Abstract

The rate of return to schooling appears to be nearly two percentage points greater for females than for males in the National Longitudinal Survey of Youth data set, despite the fact that females tend to earn less, both absolutely and controlling for personal characteristics. A survey of previous studies reporting wage equations reveals that a higher return to female schooling appears to be the norm, although it has not attracted comment. This paper considers various explanations. The most important involves the detrimental impact of discrimination and other factors that cause women to accept wage offers that undervalue their characteristics. It is hypothesized that the better educated is a woman, the more able and willing she is to overcome these handicaps and compete with men in the labour market, and an index of discrimination disaggregated by years of schooling is constructed using Oaxaca decompositions. This index is indeed negatively correlated with schooling and it accounts for about one half of the differential in the male and female schooling coefficients. Next considered is the possibility that part of the differential could be attributable to male-female differences in the quality of educational attainment, as proxied by their academic outcomes in high school. The NLSY females did indeed perform better than the males, but there is little association between academic attainment and Earnings and allowing for it made no difference to the estimate of the differential in the returns to schooling. The third explanation considered is that women choose to work in sectors where education is relatively highly valued. Controlling for this effect does indeed account for much of the remaining differential.

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Why is the Rate of Return to Schooling Higher For Women Than For Men?

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1. Introduction

Differentials in earnings by sex and ethnicity, persistent despite legislation against discrimination, have provoked a large and growing investigative literature (for surveys, see Lloyd and Neimi, 1979; Treiman and Hartmann, 1981; Madden, 1985; Cain, 1986; Gunderson, 1989; Blau and Kahn, 1992; Altonji and Blank, 1999; Blau and Kahn, 2000). The standard approach to the analysis of the determinants of earnings differentials, the Blinder–Oaxaca decomposition, involves the fitting of a Mincerian semilogarithmic wage equation

$$\ln Y = b_0 + \sum_{i=1}^{k} b_i X_i + u \tag{1}$$

where Y is a measure of earnings, the X_i are a set of k personal and labour market characteristics, and u is a disturbance term. In the case of sex differentials, the function is fitted for male and female samples separately. Using superscripts m and f for males and females, and noting that the fitted equations pass through the sample means,

$$\overline{\ln Y}^{m} = \hat{b}_0^{m} + \sum_{i=1}^{k} \hat{b}_i^{m} \overline{X}_i^{m} \tag{2}$$

$$\overline{\ln Y}^{f} = \hat{b}_0^{f} + \sum_{i=1}^{k} \hat{b}_i^{f} \overline{X}_i^{f} \tag{3}$$

Subtracting (3) from (2), the difference can be written

$$\overline{\ln Y}^{m} - \overline{\ln Y}^{f} = \sum_{i=1}^{k} \hat{b}_i^{m} \left(\overline{X}_i^{m} - \overline{X}_i^{f}\right) + \left(\hat{b}_0^{m} - \hat{b}_0^{f}\right) + \sum_{i=1}^{k} \overline{X}_i^{f} \left(\hat{b}_i^{m} - \hat{b}_i^{f}\right) \tag{4}$$

(Blinder, 1973; Oaxaca, 1973). The first term is said to measure that part of the earnings gap attributable to differences in characteristics and the other two that part attributable to discrimination. The first term is then typically decomposed into subcomponents attributable to individual characteristics. Some well-known caveats should be noted.

First, an alternative and in general different decomposition may be obtained by weighting the differences in characteristics by the female coefficients and the difference in coefficients by male characteristics. Under the hypothesis that the female coefficients would be the same as those for males in the absence of discrimination, it may be argued that the first decomposition is to be preferred. However, this is at best an approximation since male coefficients as well as the female ones may be affected by discrimination in the labour market. Various modifications to the decomposition have been proposed to address this issue (Reimers, 1983; Cotton, 1988; Neumark, 1988; Oaxaca and Ransom, 1994, 1999). Second, some of the characteristics, particularly investments in human capital such as schooling and training, may be endogenous and differences in characteristics may therefore partly be attributable to anticipated discrimination (Oaxaca, 1973). Third, it may be argued that some of the differences in coefficients may be attributable to factors other than discrimination. In particular, women may have tastes for certain kinds of work that cause them to be concentrated in relatively poorly paid occupations.[1] Circumstances may also be a factor, especially in the case of women with children, who may be willing to accept a wage offer that undervalues their characteristics if the job fits well with other responsibilities.

In the case of the United States, where the difference in earnings has been diminishing (O'Neill and Polachek, 1993; Blau and Kahn, 2000), the component attributable to differences in characteristics is now considered to be relatively small. Accordingly interest has become focused on the component attributable to differences in coefficients. Some authors, following the example of Blinder (1973), have attempted to decompose this part of the gap into subcomponents attributable to differences in the coefficients of individual characteristics and to the difference in the intercepts ('pure discrimination'), but here there is an asymmetry in the analysis, for while the decomposition of the differences in characteristics component is uncontroversial, a parallel decomposition of the discrimination component is illegitimate (Jones, 1983; Oaxaca and Ransom, 1999). Nevertheless, the signs of the differences in the coefficients are of interest and, given that the discrimination component favours males, one might expect that for the two most important variables, schooling and work experience, the female coefficients will be smaller. While this appears to be the case for work experience, it surprisingly does not appear to be true for schooling. Indeed, if anything, there appears to be a tendency for the estimated schooling coefficient to be *larger* for females.

The objective of Section 2 is to document this tendency. Section 3 suggests some factors that may be responsible for it. Section 4 uses data from the NLSY to test these hypotheses. Section 5 offers some conclusions.

[1] However, it may not be possible to make a clear distinction between tastes and discrimination, given that

2. Previous Findings[2]

Although the studies that have investigated male and female earnings are now legion, the number that actually report separate schooling coefficients (here described as the rate of return[3]) is much smaller. Many of those focusing on the returns to education with data on both sexes have fitted pooled regressions, allowing for a sex differential by including a simple dummy variable with no interactive term. Where separate regressions have been run, schooling is often among the unreported controls. In the case of the subliterature on the earnings gap, the single most plentiful source of studies with separate regressions, the regression coefficients are sometimes not reported at all. And when the schooling coefficients are reported, there is seldom enough information to determine whether they are significantly different, exceptions being Madden (1978) and Angle and Wissman (1981).

The present survey, summarized in Table 1, is confined to US studies with the Mincerian semilogarithmic specification of the wage equation. It is generally impossible to infer rates of return from studies that have used linear specifications and for that reason a number of widely cited studies (Cohen, 1971; Suter and Miller, 1973; Featherman and Hauser, 1976; Roos, 1981; Grubb, 1993) are not included. The linear Model is in any case a misspecification (Heckman and Polachek, 1974; Dougherty and Jimenez, 1991).

Also excluded are those studies that include occupation and/or industry variables as explanatory variables in the regression specification (Blinder, 1973; Treiman and Terrell, 1975; Loury, 1990; O'Neill and Polachek, 1993; Oaxaca and Ransom, 1994). Such specifications, dubbed 'full-scale' by Oaxaca (1973), have been widely used in the earnings gaps literature as a means of assessing how much of the male-female earnings gap is attributable to occupational segregation. Since much of the impact of schooling on earnings is mediated by occupational attainment, the interpretation of the schooling coefficient in a specification controlling for occupational attainment is different and narrower than in a specification that does not.

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tastes may be influenced by anticipated discrimination (Daymont and Andrisani, 1984; Gunderson, 1989).

2 Brief descriptions of the data sets, with web references, are provided in Appendix A.

3 This involves some licence since the coefficient can be described as an estimate of the internal rate of return only under restrictive assumptions. See, for example, the appendix in Dougherty and Jimenez (1991).

Studies using data from the Panel Study of Income Dynamics (PSID)

Corcoran and Duncan (1979), using data from the 1976 round of the PSID to investigate the contributions of differences in work histories and training to the male-female earnings gap, found male and female schooling coefficients of 0.059 and 0.077 for whites and 0.061 and 0.076 for blacks. Wellington (1993), using data for whites from the 1976 and 1985 rounds to investigate how differences in work histories and training have affected the wage gap over time, found male and female schooling coefficients of 0.049 and 0.074 in 1976, and 0.062 and 0.079 in 1985. Blau and Kahn (1997), using data from the 1980 and 1989 rounds to investigate the causes of trends in the male-female earnings gap, found male and female schooling coefficients of 0.066 and 0.084 in 1980 and 0.090 and 0.083 in 1989. They included in their specification dummy variables for a college degree and an advanced degree. The female coefficients were greater for these in both years, especially in 1989.

Taken together, these studies indicate a gap approaching two percentage points at least until the end of the 1980s.

Studies using data from the National Longitudinal Study of the High School Class of 1972 (NLS72) and High School and Beyond (HS&B)

Daymont and Andrisani (1984), using earnings data from the 1979 round of the NLS72 to investigate how much of the male-female earnings gap could be attributed to work preferences and choice of college major, found higher female coefficients for the dummy variables for master's and PhD degrees.

Altonji (1993), using earnings data from the 1979 and 1986 rounds of the NLS72 to investigate how expected returns to education affect the decision to stay in school and choice of college major, found that the female coefficients are greater than the corresponding male coefficients for 17 out 18 educational dummy variables: two partial college dummy variables (less than two years, and two or more years), ten college degree variables classified by area of major field, and six advanced degree variables classified by area of major field. The exception was an advanced degree category where the male coefficient is marginally greater. Altonji notes the pattern of greater female coefficients and states that it is consistent with the findings of other studies.

Grogger and Eide (1995), using pooled data from the NLS72 and HS&B for 1977, 1978, 1979 and 1986 (NLS72) and 1986 (HS&B) to investigate the determinants of the rise in the college wage premium in the 1980s, found that the returns to partial college education, a college degree, and a postgraduate degree were all much greater for women for new entrants to the labour force in both 1978 and 1986. However they also found that the male disadvantage diminishes with work experience. Allowing for an interaction between the returns to education and the returns to work experience, their estimates imply that the gap is eliminated for college graduates after ten years of experience and for postgraduates after six.

Kane and Rouse (1995), using NLS72 data to 1986 to compare the returns to two-year college with those to four-year college, found male and female schooling coefficients of 0.042 and 0.064 for the former and 0.046 and 0.062 for the latter.

Murnane, Willett, and Levy (1995), using NLS72 data for 1978 and HS&B data for 1986 to investigate the increase in the impact of cognitive skills, as represented by the mathematics test score, on wages, found male and female schooling coefficients of 0.013 and 0.037 for the earlier date and 0.021 and 0.037 for the later one.

Brown and Corcoran (1997), using NLS72 data to 1986 to investigate how choice of college major affects the male-female earnings gap, divided the NLS72 respondents into three groups, high school graduates, partial college, and college graduates, and used earnings data from the 1986 round to fit wage equations for each group separately. For the high school graduate group there was no variation in years of schooling. For the partial college group, they found male and female schooling coefficients of 0.035 and 0.052. For the college graduates, introducing dummy variables for a master's degree and for a PhD, they found a higher female coefficient for the former and a lower one for the latter. They also use data from the 1984 round of the Survey of Income and Program Participation (SIPP) to fit wage equations, dividing the sample into those who did not graduate from high school, those who did but did not graduate from college, and college graduates. The male and female years of schooling coefficients for the first subsample were 0.019 and –0.012, for the second 0.042 and 0.051, and for the third 0.031 and 0.037.

Loury (1997), using the NLS72 data to 1979 and HS&B data to 1986 to investigate how changing rewards to college majors contributed to the reduction in the male-female earnings gap among young college-educated adults in the 1980s, and defining dummy variables for partial college and for college graduates, found that the female coefficients were greater than the male ones for both variables for both years.

The findings of those studies with years of schooling coefficients appear to be in line with those using the PSID. The estimates of Murnane, Willett and Levy are relatively low, but they included a mathematics test score among their controls. Dropping the score, the estimates become closer to those in other studies, with the female coefficients two percentage points higher than the male ones. The studies using dummy variables are harder to assess but, with the exceptions of Altonji's advanced degree and Brown and Corcoran's high school drop-outs (SIPP) and PhD category (NLS72), they uniformly report higher coefficients for females than for males.

Studies using the National Longitudinal Surveys of Labor Market Experience

Mincer and Polachek (1974), using 1967 data from the National Longitudinal Survey of Mature Women (NLS–MW) to investigate the determinants of female earnings, found a schooling coefficient of 0.077 for single women and 0.063 for married women. For comparison, they fitted a parallel wage equation for males using data from the Survey of Economic Opportunity (SEO) and obtain a coefficient of 0.071.

Madden (1978), using 1969 data from the National Longitudinal Surveys of Young Men and Young Women (NLS–YM, NLS–YW) to investigate how male-female differences in the returns to education affect male-female educational attainment, found male and female schooling coefficients of 0.046 and 0.093 for whites and 0.050 and 0.075 for blacks. Unusually, she performs formal tests of the difference, finding it to be significant at the 0.1 percent level for whites and the 5 percent level for blacks.

Angle and Wissman (1981), using NLS–YM and NLS–YW data to 1975 to investigate the impact of choice of college major on the male-female earnings gap, and restricting the sample to those who had at least some college, found male and female schooling coefficients of 0.040 and 0.076, the latter estimated via a female*years of schooling interactive variable, whose coefficient was significant at the 5 percent level and perhaps higher.[4] Using dummy variables, they also find higher female coefficients for associate's and bachelor's degrees, but a lower one for a master's degree.

Rumberger and Daymont (1984), using National Longitudinal Survey of Youth 1979– (NLSY79) data for 1980 to investigate the impact of high school curriculum on labour market outcomes, found male and female schooling coefficients of 0.047 and 0.055. Their sample was restricted to those who had not completed a year of college and the respondents were very young (aged 18–22).

[4] Angle and Wissman do not report standard errors and consider only the 5 percent significance level.

Neumark (1988), using NLS–YM and NLS–YW data for 1980 to fit a theoretical Model of employers' discriminatory behaviour, found male and female schooling coefficients of 0.062 and 0.072.

Kane and Rouse (1995) replicated their NLS72 analysis using NLSY79 data for 1990, defining seven educational dummy variables, and found that, with the exception of that for 'other degree', all were higher for males.

Duncan (1996), using NLSY79 earnings data from all the rounds from 1979 to 1988, found male and female schooling coefficients of 0.032 and 0.067 for whites and 0.033 and 0.057 for blacks. He reports that the difference in the case of whites was significant at least at the 5 percent level. The difference was not tested for blacks.

Kane and Rouse's study is one of the few that does not follow the pattern of higher schooling coefficients for females and, as far as comparisons are possible, it conflicts with the findings of the present study. Mincer and Polachek's findings also appear to conflict, at least for married women, but the comparison may be affected by their use of different data sets for males and females. The findings of Angle and Wissman are remarkable because it was evident that they were wholly unanticipated by the authors. The introduction to their paper confidently asserts that 'it is safe to say' that the literature points to a higher schooling coefficient for males and they argue that discrimination in the labour market should be expected to have this effect.

Studies using other US data sets:

Oaxaca (1973), using data from the 1967 Survey of Economic Opportunity to illustrate his decomposition of the male-female earnings gap, fits two specifications, a 'personal characteristics' Model and a 'full-scale' Model where industry and occupation dummies were added. In both, schooling was included with a quadratic term. Confining attention to the former, he finds implicit coefficients for 12 years of schooling of 0.046 for white males and 0.015 for white females, the corresponding estimates for blacks being 0.007 and 0.016.

Malkiel and Malkiel (1973), using data from a single large firm, found a smaller schooling coefficient for females in 1969 with a narrow regression specification but larger coefficients for females in 1966, 1969, 1970, and 1971 with an expanded specification with further personal characteristics. Although widely cited, this was a very specialized study, the respondents being occupationally homogeneous (professionals with technical or scientific expertise requiring advanced training) and the sample very small (159 males and 113 females).

Rosenzweig (1976), using data from the 1 in 10,000 sample of the 1970 Census to illustrate how the use by Blinder (1973) of age instead of potential work experience as a control affects the decomposition of the male-female earnings gap, found male and female schooling coefficients of 0.078 and 0.116. King (1977), using data from one of the 1 in 100 public use samples of the same census to investigate the impact of occupational segregation on female experience-earnings profiles, found male and female schooling coefficients of 0.062 and 0.025. His sample was confined to those in professional occupations and the reported R^2 were very low, 0.017 for males and 0.012 for females.

Gerhart (1990), using data from a single large firm, 1976–1986, found all 6 current salary dummy coefficients higher for females, and 4 of 6 starting salary coefficients.

Hersch (1991), using data from eighteen firms in Eugene, Oregon, found male and female schooling coefficients of 0.041 and 0.056 in a conventional wage equation specification. Adding job characteristics, some of which were endogenous to schooling, reduced the coefficients to 0.030 and 0.041, respectively. Her sample was relatively small (631 respondents) and non-random.

Barron, Black, and Lowenstein (1993), using data from the 1982 (second) wave of the Employment Opportunity Pilot Project (EOPP) to investigate how much of the male-female earnings gap is attributable to females having relatively low labour force attachment and receiving relatively little on-the-job training, estimated wage equations for starting salaries and experienced wages. They found that males had higher coefficients for the three schooling dummy variables for starting wages and for two of the categories for experienced wages.

Card (1999), using March 1994–1996 CPS data to illustrate how the measure of earnings (hourly or annual) affects the estimates of the schooling coefficients, found (hourly) male-female schooling coefficients of 0.100 and 0.109. This was a very large sample and the difference, though small, was highly significant.

The studies summarized in this category are obviously more heterogeneous than those in the preceding ones and while there is some evidence of higher female coefficients, it is not strong. Three of the studies (Malkiel and Malkiel, 1973; King, 1977; and Gerhart, 1990) used the logarithm of annual, rather than hourly, earnings as the dependent variable and so the coefficients reflect the effect of schooling on hours worked as well as earnings per hour. As Card (1999) has shown, this is likely to inflate female schooling coefficients more than male ones.

Studies using non-US data

Given the institutional differences between the labour market in the US and labour markets in other countries, one would not anticipate that a feature of the US market would necessarily be found elsewhere. However, two recent surveys suggest that the phenomenon may not be confined to the US. Trostel, Walker, and Woolley (2002) estimate the returns to schooling in 28, mostly European, countries with data derived from a common survey instrument and found that the female schooling coefficient was higher in 24. Psacharopoulos and Patrinos (2002) list 95 estimates of male and females schooling coefficients from 49 countries at different dates. Of these 63 are greater for females, 3 equal, and 23 greater for males.

In conclusion, although there are exceptions, the majority of the studies lend support to the view that the schooling coefficient has been higher for females in the US, and seemingly elsewhere as well.

3. Possible Causes of the Effect

Sample selection bias

The starting point for the present discussion is an assumption that, in the context of a comparison of male and female earnings, the Mincerian wage equation (1) is only part of the story and needs supplementation. A common development in the literature is to treat the wage equation as the second stage of a two-stage Model of labour force participation and earnings. If labour force participants differ from non-participants, OLS estimates are likely to be inconsistent. Given that most males do participate, while many females do not, the strength of the bias could be different for the sexes and this might be a factor responsible for part of the difference in the OLS coefficients.

The first stage Models the decision to participate, relating the net benefit of participating, B^*, a latent variable, to a set of m variables Q_j and a random term e:

$$B_i^* = d_0 + \sum_{j=1}^{m} d_j Q_{ji} + e_i \qquad (5)$$

Potential earnings, Y^*, are given by the Mincerian relationship

$$Y_i^* = b_0 + \sum_{j=1}^{k} b_j X_{ji} + u_i \qquad (6)$$

with actual earnings, Y_i, equal to Y_i^* if $B_i^* > 0$ and Y_i not being observed if $B_i^* \leq 0$. It can be shown (Heckman, 1976) that

$$E(Y_i \mid B_i^* > 0) = E\left(Y_i \mid e_i > -d_0 - \sum_{j=1}^{m} d_j Q_{ji} \right) = b_0 + \sum_{j=1}^{k} b_j X_{ji} + \frac{s_{ue}}{s_e} l_i \qquad (7)$$

where s_{ue} is the population covariance of u and e, s_e is the standard deviation of e, and l_i, known as the inverse of Mill's ratio, is the ratio of the marginal density and cumulative functions of $e\ /s_e$?the standardized distribution of e, evaluated at e_i. It follows that a conventional wage regression that lacks the final term in (7) will be subject to omitted variable bias if the unobserved factors in the participation equation and the wage equation are related and s_{ue} is not equal to zero.

Despite this, even in recent years the wage equation studies that have taken account of potential sample selection bias are outnumbered by those that have not. The few that have reported the impact on the estimates of the regression coefficients have found it to be small. Thus Kenny et al (1979), using a sample of 1,373 white males from the Project Talent data set, found that allowing for selection bias and endogeneity reduced the years of college coefficient from 0.041 to 0.038. Heckman (1980), using a sample of 1,735 white, married women from the NLS–MW, found that allowing for selection bias increased the schooling coefficient from 0.076 to 0.078. Blau and Beller (1988), using CPS data for 1971 and 1981, found that allowing for sample selection bias has no effect on female coefficients in 1971 or male coefficients in 1981. However it increased the male schooling coefficient in 1971 from 0.065 to 0.069 and increased the female coefficient in 1981 from 0.054 to 0.061. Wellington (1993), using PSID data on white men and women in 1976 and 1985, did not find significant selectivity bias in the wage equations for either year and found minimal changes in the schooling coefficients when allowance was made for it.

Schooling and discrimination

A second extension to the basic Mincerian Model is to allow for the possibility that schooling may have two effects on earnings, at least for females: a direct human capital effect, and an indirect, anti-discrimination effect. It is possible that the impact of discrimination may not be uniform in the labour market and that, in particular, it may be inversely related to the level of schooling. There are two reasons for hypothesizing this. First, the better educated the individual, the more likely is he or she to have a degree or other formal qualification that would help to standardize wage offers regardless of sex. Second, the better educated a woman, the less likely is she to be tolerant of discrimination. Failing to allow for a negative association between discrimination and schooling could cause the female schooling coefficient to be overestimated. Similar arguments may be made with respect to that component of the unexplained earnings gap attributable to tastes or circumstances. The better educated a woman the more likely is she to be willing to seek employment outside the low-paying traditionally female occupations and the less likely is she to be competing for jobs where she would be penalized for a lesser ability to perform physically demanding tasks or for a greater aversion to poor working conditions or antisocial hours. At the same time, the better educated a woman and the greater her potential earnings, the more able she is to pay for child care and other services that allow her to seek a wage offer that fully values her characteristics. The impact of these factors may be inversely related to the level of schooling, and again, failing to allow for them could cause an upward bias in the estimated female schooling coefficient.

Quality of the schooling investment

A third possible explanation of the differential in the male-female schooling coefficients is that there may be a difference in the quality of male and female education. It is not suggested that there is any difference in quality as conventionally measured in terms of school resources, and in any case micro-level studies have failed to find significant effects of school resources on earnings, at least in recent times (Betts, 1995, 1996). However, if females tend to be more motivated students than males and extract more from their time in school, measuring schooling in terms of years of enrolment may mask systematic differentials in the quality of the schooling investment, and if the quality of the investment is correlated with years of enrolment, its omission from the regression specification could cause differential biases in the male and female schooling coefficients.

No previous study has addressed this issue directly, and only two have reported results that permit a secondary analysis. Grogger and Eide (1995), in their analysis of factors affecting changes in the college wage premium over time using NLS72 and HS&B data, present results that support the hypothesis that females do better in high school but cast doubt on the hypothesis that this contributes to their higher returns to years of schooling. Grogger and Eide include in their wage equations dummy variables for respondents reporting mostly As and Bs in high school, and for respondents reporting mostly Bs and Cs. In the NLS72 data set, 39 percent of females reported As and Bs, as opposed to only 19 percent of males. In the HS&B data set, the figures were 43 percent and 26 percent, respectively. Combining As and Bs with Bs and Cs, the figures are 87 percent for females and 70 percent for males in the NLS72 data set and 84 percent and 69 percent, respectively, in the HS&B data set. All the entry-level schooling coefficients are greater for females than for males, and all are reduced when the attainment dummy variables and test scores (scores on mathematics and vocabulary tests and a 'mosaic' test of perceptual speed and accuracy) are introduced. However, the differences in the male and female schooling coefficients are unaffected.

At the college level, Loury (1997), in her analysis of the impact of college major on the male-female earnings gap among college graduates, using the NLS72 and HS&B data sets, presents two sets of results for each: one with dummy variables for partial college and college graduate as the only schooling variables, the second adding the respondent's college grade point average, dummy variables for business, engineering and other majors, and a dummy variable for quality of college. In the NLS72 data set, the mean GPA was 2.76 for males and 3.03 for females, the difference being highly significant. In the HS&B data set, the corresponding figures were 2.81 and 2.89, respectively, the difference still being significant. However the impact of its introduction on the schooling coefficients is made unclear by the simultaneous inclusion of the college major dummy variables and the college quality variable. The introduction of the GPA and these variables causes the partial and complete college coefficients to fall for both males and females in both data sets. In the case of the NLS72 data set, the female premium for incomplete college is unaffected and that for complete college increases a little. In the case of the HS&B, the female discount for incomplete college becomes a premium, while the female premium for complete college decreases a little.

A number of studies have documented male-female differences in high school curriculum, measured in terms of semester hours or credits, and their impact on earnings,[5] but none has investigated the effect of introducing such controls on the schooling coefficients themselves.

At the college level, a large number of studies have investigated the contribution of the choice of college major to the male-female earnings gap but only two, in addition to Loury (1997), report wage equations with and without college major controls. Angle and Wissman (1981), using NLS–YM and NLS–YW data, report figures that show that introducing college majors as controls has no impact on the female schooling premium in the case of their omitted category, 'other' major fields and incomplete college, and actually increases the differential in the case of the other five categories. Grogger and Eide (1995) find that when they replace their college graduate dummy by six such dummies interacted with major field of study, the female premium increases for science majors and falls for engineering majors.

Occupational choice

A further possible reason for a differential in the male-female schooling coefficients is that females may be under-represented in jobs where schooling is a relatively unimportant factor in the determination of earnings. For example, they may be under-represented among union workers, where schooling is subordinated to seniority as a determinant of earnings, or in self-employment where entrepreneurial skills are relatively highly valued. An alternative is to take an occupational approach. There is a consensus in the literature that most of the male-female earnings gap is attributable to the tendency for women to be segregated in occupations with relatively low pay (Treiman and Hartmann, 1981; Cain, 1986; Gunderson, 1989; Chauvin and Ash, 1994; Altonji and Blank, 1999). In the case of the differential in the male-female schooling coefficients, it was hypothesized that there might also be an occupational effect in that the value of schooling may vary among occupations. The fact that 'female' occupations pay relatively poorly does not exclude the possibility that, within them, education is valued relatively highly. In particular, it was hypothesized that schooling is relatively unimportant in managerial and skilled manual occupations where females are under-represented.

[5] See Brown and Corcoran (1997) for a recent study of this kind. Earlier studies have mostly been contributions to the repeated evaluation of vocational education in the 1970s and the 1980s, the consensus being that vocational education does not have a direct impact on labour market outcomes (Rumberger and Daymont, 1984).

Measurement error and endogeneity of schooling and work experience

Male-female differentials in the endogeneity of schooling or work experience, or differentials in measurement error in them, could in principle account for part of the differential in the estimates of their schooling coefficients. These possibilities will not be pursued here. No study to date has examined the impact of possible differential endogeneity, although the differences in the distributions of male and female schooling suggest that it may exist. Differential endogeneity of work experience could also impact on the schooling coefficients. However, those studies that have estimated the impact of work experience endogeneity on female schooling coefficients have found it to be small.[6] With regard to measurement error, there is in general no reason to suppose that this affects estimates of male and female schooling differently. However, in the case of work experience, the female measure is likely to be subject to relatively large conceptual measurement error. Women have a greater propensity to interrupt their employment, and there is evidence that work experience prior to an interruption has less labour market value than experience since the most recent interruption (Corcoran, 1978). Accordingly the female experience coefficient may be subject to a relatively large downwards bias. If there is a negative correlation between schooling and work experience, a relatively large downwards bias in the female experience coefficient could in turn give rise to a relatively large downwards bias in the female schooling coefficient and as a consequence the extra return to female schooling would actually be underestimated.[7]

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[6] Mincer and Polachek (1974) found that allowing for endogentiy reduced the female schooling coefficient form 0.053 to 0.048. However, Sandell and Shapiro (1978), reworking their analysis found a smaller reduction, from 0.061 to 0.058. Heckman (1980), simultaneously allowing for selectivity, found that allowing for work experience endogeneity reduced the schooling coefficient from 0.078 to 0.076.

[7] In the model

$$Y = b_0 + b_1 X_1 + b_2 X_2 + u$$

where X_1 is subject to measurement error with expected value 0 and variance s_w^2, it can be shown that the limiting value of the OLS estimator of b_2 is

$$\text{plim}\,\hat{b}_2 = b_2 + \frac{b_1 s_w^2 s_{X_1X_2}}{s_{X_1}^2 s_{X_2}^2 - \left(s_{X_1X_2}\right)^2}$$

where $s_{X_1}^2$ and $s_{X_2}^2$ are the population variances of X_1 and X_2 and $s_{X_1X_2}$ i s their population covariance. Given that schooling and work experience tend to be two of the most important variables in wage equations, this relationship may be a guide to the behaviour of the schooling coefficient, despite the multiplicity of additional variables. If $s_{X_1X_2}$ is negative, the bias will be downwards. If work experience is subject to

4. Evidence from the National Longitudinal Survey of Youth 1979 –

Data and estimation method

The data set used for the present analysis is the NLSY 1979-, a panel study sponsored by the Bureau of Labor Statistics and managed by the Center for Human Resource Research at the Ohio State University. It consists of a nationally representative core sample of approximately 6,000 individuals aged 14–21 in 1979, the base year, and supplementary oversamples of minorities, poor whites, and those serving in the military. The survey was fielded annually until 1994 and since then it has been fielded biennially. The data used in the present analysis were taken from the core sample, the hourly earnings and other work variables as pooled data for the current or most recent job at the 1988, 1992, 1996, and 2000 interviews, with earnings being converted into 1996 constant dollars using the Urban Consumer Price Index. Observations were dropped if hourly earnings were less than $2.50 or more than $100, if the respondent was currently attending school, or if transcript data had not been collected or were incomplete.[8] Those working fewer than 30 hours per week were dropped from the wage equations. Table 2 presents summary statistics for the key variables. Altogether in the wage equations there were 10,182 observations relating to 3,527 individuals. To take account of the fact that there were multiple observations for most respondents, the model was fitted using random effects, the appropriate procedure in the case of a random sample from a large population (Baltagi, 2001; Hsiao, 1986).

Initial specification

Column 1 of Table 3 shows the result of an initial regression of the logarithm of hourly earnings on a female dummy variable, years of schooling, and a set of control variables. The latter comprised work experience and its square, tenure with the current employer and its square, dummy variables for black and hispanic ethnicity, a dummy variable for being married with spouse present, the arithmetic reasoning, word knowledge, paragraph comprehension, numerical operations and coding speed test scores from the Armed

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greater measurement error for females than for males, the differential in the male-female schooling coefficients will be underestimated.

[8] Following Rumberger and Daymont (1984), a transcript was deemed incomplete if it did not show at least three credits in each grade of high school attended.

Services Vocational Aptitude Battery, dummy variables for living in the country or on a farm when aged 14, a dummy variable for the purchase of magazines by anyone in the family when the respondent was aged 14, a dummy variable for living in the north-east, north-central, or west census regions, dummy variables for living in an urban area, and the local unemployment rate. All of the control variables were interacted with the female dummy variable.

The regression indicates years of schooling coefficients of 0.0509 for males and 0.0694 for females, respectively, the differential of 0.0185 being significant at the 1 percent level.

To investigate the stability of the differential, the female and years of schooling interactive term was replaced by triple interactive terms for the four sample period years. The differential declines from 1988 to 2000 but remains highly significant (Column 2).

Sample selection bias

Column 3 of Table 3 presents the results of re-estimating the Model allowing for selectivity. The explanatory variables in the probit regression comprised all of those in the wage equation with age, a dummy variable for having a child aged less than 6 in the household, and another dummy variable for having a child aged less than 16 but not less than 6 in the household, each with female interactive terms, added as identifying variables. The Model was fitted using maximum likelihood estimation with the inverse of Mill's ratio interacted with the female dummy variable. There is some evidence of selectivity for males, and more for females, the reduction in the negative coefficient of the female dummy variable indicating that to a large extent the latter reflects the impact of selectivity, rather than being female *per se*. The estimate of the schooling coefficient for males rises to 0.0553 and that for females rises to 0.0708. As a consequence the differential in the schooling coefficients is reduced to 0.0155. In line with most previous studies, the impact of adjusting for sample selection bias does not appear to be dramatic and as will be seen it attenuates with further changes to the specification of the model.

Variation in the impact of discrimination, tastes, and circumstances

To investigate the relationship between the impact of schooling on earnings and the impact of discrimination, tastes, and circumstances (henceforward DTC), a Oaxaca decomposition of the earnings gap was performed for each year of schooling, with those respondents with fewer than 11 years, or more than 17, being grouped into single categories. Table 4 shows the mean earnings of males and females by years of schooling, female earnings adjusted for differences in coefficients,[9] and the unexplained part of the decomposition attributed to DTC. The latter on the whole varies inversely with years of schooling as hypothesized, the greater the education of a female, the more willing or better placed she is to compete with males in the labour market.

Column 4 of Table 3 shows the results of fitting the Model with this index of DTC added to the regression specification. Column 5 additionally allows for selectivity. The index of DTC has a highly significant negative coefficient. A comparison of columns 1 and 4, or columns 3 and 5, suggests that failure to allow for variation in DTC has no effect on the male schooling coefficient but biases upwards the female coefficient by approximately a percentage point, accounting for the greater part of the differential in the schooling coefficients.

Quality of the schooling investment

Any measure of academic attainment is inherently arbitrary but for the present purposes the type of courses taken in high school, and the grades earned on those courses, will be taken as a proxy. High school transcripts were collected for most of the civilian NLSY respondents directly from their schools in a supplementary survey undertaken in three rounds over the period 1980–1983.[10] The information collected includes the names of the courses, letter grades and numbers of credits. Table 5 presents summary statistics for credits and grade points in mathematics, English, science, other academic subjects, and vocational subjects for the present sample, grade points being defined as credits weighted by grade with A being given a value of 4, B 3, etc.[11] The table shows that, apart from vocational courses, where females earned more credits than males, the distributions of

[9] To make the decompositions comparable with those in most other studies, male coefficients were used for valuing characteristics. The analysis was repeated using Reimers' decomposition instead (characteristics valued by the average of the male and female coefficients). The results, presented in Appendix B, are very similar.

[10] For details, see http://www.bls.gov/nls/79guide/1999/nls79g4b.pdf

[11] The courses were coded as follows: mathematics: 1101–1199; English: 501–557; science: 1501–1599; other academic 558–699 (language arts other than English), 601–699 (languages), 1501–1599 (social science); vocational 101–199 (agriculture), 401–499 (distributive trades), 701–799 (health), 901–999 (home economics), 1401–1499 (office), and 1601–1799 (technical and trades). The remaining courses in art, physical education, etc., were not coded.

credits were similar for the two sexes. However females earned uniformly higher average grades, and hence grade points, for the five subject areas. The correlation coefficients for grade points and years of schooling for the academic subject areas were in the range of 0.57 to 0.64 for males and 0.46 to 0.58 for females, implying that if academic grade points had an effect on earnings, omitting them would lead to an upwards bias in estimates of the returns to years of schooling. The vocational subject area had correlations of –0.12 and –0.14 for the two sexes.

The results of introducing the grade point variables, with interactive terms to allow for differences in their impact for males and females, are shown in Table 6. Apart from a positive effect for science for males significant at the 5 percent level, there is no evidence that academic attainment in any discipline impacts on earnings. Replacing grade points by raw credits does not alter this conclusion. As a consequence, the introduction of grade points has no systematic effect on the coefficients of the other variables.

These negative findings are reported because, as Altonji (1995) notes, the number of studies that have attempted to relate high school transcript data to labour market outcomes is relatively small. Apart from some early studies investigating the impact of vocational education, the NLSY transcript data, in particular, have been little exploited. A contributory factor to the neglect my be the fact that the unstructured coding of the NLSY transcript data must have presented a daunting task to analysts in the early years of the data set.

The present findings appear to be in line with the literature that exists. Altonji (1995), using transcript data from the NLS72, divided the curriculum into science, foreign languages, social studies, English, mathematics, industrial arts, commercial studies, and fine arts, and found that only foreign languages had a consistently significant effect across a variety of alternative specifications. He concludes that the estimated collective impact on earnings of the different subject areas cannot explain the impact of a year of high school on earnings. His specification combines males and females with a sex dummy variable. Brown and Corcoran (1997), using SIPP data, found a significant positive effect for geometry/trigonometry for males and languages for females, and a significant negative effect for industrial arts for females, for their subsample of high school graduates not college graduates. The effects disappear for college graduates, with college majors added as controls. Using NLS72 data, they found positive significant effects for languages and mathematics for males and languages for females, and a significant negative effect for commercial studies for females, for their subsample of high school graduates with no college. Again, the effects disappear for those with partial or complete college, with college majors added as controls. Earlier studies have mostly been contributions to the repeated evaluation of vocational education in the 1970s and the 1980s and have made a simple distinction between academic and vocational categories, the consensus being that vocational education does not have a direct impact on labour market outcomes (Rumberger and Daymont, 1984).

Job characteristics and occupational choice

The issue of the impact of job characteristics on the differential in the male-female schooling coefficients will be approached in two steps, first introducing two general job characteristics — class of worker and mode of pay determination — and then detailed occupation. All job characteristic variables are endogenous, but the endogeneity of the general job characteristics is usually ignored in practice. The class of worker categories are employment in the private sector (the omitted category), employment by government or a non-profit organization, and self-employment. The mode of pay determination categories are pay not determined by collective bargaining (the omitted category), and pay determined by collective bargaining. For linguistic convenience, these categories will be referred to as nonunion and union. Because unions for blue collar and white collar workers tend to have different characteristics, separate dummy variables were defined for them. Table 7 presents summary statistics.

It was hypothesized that part of the differential in the male-female schooling coefficients might be attributable to the returns to schooling being relatively low in categories where females are under-represented. In particular, it was anticipated that this might be the case for blue collar union workers, given the influence of skills and seniority in their pay determination, and the self-employed, where entrepreneurial skills are important. Accordingly, the class of worker and union dummy variables were interacted with the years of schooling variable in an initial specification. To allow for male-female differences, the job characteristic dummy variables and their interactives with schooling were further interacted with the female dummy variable in an expanded specification. Both specifications included all the previous control variables.

The resulting schooling coefficients, shown in Table 8, indicate that, in the initial specification (Columns 1 and, with allowance for selectivity, Column 3), compared with the nonunion private sector, returns to schooling are lower for government workers, the self-employed, and for union workers, and this could account for part of the male-female differential in the schooling coefficients found in previous specifications. Column 2 and, with allowance for selectivity, Column 4, present the results of the expanded version with the addition of the female interactive variables. For the private, nonunion sector, the schooling coefficient differential is negligible. Estimates for the other categories are erratic, reflecting the relatively small number of observations in them.

In the second step, the procedure used for the general job characteristics was extended to include occupations, with one difference. In the case of the job characteristics, there was a natural choice of omitted category, private sector nonunion employment. When occupations are included, there is no natural omitted category. Table 9 provides summary statistics on employment by occupation and sex. Table 10 shows the schooling coefficients for each occupation (Columns 1 and, allowing for selectivity, Column 3). Column 2 and, with allowance for selectivity, Column 4, present the results of adding the female interactive variables.

The initial speculation that part of the differential in the male-female schooling coefficients could be due to females being underrepresented in management, where schooling might be relatively unimportant, is not borne out since the under-representation of females is minor and the schooling coefficient is actually relatively large. The likely explanation is that administrators rather than entrepreneurs dominate this employment category, at least in the NLSY cohort. However the other initial hypothesis appears to be correct. Females are underrepresented among the five categories of manual worker, which account for 44 percent of male employment and only 10 percent of female employment, and the returns to schooling in these occupations are indeed low. Using the coefficients in Column 1, the weighted average of the return to schooling across occupation is 0.042 for males and 0.051 for females. Thus it appears that occupational segregation can account for nearly a percentage point of the differential in the male-female schooling coefficients. Introducing female interaciton terms does not alter this conclusion. The estimate for males is the same and that for females is 0.049.[12] Allowing for selectivity has very little effect.

[12] The F statistic for the joint explanatory power of the occupation-schooling-sex triple interactives, 1.83, is just significant at the 5 percent level (critical value 1.80). However this appears to be attributable to the heterogeneity of the service category.

5. Conclusions

The survey in Section 2 found that estimates of the returns to schooling in the US tend to be higher for females than for males, despite the fact that females tend to earn less, both absolutely and controlling for personal characteristics. Certainly this is so in the case of the NLSY cohort, where the initial estimate of the differential in the male-female schooling coefficients was 0.0185. This may actually have been an underestimate of the extra return to schooling of females, given the likelihood of a downward bias imparted by measurement error in their work experience.

The differential in the schooling coefficients falls to 0.093 with the introduction of the index of discrimination, tastes, and circumstances, suggesting that a minor but important side-benefit of schooling for females is that it makes them more able and willing to compete with males in the labour market.

The hypothesis that part of the differential in the male-female schooling coefficients could be attributable to male-female differences in the quality of educational attainment, as proxied by their academic outcomes in high school, was not sustained. The NLSY females did indeed perform better than the males, but there was virtually no association between academic attainment and earnings and allowing for it made little difference to the estimate of the differential in the schooling coefficients.

The effect of occupational choice was investigated in two stages. The introduction of sector of employment and collective bargaining variables virtually eliminated the differential in the male-female schooling coefficients for the private sector. When occupational categories were added to the specification, it was found that part of the differential in the male-female schooling coefficients is attributable to occupational segregation, males being disproportionately attracted to manual occupations where schooling is poorly valued and females being disproportionately attracted to professional ones where it has a high return.

Initially it appeared that to a minor extent part of the differential in the male-female schooling coefficients might be explained by selection bias. However, after the index of discrimination, tastes, and circumstances and the job characteristics variables had been added to the specification, allowing for selectivity made very little difference to the schooling coefficients.

Table 1
US studies with male and female schooling coefficients

Study	Data	DV	Controls	Findings
Altonji (1993)	NLS72 1977–1986 *n* = 38,595 (no M/F breakdown)	H	we, we^2, fb, eth, abil, reg	Higher female coefficients for both partial college dummy variables, all 8 college degrees, and 5 of the 6 advanced degrees.
Angle and Wissman (1981)	NLS Young Men and Young Women. M 2,831, F 1,677	H	age, fb, eth	M 0.040, F 0.076. Female BA and MA dummy variable coefficients higher but PhD coefficient lower. Sample restricted to respondents who had at least some college.
Barron, Black, and Lowenstein (1993)	EOPP 1982 M 683 F 578	H	pwe, tr, part-time, empsize	Dummy variable coefficients lower for females for three categories, starting wages, lower for two categories, experienced wages.
Blau and Kahn (1997)	PSID 1980 and 1989 1980: M 1,784, F 1,081. 1989: M 1,591, F 1,149	H	we, we^2, eth	1980: M 0.066, F 0.084, female college and advanced degree dummy coefficients also higher. 1989: M 0.090, F 0.083, female college and advanced degree coefficients much higher.
Brown and Corcoran (1997)	SIPP for 1984 and NLS72 1986 SIPP: subsamples from M 8,695, F 7,171 NLS72 subsamples from M 2,635, F 2,359	H	SIPP: we, we^2, ten, ten^2, tr, eth, mar, child, hsc, reg. NLS72: we, ten, tr, eth, mar, abil.	SIPP: females have higher coefficients for college and postgraduate years of schooling. Lower coefficient for high school years of schooling. NLS72: partial college years-of-schooling coefficient higher for females.
Card (1999)	CPS March 1994–1996. M 102,639, F 95,309	H	pwe, pwe^2, pwe^3, eth	M 0.100, F 0.109
Corcoran and Duncan (1979)	PSID 1977. Whites M 2,250, F 1,326. Blacks M 895, F 741	H	we, we^2, ten, attach, reg	Whites: M 0.059, F 0.077. Blacks: M 0.061, F 0.076
Daymont and Andrisani (1984)	NLS72 1979 M 1,482, F 1,353	H	we, weeks, mar, prefer	Relative to college graduates, females have higher coefficients for master's and PhD.
Duncan (1996)	NLSY 1979-1988 M 34,333, F 30,578	H	we, we^2, mar, reg, urban, hours per week	Whites: M 0.032, F 0.067 Blacks: M0.033, F 0.057
Gerhart (1990)	Single large firm, 1976–1986. M 3,564, F 1,053	A	pwe, pwe^2, perform, ten	For females, all 6 dummy coefficients higher for current salary regressions, and 4 of 6 coefficients higher for starting salary regressions.
Grogger and Eide (1995)	NLS72 1977–1986 and HS&B 1986 (pooled). M 19,597 F 16,223.	H	we, eth, fb, abil, hsc.	Relative to high school graduates, females have higher coefficients for partial college, college graduates, and postgraduate degree, the differentials falling with experience.
Hersch (1991)	Sample from 18 firms in Eugene, Oregon M 414, F 217	H	we, we^2, ten, ten^2, eth, mar, child	M 0.041, F 0.056
Kane and Rouse (1995)	NLS72 1986 and NLSY 1990. NLS72: M 3,249, F 3,514. NLSY: M 2,271, F 2,277	H	NLS72: we, we^2, eth, fb, abil, reg. NLSY: we, age, eth, fb, abil, reg.	NLS72: 2-year college M 0.042, F 0.064; 4-year college M 0.046, F 0.062. NLSY: lower dummy variable coefficients for females for all educational categories except other degree.
King (1977)	1970 Census 1 in 100 M 4,253, F 4,483	A	pwe, eth, weeks	M 0.062, F 0.025. Sample confined to those in professional occupations.

Loury (1997)	NLS72 1979, HS&B 1986. NLS72: M 1,384, F 1,184. HS&B: M 732, F 915	W	Both data sets: ten, weeks, mar, union, hsc	Higher coefficients for females for partial and complete college, both data sets.
Madden (1978)	NLS Young Men and Young Women. Whites: M 1,074, F 1,473; Blacks M 453, F 583	H	we, ten, eth, fb, reg, abil, mar, weeks	Whites: M 0.046, F 0.093; Blacks M 0.050, F 0.075
Malkiel and Malkiel (1973)	Single large firm. M 159, F 113 1966-1971	A	Narrow: we, we^2 Expanded: we, we^2, PhD, mar, pub	Narrow specification, 1969: M 0.091, F 0.078. Expanded specification, 1966 M 0.033, F 0.059; 1969 M 0.041, F 0.043; 1970 M 0.042 F 0.046; 1971 M 0.036 F 0.044, with lower coefficients for females for a PhD dummy variable.
Mincer and Polachek (1974)	NLS Mature Women and SEO, both for 1966 (*n* not stated).	H	Married M: we, we^2. Married F: estimated we, ten, home time. Single F: we, we^2, ten	Married M 0.071, married F 0.063, single F 0.077
Murnane, Willett, and Levy (1995)	NLS72, 1978 and HS&B, 1986. NLS72 M 4,114, F 3,925. HS&B M 1,980, F 2,163	H	we, eth, fb, part-time	NLS72: M 0.013, F 0.037 HS&B: M 0.021, F 0.037
Neumark (1988)	NLS Young Men and Young Women, 1980 M 1,819, F 1,505	H	we, age, eth, mar, reg, urban, union	M 0.062, F 0.072
Oaxaca (1973)	SEO 1967. Whites: M 8,123, F 4,962. Blacks: M 3,897, F 3,502	H	pwe, pwe^2, mar, urban, reg, part-time	Used a quadratic for years-of-schooling Whites: females have lower implicit coefficients Blacks: females have higher implicit coefficients.
Rosenzweig (1976).	1970 Census, 1 in 10,000 sample M 3,251, F 375	H	pwe, pwe^2	M 0.078, F 0.116
Rumberger and Daymont (1984)	NLSY79, 1980 M 713, F 648	H	pwe, eth, mar, abil, child	M 0.047, F 0.055. Sample restricted to those who did not complete a year of college.
Wellington (1993)	PSID 1976, 1985. 1976: M 1,535 F 1,002. 1985: M 1,901 F 1,544	H	we, we^2, ten, tr, home time, attach, reg, urban	1976: M 0.049, F 0.074 1985: M 0.062, F 0.079 Whites only.

Notes:

Data: M male, F female, with numbers of observations. For data set abbreviations, see section 3.

DV (dependent variable): H hourly wage, W weekly earnings, A annual earnings, in each case logarithmic.

Controls: we: work experience; pwe: potential work experience; ten: tenure; tr: training; fb: family background; eth: ethnicity; mar: married; child: number of children; abil: ability; hsc: high school courses or grades; reg: region; urban: resides in urban area, or SMSA, or city size; weeks: weeks worked in the year; attach: indicator of labour market attachment; prefer: respondent's preferences for job characteristics; perform: job performance; pub: significant publications record; part-time: part-time job, union: wage determined by collective bargaining; empsize: employer size.

Findings: M male, F female, numbers refer to years-of-schooling coefficients.

Table 2
Summary statistics

	males	females
Mean years of schooling	13.53	13.80
Percent working full-time	87.0	63.8
Mean of logarithm of hourly earnings	2.62	2.40
Geometric mean of hourly earnings	13.74	10.97
Number of observations	6,477	7,133

Table 3
Wage equations, dependent variable logarithm of hourly earnings

	(1)	(2)	(3)	(4)	(5)
Female	–0.4530** (0.1073)	–0.5311** (0.1086)	–0.2369* (0.1170)	–0.2572* (0.1228)	–0.0150 (0.1361)
Schooling	0.0509** (0.0041)	0.0509** (0.0041)	0.0553** (0.0047)	0.0509** (0.0041)	0.0554** (0.0047)
Schooling* female	0.0185** (0.0060)	–	0.0155* (0.0063)	0.0093 (0.0066)	0.0054 (0.0070)
Schooling* female*1988	–	0.0267** (0.0061)	–	–	–
Schooling* female*1992	–	0.0211** (0.0060)	–	–	–
Schooling* female*1996	–	0.0178** (0.0060)	–	–	–
Schooling* female*2000	–	0.0168** (0.0060)	–	–	–
Index of DCT ('discrimination')	–	–	–	–0.3898** (0.1188)	–0.4584** (0.1550)
Inverse of Mill's ratio	–	–	–0.0852* (0.0364)	–	–0.0872* (0.0353)
IMR*female	–	–	–0.1579** (0.0551)	–	–0.1533** (0.0529)
R^2	0.3895	0.3924	–	0.3905	–
c^2	–	–	4.92*	–	5.95*
n	10,182	10,182	13,610	10,182	13,610

*, ** significant at the 5 and 1 percent levels. Standard errors in parentheses. For controls, see text.

Table 4
Logarithm of hourly earnings by years of schooling

years of schooling	males	females	females, adjusted	difference	*n*, males	*n*, females
<11	2.2315	2.0191	2.2380	0.2189	219	99
11	2.2861	2.0239	2.3243	0.3005	186	56
12	2.4867	2.2154	2.4697	0.2543	2,524	1,914
13	2.6272	2.2872	2.5542	0.2669	474	414
14	2.6441	2.4153	2.6076	0.1923	503	466
15	2.6469	2.4450	2.6582	0.2131	192	253
16	2.8699	2.6652	2.8446	0.1794	996	856
17	2.8820	2.6355	2.7827	0.1472	160	186
18	3.1000	2.7577	3.0568	0.2991	192	175
>18	3.0594	3.0583	2.9135	–0.1448	188	129

Table 5
High school credits and grade averages

	males			females		
	credits	average grade	grade points	credits	average grade	grade points
mathematics	2.35	2.45	5.75	2.20	2.59	5.72
English	3.25	2.32	7.55	3.31	2.71	8.97
science	1.87	2.49	4.65	1.81	2.64	4.77
other academic	4.01	2.53	10.12	4.25	2.78	11.79
vocational	2.12	2.68	5.68	3.46	2.84	9.81
n	1,664	1,664	1,664	1,816	1,816	1,816

Table 6
Wage equations with curriculum variables

	(1)	(2)
Female	–0.1941	–0.0131
	(0.1074)	(0.1460)
Schooling	0.0463**	0.0530**
	(0.0048)	(0.0056)
Schooling*female	0.0098	0.0036
	(0.0073)	(0.0080)
Discrimination	–0.4007**	–0.4654**
	(0.1188)	(0.1539)
Math	0.0007	–0.0005
	(0.0032)	(0.0033)
Math*female	0.0064	0.0073
	(0.0046)	(0.0046)
English	–0.0058	–0.0063
	(0.0032)	(0.0034)
English*female	0.0029	0.0032
	(0.0044)	(0.0044)
Science	0.0066*	0.0073*
	(0.0032)	(0.0034)
Science*female	–0.0048	–0.0058
	(0.0045)	(0.0046)
Other academic	0.0028	0.0014
	(0.0018)	(0.0018)
Other academic* female	–0.0013	–0.0000
	(0.0025)	(0.0025)
Vocational	–0.0021	–0.0022
	(0.0013)	(0.0014)
Vocational*femalee	0.0024	0.0022
	(0.0018)	(0.0018)
IMR	–	–0.0959**
		(0.0329)
IMR*female	–	–0.1385**
		(0.0514)
R^2	0.3928	–
c^2	–	8.22**
n	10,182	13,610

*, ** significant at the 5 and 1 percent levels.
Standard errors in parentheses. For controls, see text.

Table 7
Summary statistics, type of employment

	proportion female	proportion of sample	proportion of males	proportion of females
Government worker	0.584	0.162	0.121	0.211
Self-employed	0.283	0.061	0.080	0.039
White-collar collective bargaining	0.470	0.109	0.105	0.115
Blue-collar collective bargaining	0.260	0.062	0.083	0.036

Table 8
Schooling coefficients from wage equations, dependent variable logarithm of hourly earnings

	(1)	———— (2) ————		(3)	———— (4) ————	
	Schooling coefficients	Schooling coefficients	Female interactives	Schooling coefficients	Schooling coefficients	Female interactives
Private, employed	0.0598**	0.0587**	0.0029	0.0703	0.0698**	0.0029
	(0.0039)	(0.0050)	(0.0079)	(0.0044)	(0.0059)	(0.0088)
Government	–0.0110*	–0.0177*	0.0104	–0.0194**	–0.0299**	0.0171
	(0.0051)	(0.0074)	(0.0102)	(0.0063)	(0.0098)	(0.0125)
Self-employed	–0.0176**	–0.0142*	–0.0105	–0.0325*	–0.0289**	–0.0125
	(0.0069)	(0.0084)	(0.0149)	(0.0133)	(0.0164)	(0.0282)
Collective (white collar)	–0.0256**	–0.0358**	0.0204	–0.0324**	–0.0463**	0.0209
	(0.0054)	(0.0076)	(0.0109)	(0.0066)	(0.0098)	(0.0128)
Collective (blue collar)	–0.0114	–0.0197	0.0239	–0.0364**	–0.0381**	0.0073
	(0.0103)	(0.0125)	(0.0221)	(0.0118)	(0.0142)	(0.0257)
R^2	0.4099	0.4106		–	–	
c^2	–	–		5.09*	3.47	
n	10,182	10,182		13,610	13,610	

*, ** significant at the 5 and 1 percent levels. Standard errors in parentheses. For controls, see text.

Table 9
Summary statistics, occupation of employment

	proportion female	proportion of sample	proportion of males	proportion of females
Professionals	0.624	0.181	0.135	0.228
Technicians	0.406	0.044	0.051	0.036
Managers	0.431	0.153	0.173	0.133
Sales workers	0.467	0.052	0.054	0.049
Clerical workers	0.822	0.177	0.062	0.293
Skilled workers	0.089	0.112	0.203	0.020
Operatives	0.367	0.071	0.088	0.052
Transport workers	0.131	0.035	0.060	0.009
Labourers	0.192	0.047	0.075	0.018
Farm workers	0.185	0.010	0.016	0.004
Service workers	0.655	0.119	0.081	0.157

Table 10

Schooling coefficients from wage equations, dependent variable logarithm of hourly earnings

	(1)	(2)		(3)	(4)	
	schooling coefficient	schooling coefficient	female interaction	schooling coefficient	schooling coefficient	female interaction
Professionals	0.0683**	0.0666**	0.0038	0.0680**	0.0667**	0.0027
	(0.0071)	(0.0104)	(0.0143)	(0.0071)	(0.0103)	(0.0142)
Technicians	0.0603**	0.0573**	0.0070	0.0627**	0.0576**	0.0111
	(0.0114)	(0.0154)	(0.0226)	(0.0113)	(0.0154)	(0.0224)
Managers	0.0769**	0.0733**	0.0097	0.0768**	0.0733**	0.0094
	(0.0064)	(0.0083)	(0.0131)	(0.0064)	(0.0083)	(0.0129)
Sales workers	0.1066**	0.1041**	0.0060	0.1020**	0.1041**	–0.0136
	(0.0125)	(0.0166)	(0.0251)	(0.0125)	(0.0166)	(0.0260)
Clerical workers	0.0414**	0.0462**	–0.0060	0.0456**	0.0465**	–0.0018
	(0.0070)	(0.0139)	(0.0161)	(0.0071)	(0.0138)	(0.0161)
Skilled workers	0.0275**	0.0229**	0.0344	0.0286**	0.0232*	0.0314
	(0.0093)	(0.0103)	(0.0212)	(0.0093)	(0.0103)	(0.0211)
Operatives	0.0004	0.0065	–0.0167	0.0026	0.0067	–0.0144
	(0.0099)	(0.0128)	(0.0198)	(0.0099)	(0.0128)	(0.0200)
Transport workers	–0.0073	–0.0156	0.1134	–0.0058	–0.0153	0.1196*
	(0.0235)	(0.0245)	(0.0556)	(0.0234)	(0.0244)	(0.0573)
Labourers	0.0062	0.0065	0.0123	0.0071	0.0066	0.0329
	(0.0138)	(0.0150)	(0.0406)	(0.0138)	(0.0150)	(0.0426)
Farm workers	0.0057	0.0056	–0.0037	0.0050	0.0055	–0.0382
	(0.0303)	(0.0308)	(0.1639)	(0.0300)	(0.0305)	(0.1664)
Service workers	0.0316**	0.0556**	–0.0522**	0.0335**	0.0558**	–0.0498*
	(0.0110)	(0.0139)	(0.0224)	(0.0109)	(0.0138)	(0.0224)
R^2	0.9774	0.9775			–	
c^2	–	–		0.16	0.17	
n	10,174	10,174		12,253	12,523	

*, ** significant at the 5 and 1 percent levels. Standard errors in parentheses. For controls, see text.

Appendix A: Information on the Data Sets

Current Population Surveys (CPS)

(http://www.bls.census.gov/cps/)

The CPS, a household survey with monthly interviews of about 50,000 households, is sponsored jointly by the Census Bureau and the Bureau of Labor Statistics. It has a panel element of short duration, housing units being interviewed for four months, left alone for eight months, and then interviewed for another four months. Detailed labour market information is collected by its March Supplement, the CPS Annual Demographic Survey, for which the sample size is increased, to about 65,000 prior to 2001 and 100,000 currently. The survey has the advantage of a larger sample size than other surveys used for labour market analysis but the disadvantage for wage equations that there is no direct measure of work experience.

National Longitudinal Study of the High School Class of 1972(NLS72) and High School and Beyond (HS&B)

(http://nces.ed.gov/surveys/nls72/; http://nces.ed.gov/surveys/hsb/)

The NLS72, conducted by the National Center of Educational Statistics (NCES) of the US Department of Education, was a panel study of a sample of 22,652 individuals originally interviewed while high school seniors in 1972. Follow-up surveys were undertaken in 1973, 1974, 1976, 1979 and, of a subsample of 12,841, in 1986. HS&B, also conducted by the NCES, comprises cohorts of high school sophomores in 1980 followed up in 1982, 1984, and 1986, and, in the case of the sophomore cohort, 1992. The attractions of these data sets for fitting wage equations are the wealth of information concerning high school courses and scores for tests of mathematics, vocabulary, and perception (the 'mosaic' test.) In the case of NLS72, further educational data are available from the survey of postsecondary transcripts undertaken in 1984.

National Longitudinal Surveys of Labor Market Experience

(http://www.bls.gov/nls/)

Sponsored by the Department of Labor, the National Longitudinal Surveys of Labor Market Experience consist of a set of six panel studies: the NLS of Young Men, 5225 men aged 14–24 interviewed from 1966 to 1981; the NLS of Older Men, 5,020 men aged 45–59 interviewed from 1966 to 1990; the NLS of Mature Women, 5,083 women aged 30–44 interviewed from 1967 until the present; the NLS of Young Women, 5,159 women aged 14–24 interviewed from 1968 to the present, the NLS of Youth 1979 (1979–present), and the NLS of Youth 1997 (1997–present). The data sets have the advantage of being nationally representative of their cohorts (in some cases with supplementary oversamples of minorities) and of being rich in data relevant to the fitting of wage equations. They have the disadvantage of being cohort-specific.

Panel Study of Income Dynamics (PSID)

(http://www.isr.umich.edu/src/psid/)

The PSID is an ongoing longitudinal survey of a nationally representative sample of US households, the number growing from 4,800 at its inception in 1968 to more than 7,000 in 2001. The study includes a supplementary oversample of low-income families. The core sample is not a random sample of the adult working population because labour market information is collected only from heads of households and, for male heads, their wives. Its inclusion of data on actual work experience is an attractive feature for fitting wage equations. For a comparison of PSID and CPS, see the appendix to Blau and Kahn (1997).

Survey of Economic Opportunity (SEO)

(no website; relationship with PSID explained at

http://www.isr.umich.edu/src/psid/overview.html#Overview; data link at

http://www.sscnet.ucla.edu/issr/da/index/techinfo/m2031.htm)

The SEO, sponsored by the then U.S. Office of Economic Opportunity with two waves in 1966 and 1967, was a nationally representative survey of some 30,000 households with a supplementary sample of low-income households. The PSID over-sample of low-income households was drawn from the SEO.

Employment Opportunities Pilot Project

(no website)

Sponsored by the National Institute of Education and the National Center for Research in Vocational Education, this was a two-wave longitudinal survey of employers in 31 counties in the US. It was designed to oversample low-income workers.

Survey of Income and Program Participation (SIPP)

(www.sipp.census.gov/sipp/)

Established by the Census Bureau in 1983, the SIPP is a nationally representative household survey, currently with about 40,000 households, with a strong panel element and interviews at four-monthly intervals. It is intended to collect data on income, labour market participation, and participation in government programs, one objective being to monitor the effectiveness of the latter and to anticipate their future cost.

Appendix B: Summary of Results Using Reimers' Decomposition

The table shows the schooling coefficients corresponding to those reported in Tables 3, 6, and 8 using an index of discrimination, tastes, and circumstance using Reimer's decomposition of the male-female earnings gap.

	Schooling	Schooling* female	Index of discrimination
Table 3, Column 1	0.0509** (0.0041)	0.0185** (0.0060)	–
Table 3, Column 3	0.0553** (0.0047)	0.0155* (0.0063)	–
Table 3, Column 4	0.0509** (0.0041)	0.0106 (0.0064)	–0.4801** (0.1329)
Table 3, Column 5	0.0554** (0.0047)	0.0081 (0.0066)	–0.4732** (0.1646)
Table 6, Column 1	0.0463** (0.0048)	0.0112 (0.0071)	–0.4918** (0.1330)
Table 6, Column 2	0.0530** (0.0056)	0.0062 (0.0078)	–0.4833** (0.1643)
Table 8, Column 2	0.0587** (0.0050)	0.0043 (0.0077)	–0.5022** (0.1320)
Table 8, Column 4 (nonunion private sector)	0.0698** (0.0059)	0.0055 (0.0086)	–0.4094* (0.1617)

References

Altonji, J.G. (1993), 'The Demand for and Return to Education when Education Outcomes are Uncertain', Journal of Labor Economics, 11(1, part 1): 48–83.

Altonji, J.G. (1995), 'The Effects of High School Curriculum on Education and Labor Market Outcomes', Journal of Human Resources, 30(3): 409–438.

Altonji, J.G. and Blank, R.M. (1999), 'Race and Gender in the Labor Market', in O. Ashenfelter and D. Card (eds.), *Handbook of Labor Economics*, Vol 3, Amsterdam: Elsevier, 3143–3259.

Angle, J. and Wissman, D.A. (1981), 'Gender, College Major, and Earnings', Sociology of Education, 54: 25–33.

Baltagi, B.H. (2001), *Econometric Analysis of Panel Data* (second edition), Chichester, UK: John Wiley.

Barron, J.M., Black, D.A. and Loewenstein, M.A. (1993), 'Gender Differences in Training, Capital, and Wages', Journal of Human Resources, 28(2): 343–364.

Betts, J.R. (1995), 'Does School Quality Matter? Evidence from the National Longitudinal Survey of Youth', Review of Economics and Statistics, 77(2), 231–250.

Betts, J.R. (1996), 'Is there a Link between School Inputs and Earnings? Fresh Scrutiny of an Old Literature', in G. Burtless (ed.), *Does Money Matter? The Effect of School Resources on Student Achievement and Adult Success*, Washington DC: Brookings Institution.

Blau, F.D. and Beller, A.H. (1988), 'Trends in Earnings Differentials by Gender', Industrial and Labor Relations Review, 41(4): 513–529.

Blau, F.D. and Kahn, L.M. (1992), 'Race and Gender Pay Differentials', in D. Lewin, O.S. Mitchell, and P.D. Sherer (eds.), *Research Frontiers in Industrial Relations and Human Resources*, Madison, WI: Industrial Relations Research Association.

Blau, F.D. and L.M. Kahn (1997), 'Swimming Upstream: Trends in the Gender Wage Differential in the 1980s', Journal of Labor Economics, 15(1, part 1): 1–42.

Blau, F.D. and Kahn, L.M. (2000), 'Gender Differences in Pay', Journal of Economic Perspectives, 14(4), 75–99.

Blinder, A.S. (1973), 'Wage Discrimination: Reduced Form and Structural Estimates', Journal of Human Resources, 8(4), 436–455.

Brown, C. and Corcoran, M. (1997), 'Sex-Based Differences in School Content and the Male-Female Wage Gap', Journal of Labor Economics, 15(3, part 1): 431–465.

Cain, G.G. (1986), 'The Economic Analysis of Labor Market Discrimination: a Survey', in O. Ashenfelter and R. Layard (eds.), *Handbook of Labor Economics, Vol 1*, Amsterdam: Elsevier Science Publishers.

Card, D. (1999), 'The Causal Effect of Education on Earnings', in O. Ashenfelter and D. Card (eds) *Handbook of Labor Economics*, *Volume 3*, New York: Elsevier.

Chauvin, K.W. and Ash, R.A. (1994), 'Gender Earnings Differentials in Total Pay, Base Pay, and Contingent Pay', Industrial and Labor Relations Review, 47(4), 634–649.

Cohen, M.S. (1971), 'Sex Differences in Compensation', Journal of Human Resources, 6(4): 434–447.

Corcoran, M. (1978), 'Work Experience, Work Interruption, and Wages', in G.J. Duncan and J.N. Morgan (eds.), *Five Thousand American Families*: *Patterns of Economic Progress*, *Volume 6*, Ann Arbor: University of Michigan, Institute for Social Research.

Corcoran, M. and Duncan, G.J. (1979), 'Work History, Labor Force Attachment, and Earnings Differences between the Races and Sexes, Journal of Human Resources, 14(1): 3–20.

Cotton, J. (1988), 'On the Decomposition of Wage Differentials', Review of Economics and Statistics, 70(2): 236–243.

Daymont, T.N. and Andrisani, P.J. (1984), 'Job Preferences, College Major, and the Gender Gap in Earnings', Journal of Human Resources, 19(3): 408–428.

Dougherty, C. and Jimenez, E. (1991), 'The Specification of Earnings Functions: Tests and Implications', Economics of Education Review, 10(2): 85–98.

Duncan, K.C. (1996), 'Gender Differences in the Effect of Education on the Slope of the Experience-Earnings Profiles', American Journal of Economics and Sociology, 55(4), 457–471.

Featherman, D.L. and Hauser, R.M. (1976), 'Sexual Inequalities and Socioeconomic Achievement in the U.S., 1962–1973', American Sociological Review, 41(3): 462–483.

Gerhart, B. (1990), 'Gender Differences in Current and Starting Salaries: the Role of Performance, College Major, and Job Title', Industrial and Labor Relations Review, 43(4): 418–433.

Grogger, J. and Eide, E. (1995), 'Changes in College Skills and the Rise in the College Wage Premium', Journal of Human Resources, 30(2): 280–310

Grubb, W.N. (1993), 'The Varied Economic Returns to Postsecondary Education', Journal of Human Resources, 28(2): 365–382.

Gunderson, M. (1989), 'Male-Female Wage Differentials and Policy Responses', Journal of Economic Literature, 27(1): 46–72.

Heckman, J.J. (1976), 'The Common Structure of Statistical Models of Truncation, Sample Selection and Limited Dependent Variables, and a Simple Estimator for Such Models', Annals of Economic and Social Measurement, 5(4): 475-492.

Heckman, J.J. (1980), 'Sample Selection Bias as a Specification Error with an Application to the Estimation of Labor Supply Functions', in J.P. Smith (ed.), *Female Labor Supply: Theory and Estimation*, Princeton, NJ: Princeton University Press.

Heckman, J.J. and Polachek, S. (1974), 'Empirical Evidence on the Functional Form of the Earnings-Schooling Relationship', Journal of the American Statistical Association, 69(346): 350–354.

Hersch, J. (1991), 'Male-Female Differences in Hourly Wages: the Role of Human Capital, Working Conditions, and Housework', Industrial and Labor Relations Review, 44(4), 746–759.

Hsiao, C. (1986), *Analysis of Panel Data*, Cambridge: Cambridge University Press.

Jones, F.L. (1983), 'On Decomposing the Wage Gap: a Critical Comment on Blinder's Method', Journal of Human Resources, 18(1): 126–130.

Kane, T.J. and Rouse, C.E. (1995), 'Labor-Market Returns to Two- and Four-Year College', American Economic Review, 85(3): 600–614.

Kenny, L.W., Lee, L.-F., Maddala, G.S. and Trost, R.P. (1979), 'Returns to College Education: an Investigation of Self-Selection Bias Based on the Project Talent Data', International Economic Review, 20(3): 775–789.

King, A.G. (1977), 'Is Occupational Segregation the Cause of the Flatter Experience-Earnings Profiles of Women?', Journal of Human Resources, 12(4): 541–549.

Lloyd, C. and Neimi, B. (1979), *The Economics of Sex Differentials*, New York: Columbia University Press.

Loury, L.D. (1990), 'Effects of Cohort Size on Postsecondary Training', in G. Burtless (ed.) *A Future of Lousy Jobs? The Changing Structure of U.S. Wages*, Washington, DC: Brookings Institution.

Loury, L.D. (1997), 'The Gender Earnings Gap among College-Educated Workers', Industrial and Labor Relations Review, 50(4): 580–593.

Madden, J.F. (1978), 'Economic Rationale for Sex Differences in Education', Southern Economic Journal, 44(4): 778–797.

Madden, J.F. (1985), 'The Persistence of Pay Differentials: the Economics of Sex Discrimination', in L. Larwood, A. Stromberg and B. Gutek (eds.), *Women and Work: an Annual Review*, Beverley Hills, CA: Sage.

Malkiel, B.G. and Malkiel, J.A. (1973), 'Male-Female Pay Differentials in Professional Employment', American Economic Review, 63(4): 693–705.

Mincer, J. and Polachek, S. (1974), 'Family Investments in Human Capital: Earnings of Women', Journal of Political Economy, 82(2, part 2): S76–S108.

Murnane, R.J., Willett, J.B. and Levy, F. (1995), 'The Growing Importance of Cognitive Skills in Wage Determination', Review of Economics and Statistics, 77(2): 251–266.

Neumark, D. (1988), 'Employers' Discriminatory Behavior and the Estimation of Wage Discrimination', Journal of Human Resources, 23(3): 279–295.

Oaxaca, R.L. (1973), 'Male-Female Wage Differentials in Urban Labor Markets', International Economic Review, 14(3): 693–709.

Oaxaca, R.L. and Ransom, M.R. (1994), 'On Discrimination and the Decomposition of Wage Differentials', Journal of Econometrics, 61(1), 5–21.

Oaxaca, R.L. and Ransom, M.R. (1999), 'Identification in Detailed Wage Decompositions', Review of Economics and Statistics, 81(1): 154–157.

O'Neill, J. and Polachek, S. (1993), 'Why the Gender Gap in Wages Narrowed in the 1980s', Journal of Labor Economics, 11(1): 205–228.

Psacharopoulos, G. and Patrinos, H.A. (2002), 'Returns to Investment in Education: a Further Update', World Bank Policy Research Working Paper 2881, Washington DC.

Reimers, C.W. (1983), Labor market Discrimination against hispanic and black Men, Review of Economics and Statistics, 65(4): 570–579.

Roos, P.A. (1981), 'Sex Stratification in the Workplace: Male-Female Differences in Economic Returns to Occupation', Social Science Research, 10(3): 195–223.

Rosenzweig, M.R. (1976), 'Nonlinear Earnings Functions, Age, and Experience: a Nondogmatic Reply and Some Additional Evidence', Journal of Human Resources, 11(1): 23–27.

Rumberger, R.W. and Daymont, T.N. (1984), 'The Economic Value of Academic and Vocational Training Acquired in High School', in M. Borus (ed.), *Youth and the Labor Market*, Kalamazoo: W.R. Upjohn Institute for Employment Research.

Sandell, S.H. and Shapiro, D. (1978), 'The Theory of Human Capital and the Earnings of Women: a Re-examination of the Evidence', Journal of Human Resources, 13(1): 103–117.

Suter, L.E. and Miller, H.P. (1973), 'Income Differences between Men and Career Women', American Journal of Sociology, 78(4): 962–974.

Treiman, D.J. and Hartmann, H.I. (eds.) (1981), *Women, Work, and Wages: Equal Pay for Jobs of Equal Value*, Washington DC: National Academy Press.

Treiman, D.J. and Terrell, K. (1975), 'Sex and the Process of Status Attainment: a Comparison of Working Women and Men', American Sociological Review, 40(2): 174–200.

Trostel, P., Walker, I. and Woolley, P. (2002), 'Estimates of the Economic Return to Schooling for 28 Countries', Labour Economics, 9(1): 1–16.

Wellington, A.J. (1993), 'Changes in the Male/Female Wage Gap, 1976–1985', Journal of Human Resources, 28(2): 383–411.

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