-96 year effect for the model in the specific column relative to the model in column 2.

Appendix Table C-1: College High School Wage Differentials by Experience and Year

	Experience Range:							
	1-5	6-10	11-15	16-20	21-25	26-30	31-35	36-40
1975-76	0.335 (0.013)	0.329 (0.016)	0.347 (0.019)	0.343 (0.020)	0.374 (0.021)	0.364 (0.022)	0.373 (0.028)	0.246 (0.041)
1979-81	0.311 (0.010)	0.276 (0.011)	0.312 (0.013)	0.288 (0.015)	0.352 (0.016)	0.301 (0.016)	0.306 (0.019)	0.237 (0.027)
1984-86	0.541 (0.011)	0.401 (0.011)	0.407 (0.012)	0.413 (0.015)	0.384 (0.017)	0.384 (0.019)	0.394 (0.020)	0.310 (0.027)
1989-91	0.541 (0.011)	0.508 (0.011)	0.445 (0.012)	0.442 (0.013)	0.428 (0.016)	0.354 (0.018)	0.380 (0.021)	0.424 (0.027)
1994-96	0.537 (0.013)	0.521 (0.013)	0.562 (0.013)	0.443 (0.014)	0.432 (0.015)	0.380 (0.019)	0.390 (0.024)	0.299 (0.031)

Notes: Standard errors in parentheses. The elements of the table are estimates of the difference in mean log wages between individuals with 16 and 12 years of education in the indicated experience range. Experience is defined as age-20 for men with 12 years of education, and age-25 for men with 16 years of education.

Appendix Table C-2: Decompositions of College-High School Wage Differentials by Experience and Year into Experience, Time, and Cohort Effects

	Unrestricted Model	No Cohort Effects	8 Oldest Cohorts Only	Models with Restricted Cohort Effects: 5 Oldest Same	8 Oldest Same
Degrees of Freedom	18	28	18	21	24
Chi-squared	23.70	261.19	34.94	27.08	40.41
(p-value)	(0.17)	(0.00)	(0.01)	(0.17)	(0.02)
R-squared	0.98	0.81	0.88	0.98	0.97
Year Effects (relative to 1975-76):					
1979-81	--	-0.040	-0.046	-0.050	-0.047
		(0.026)	(0.015)	(0.011)	(0.011)
1984-86	--	0.087	0.041	0.041	0.047
		(0.026)	(0.015)	(0.016)	(0.012)
1989-91	--	0.124	0.061	0.046	0.055
		(0.026)	(0.016)	(0.022)	(0.014)
1994-96	--	0.143	0.046	0.038	0.051
		(0.027)	(0.019)	(0.029)	(0.016)
Fraction of 1975-96 Change Explained[a/]	--	0.00	0.68	0.73	0.64
Cohort Effects (entry year):					
1960-64	--	--	--	-0.028	--
				(0.017)	
1965-69	--	--	--	0.003	--
				(0.022)	
1970-74	--	--	--	0.013	--
				(0.028)	
1975-79	--	--	--	0.039	0.028
				(0.035)	(0.011)
1980-84	--	--	--	0.162	0.147
				(0.042)	(0.015)
1985-90	--	--	--	0.168	0.149
				(0.050)	(0.019)
1991-95	--	--	--	0.171	0.148
				(0.058)	(0.026)

Notes: Standard errors in parentheses. Models are fit by weighted least squares to the 40 college-high school wage gaps shown in Table C-1. Weights are inverse sampling variances of the estimated wage gaps. See note to Table 2.
[a/] This row shows the proportional reduction in the 1994-96 year effect for the model in the specific column relative to the model in column 2.

Table 1: College High School Wage Differentials by Age and Year

	Age Range						
	26-30	31-35	36-40	41-45	46-50	51-55	56-60
1975-76	0.115 (0.013)	0.217 (0.016)	0.296 (0.019)	0.323 (0.020)	0.351 (0.021)	0.369 (0.023)	0.401 (0.030)
1979-81	0.100 (0.010)	0.149 (0.011)	0.231 (0.014)	0.262 (0.016)	0.308 (0.016)	0.317 (0.017)	0.337 (0.020)
1984-86	0.256 (0.011)	0.284 (0.011)	0.292 (0.013)	0.347 (0.016)	0.352 (0.018)	0.400 (0.020)	0.408 (0.022)
1989-91	0.299 (0.011)	0.378 (0.011)	0.363 (0.012)	0.370 (0.013)	0.372 (0.016)	0.336 (0.019)	0.436 (0.022)
1994-96	0.317 (0.013)	0.419 (0.013)	0.436 (0.013)	0.417 (0.015)	0.363 (0.016)	0.349 (0.020)	0.396 (0.025)

Notes: Standard errors in parentheses. The elements of the table are estimates of the difference in mean log wages between individuals with 16 and 12 years of education in the indicated years and age range. Samples contain a rolling age group. For example, the 26-30 year old group in the 1979-81 sample includes individuals 25-29 in 1979, 26-30 in 1980, and 27-31 in 1981.

Table 2: Decompositions of College-High School Wage Differentials by Age and Year into Cohort, Age, and Time Effects

	Unrestricted Model	No Cohort Effects	7 Oldest Cohorts Only	Models with Restricted Cohort Effects: 4 Oldest Same	7 Oldest Same
Degrees of Freedom	15	24	14	17	20
Chi-squared (p-value)	11.36 (0.73)	216.63 (0.00)	13.56 (0.48)	11.37 (0.84)	16.58 (0.68)
R-squared	0.99	0.84	0.98	0.99	0.99
Year Effects (relative to 1975-6):					
1979-81	--	-0.048 (0.026)	-0.060 (0.010)	-0.060 (0.009)	-0.065 (0.008)
1984-86	--	0.055 (0.027)	0.008 (0.011)	0.015 (0.014)	0.004 (0.009)
1989-91	--	0.100 (0.027)	0.016 (0.012)	0.027 (0.020)	0.010 (0.010)
1994-96	--	0.128 (0.027)	-0.003 (0.015)	0.025 (0.027)	0.002 (0.012)
Fraction of 1975-96 Change Explained[a/]	--	0.00	1.02	0.80	0.98
Cohort Effects:					
1935-39	--	--	--	-0.032 (0.014)	--
1940-44	--	--	--	-0.024 (0.019)	--
1945-49	--	--	--	-0.025 (0.025)	--
1950-54	--	--	--	0.029 (0.032)	0.062 (0.011)
1955-59	--	--	--	0.108 (0.038)	0.147 (0.011)
1960-64	--	--	--	0.145 (0.045)	0.189 (0.014)
1964-69	--	--	--	0.160 (0.053)	0.210 (0.019)

Notes: Standard errors in parentheses. Models are fit by weighted least squares to the 35 age-group by year college-high school wage gaps shown in Table 1. Weights are inverse sampling variances of the estimated wage gaps. Model in column (1) is not identified. All models include age effects

[a/] This row shows the proportional reduction in the 1994-96 year effect for the model in the specific column relative to the model in column 2.

Table 3: Alternative Decompositions of College-High School Wage Differentials by Age and Year into Cohort, Age, and Time Effects

	Specification:			
	(1)	(2)	(3)	(4)
Degrees of Freedom	19	22	18	21
Chi-squared	16.56	17.92	15.91	17.58
(p-value)	(0.62)	(0.71)	(0.60)	(0.68)
R-squared	0.99	0.98	0.99	0.98
Year Effects (relative to 1975-6):				
1979-81	-0.065	-0.068	-0.060	-0.065
	(0.009)	(0.007)	(0.010)	(0.008)
1984-86	0.004	0	0.005	0
	(0.009)		(0.009)	
1989-91	0.010	0	0.013	0
	(0.010)		(0.011)	
1994-96	0.002	0	0.007	0
	(0.013)		(0.013)	
Loading Factors on Year Effects for Age Groups:[a]				
Age 26-30	1.04	1.10	1.11	1.14
	(0.24)	(0.24)	(0.28)	(0.26)
Age 31-35	1.00	1.00	1.22	1.14
	--	--	(0.25)	(0.23)
Age 36+	1.00	1.00	1.00	1.00
	--	--	--	--
Cohort Effects:				
1950-54	0.063	0.067	0.060	0.066
	(0.010)	(0.011)	(0.010)	(0.009)
1955-59	0.147	0.147	0.141	0.148
	(0.011)	(0.008)	(0.013)	(0.009)
1960-64	0.189	0.194	0.182	0.192
	(0.015)	(0.010)	(0.018)	(0.011)
1964-69	0.209	0.209	0.201	0.208
	(0.020)	(0.014)	(0.023)	(0.014)

Notes: Standard errors in parentheses. Models are fit by weighted nonlinear least squares to the 35 age-group by year college-high school wage gaps shown in Table 1. Weights are inverse sampling variances of the estimated wage gaps. All models include unrestricted age effects, and assume that oldest 7 cohorts (born from 1915 to 1949) have the same cohort effect.

[a] Year effects are allowed to affect different age groups differently, with the normalization that groups age 36 and older are affected similarly.

Table 4: College Completion Rates by Age and Year, Adult Men

	Age Group:						
	26-30	31-35	36-40	41-45	46-50	51-55	56-59
A. Fraction of Age Group with College or Higher Education							
1975-76	0.287	0.265	0.221	0.210	0.174	0.150	0.129
	(0.004)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)
1979-81	0.249	0.306	0.267	0.239	0.217	0.199	0.171
	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
1984-86	0.238	0.281	0.323	0.284	0.248	0.225	0.194
	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
1989-91	0.237	0.241	0.291	0.324	0.288	0.248	0.228
	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)
1994-96	0.261	0.257	0.259	0.294	0.339	0.290	0.253
	(0.004)	(0.003)	(0.003)	(0.004)	(0.004)	(0.005)	(0.005)
B. Fraction of College Graduates with Some Post-Graduate Schooling							
1975-76	0.412	0.514	0.496	0.510	0.421	0.471	0.460
	(0.009)	(0.010)	(0.012)	(0.013)	(0.014)	(0.015)	(0.018)
1979-81	0.392	0.470	0.551	0.531	0.510	0.440	0.456
	(0.007)	(0.006)	(0.008)	(0.009)	(0.010)	(0.010)	(0.011)
1984-86	0.346	0.419	0.475	0.546	0.518	0.515	0.488
	(0.007)	(0.007)	(0.007)	(0.008)	(0.010)	(0.010)	(0.011)
1989-91	0.296	0.365	0.419	0.469	0.502	0.485	0.503
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.010)	(0.011)
1994-96	0.220	0.299	0.338	0.372	0.428	0.457	0.445
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.009)	(0.011)

Notes: Standard errors in parentheses. Samples include men (workers and nonworkers) in indicated age ranges. See notes to Table 1.

Table 5: Estimated Models for the College-High School Wage Gap, By Cohort and Year

	(1)	(2)	(3)	(4)	(5)	(6)
Relative Supply	--	-0.208 (0.017)	-0.222 (0.034)	-0.207 (0.017)	-0.208 (0.017)	-0.202 (0.019)
Trend (1975=0)	--	--	0.013 (0.001)	0.010 (0.001)	0.010 (0.001)	0.010 (0.001)
Year Effects:						
1979-81	-0.048 (0.026)	-0.017 (0.010)	0.0	-0.068 (0.007)	-0.068 (0.008)	-0.068 (0.008)
1984-86	0.055 (0.027)	0.101 (0.011)	0.0	0.0	0.0	0.0
1989-91	0.100 (0.027)	0.152 (0.011)	0.0	0.0	0.0	0.0
1994-96	0.128 (0.027)	0.206 (0.012)	0.0	0.0	0.003 (0.010)	0.0
Cohort Effects:						
Born 1960-64	--	--	--	--	--	0.008 (0.013)
Born 1965-69	--	--	--	--	--	0.014 (0.018)
Chi-squared	216.63	29.98	130.05	30.10	29.98	29.23
Degrees of freedom	24	23	25	24	23	26
P-value	0.00	0.15	0.00	0.18	0.15	0.30

Notes: Standard errors in parentheses. Models are fit by weighted least squares to the 35 age-group by year college-high school wage gaps shown in Table 1. All models include unrestricted age effects (not reported). Year effects shown as 0.0 with no standard error are set to 0.

Table 6: Estimated Models for the College-High School Wage Gap, By Cohort and Year

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Relative Supply	-0.207 (0.017)	--	-0.200 (0.019)	--	-0.193 (0.020)	-0.195 (0.019)	-0.194 (0.019)
Trend (1975=0)	0.010 (0.001)	--	--	--	--	0.009 (0.001)	0.009 (0.001)
Relative Use of Computers	--	0.627 (0.254)	0.099 (0.118)	--	--	--	-0.051 (0.111)
Relative Union Rate	--	--	--	0.814 (0.243)	0.176 (0.126)	0.160 (0.118)	0.207 (0.158)
Year Effects:							
1980	-0.068 (0.007)	-0.048 (0.024)	-0.019 (0.011)	-0.049 (0.022)	-0.020 (0.010)	-0.060 (0.009)	-0.065 (0.014)
1985	0.0	-0.128 (0.078)	0.071 (0.038)	-0.032 (0.034)	0.079 (0.019)	0.0	0.0
1990	0.0	-0.129 (0.095)	0.115 (0.047)	-0.019 (0.042)	0.123 (0.024)	0.0	0.0
1995	0.0	-0.166 (0.122)	0.157 (0.060)	-0.026 (0.051)	0.167 (0.030)	0.0	0.0
Chi-squared	30.10	171.15	29.06	145.43	27.56	27.97	27.72
Degrees of freedom	25	23	22	23	22	24	23
P-value	0.22	0.00	0.14	0.00	0.19	0.26	0.23

Notes: Standard errors in parentheses. Models are fit by weighted least squares to the 35 age-group by year college-high school wage gaps shown in Table 1. All models include unrestricted age effects (not reported). See notes to Table 5.

Table 7: Estimated Models for the College-High School Wage Gap, Using Data for 1985-95 Only

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Relative Supply	-0.208 (0.031)	--	-0.196 (0.038)	--	-0.229 (0.039)	-0.224 (0.036)	-0.224 (0.062)
Trend (1975=0)	0.011 (0.001)	--	--	--	--	0.014 (0.004)	0.014 (0.005)
Relative Use of Computers	--	-0.699 (0.254)	-0.165 (0.223)	--	--	--	0.005 (0.320)
Relative Union Rate	--	--	--	1.059 (0.878)	-0.490 (0.510)	-0.430 (0.475)	-0.438 (0.759)
Year Effects:							
1990	0.0	0.097 (0.031)	0.063 (0.018)	0.004 (0.039)	0.071 (0.022)	0.0	0.0
1995	0.0	0.200 (0.066)	0.133 (0.039)	-0.012 (0.075)	0.149 (0.046)	0.0	0.0
Chi-squared	15.17	51.59	14.32	61.50	13.83	14.12	14.12
Degrees of freedom	12	11	10	11	10	11	10
P-value	0.23	0.00	0.16	0.00	0.18	0.23	0.17

Notes: Standard errors in parentheses. Models are fit by weighted least squares to the 35 age-group by year college-high school wage gaps shown in Table 1. All models include unrestricted age effects (not reported). See notes to Table 5.

Table 8: Estimated Models for the College-High School Wage Gap, By Cohort and Year

	Based on Wage Gaps from Table 1 (1975-95):			Based on Wage Gaps from Table B-1:			
				1975-95	1959-95		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Relative Supply	-0.207 (0.017)	-0.209 (0.030)	-0.207 (0.017)	-0.203 (0.022)	-0.195 (0.019)	-0.179 (0.027)	-0.180 (0.020)
Trend (1975=0)	0.010 (0.001)	0.033 (0.008)	0.010 0.005	0.011 (0.001)	0.010 (0.001)	0.031 (0.004)	0.017 (0.004)
Aggregate Supply Index	--	-0.705 (0.261)	-0.003 (0.173)	-	--	-0.592 (0.112)	-0.206 (0.105)
Year Effect for 1980	-0.068 (0.007)	0.0	-0.068 (0.009)	-0.075 (0.010)	-0.089 (0.010)	0.0	-0.074 (0.012)
Chi-squared	30.1	100.64	30.1	37.07	81.14	142.03	73.66
Degrees of freedom	25	25	24	25	39	39	38
P-value	0.22	0.00	0.18	0.06	0.00	0.00	0.00

Notes: Standard errors in parentheses. Models in columns 1-3 are fit by weighted least squares to the 35 age-group by year college-high school wage gaps shown in Table 1. Models in columns 4-7 are fit by weighted least squares to the wage gaps shown in Table B-1. The estimates in column 4 use only the 35 gaps for 1975-1995; the estimates in columns 5-7 use 49 gaps for 1959, 1970, and 1975-95. All models include unrestricted age effects. The aggregate supply index is based on the ratio of college-equivalents to high school equivalents among the combined male-female labor force in the particular year.